Comparative essays in labour market outcomes.

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COMPARATIVE ESSAYS IN LABOUR MARKET OUTCOMES

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SUMMARY

This thesis consists of three essays which provide a detailed empirical investigation of the returns to education, gender wage gap and public-private wage differential in Bulgaria, Serbia, Russia and Tajikistan – countries that have received little attention in the literature. The studies are based on rich data sets which allow the most up-to-date analysis of the specific labour market outcomes. All three essays go a step further than the existing empirical literature since in each one the quantile regression results showed a much broader picture than the ones based on central tendency measures such as Ordinary Least Squares (OLS).

The first essay looks at what had happened to the returns to human capital in Bulgaria over the period from early 1986 pre-transition to 2003. The study also contributes to the literature by estimating returns to education across the entire wage distribution, providing further evidence from Serbia, Russia and Tajikistan. Moreover, it deals with endogeneity and sample selection biases in a quantile regression framework.

The second essay estimates gender wage gaps in the selected countries by applying a decomposition method that simulates marginal distributions from the quantile regression process. The study seeks to extend the popular Machado and Mata (2005) distributional approach by addressing the 'index' number problem suggested by Neumark (1988) and Oaxaca and Ransom (1994 and 1998). The gender wage gap decomposition is performed for each quantile of the earnings distribution by using the pooled wage structure as a non-discriminatory structure and giving a much richer picture of the influence of the covariate and coefficient effects.

The third essay provides a comprehensive empirical study on the public-private wage differential in Bulgaria, Serbia, Russia, and Tajikistan. The study seeks to understand whether the differential in the public-private sector payment is explained by differences in workers characteristics or the difference in the returns to these characteristics. The endogenous sector choice is also considered. The study further analyses what has happened to the public sector hourly earnings differential at different points in the conditional earnings distribution and over time by adapting the Donohue-Heckman time-wise decomposition.

DECLARATION

This work has not been previously accepted in substance for any degree and is not being concurrently submitted in candidature for any degree.

Signed ______________________________ (candidate)
Date 02/11/2012

STATEMENT 1

This thesis is the result of my own investigations, except where otherwise stated. Where correction services have been used, the extent and nature of the correction is clearly marked in a footnote(s).

Other sources are acknowledged by footnotes giving explicit references. A bibliography is appended.

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STATEMENT 2

I hereby give consent for my thesis, if accepted, to be available for photocopying and for inter-library loan, and for the title and summary page to be made available to outside organisations.

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Material from the Living Standard Measurement Surveys has been made available from the World Bank which bears no responsibility for this analysis or interpretation of the data.

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ABBREVIATIONS

2SLS	Two stage Least Squares
BIHS	Bulgarian Integrated Household Survey
CIS	Commonwealth of Independent States
CEE	Central and Eastern Europe
CDF	Cumulative Density Function
EBRD	European Bank for Reconstruction and Development
ECA	Europe and Central Asia
ETF	European Training Foundation
EU	European Union
FRS	Family Resources Survey
FSU	Former Soviet Union
GDP	Gross Domestic Product
GHS	General Household Survey
GNP	Gross National Product
GSS	General Social Survey
GSOEP	German Socio-economic Panel
ISCED	International Standard Classification of Education
IIA	Independence of Irrelevant Alternatives
IV	Instrumental Variable
IMF	International Monetary Fund
QR	Quantile regression
LAD	Least Absolute Deviation
LATE	Local Average Treatment Effect
LFS	Labour Force Survey
LPM	Linear Probability Model
LSMS	Living Standard Measurement Survey
NACE	Classification of Economic Activities in the European Community
NATO	North Atlantic Treaty Organization
NGO	Nongovernment organisation
OECD	Organisation for Economic Cooperation and Development
OLS	Ordinary Least Squares
ONS	Office for National Statistics
RLMS	Russian Longitudinal Monitoring Survey
SFRY	Socialist Federal Republic of Yugoslavia
VTsIOM	Russian Centre for Public Opinion Research
WB	World Bank
UNESCO	United Nations Educational, Scientific and Cultural Organisation
UNDP	United Nations Development Programme
UNICEF	United Nations Children's Fund
USSR	Union of Soviet Socialist Republics

CHAPTER ONE

Research Introduction and Overview

1.1. Background and motivation of the study

It is well known that education and earnings are the main determinants of good labour market outcomes for individuals. Many empirical studies of the relationship between education of individuals and their income show that better educated workers earn higher wages in the labour market (Ashenfelter and Rouse, 1998). This is a conventional view of return to education. Estimate of returns to education are useful indicator of the productivity of education and the incentive for individuals to invest in their own human capital. Like other investments it is natural to ask is it profitable for individuals to invest in education. The answer to this question, if we have labour markets such as those in transition, is not straightforward. Transition countries could be considered as a special case when studying returns to education. Under Soviet rule wages were compressed and differences in wages, depending on the educational level, were low. As a result of the transition process, education became more highly rewarded. Structural transformations, disruptions and economic disequilibria were important factors for the relatively rapid increase in the returns to education. Lei (2005) finds that an increase in the returns to education, especially for women, is a common feature of transitional economies. Relatively little is known about the late transition years and especially the period of EU accession, which was characterised by the progressive implementation of a free market economy in the countries of Central and Eastern Europe. From this point of view, it is interesting to estimate returns to education and to observe possible changes in the rates of return to education during the period of transition.

Trying to quantify the impact of education by measuring the relationship between the number of years of schooling and earnings, however, has many shortcomings. Obtaining accurate and credible measures of returns to schooling involves minimizing biases caused by omitted variables, endogeneity and sample selectivity. Blackburn and Neumark (1995) remarked that one of the longest-running debates in empirical labour economics regards the biases in Ordinary Least Squares (OLS) estimates of the economic return to schooling. It is commonly assumed that the most important unobserved component of earnings is innate ability. However, another bias that can contaminate OLS estimates is endogeneity which arises because people with higher marginal returns to education choose higher levels of schooling. Finally, the principle of comparative advantage suggests that there should be some correlation between the decision to participate in the labour market and the size of the reward for educational investments. If this is the case, then not accounting for participation would provide an unrepresentative estimate of the rate of returns to education. Ideally the way we measure the returns to education would account for all these estimation difficulties.

One further conclusion from the existing literature is that men have generally benefited more from the transition process than women because they differ in their ability to handle the uncertainty inherent in a market economy. The downturn in economic development experienced by most transition countries during the 1990s was counterproductive to the provision of equal opportunities for men and women (Malysheva and Verashchagina, 2008). The most evident effect of transition was a decline in female labour force participation as well as a concentration of women into low-paid jobs (Brainerd, 2000). Although the gender wage gap has obviously narrowed over the last few years, earnings differential between men and women remains a feature of transition countries. According to the 2009 Human Development Report, the ratios of estimated female to male income is 0.64 in Russia, 0.65 in Bulgaria, 0.59 in Serbia, and only 0.47 in Tajikistan. According to the World Bank 2005 Report, the gender gap in Tajikistan is the worst in the region.

Nowadays, gender disparities exist, even in countries without glaring male domination and measuring these disparities is a necessary step towards implementing corrective policies. In light of the international awareness of gender issues, it is surprising that no country has yet managed to eliminate the gender wage gap. Those that have succeeded in narrowing the gap are the Nordic countries, with Sweden standing out as the most advanced in the world, followed by New Zealand, Canada, United Kingdom, Germany and Australia, all of which have made considerable progress in removing obstacles to the full participation of women in their societies (Claros and Zahidi, 2005). Due to the fact that the issue of the gender wage gap has continued to attract considerable political attention it is useful to consider how some countries in transition have addressed the unequal pay of men and women. The interest in this issue becomes even more relevant given the fast changing nature of the labour market in transition economies.

Perhaps the most common approach to identifying and quantifying the gender wage gap is the traditional technique of decomposing inter-group differences in mean levels of the outcome into a part due to different observable characteristics or 'endowments' across groups (the characteristic effect) and a part due to differences in the returns that workers receive for a given set of characteristics (the coefficient effect). One important choice to be made in the Oaxaca (1973) and Blinder (1973) decomposition is the assumption made concerning what wages would have been in the absence of discrimination, that is, the non-discriminatory wage structure. Symptomatic of the 'index number problem', results vary depending on the choice of male or female prices for weighting the decomposition equation. Various studies have emerged: adopting either the male wage or the female wage structure as the reference wage structure (Oaxaca 1973); adopting a weighted average of the male and female wage structure with the proportion of each subgroup in the population used as weights (Cotton 1988); or adopting the pooled wage structure (Neumark, 1988, Oaxaca and Ransom, 1994).

More detailed explanations of the wage gap also include utilizing quantile regression technique to understand the contribution of observed characteristics at different part of the wage distribution. Machado and Mata (2005) for example, develop a technique based on quantile regression to study the various source of wage inequality through a counterfactual density analysis. Hence the wage gap at each quantile in the distribution can be decomposed into a portion due to the differences in observed characteristics; a portion due to the differences in the returns to those characteristics; and a portion due to differences in residual wages. However, this technique shares all the problems of the original Oaxaca-Blinder decomposition, e.g., a potential sensitivity of the results in respect of the choice of the reference group. This is the starting and most challenging point for the extension of the Machado and Mata (2005) quantile regression approach and addressing the 'index' number problem along the line suggested by Neumark (1988) and Oaxaca and Ransom (1988 and 1994).

The current thesis also considers the fact that gender pay differences are just one of the many ways in which wage rates differ between individuals. The sector of employment might affect employee earnings in a number of ways and wage differentials might exist between public and private sector workers. Most studies of developed countries reveal that public sector workers are likely to get a positive wage premium relative to private sector workers. For example, Fogel and Lewin (1974) argue that public sector employers frequently pay more than necessary to attract a work force at the low and middle-skill ranges, and generally pay less than necessary to attract employees of average quality at the upper managerial and professional levels. This empirical conclusion corresponds with the institutional framework within which the public sector wages prevail in the private sector. Higher job security and unionization of public sector workers strengthens their bargaining power, easing access to rents and allowing a higher wage premium (Gimpelson and Lukiyanova, 2009). Generally, if the public sector overpaid in comparison with the private sector, employees in the private sector may decide to queue for the relatively high paying jobs in the public sector. If the government underpaid in comparison with the private sector, it would not find skilled employees.

During the transition process, the size of the public sector has been typically reduced and the labour market has changed from an exclusively formal market with public institutions to an extremely polarized market, with pre-transition formal employment in the public sector coexisting with an emerging and often highly informal and unregulated, private sector (Valmori, 2008). In all transition countries there is clear evidence of emerging non-state activity. In Bulgaria, for instance, after the crisis of 1996 and 1997, stabilization and mass privatization, the private sector became a larger part of the economy (Falaris, 2004). Employment in the private sector rose from 19% in 1986 to 75% in 2008. Similarly, the proportion of the private sector employees increased sharply in Serbia, Russia and Tajikistan. Given the differences in wage setting procedures and changes in the sector of employment during transition, it is important to investigate whether an identical employee working in the same job in the public and in the private sector earn the same amount. This strand of the literature has been largely underdeveloped for the selected countries, which is surprising given the current interest in public wage setting behaviour in these countries.

All these issues associated with the labour market outcomes constitute the broad starting point of this thesis. The collection of empirical investigations in these issues compromises three labour market essays. Each essay explores a different aspect of the labour market and aims to provide new evidence on interesting empirical patterns within the labour markets of a number of transition countries.

Specifically, the thesis is motivated by the following major concerns. First, cross-country empirical studies that evaluate rates and trends in returns to education in transition economies are limited. An interesting question is, therefore, whether the estimated returns to education have varied significantly over time. Second, despite the remarkable advances in micro-econometrics, many issues regarding estimation biases are far from being resolved. Substantial attention in the literature has been devoted to the problem of sample selection bias in the estimation of the returns to education, when wages are only observed as conditional on employment. Furthermore, if education is endogenous then estimation by OLS will yield biased estimates. Therefore, it is important to estimate returns to education in the transition countries thus paying particular attention to a number of important estimation difficulties – endogeneity, heterogeneity and sample selection biases. Third, the fast changing nature of the labour market in transition economies makes gender wage differentials an interesting topic for research, and understanding the effect of the transition process may have an essential key to understanding economic inequality. This thesis seeks to present a richer description of the differential by using a quantile regression approach and providing evidence from four countries. Finally, little research has been provided about the public-private wage differential in the case of transition and in particularly about the evolution of the wage differential. The development of a private sector is generally viewed as a key part of a successful transformation to a market economy (Hyder and Reilly, 2005). It is appealing, therefore, to examine what has happened to the public sector pay in the countries of interest.

The thesis considers the case of four countries – Bulgaria, Serbia, Russia, and Tajikistan – formerly socialist economies that are culturally and economically quite different. The study is comparative in the sense that the evidence available for each country will be compared with the rest of the examined economies. When attempting to make such comparison, it is necessary to discuss how the countries have been chosen. This can be answered from several points of view. Firstly, the interest in transition countries has developed since the beginning of reforms from planned to market economy and only a few empirical studies have investigated these issues in the context of transition. Secondly, the experience of the Eastern European and Central Asian economies is an interesting case study as the dramatic political, economic and social changes experienced during the transition provides a natural 'laboratory', with tremendous variation in key variables (Svejnar, 1999). At the beginning of the 1990s all former communist countries possessed a highly educated labour force and were expected to successfully utilize this precious resource. Among the numerous changes, transition brought inequality between men and women in terms of employment and earnings. It is then natural to ask how the position of women in the labour market has been affected by such changes. Finally, the focus on the transition countries was motivated by the fact that international institutions, such as the World Bank, were actively involved in supportive reforms in these countries. It is hoped therefore, that the analysis will be of interest to international policy makers.

The four countries offer a range of economic and labour market conditions and have experienced different performance over the past decade. Specifically, these economies differ dramatically from one another in terms of initial conditions, policies and outcomes. For example, the common perception in the literature views Bulgaria and Serbia as Balkan countries where reforms have proceeded more slowly compared to the advanced reformers of Central Europe. Serbia was confronted with a delayed transition and an erosion of skills. It differs from Bulgaria with respect to the speed of educational reforms and their impact on the individual labour market outcomes (Arandarenko *et al.*, 2006). The choice of Serbia is based on the fact that as a country its initial conditions were very different from the other transition countries. Serbian transition commenced 10 years after the first East European countries, offering an early experience on which the country could rely when designing its transition process (Stålhandske, 2007). In the period from 1991-2000, when other countries in transition were strengthening their economic systems, Serbia went through a five years civil war, isolation and sanctions imposed by the international community, hyperinflation, escalation of terrorism, secessionism in Kosovo and Metohia and NATO bombing.

Russia is a country where the communist experiment had the longest history (Atkinson and Micklewright, 1992). It is an interesting case with its transition away from a planned economy, characterised by over-education and over-employment to a market based economy. The study by the IMF, World Bank, OECD and EBRD (1991) shows that the move to a market economy and the removal of government interventions imposed substantial hardship on many groups in the Russian population. Following the end of central planning, wage inequality in the country had risen far more than in the Central and Eastern European (CEE) transition countries (Lehman and Wadsworth, 2001). According to estimates based on official statistics, the Gini coefficient for wages in Russia rose from 0.2 before transition to around 0.5 in 1996 (Flemming and Micklewright, 1999). In contrast, over the same period, the estimated Gini index for wages in the CEE countries rose from 0.2-0.25 to levels in the range of 0.3-0.35 (Lehman and Wadsworth, 2001). Given these facts, the interest in Russia becomes even more relevant.

Finally, as part of the Soviet Union, Tajikistan differs dramatically from other countries included in the thesis. While Russia has enjoyed rapid economic growth and has a promising future in terms of strong fundamentals as a middle-income country, being well endowed with human capital and abundant natural resources, in Tajikistan nearly two-thirds of the population continue to live in poverty. A comprehensive study on Tajikistan is of particular interest, because the transition process has witnessed a process of political liberalisation, which in association with an underlying Islamic culture has allowed the resurgence of more particular attitudes to female labour force participation. In addition, there are shortcomings in studies and available information on Tajikistan.

In selecting these four countries, others which could put forward claims for attention have had to be ignored. An obvious choice would have been some Central

European countries, such as the Czech Republic, Hungary and Poland, which would have provided useful comparisons. However, in order to keep the thesis manageable, we had to restrict ourselves to the four countries – Bulgaria, Serbia, Russia and Tajikistan.

1.2.Research objectives

The overarching aim of the thesis is to provide a detailed empirical investigation of the returns to education, gender wage gap and public-private wage differential in four transition countries. To achieve this aim, the empirical chapters develop a body of evidence. In fulfilling the aim a number of objectives were considered to be important:

1. To provide a comprehensive review of the literature related to the main aspects of the research and classify the main studies concerning the returns to education, gender wage gap and public-private wage differential with a special focus on transition countries.

2. To analyse and estimate returns to education in the countries of interest, first by answering the question: What has happened to the returns to human capital in Bulgaria before and during transition from a centrally planned to a market economy? Examining this issue, several methodological and empirical limitations will be faced. The study deals with endogeneity and sample selectivity issues. Further, the extent to which the return to education varies across the conditional earnings distribution will be explored by utilising data from Bulgaria, Serbia, Russia and Tajikistan.

3. To explain gender wage differential across the entire earnings distribution within the broader context of changes in the wage structure of Bulgaria, Serbia, Russia and Tajikistan. The study assesses whether the size of the gender wage differential has been widening or narrowing across the distribution in two separate time periods. The study seeks to extend the popular Machado and Mata (2005) decomposition procedure by utilising the

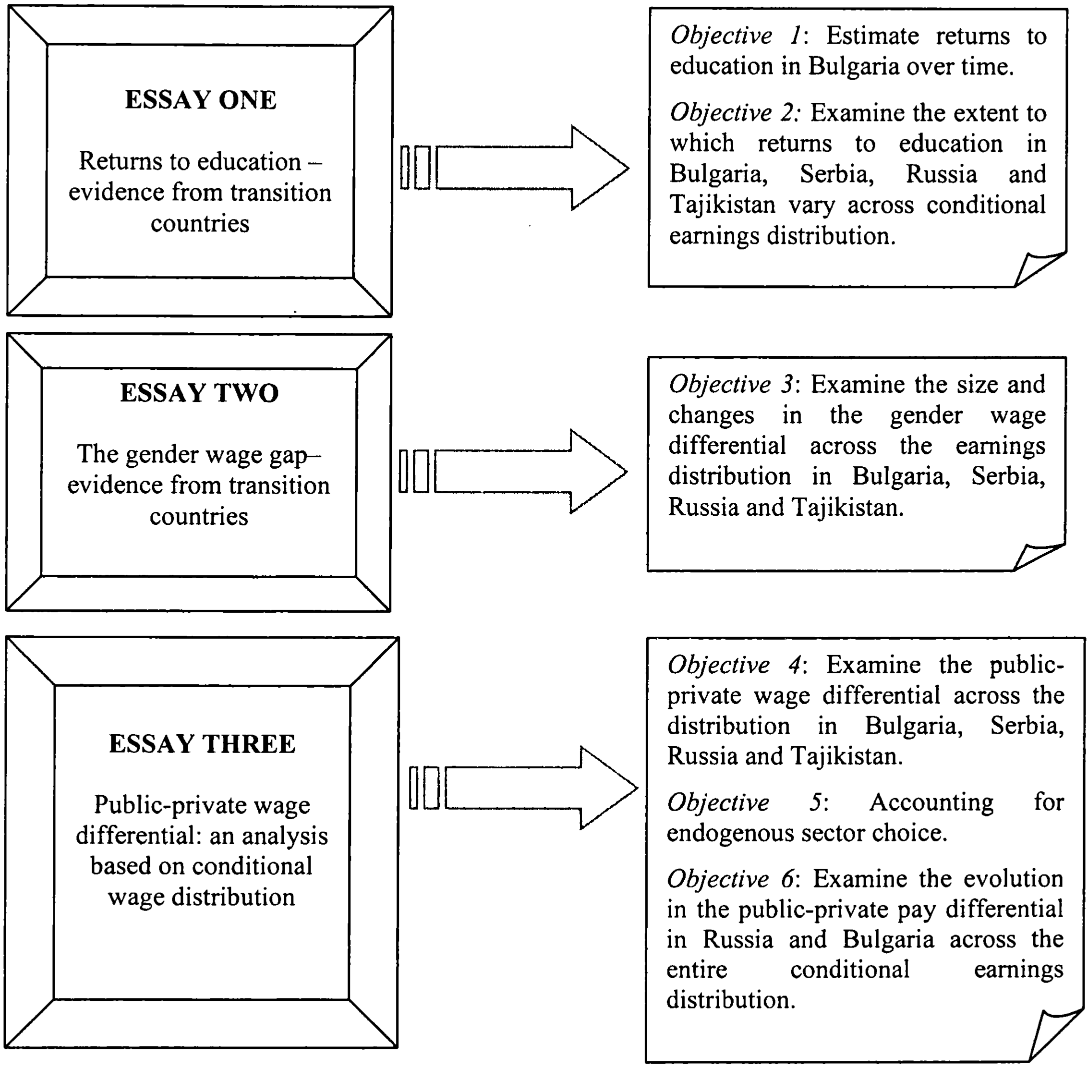
coefficient estimates from the pooled wage regressions as a non-discriminatory wage structure.

4. To examine the public-private wage differential across the entire earnings distribution in transition and extend the analysis by providing empirical evidence on the evolution of the public sector pay gap. The study adapts the Donohue and Heckman (1991) time wise decomposition within the Machado and Mata (2005) distributional approach to account for changes over two time periods. Another key objective is to control for the endogenous sector choice and to address the issue by extending the decomposition technique to account for selection across the entire wage distribution.

5. To summarise and evaluate the main findings of the research, draw conclusions based on these findings and implications for the policy makers and economists, and identify directions for future research in this field.

For ease of reference Figure 1.1 below summarises the structure and maps the main aims of the present thesis.

FIGURE 1.1: Research outlines and objectives



1.3. Data and methodology

The objectives are achieved through use of rich data sources and application of the latest econometrics methods. The main difficulty in the study of evolution in the returns to education in Bulgaria is the lack of reliable data over time. Gathering data for the early transition and especially for the communist period is extremely difficult. The first empirical analysis is based on five cross-sectional household surveys. One is administrated before transition and the others after the initial economic reforms were held: the 1986 Town and Village Survey and Bulgaria Integrated Household Surveys (BIHS) collected in 1995, 1997, 2001 and 2003. The empirical results in Chapter Five are based on data from Bulgaria, Serbia, Russia and Tajikistan, collected in 2003. The gender wage decomposition analysis in

Chapter Seven and the public-private wage differential estimates in Chapter Nine are conducted using data from the same four transition countries.

The empirical research involves the use of the latest applied econometrics techniques, such as parametric and semi-parametric two-step estimates, instrumental variables (IV), quantile regression decomposition techniques and probability models. In particular, to address the problem of endogeneity bias in the returns to education in Bulgaria over time, an IV approach has been applied. Identification in the endogeneity adjusted two-step estimations is achieved by making use of a mother's education as an instrument, assuming that there is a high correlation between children's educational attainments and those of their parents. Further, the question of the impact sample selection has on the returns to education is addressed by adopting Heckman's (1979) two-step procedure and a semi-parametric two-step estimator. The distributional approach is based on the use of a quantile regression (QR) method. As the bivariate normality assumption between the error terms in the earnings and participation equation does not hold in the QR model, following Buchinsky (1998) we use a higher order correction term to account for sample selection. To control for endogeneity bias in a QR framework, the control function approach proposed by Lee (2007) was adopted and regional variations in average schooling used as an instrument. The approach corrects for endogeneity by adding a residual power series as additional explanatory variables in the earnings equation. In line with more recent studies on gender wage differentials, the QR method is used as it allows estimation of the marginal effect of the covariates on the wages at various points of the distribution.

The gender wage gap across the earnings distribution is decomposed by utilising and extending the Machado and Mata (2005) technique accounting for the 'index' number problem suggested by Neumark (1988) and Oaxaca and Ransom (1988 and 1994). From our point of view, it is important that the estimation method chosen allows us to take account of the differences across the wage distribution and the non-discriminatory wage structure found from a pooled sample of males and females.

The same decomposition technique is applied to examine the public-private wage differential in Chapter Nine. The evolution in the public sector hourly earnings differential at different points in the conditional earnings distribution is examined by adapting the Donohue-Heckman time wise decomposition. In this analysis the public sector pay gap is estimated separately for both males and females. To estimate the wage equation intercept, which is needed for the Oaxaca-Ransom wage decomposition consistency, while accounting for sector choice, we follow the procedure suggested by Andrews and Schafgans (1998).

1.4. Structure of the study

To achieve our aims, the three main essays on labour economics issues develop a body of evidence. All three essays go a step further than the current literature on the returns to education, gender and the public sector wage gap by developing a quantile regression approach which provides a much broader picture than the results based solely on OLS estimates. Each essay explores a different aspect of the labour market outcomes but they are structured in a similar manner: providing motivation, highlighting the most important elements of the literature, developing a methodology, and presenting and discussing the main results. The layout of the thesis is as follows:

Chapter Two surveys the literature on the returns to education and considers the main issues involved in estimating the returns to education. It also examines how major estimation biases can be potentially overcome. Specific emphasis is placed on previous empirical studies in transition countries. Similarly, the chapter reviews the most important literature related to the methodology used in order to put the research into perspective.

Chapter Three provides an overview of the education systems in each of the selected transition countries, and highlights their most important characteristics. It also provides a brief background of the Bulgarian, Serbian, Russian and Tajik economies by reviewing their labour market performance and some macroeconomic indicators.

Chapter Four looks at what had happened to the returns to human capital in Bulgaria over the period from early 1986 pre-transition to 2003. The study deals with endogeneity and selectivity bias by adopting different econometric approaches.

Chapter Five considers the extent to which returns to education vary across the earnings distribution by using cross-sectional data from Bulgaria, Serbia, Russia and Tajikistan over the same time period. The study examines the empirical implications of allowing for endogeneity and sample selection bias in a quantile regression framework by utilising Buchinsky (1998, 2001) power series estimator and Lee's (2007) control function approach.

Chapter Six outlines theoretical and decomposition analysis of the wage gap between males and females. It sets out the conventional econometric approach to the analysis of unequal pay and introduces the methodologies used in Chapter Seven. Discussion in the literature over the techniques of earnings decomposition starts from the main estimation problems and methodologies used. The emphasis is on the more recent decomposition techniques. The Chapter concludes with the motivation for the chosen decomposition technique.

Chapter Seven presents the empirical analysis of the gender wage gap in Bulgaria, Serbia, Russia and Tajikistan. The study applies a quantile regression procedure to simulate marginal earnings distributions and then assesses whether the gender wage gap in the transition countries has been widening or narrowing across the entire distribution over two time periods. The empirical study is comparative in the sense that we seek to compare the evidence from these four transition countries. The analysis of the gender wage differential is based on personal characteristics only. The emphasis is on human capital, rather than occupational characteristics. The specification of the earnings equation is based on a pure human capital model where the only determinants of differences in wages are variations in males' and females' specific human capital characteristics.

Chapter Eight overviews the empirical literature related to the public-private wage differential. It provides some theoretical explanations for the existence of earnings differential between public and private sector and discusses the main estimation methodologies.

Chapter Nine investigates public-private pay differential across the conditional earning distribution in the same four transition economies. The endogeneity of sector choice is considered. The study is further extended by using data from Bulgaria and Russia to answer the question, what has been happening to the public-private wage differential at different points in the conditional earnings distribution over time. It also adopts the Donohue and Heckman (1991) time-wise decomposition framework to identify the main factors contributing to the change in the ratio of public to private sector earnings over the same time period.

Chapter Ten highlights the main research findings emanating from this thesis, and provides policy implications and recommendations. This Chapter also discusses limitations of the analysis and identifies some suggestions for future research.

ESSAY ONE

RETURNS TO EDUCATION – EVIDENCE FROM TRANSITION COUNTRIES

CHAPTER TWO

REVIEW OF METHODOLOGICAL ISSUES AND EMPIRICAL LITERATURE

2.1. Introduction

The measurement of the return to education has been the object of extensive empirical research and has been the focus of considerable debate in the literature. If estimation of the returns to education is measured perfectly, with no measurement errors, and there is no endogeneity between earnings, schooling and other covariates, then the returns to education will be consistent – that is its probability limit will be equal to the average marginal return to schooling. However, a dominate feature of the simple Ordinary Least Squares (OLS) human capital earnings function is that such estimates may be misleading if individuals' schooling level is measured erroneously, or it may be correlated with the regression error and the OLS estimate of the schooling coefficient will be biased. The empirical literature focuses on ability bias, endogeneity, sample selectivity, and measurement error as the central problems in estimating the rate of return to education. Unmeasured or erroneously measured human capital components summarized by "ability bias" play an important role and the main econometric problem is of omitted variable bias in the earnings equation. Furthermore, if education is endogenous then estimation by least squares will yield biased results. Finally, the sample of population that work in the labour market, and so are included in the wage regression, may be biased, with potential high earners being overrepresented. Therefore, self-selection into paid employment could be another source of potential bias in the estimates.

There have been a number of approaches in the literature to deal with these problems. One approach attempts to control for unobserved characteristics that may bias conventional OLS by including ability proxies (Bedi and Gaston, 1999). Different set of studies have relied on using data for twins to estimate returns to education. Another broad approach relies on constructing a 'selectivity correction' term from a schooling attainment equation and including the correction term in the earnings equation to obtain a consistent estimate of the rate of return to education. More convincing approach relies on using exogenous 'natural' variation in educational attainment to provide instrumental variables (IV) estimates of the returns to education.

In this Chapter the main problematic issues involved in estimating the rate of return to education are discussed. In Section 2.2 we consider the main theoretical framework and empirical difficulties in estimating the return to education. We examine how potential biases can be overcome. Section 2.3 provides an overview of the international literature related to the rate of returns to education, with a special focus on previous studies that have examined return to education in transition countries. We divide the empirical literature review into parts, based on the methodology that they employ to tackle the estimation problems. Finally, concluding comments based on the literature review are provided in Section 2.4.

2.2. Theoretical framework and estimation problems

The human capital theory was first announced in 1960 by Theodore Schultz, but the starting point of most works on the estimation of rates of return to education itself took place four years later with Becker's monograph Human Capital (1964). Estimation of human capital model was further enhanced by Mincer (1974), who was one of the first to apply human capital concepts directly to the personal distribution of earnings and used a standard earnings function to estimate the returns to education and experience. Mincer's model focuses on the life-cycle dynamics of earnings and on the relationship between observed earnings, potential earnings and human capital investment, both in terms of formal schooling and job investment, where the optimal schooling decision is reached by balancing benefits and costs of alternative schooling choices. Investment in human capital takes two complementary forms: formal schooling measured by years of school completed, and work experience measured by potential years in the labour force subsequent to the completion of schooling. Consequently, post-school work experience can also be viewed as an investment in future earnings capacity (Hill, 1981).

Mincer's model of earnings (1974) is one of the most commonly estimated relationships in labour economics. The usefulness of his theory lies in the extent to which it provides a unified interpretation of the detailed empirical characteristics of the earnings distribution (Hill, 1981). In one equation, Mincer's framework captures two economic concepts: a pricing equation, or hedonic wage function, revealing how the labour market rewards productive attributes like schooling and work experience; and the rate of return to schooling which can be compared with the interest rate to determine the optimality of human capital investments (Heckman *et al.*, 2003).

The Mincer specification is given by the following equation[1]:

$$\ln[Y(s,x)] = \alpha_0 + \rho_s s + \beta_0 x + \beta_1 x^2 + \varepsilon \tag{2.1}$$

where $Y(s,x)$ is the wage or earnings at schooling level s and work experience x, ρ_s is the 'rate of return to schooling' assumed to be the same for all schooling levels, and ε is a random variable with $E(\varepsilon \mid s,x) = 0$. While ρ_s may vary among individuals, it is assumed to be uncorrelated with s. This model is motivated by two conceptually different theoretical frameworks used by Mincer (1958, 1974), which briefly are reviewed in this section.

The first model (Mincer, 1958) uses the principle of compensating differences to explain why individuals with different levels of education receive different earnings over their lifetimes. The model considers that individuals have identical abilities and opportunities, that there is no uncertainty, credit markets are perfect, but occupations differ in the amount of training they require. Education is costly because individuals forego earnings while in education (Heckman, 2003). However, the model entails no direct costs. Individuals are assumed to be identical and they require a compensating differential to work in occupations that require a longer training period (Heckman *et al.*, 2005). The size of compensating differential is determined by equating the present value of earnings streams net of the costs

[1] Psacharopoulos (1981) and Psacharopoulos and Patrinos (2004) provide surveys of an enormous Mincer based earnings literature.

associated with different levels of investment (Heckman, 2003). Further, Heckman *et al.* (2006) argue that this framework ignores uncertainty about future earnings as well as nonpecuniary costs and benefits of schooling and work.

If we let $Y(s)$ show the annual earnings of an individual with s years of education, which are assumed to be constant over his/her lifetime; r an externally determined interest rate; and T the length of working life, which is assumed not to depend on s, the present value of earnings associated with schooling level s is:

$$V(s) = Y(s)\int_s^T e^{-rt}dt = \frac{Y(s)}{r}(e^{-rs} - e^{-rT}) \qquad (2.2)$$

An equilibrium characterized by heterogeneous schooling choices requires that individuals are indifferent between schooling levels. The allocation of people to different schooling levels, therefore, is driven by demand conditions. Equating the earnings streams associated with different schooling levels and taking logarithms yields the following equation:

$$\ln Y(s) = \ln Y(0) + rs + \ln((1 - e^{-rT})/(1 - e^{-r(T-s)})) \qquad (2.3)$$

The final term on the right hand side of (2.3) is an adjustment for finite life, which converges to zero as T gets large[2]. Specifically, Mincer (1958) observes that for large T, the coefficient of years of schooling equals the interest rate r. People with more education receive higher earnings. Because the internal rate of return to schooling represents the discount rate that equates lifetime earnings streams for different education choices, it will also equal the interest rate in this model. Therefore, ρ_s in equation (2.1) yields an estimate of the internal rate of return and when $\rho_s = r$ the education market is in equilibrium. If $\rho_s > r$, there is under-investment in education (Heckman *et al.,* 2006).

The second model used by Mincer (1974), and now widely applied, is motivated by entirely different assumptions from his earlier model. The second model focuses on

[2] This term also disappears if the retirement age - T is allowed to increase one-for-one with s, so post-school working life is the same for persons of all schooling levels.

the life-cycle dynamics of earnings and on the relationship between observed earnings, potential earnings, and human capital investment. Persons are ex-ante heterogeneous, so the compensating differences motivation of the first model is absent. At the same time, no explicit assumptions are made about the background of the economic environment (Heckman *et al.*, 2003).

Mincer (1974) specifies observed earnings as a function of potential earnings net of human capital investment costs, where potential earnings at any time depend on investments in previous time periods (Heckman, 2003). Let P_t be potential earnings at time t, investment in training can be expressed as a fraction k_t of potential earnings invested, i.e. $C_t = k_t P_t$ and let ρ_t be the average return to training investments made at age t. Then:

$$P_{t+1} = P_t + C_t \rho_t = P_t(1 + k_t \rho_t) \tag{2.4}$$

Repeated substitution yields $P_t = \prod_{j=0}^{t-1}(1 + \rho_j k_j) P_0$

Formal schooling is defined as years spent in full-time education $(k_t = 1)$, which is assumed to take place at the beginning of life and to yield a rate of return ρ_t which is constant across all years of schooling $(\rho_t = \rho_s)$. Assuming that the rate of return to post-school investment ρ_t is constant and equal to ρ_0 one can write:

$$\ln P_t = \ln P_0 + s \ln(1 + \rho_s) + \sum_{j=s}^{t-1} \ln(1 + \rho_0 k_j) \tag{2.5}$$

which yields the approximate relationship (for small ρ_s and ρ_0)

$$\ln P_t \approx \ln P_0 + s\rho_s + \rho_0 \sum_{j=s}^{t-1} k_j \tag{2.6}$$

To establish a relationship between potential earnings and years of labour market experience, Mincer (1974) follows Ben-Porath (1967) and assumes a linearly declining rate of post-school investment:

$$k_{s+x} = k(1 - \frac{x}{T}) \tag{2.7}$$

where $x = t - s \geq 0$ is the amount of work experience at age *t*. The length of working life, *T*, is assumed to be independent of years of schooling. Under these assumptions, the relationship between potential earnings, schooling and experience is given by:

$$\ln P_{x+s} \approx [\ln P_0 - k\rho_0] + \rho_s s + (\rho_0 k + \frac{\rho_0 k}{2T})x - \frac{\rho_0 k}{2T}x^2 \tag{2.8}$$

Observed earnings equal potential earnings less investment costs, producing the following relationship for observed earnings:

$$\ln Y(s,x) \approx \ln P_{x+s} - k(1 - \frac{x}{T}) =$$

$$= [\ln P_0 - k\rho_0 - k] + \rho_s s + (\rho_0 k + \frac{\rho_0 k}{2T} + \frac{k}{T})x - \frac{\rho_0 k}{2T}x^2$$

$$= \alpha_0 + \rho_s s + \beta_0 x + \beta_1 x^2 \tag{2.9}$$

Thus, we arrive at the standard form of Mincer's earnings model (equation (2.1)) which regresses log earnings on a linear term in years of schooling and linear and quadratic terms in years of labour market experience (Heckman *et al.*, 2003).

The Mincer model (1958) provides an equilibrium-based model of earnings determination and an easy way to estimate the internal rate of return to schooling. However, Heckman (2008) criticise Mincer's model because of the assumptions that individuals have identical abilities and opportunities, credit markets are perfect, there are no taxes and the environment is perfectly certain. More specifically, in the model there is no loss of working life with additional years of schooling, earning functions are multiplicatively separable in experience and schooling, marginal returns equal average returns[3], which together with multiplicative separability implies linearity of log earnings in schooling over the observed range of schooling choices. Mincer also neglects major determinants of actual returns, such as the direct and indirect costs of schooling, taxes, the length of an individual's working life and uncertainty about future returns at the time the schooling decision is made

[3] Carneiro *et al.* (2006) present evidence that the marginal return does not equal the average return. See also Heckman *et al.* (2006).

(Heckman *et al.*, 2003). In most applications of Mincer model, it is assumed that the intercept and slope coefficient in equation (2.1) are identical across individuals. This implicitly assumes that P_0, k, ρ_s and ρ_0 are the same across persons and do not depend on the level of schooling.

Mincer formulates a more general model that allows for possibility that k and ρ_s differ across persons, which produce a random coefficient model:

$$\ln Y(s_i, x) = \alpha_{0i} + \rho_{si} s_i + \beta_{0i} x_i + \beta_{1i} x_i^2 + \varepsilon_i \tag{2.10}$$

Letting $\alpha_0 = E(\alpha_{0i}),\ \rho_s = E(\rho_{si}),\ \beta_0 = E(\beta_{0i}),\ \beta_1 = E(\beta_{1i})$ we may write this expression as:

$$\ln Y(s,x) = \alpha_0 + \rho_s s + \beta_0 x + \beta_1 x^2 + [(\alpha_{0i} - \alpha_0) + (\rho_{si} - \rho_s)s + (\beta_{0i} - \beta_0)x + (\beta_{1i} - \beta_1)x^2]$$

where the terms in brackets are part of the error.
Mincer (1974) initially assumes that $(\alpha_{0i} - \alpha_0), (\rho_{si} - \rho_s), (\beta_{0i} - \beta_0), (\beta_{1i} - \beta_1)$ are independent of (s, x), although he relaxes this assumption in later work.

The empirical approximation of the human capital theoretical framework, adopted by most researchers, is the familiar semi-logarithmic form of an earnings equation, where log earnings is modelled as the sum of a linear function of years of education and a quadratic function of years of potential experience[4]:

$$\log Y_i = \alpha_i + \beta_1 s_i + \beta_2 x_i + \beta_3 x_i^2 + \varepsilon_i \tag{2.11}$$

where $\ln Y_i$ is the natural logarithm of the observed wage for individual *i*; α_i is a constant term; s_i is the number of years in schooling for individual *i*; β_1 is the average rate of return to one additional year of schooling; x_i is post-schooling experience of the individual worker which is entered in linear and quadratic forms; and ε_i is assumed to be normally distribute with mean zero and variance σ_i^2.

[4] Although a human capital earnings function is typically associated with the work of Mincer, other authors also made important contributions to that literature.

Some points regarding Mincer specification should be kept in mind. First, both the theoretical model and its empirical representation assume no interpersonal variation in the rates of return (Hill, 1981). Second, this specification is often estimated by an OLS estimator, which suffers from a number of potential biases that depending on their origin may exert either downward or upward pressure on the OLS estimates (Card, 1999). For example, the basic Mincerian specification assumes that earnings are linear in education, so that each year of education adds the same percentage amount to earnings irrespective of the particular year of education. In addition, schooling is often treated as exogenous, although education may be an endogenous choice variable in human capital theory (Harmon and Walker, 2001). Education is an optimizing investment decision based on future earnings and current costs – that is the difference in earnings from undertaking and not undertaking education and the total cost of education including any foregone earnings. If investment in education results in higher earnings, higher earnings may also result in an increase in the amount of education consumed, which is more likely the higher the ability of the individual. This is referred in the literature as endogeneity bias. In addition, in the Mincer specification, the disturbance term captures unobservable individual effects and these individual factors may also influence the schooling decision, and thereby induce a correlation between schooling and the error term in the earnings function. A common example is unobserved ability. The role of unobserved ability in affecting earnings is a subject that has produced a large body of literature[5]. According to Griliches (1977), ability bias arises because the estimation procedure is unable to separate the contribution of unobserved ability to productivity from that made by education, and attributes it all to education. Heckman and Vytlacil (2001a), show that education and cognitive ability are so strongly associated that the wage effects of the two cannot be separated.

The early efforts of Becker (1964) and Denison (1964) recognize that the simple earnings-by-education relationship typically overestimate the pure contribution of education because of an correlation between measured ability and years in school completed. The authors suggest that the observed income by education wage gap should be reduced by about one-third. Becker (1964) argues that the true rate of

[5] See, for example, Griliches (1970) and Hause (1972).

return to education is grossly overestimated because persons differing in level of education also differ in terms of other characteristics that cause their incomes to differ systematically. Unless there is a sufficiently large negative correlation between ability and opportunity, the model predicts that those workers who choose more schooling will also tend to have higher level of ability. The failure to control for ability in a wage regression should therefore, result in upward-biased estimates of the returns to schooling (Blackburn and Neumark, 1995).

In this context, the problem resolves around the definition and measurement of ability. In common usage, 'ability' means the power to perform something. In Becker's definition, this power is the capacity to increase future earnings by current investments. However, such capacity could be due to a combination of genetic factors, previous investment, experience, and other elements that might be difficult to unravel theoretically or statistically. Hause (1972) argues that there are some difficulties in going from this definition to an empirical study of schooling and its return. Difficulties arise in finding a data source that contains a joint distribution of ability, schooling and earnings (Hill, 1981). Most data sets do not contain information on ability; so that ability bias is, in this case a type of omitted variable bias[6]. Other authors measure ability by an intelligence or achievement tests. However, Hill (1981) argues that intelligence is not fixed and is not independent of schooling and other influences. In fact, intelligence can be affected by schooling and in turn, affects the amount of learning achieved in a given school situation (Hause, 1972).

On the other hand is the effect of ability on earnings over time for a given level of educational attainment. If this is true there would be no tendency for the coefficient on ability to attenuate with time in the labour force (at least over the first half of the life cycle of earnings). However, there may be a tendency for the effect of ability to increase with education, especially at high levels of education. Behind this hypothesis lies the idea of able people being more effective than the less able in raising productivity through job experience. That is measured ability and learning in the labour force are complements in the production of earnings. The persistence

[6] One exception is the National Child Development Survey, which analyses periodically a cohort of individuals born in 1958. The survey includes measures of both reading and mathematical ability.

or increase in the ability coefficient over time implies the existence of productivity differentials in earnings even after individuals reach peak-level earnings. Such differentials are compatible with the large observable dispersion in personal earnings, even after standardizing for age, schooling and occupation (Hause, 1972).

There have been numerous approaches in the literature to deal with omitted ability bias. First, measures of ability have been incorporated to proxy for unobservable effects. The inclusion of direct measures of ability would reduce the estimated education coefficient if it acts as a proxy for ability, since the coefficient on education then captures the effect of education alone since ability is controlled for. Another approach includes an explicit proxy for ability, for example from an intelligence quotient (IQ) test (Griliches and Mason, 1972; Griliches, 1977). Some data sets have a direct measure of IQ, but evidence regarding the relationship between intellectual capacity and earnings is not entirely consistent (Zax and Rees, 2002). Other studies simply find detailed test score information and include these test score variables in the earnings equation to obtain estimates of the returns to education after directly controlling for measured cognitive ability (Tobias, 2003) Specifically, if more able individuals 'convert' schooling into human capital more efficiently than the less able, this would raise the return to schooling for the more able. On the other hand, the more able may have higher opportunity costs since they may have been able to earn more in the labour market.

Second, one might exploit within twin differences in earnings and education if we accept the assumption that unobserved effects are additive and common within twins (Harmon *et al.,* 2001). The fixed effect estimator on a sample of identical twins can be used, from which inherent ability and family background effects are assumed to be the same. This procedure is perhaps more plausible for identical twins who share the same genetics and the same family background. If monozygotic twins indeed share the same level of ability, then estimating the difference in returns to education between twins will eliminate ability bias. However, twins represent a quite distinct population group and often sample size is small. Even where the sample size is larger as in Miller *et al.* (1995) a substantial proportion of the sample of twins report the same level of education. As a result, earnings function is sometimes estimated on siblings, father-son or mother-daughter paired samples, using fixed-effects or first differencing approaches. By introducing sub-samples of households with at least two individuals of a given gender in employment, the fixed-effects method effectively controls for all household variables that are common across these individuals within a household.

A final approach deals directly with a simultaneous relationship between schooling and earnings by specifying a two-stage equation system and identifying variables that affect schooling but not the dependent variable other than through the variable that is being instrumented (Harmon *et al.*, 2003). If one finds such a variable(s), then it can be used to construct instrumental variable estimates of the return to education. More specifically, the idea is to instrument schooling by decomposing the earnings function into a linear function expressed as:

$$s_i^* = \pi Z_i + \eta_i \tag{2.12}$$

where Z_i is a vector of exogenous variables that influence the educational decision. The next equation is estimated by substituting the fitted values from the first-stage regression of s_i^* on Z_i .

$$\ln Y_i = \alpha_i + \beta_1 \hat{s}^*{}_i + \beta_2 x_i + \beta_3 x + \varepsilon_i \tag{2.13}$$

Adding $\hat{s}_i^*$ is sufficient to eliminate the endogeneity bias arising from the possible correlation between schooling and unobservable influences on earnings. By construction, $\hat{s}^*$ is correlated with actual schooling and will capture the effect of education on earnings. However, $\hat{s}^*$ is by construction uncorrelated with the error term in the earnings function. Consequently the estimated return β_1, based on predicted schooling, is unbiased. The method described is the Two-Stage Least Squares (2SLS), which is a special case of the instrumental variable (IV) approach.

A consistent estimator for the returns to education also depends on the statistical identification of the model. In order to identify the effect of schooling on earnings we need to find a variable which is a determinant of schooling and can be omitted from the earnings function. However, searching for such instruments is potentially difficult and sometimes they are only weakly correlated with the endogenous variable in question. Using such variables as instruments is likely to produce estimates with large standard errors. Some authors highlight that many existing IV studies have been undermined by a lack of precision in their first stage estimates[7]. For example, Bound and Jaeger (1996) show how quarter of birth interactions with state and year, used in Angrist and Krueger (1991) study, form weak instruments causing IV to be more biased than the OLS.

Furthermore, the level of individual earnings is not wholly determined by education. A short list of other factors will include: location, occupation and industry, gender, race, age, physical condition, and intelligence, including both psycho-motor and intellectual skills. A simple bivariate analysis will always lead to an overestimation of the influence of education on income[8]. Many studies use age as a substitute, or proxy, for job tenure because job tenure is not readily available in most survey data sets. Age is typically a measurement-error-free variable which can be observed for all individuals. However, earnings regressions that contain supervisory status or occupations tend to produce lower returns to schooling due to occupational outcomes which are also correlated with schooling (Vernon, 2002).

Another issue in the returns to education literature is that married women typically experience breaks in their labour force participation, and most cross-sectional data do not provide information on either the amount of time spent outside the labour market or on the timing of these breaks. Researchers are forced to proxy, such as amount of potential experience estimated from age and educational level, which may fails to capture the sample variation in actual experience. Some cross-sectional or longitudinal studies in which information on actual experience is recorded (Malkiel and Malkiel, 1973; Mincer and Polachek, 1974) suggest that the estimated return to experience is seriously biased downwards when the incorrect measure is used (Zabalza and Tzannatos, 1985). However, even when detailed information on labour market experience is available, this variable should not necessarily be considered as exogenous. Experience is made up of accumulated participation and participation depends among other factors on earnings. For men, this is probably not very important, given that most men participate continuously irrespective of the level of their earnings. However, for married women, the strong earnings

[7] See Staiger and Stock (1997) and Bound *et al.* (1995) for more details.

[8] For in depth surveys of these issues see Card (1995, 1999) and Angrist and Krueger (1991).

relationship might seriously bias the coefficient on experience, particularly when earnings are estimated using the OLS technique.

The Mincer equation also includes the experience in quadratic form in order to capture the inverted U-shape nature of the relationship between the logarithm of earnings and experience or age. In some specifications tenure is added, in linear and quadratic form, to capture the acquisition of specific as opposed to general skills. Murphy and Welch (1990) examine in detail whether the standard quadratic specification in years of potential experience adequately captures the empirical experience-earnings profile. They conclude that a quadratic function is not flexible enough to capture the main features of the experience-earnings profile. According to them the main problem is that the quadratic function understates earnings growth over the first 10 to 15 years of individual career.

Furthermore, the question of measurement error also needs to be addressed. Ashenfelter and Krueger (1994) introduce the idea that individuals differ in their propensities to over-report or under-report schooling levels. Card (1999) suggests that research over three decades points to the reliability of self-reported schooling being no more than 90%. Often, the only data available to measure education is years of schooling. However, it can be questioned whether years of schooling adequately measures 'total education'. If schooling variable are subject to survey measurement errors, then conventional OLS estimates are biased. Griliches (1977) shows, that if the measure for education is imperfect, OLS estimates can have a large downward bias. Moreover, the bias is magnified (even if the error of measurement is small) when more variables are included in the model.

Additionally, more recent literature recognizes that returns may vary across schooling levels and across persons of the same schooling level[9]. Heterogeneity in the regression coefficients is another source of potential bias in OLS estimates. One of the properties of OLS estimation is that the regression line crosses through the mean of the sample (Harmon *et al.*, 2001). People differ with respect to their marginal return to education, their marginal cost for education and their tastes or

[9] See Heckman *et al.* (1997), Dearden (1999a) and (1999b) and Blundell *et al.* (2001).

choices[10], hence the return to education is not a single parameter but can potentially differs according to differences in individuals' family backgrounds. Some individuals may be able to access funds from family or other sources in order to acquire additional education, while others are unable to do so. Theoretically, heterogeneity is an important problem. Small differences in the estimates of the return to schooling in a given year can generate large differences in estimates of lifetime earnings and of lifetime return to schooling. Therefore, unobserved heterogeneity might induce a dependency between the education variable and the error term. In this context, Card (1999) argues that people with higher returns to education tend to acquire more schooling and hence a cross-section regression of earnings on schooling yields an upwardly biased estimate of the average marginal return to schooling. An alternative methodology is available to OLS known as quantile regression (QR) which allows us to estimate the return to a particular level of education within different quantiles of the wage distribution.

In approaching this empirical measurement problem, two fundamentally different views of the labour market have been taken. The first view, associated with Griliches (1977), adopts an efficiency unit view of the labour market. Human capital is homogeneous but people possess different amounts of it (Carneiro *et al.*, 2001). This literature therefore focuses on ability bias and measurement error as the central problems in estimating the rate of return to schooling (Heckman and Vytlacil, 2001). The second view, associated with Roy (1951) and Rosen and Wills (1979), focuses on the choice of schooling and emphasizes comparative advantage in the labour market with heterogeneous human capital as a guiding principle. This view point is at odds with an efficiency unit point of view and highlights the need to address self-selection issues when considering the most appropriate estimator.

In estimating earnings function there is an additional complication and this is the problem of sample selectivity issues. It is argued that working women and men may not be randomly sampled from the overall population. Consequently, the sample of the population that work and so are included in the earnings regression, may bias

[10] In practice, returns to education can vary across people due to a number of unobserved factors, such as ability, motivation, and ambition, as well as differences in the interest rate faced by different individuals (Card, 2001).

the estimates with potential high earners being overrepresented[11]. More specifically, because offered wages are only observed for individuals who are employed, these methods require the estimation of the covariance between the error term in the wage offer equation and the error term in the equation determining the probability of employment. This covariance is the sample selection effect (Ermisch and Wright, 1994). The sample selection correction procedure may be especially important for low-income countries, because they often have a higher share of family workers and self-employed workers, and for women, because the proportion participating in the labour market may be small. Correcting for sample selection bias sometimes substantially changes the estimated rates of return to schooling.

Gronau (1974) first noted the problem posed by sample selection bias in the context of an earnings regression for women, in which wage data is only available for those who received acceptable wage offers. In response to the problem Heckman (1974) produced a maximum likelihood estimator and a two-step procedure that are able to recover unbiased estimates of the equation of interest in the presence of sample selection. Heckman (1976) argues that observed earnings are not a true reflection of the underlying earnings distribution because of the existence of selection bias. Indeed, search theory holds that unemployed people only accept a job if their earnings are higher than their reservation wage. In addition to the parametric two-step method proposed by Heckman (1976), Cosslett (1991), Gallant and Nychka (1987), Powell (1987), Ahn and Powell (1993) and Newey (1991) propose semi-parametric estimators for sample selection models, and more recently, Das *et al.* (2003) propose a fully non-parametric estimator for this model.

Sample selection is assumed to depend on a linear index and is controlled for by including the inverse Mill's ratio of the index as additional 'regressor' in the outcome estimation. The inverse Mills ratio is a nonlinear transformation of the Probit index and is a decreasing function of the probability of selection. It also indicates whether there is significant correlation between the error terms in the wage equation and the selection equation, which would otherwise bias OLS

[11] Working men and women are unrepresentative of the male and female population and policy inferences concluded from regressions for workers may be invalid, both for non-working men and women and for working men and women.

estimates (equivalent to an omitted variable problem). Dolton and Makepeace, (1986) argue that sample selection bias is an indication that the observed pattern of female earnings is lower than that which would have been observed for the average member of the whole female sample had she continued to work.

Despite the clear definition of this covariance, there have often been problems associated with its interpretation. In particular, a negative covariance has been viewed as being problematic, resulting from misspecification of the wage equation (Steinberg, 1989). The sign of sample selectivity term depends in part on the variables included in the earning and sample selection equations. A positive and significant coefficient on the inverse Mills ratio may imply that those who participate in the labour market have higher conditional mean earnings than those that do not. In this context, Dolton and Makepeace (1987) conclude that the notation of a correct sign for the selection coefficient is difficult to resolve on economics ground. They indicate that the positive Mills sign would suggests that increased earnings accrue to the type of person who is more likely to have made the participation choice compared with someone who is less likely to have made that choice. Alternatively, Blackaby *et al.* (2002) argue that a negative and significant selectivity term suggest that if those out of work were to find work they would have higher earnings than individuals with similar characteristics already in jobs. This is compatible with such individuals setting higher reservation wages, which is consistent with their lower employment probability.

The first step in a sample selection procedure is to specify an employment participation equation for women or men in a form of Probit function on a sample of employed and non-employed. Then selectivity bias can be corrected by including an additional regressor in the earnings function related to the probability of participation. Importantly, the Heckman correction requires a valid exclusion restriction. In this regards, Wooldridge (2006) shows two important implications. First, any element that appears as an explanatory variable in the first-step regression equation should also be an explanatory variable in the selection equation. Second, one should have at least one element in the selection equation that is not in the regression equation. Concerns that distributional assumptions made in Heckman two-step procedure were too strong lead to development of a semi-parametric technique that does not require distributional assumptions for identification. These methods, however, require that all employment and wage regressors are exogenous.

There are other criticisms of the Heckman two-step approach. For example, Nelsons (1984) shows that the inverse Mills ratio is often highly collinear with the explanatory variables and hence, estimates of returns to education tend to be unstable, non-robust and sensitive to minor changes in specification of both the earnings and participation functions. Nelson (1984) Monte Carlo results show that the standard errors of the estimated returns to education can be very large when the degree of collinearity is high. However, such comments are of second order of importance when set against the unwarranted bivariate normality assumption discussed in the previous paragraph (Moffitt, 1999). These two issues are related and both have been addressed by developing semi-parametric methods which do not rely on arbitrary distributional assumptions[12].

A number of studies have attempted to estimate returns to education for males and females separately to see if these returns differ according to gender[13]. Gender differences in the returns to education may arise for several reasons: differences in the opportunity costs of schooling for males and females; gender differences in traits, or strengths that are valued differently by the market; gender differences in jobs; relative scarcity; and sex discrimination in the labour market. A priori, none of these factors suggest a higher rate of return to education for males or females. However, Madden's (1987) argues that women's experience of higher average wage returns to education in the US is consistent with their higher average years of school attendance. Similarly, Altonji (1993) investigates how expected returns to education affect the decision to stay in school and the choice of college major. He finds female coefficients to be greater than the corresponding male coefficients for 17 out of 18 educational dummy variables. Elsewhere, Behrman and Deolalikar (1990)[14] argue that in Indonesia a possible reason for a higher return to education for females may be due to the manufacturing technology adopted in the country. If factories are structured in a way that physical strength is important to productivity,

[12] See Powell (1993) and Vella (1998).
[13] See Madden (1978) and Angle and Wissman (1981).
[14]Authors show that the return to university degree is 25% higher for females than for males.

the earning premium for men in unskilled factory positions and with low schooling would be considerable.

Finally, some authors stress the effect of schooling quality on earnings. Bedi and Edwards (2002), for example, examine the economic effects of school quality and find that men educated in countries with better quality schooling earned significantly higher incomes than men educated in low-quality schools. Similarly, Johnson and Stafford (1973) compute rates of return to quality as well as quantity, using the average price-deflated expenditure per pupil as a quality index. They find that expenditure per pupil affects years of schooling attained. Several other papers examine the impact of school quality on earnings for developed countries (mainly the United States). These papers provide mixed evidence. Some report strong effects of school spending on students' subsequent earnings while another find little or no effect[15].

2.3. Selective review of empirical studies

2.3.1. Instrumental variable (IV) literature

A general pattern that emerges in all IV studies is that the IV estimates of the returns to education tend to be larger than the OLS estimates[16]. This is true across studies that use different data sets and different instruments. By going from OLS to IV, the estimates often increase by more than 30% and in some cases by close to 100%, indicating that OLS estimates underestimate the true value of returns to schooling.

There are several explanations for this and measurement error has been cited as an explanation to reconcile these apparently different results[17]. Attenuation bias caused by the measurement error of schooling reduces OLS estimates. More specifically, if an individual's schooling level is measured erroneously and the true value of the returns to schooling is positive, the OLS estimate will be biased toward zero. If

[15] Examples of studies that find a positive and significant school quality effect on earnings include Rizutto and Wachtel (1980) and Card and Krueger (1992a,b). Papers that find insignificant school quality effects include Ribich and Murphy (1975), Betts (1995), and Grogger (1996). Heckman *et al.* (1995) also report insignificant effects, although the paper does report a positive effect of school quality on earnings for college graduates.

[16] Another pattern is that IV estimates are also more imprecise than OLS estimates of the return to schooling.

[17] Schooling is measured with error.

attenuation bias is relatively larger, the resulting IV estimates will be higher than the OLS estimates. Based on Card (1999), measurement error bias itself can explain 10% gap in the estimated returns between OLS and IV estimations. Generally, the two biases that exist simultaneously in applying the OLS estimation are the upward bias caused by omitted ability variables and the downward bias caused by measurement error in schooling. The result of the IV estimates depends on the relative magnitudes of the omitted ability and attenuation biases. If the instruments are not correlated with the measurement error in the schooling level, then IV estimates will be free of both biases.

However, one of the criticism of the IV estimates resolves around the concern that the instrument may not be truly independent of the earnings residual. If the instrument is positively correlated with the earnings, the IV estimator may be biased upward. Second, an instrument may be so weakly correlated with the troublesome variable that in practice it will not overcome the bias of the OLS. Thus, when the instrumental variable is weak, the bias in the estimates can be very large.

The recent literature has suggested two different interpretations of the results obtained using IV methods. Card (1999) shows that from a simple model of endogenous choice, the return to education is not a single parameter in the population but a random variable that may vary with other individual characteristics. In other words, the differences between OLS and IV estimates might be caused by heterogeneity in returns to schooling (Bhalotra and Sanhueza, 2004). Card (1999) also suggests that the instrument probably influences the educational decision of individuals with high marginal returns and hence high discount rates. Consequently if IV relies on 'interventions' that affect the schooling choice of children from relatively disadvantaged family backgrounds (high discount rates) then their marginal returns to schooling will be higher than the average return to schooling of the population as a whole. The second interpretation is based on the evaluation of 'treatment effect' (Heckman, 1997). Angrist *et al.* (1996) show that the only treatment effect that IV can consistently estimate is the Local Average Treatment Effect (LATE) that is the average treatment effect (average return to education) for those who change treatment status (educational choice) because they act in accordance with the assignment to treatment mechanism (instrument). Different instruments should estimate different returns to education associated with different subgroups in the population. For example, IV estimates of the returns to schooling based on college proximity as an instrument should be interpreted as the average return to education for a person that acquires an additional year of education only because they live close to college, but would drop out if no college had been nearly.

A common technique of the IV literature is the use of family background variables as instruments. One such instrument is mother's or father's education. Card (1999) observes that there is a tradition of utilising family background data such as mother's and father's educational attainment to control for unobserved ability. There is high correlation between children's' educational attainments and those of their parents and it is widely accepted that educated parents may motivate their child to study harder. Additionally, it is assumed that family background has no direct effect on earnings, but only affects earnings through its effect on education. Using data from 1972-1976 General Social Survey for the US, Card (1999) shows that each additional year of schooling of either parent increases schooling completed by their children by about 0.4 years. He also states that around 30% of the observed variation in education among US adults is explained by parental education.

Family background variables have been used to measure both demand for schooling shift factors and supply shift factors. While the true factors must be independent, family background measures may be correlated with the ability and child's human capital. It is not possible to separate out these effects with the measures available for individuals and families. A further complication in the use of family background variables in earnings equation comes from interacting ability measures with years of education (Fleisher, 1977; Altonji and Dunn, 1994). In these applications, family background variables, as proxies for ability, are used to control for measurable and systematic components affecting the rate of return across individuals (Griffin and Ganderton, 1996). Ashenfelter and Zimmerman (1997) for example, use father's and brother's education as a background variables . The use of parental education as an instrument increases the rate of return by 15% above the corresponding OLS estimates. For men, the addition of parent education to the vector of explanatory variables lowers the OLS estimate of the returns to education by 5% to 10%. For women, the addition of mother's education has no effect on the estimates[18].

Behrman and Taubman (1977) argue that ability and other unobservable effects may be regarded as the consequence of genetic and environmental contribution of the family. If it is assumed that the 'unobserved component' is a pure 'family effect' which captures these genetic and environmental effects, then data on sibling, especially twin data, may be used to control for these unobservables and permit an unbiased estimate of the return to education. The key idea behind this strategy is that some of the unobserved differences that bias cross-sectional comparison of education and earnings are reduced or eliminated within families (see Ashenfelter and Card, 1999). In this case, differences of levels of schooling for the twins or siblings can be exploited to estimate the effect of education on the wage. Behrman and Taubman (1977) argue that since brothers come from the same economic and social background, and presumably differ less in terms of innate ability than typical elementary, high-school and college persons, many kinds of ability often considered important in explaining earning differentials would be held constant. On the other hand, some brothers may become relatively well-educated precisely because of unusual ambition and other kinds of ability rather than because of interest, luck, or other factors uncorrelated with earnings. Card (1999) gives an overview of several studies that use twin data. He concludes that under the assumption that identical twins have identical abilities the within-family estimator gives a consistent estimate for the average marginal returns to education. A drawback of these methods is the possible lack of generalization to non-twins and the potential failure of the identical abilities assumptions for identical twins and siblings. If the assumption does not hold, twin studies might overestimate the effect of education on earnings (Ebber, 2004).

The empirical literature surveyed may be characterized as ranging between two extreme positions, one saying that family effects lead to serious overestimation of

[18] For a more detailed discussion we refer to Card (1999).

the true returns to education, which should be estimated solely from differences between siblings, preferably monozygotic (MZ) twins (Gronau, 2005). The other extreme would take the position that family effect works almost entirely via education and hence causes little bias in the estimated coefficient, and that the decline observed in within-sibling estimates is the result of aggravating other problems, such as errors in measurement and simultaneity, and is not a reflection of true family effects (Griliches, 1979).

Other studies exploit natural variation in data caused by exogenous influences on the educational decision. For example, Angrist and Krueger (1991) exploit natural variation in factors affecting the schooling decision, such as interactions between quarter of birth and compulsory schooling laws, in order to create instruments for education that are uncorrelated with ability. The underlying idea is that a person who has been born early in the year (the first quarter) reaches the minimum school leaving age after a smaller amount of schooling than persons born later in the year. The actual amount of educational attainment is related to the quarter in which they were born, while at the same time there seems no reason to believe that quarter of birth has an independent effect on earnings (Angrist and Krueger, 1991).

Studies in different countries have generally found that increasing the school-leaving age boosts educational attainment, including in Britain (Harmon and Walker, 1995), Canada (Oreopolous, 2003), Norway (Aakvik *et al.,* 2003) and Sweden (Meghir and Palme, 2003). In Germany, the results are more mixed. For instance, Pischke (2004) finds that an increase in school-leaving laws boosted educational attainment for the cohort born 1930-1960. However, it is unknown whether this difference is due to the impact of school starting and leaving ages, or to the age of the two cohorts.

Angrist and Krueger (1991) study has been criticized by Bound *et al.* (1995) with the argument that quarter of birth may have an impact on earnings other than through the effect of education. Moreover, Bound *et al.* (1995) point out that Angrist and Krueger's IV model includes large number of weak instruments and this will result in large inconsistency in the estimates. Bound *et al.* (1995) show that when the correlation between the instrument and endogeneous variable is weak, potential bias arises from any small correlation between the instrument and the error term. This is confirmed by the findings of Staiger and Stock (1997), who compute a variety of asymptotically valid confidence intervals for standard IV and limited information maximum likelihood (LIML) estimates. A second criticism of Angrist and Krueger's findings, raised by Bound and Jaeger (1996), is that quarter of birth may be correlated with unobserved ability differences.

In a later study, Angrist and Krueger (1992) propose the idea that because college enrolment led to draft exemptions, potential draftees for the Vietnam campaign had this exogenous influence on their schooling decision. The instruments are based around numbers assigned on the basis of month and day of birth from which a 'draft lottery' was conducted. As in other studies, the IV results are higher than the OLS, but the difference is insignificant, perhaps reflecting later work that suggested the instruments were only marginally significant in influencing the education decision (Bound *et al*, 1995).

Furthermore, Card (1993) exploits data on proximity to educational institutions assuming that people living near a college are more likely to avail themselves of the facility than someone living farther from college. College proximity may have a direct effect on earnings, since families that place a strong emphasis on education may decide to live near a college, while their children may have higher abilities and motivation to achieve labour market success (Verbeek, 2000). Instruments based on schooling reforms, such as changes in compulsory school attendance laws are biased upwards compared to OLS because of unobserved differences between the characteristics of the treated and non-treated group, since these reform treatments are often non-random (Card, 1999). For example, factors like compulsory schooling or schooling availability are most likely to affect individuals who might otherwise have relatively low educational level. If because of potential heterogeneity these individuals have higher than average marginal returns to education, then instruments based on these variables tend to recover the returns to education for a subset of individuals with higher returns to education, resulting in IV estimates which are larger than the OLS estimates.

A somewhat different approach is used by Duflo (2001), where returns to education estimates are based on the exposure of individuals to a massive investment program in education in Indonesia in the early 1970's. Individuals were assigned to the treatment on the basis of their date of birth and the district in which they lived. Similarly, Denny and Harmon (2001) apply the same approach in Ireland.

Another alternative is to group observations according to childhood smoking behaviour on the grounds that those who choose to smoke at an early age have a higher discount rate in terms of time preference. That is they place more weight on satisfying current wants at the expense of the future (Evans and Montgomery, 1994; Chevalier and Walker, 1999). The intuition for that instrument comes from the acknowledgement that just as schooling is non-randomly assigned across the population, the decision to engage in (un)healthy habits is non-randomly distributed. Evans and Montgomery (1994) note that 'more educated people have better health and better health habits'. There is still a debate as to whether or not this education-health relationship is causal i.e. more educated people learn the dangers of poor health habits and are thus less likely to engage in them. On the other hand, smoking at age 16 is not correlated with current earnings, but is correlated with the educational choice. The decision that an individual makes at age of 16 on whether to continue in education or not is likely to be significantly affected by his/her discount rate, whether that is because of access to financial resources or because of the individual's rate of time preference. The authors find that concurrence in the timing of smoking and school leaving decisions generate a statistically precise and quantitatively large correlation between years of education and early smoking.

An alternative to the instrumental variable approach, as already mentioned, is the study of education and earnings outcomes for twins, siblings or mother-daughter/ father-son pairs[19]. The assumption made in this literature is that differencing eliminates bias due to unobserved ability, but potentially exacerbates measurement error, so that instrumenting differenced schooling eliminates the resulting attenuation towards zero. The weakness in the method is that differencing may not

[19] Since our empirical results are not based on a twin sample, we do not provide detailed discussion on this literature.

remove all of the ability bias if there is some individual component to monozygotic (MZ) ability that is not removed by differencing. Most recent studies on twins follow the Ashentfelter and Krueger (1994) study, which uses the Princeton Twins Survey collected at an annual twin's festival in 1991 in Twinsburg, Ohio. Their results indicate that each year of education completed increases worker's earning on average by 12-16%. Ashentfelter and Krueger (1994) also find weak evidence that unobserved ability may be negatively related to the level of education. The survey of twins by Griliches (1979) concludes that the estimated returns to education, where ability bias is purged via differencing within twin pairs, is typically lower than the estimated return from the whole sample (i.e. without differencing).

2.3.2. Semi-parametric literature

A major feature of the recent literature is the development of alternatives to the Heckman selectivity estimator that do not rely on normality and linearity assumptions. The resulting semi-parametric models combine components of parametric and non-parametric models. Hence, they possess the virtue of easy interpretability of parametric models and the flexibility of nonparametric models (Lei, 2005).The semi-parametric estimates avoid both the assumptions of parametric models and the functional form since they do not assume a known link function, and also solve the curse of dimensionality of nonparametric models[20]. A large number of estimators have been proposed to deal with selection issues. For instance, Newey (1991) proposes a semi-parametric two-step estimator based on a series expansion of the inverse Mill's ratio. He suggests estimating a linear index semi-parametrically to allow for non-Gaussian errors in the first-step regression. An important feature of these new models is that they relax the assumption of joint bivariate normality in the error terms, which is assumed in Heckman's two-step approach. Many of these new models use nonparametric method for binary regression models. They include Manski's maximum score method (1975, 1985), nonparametric maximum likelihood estimation (Cosslett, 1983), weighted average derivatives (Stoker, 1986; Powell, 1987), and kernel estimation (Ichimura, 1988; Klein and Spady, 1993). Tobias (2003) explores the ability-earnings relationships

[20] The precision of non-parametric estimator decreases rapidly as the number of continuously distributed components of independent variables increase.

using flexible econometric techniques that do not require strict functional form assumptions. Estimating the returns to schooling to college and non-college groups, he concludes that there is no evidence of non-linearity in the ability-earnings relationship. Returns to education increase for the more able and the college earnings premium grows over time for individuals at all points in the earnings distribution. Tobias (2003) also finds evidence of diminishing returns to ability for those with 12 or fewer years of education and weak evidence of increasing returns to ability for those with at least some form of college education. He concludes that the most able individuals go to college, while the least able do not, and thus problems necessarily arise when trying to estimate the returns to education over the tails of the ability distribution.

2.3.3. Quantile regression (QR) literature

Since the introduction of the quantile regression (QR) model by Koenker and Basset (1978), several authors have used this technique to estimate the earnings effects of education at different points of the earnings distribution[21]. This technique has an advantage over traditional mean regression techniques in the OLS because it implicitly allows for workers' heterogeneity. More specifically, the idea with the QR is to compare the rate of return at different points in the distribution, for example the bottom with the top. By estimating a family of conditional quantile earnings function one can see how the return to education varies across the whole distribution. For instance, Bushinsky (1994) and Autor *et al.* (2006) for US, Abadie *et al.* (1998) for Spain, Machado and Mata (2001, 2005) and Hartog *et al.* (2001) for Portugal, report that returns to education tend to be increasing when moving up along the earnings distribution. This implies that conditional on observable characteristics, more educated workers display more wage dispersion. Moreover, higher returns for the higher quantiles suggest that more able individuals profit more from additional educational investment. One factor behind the distribution of earnings is the distribution of inherent ability so that 'lower ability' individuals are more likely to predominate in the bottom half of the distribution (Harmon *et al.*, 2001).

[21] In the present thesis we use the words quantile and percentile interchangeable. Note that, for example, a 0.5 quantile is equivalent to 50-percentile, which is the median.

Following Arias *et al.* (2001) if ability and education are complements in the human capital generation process then education has an additional indirect effect on human capital that increases its otherwise constant contribution to earnings. Using British data Harmon *et al.* (2003) find evidence of complementarity based on QR estimates. The return to education is higher for those at the very top of the earnings distribution compared to those at the very bottom. Similarly, Mwabu and Schlultz (1996) use QR methods on a sample of South African men and obtain different returns across quantiles. They interpret an increasing return to education as an indicator of complementarities between ability and education. However, in their study they find that returns to education in Africa decreases across the earnings distribution, which they interpret as substitutability. Elsewhere Heckman and Vytlacil (2001) note that educational sorting by ability may make it difficult to identify the effects of the two variables. Their evidence suggests that ability and schooling appear to be inseparable even if ability is perfectly observed. If educational sorting affects measured ability then this will exacerbate the identification problem. The strategies for coping with these problems leads to different interpretations of the role ability has in explaining increasing returns to education across the distribution.

Similarly, Martins and Pereira (2004) use comparable data and a common wage equation to calculate quantile returns to years of education in sixteen European countries with the purpose of analysing the behaviour of income variations at different points of the conditional distribution. By applying QR technique, they compare the returns for individuals with different ability, attempting to explain the effect of education on wage inequality. Martins and Pereira (2004) find that in most countries education has a positive impact on within-groups dispersion (in line with Buchinsky, 1994). Budria (2006) also exploits the QR technique to calculate returns to lower secondary, upper secondary and tertiary education in nine European countries. As in previous studies, returns to tertiary education in most countries increase when moving from lower to the upper quantiles. In another study, Patrinos et al. (2006) investigate the pattern of returns to education along the conditional wage distribution across a mix of East Asian and Latin American countries. They also analyse the existence of a relationship between a country's development stage and the pattern of returns to education. Their results indicate that returns decrease with quanitles in low-income East Asian countries, while exhibit a clearly opposite pattern in the case of Singapore, which is high-income country. The authors argue that such patterns are due to more job mobility in developed countries, allowing individuals to improve their position by changing jobs.

Works on the returns to education using QR technique in transition countries are limited. For example, Flabbi *et al.* (2008) estimate earnings at different point of the conditional earnings distribution and investigate the heterogeneity among workers in eight transition countries (Bulgaria, Czech Republic, Hungary, Latvia, Poland Russia, Slovak Republic and Slovenia). Their results indicate evidence of heterogeneity in returns to education across countries as well as within countries. In another study, Skoufias (2003) finds that there are no significant differences in the returns to human capital at the median and at the upper and lower tails of the distribution among male and female workers employed in the Romanian public and private sectors.

2.3.4. Previous studies on returns to education in transition countries

Despite the growing number of studies, there are still several important limitations to the research undertaken on the returns to education in post-communist countries (Arandarenko, 2006). Due to the lack of available micro-data relatively very few studies have analysed the returns to education before and after radical transformations in policy. One exception is Andrén *et al.* (2005) who examine the effect of education on earnings under socialism and transition in Romania using retrospective data from 1950 and a 1994 survey. Because of concerns about possible measurement error, the authors estimate the earning function using Least Absolute Deviations (LAD) as well as OLS. Under both estimation methods, they find a small and statistically significant impact of schooling under central planning: a constant 3-4% premium associated with an additional year of education from 1950 through 1989. These results support previous findings that the educational earnings premium was low under central planning and that it grew substantially during the transition years.

Except for Hungary which operated under a considerable less rigid economic regime than most of the rest of Central and Eastern Europe and the former Soviet Union, the returns to education in transition economies was typically around 3-4%. Ten years later the rate of return to education rises to about 8% in most transition countries. Fleisher *et al.* (2005) show that the rapid increase in returns to education during the early reform period reflects the ability of highly educated individuals to respond to changing opportunities in a disequilibrium situation. The authors find that both the speed of economic transformation from a planning to market economy and the degree of economic disequilibrium helped to explain differences in the rate of return to education over time and across countries. Similarly, Bevc (1993) estimates the rate of return to education in former Yugoslavia in 1970s and 1980s. She finds that the rate of return to education in socialist Yugoslavia was lower than in other developing countries, leading to a smaller difference in lifetime earnings by level of education. These results, however, do not necessarily imply that education was of lower importance for the economic development in the former Yugoslavia.

Empirical evidence for Central Eastern European economies, suggests that returns to education increased as the market reforms took place. Table 1.1 represents a summary of some of the main studies. Chase (1998), for example, uses the 1984 Social Stratification Survey for the Czechoslovak Socialist Republic and finds a low return to education of 2.5% for men and 4% for women, which increased to 5-6% by 1993. Similarly the return increased for all levels of schooling except for post graduate education. Filer *et al.* (1999) report additional increases in return in both the Czech and Slovak republics to 8-9% by 1997. Münich *et al.* (2004) show that the majority of the increase in the rate of return to education for women in the Czech Republic took place in the first phase of transition (1986-1996), and that no further change was observable in the later phases (1996-2002). Rutkowski (1996) also reports an increase in the returns to education in Poland, from about 5% in the late 80s to over 7% in the mid-1990s. Using data for Slovenian workers over the 1987-1991 periods, Orazem and Vodopivec (1995) explore changes in the structure of earnings and employment produced by transition. They find that the average return to years of education rises dramatically following transition. Moreover, changes in the relative returns to education are virtually identical for men and women. However, individuals with four years of university education gained the most in relative earnings, closely followed by those with two years of university training. This supports the findings of Abraham and Vodopivec (1993) who argue that higher ability workers are better able to switch jobs. Stanovnik (1997) finds a similar pattern for Slovenia, Jones and Ilaypruma (1994) for Bulgaria, and Maurer-Fazio (1999) for China. A survey by Psacharopoulos (1994), which compares returns to human capital across a large number of countries in the early transition phase, finds estimates of less than 9% in most of the former communist countries.

In a related study, Flanagan (1998) examines the effect of changes in the economic system on the incentive to invest in human capital before and after November 1989 in the Czech Republic. He finds a positive but low return to education prior to 1989. However, the estimated return increase to 5.7% between 1988 and 1996. Halpern and Kōrosi (1997) estimate the return to education using 1986, 1989, 1992, 1993 and 1994 using Hungarian survey data. They find that the estimated return to university education relative to primary (8 years) schooling rises from 45% in 1986 to 62% in 1989 and remains between 56% and 61% thereafter.

Krueger and Pischke (1992) and Bird *et al.* (1994) estimate Mincer earning functions for East Germany before and during the transition. Interestingly, Krueger and Pischke (1992) find that the rate of return to education in East Germany fell from 0.077 in 1988 to 0.062 in 1992. Similarly, Bird *et al.* (1994) use the German Socio-Economic Panel, find that the point estimate of the rate of return to education was 0.044 in 1989, and fell to 0.041 by 1991. This decline is not statistically significant but the rate is significantly lower than the rate of 0.067 found in a 1989 survey: the 1989 Survey of the Federal Republic of Germany.

Other empirical studies provide evidence that education, obtained under central planning, is not necessarily appropriate for the new market environment. Whether the skills acquired under communism can indeed be adapted to the needs of a western market economy is a question of prime importance. Most, albeit not all, empirical studies demonstrate that older generation fared worse during the transition phase. For Poland, Rutkowski (1996) and Puhani (1997) provide evidence of falling return to experience in 1982-1992 and 1992-1995. In the former Czechoslovakia, Vecernik (1995), Flanagan (1995), and Chase (1998) observe steeply declining returns to education in the early transition period. Similar results, can be found in Kertesi and Köllő, (1999) for Hungary; and Steiner and Bellman (1995), Krueger and Pischke (1992) and Burda and Schmidt (1997) for the former German Democratic Republic. A study by Sabirianova (2003) shows that returns to education in transition economies has generally evolved in two distinct phases. An early phase characterising movement away from wage setting development by a wage grid towards payments that better reflect the relative marginal products of skilled and unskilled workers. In the second phase, the relative earnings more closely match and follow the path of relative marginal products of both skilled and unskilled workers (Sabirianova, 2003).

Due to data limitations, very few studies examine the Soviet labour market prior to 1992. Although a Soviet Family Budget Survey has been conducted by the Government Statistics Agency continuously since the 1950s, the results of the survey were unavailable to researchers, and were used primary for the purposes of preparing aggregate data. Most foreign studies of the Soviet labour market are based on surveys in which individuals do not have an incentive to misreport their income. Obviously, the sample characteristics of these surveys tend to differ from those of the overall population (Vernon, 2002). Strumilin *et al.* (1966) provide the first empirical estimates of the return to education in Russia using the discounted earnings procedure. Strumilin's (1966) aim was to advocate an investment program in education. He analysed two samples, one of 2,602 machine operators (from the year 1919) and the other of 2,307 white collar workers, in order to establish a relationship between age, job experience, years of formal education and skill. His results show that the return to skill-level increases with age, reaching a maximum value at the age of 32 years for machine operators and 34 years for white-collar workers.

Later, Gregory and Kohlhase (1988) use a sample of over 2,700 Jewish immigrants to the United States in the period 1979-1982. They report returns to university education in the Soviet Union relative to secondary education of around of 22% and a return to experience of around 2-3%. However, the study has been criticised because of the unrepresentative nature of the data – all of the respondents were either from medium or large cities and they were all Jewish. Another study by Katz

(1999) uses a survey conducted in 1989 of a single city, Taganrog, whose economy depends almost entirely on heavy industry. Katz reports an average 23-35 % return to higher education for men and 14-32 % return for women, which is in line with the results of Ofer and Vinokur (1992) (approximately 29 % for men and 32 % for women). Elsewhere, Nesterova and Sabirianova (1998) use the Russian Longitudinal Monitoring Survey (RLMS) to estimate the relationship between the logarithm of hourly earnings from all jobs on a large set of controls, including supervisory responsibilities, industry, occupations and local labour market conditions. They find an average rate of return to schooling of around 6-8% in 1994-1996. Typically these estimates are higher than the returns to education computed at the beginning of the economic reforms (1992) which tends to support the argument that a transition to market economy shifts returns in favour of education.

Brainerd (1998) finds an increase in returns to education over the period 1991 to 1994. More specifically, it increased from 3.1% to 6.7% for men and from 5.4% to 9.6% for women between 1991 and 1994. Compared to secondary level qualifications, the return to university qualification increased from 8.3% to 21.6% for men and from 5.4% to 9.6% for women over the same period. Newell and Reilly (1999) also estimate an earnings function for Russia using the 1992-1996 RLMS data and report a lower return to a year of university education 4.2% in 1992 compared to higher returns of 9.7% in 1994 and 6.3% in 1996. In a similar context, Münich *et al.* (2000) finds that the communist wage grid resulted in a low rate of return to education in the Czech Republic, but the rate of return increased dramatically during transition. Münich *et al.* (2000) conclude that the communist system succeeded in maintaining small earnings differentials. A later study by Clark (2003) for Russia in the years 1994 to 1998, show a significant and positive return to education and training, which are comparable in magnitude to those in other transition countries. He finds significant returns of 6% to 13 %. Clark (2003) also finds significantly higher returns to education for workers in the non-state sector compared to the state sector.

Cheidvasser and Benitez-Silva (2007) find no improvement in the returns to education in Russia over the period 1992 to 1998. Pastore and Verashchagina

(2006) estimate Mincer earnings functions based on the 1996 and 2001 waves of the Belarusian Household Survey. Their results suggest that despite the more gradual pace of reform, the rate of return to education was not lower in Belarus compared to other transitional countries. Obtaining a university degree provided an increase in monthly earnings from the main job of about 70% in 1996 and 75% in 2001. This amounts to an annual rate of return to tertiary education of between 10.1% and 10.7%, which is higher than in most other transition countries. However, the rate of return to vocational education did fall to 6.3%. Gorodnichenko and Sabirianova (2005) also measure the returns to schooling in Russia and Ukraine between 1985 and 2002. They apply a semi-parametric approach and construct counterfactual earnings distribution for university and secondary education graduates in Ukraine using the distribution of Russian characteristics, returns and unobservable. Their key finding is that after the breakup of the Soviet Union, the returns to education diverge significantly between Russia and the Ukraine. By 2002, the estimated returns to education were almost two times higher in Russia (9.2%) compared to the Ukraine (4.5%).

Relatively few empirical studies exist for Bulgaria, mainly due to poor data. However, Jones and Ilayperuma (1994) find that the return to higher education in Bulgaria increased significantly between 1989 and 1992 for women, while somewhat surprisingly for men no education effect was found in 1989 and only a relatively small effect was found in 1991. The results are unaffected by using monthly or hourly earnings as the dependent variable. Jones and Simon (2005) estimate Bulgarian earnings determination under planned and early transition. Their main finding is that the human capital variables are more important during the early transition phase than during the planning period. For example, the return to a year of education is 5.4% during the transition period but less than 2% at the end of the planning period. Using Living Standards Measurement Surveys (LSMS) for 2002-2003, Arandarenko *et al.* (2006) estimate returns to specific types of educational degree in Bulgaria, Romania and Serbia. Their results indicate that in Bulgaria the return to education rises consistently for both males and females. However, in Serbia and Romania there are higher rewards to education for men than for women. In both of these countries, each successive educational level for men results in ahigher earnings premium, but for women only is statistically significant. While

Bulgarian university degree holders enjoy the highest marginal return, 85.7% for men and 87.9% for women, in Serbia and Romania male university degree holders earn 65% and 60.2% more than those with the lowest educational attainment. In all three countries, males who graduated from vocational schools earn much less than those who obtained general secondary degree, supporting the view that the skills received in such schools are too narrow and poorly suited to the demanded of a modern labour market. Using data spanning the transition period 1995-2003, Kovacheva (2011) finds that the returns to education in Bulgaria increased over time, with younger cohorts receiving higher returns to education.

Relatively few studies estimate the rate of return to education in Tajikistan and Kazakhstan. Some evidence is found in Rama and Scott (1999) who show a positive return to schooling, at least in the early reform period. Using the 1996 Kazakhstan LSMS, Rama and Scott (1999) estimate a rate of return of about 7-8%. Additionally, Verme (2001) estimates sector earnings equation in Kazakhstan and finds that returns to education are 7.4% in the public sector, 3.4% in the private sector and 1.7% for the self-employed. A recent study by Arabsheibani and Mussorov (2007), however, suggests that returns to education in Kazakhstan are higher than the estimates found in Rama and Scott (1999). The results are also higher for women. Specifically, for men, the IV estimates are higher than OLS by around 3%, and for women the return to education rises to 13.7% after correction for both sample selection and endogeneity. The authors conclude that men have more opportunities to earn a higher salary with a lower education compared with women.

Banzragch (2010) uses Mongolian and Tajikistan Living Standard Measurement Surveys to compute the rates of returns to schooling in these two countries. The estimated rate of return to schooling for Mongolia in the early 2000s, ranges from 5.6% to 6.5% for wage earners and over 7% for self-employed individuals. The estimated rate of return to schooling for Tajikistan in 1999 is 9.7% which declines to 4.1% in 2003 and 2007. Banzragch (2010) indicates that these rates of return are lower than the results found for other transition economies and argues that the decline in the estimated return to schooling in Tajikistan was the result of devastating civil war.

With the exception of East Germany, Tajikistan and to some extent men in Bulgaria, the studies reviewed in this section, suggest a positive and increasing rate of returns to education during the transition period. The studies also indicate that women enjoyed a higher rate of return on education than men under the communism. However, relatively few studies have examined returns to education across the entire earnings distribution in Bulgaria, Serbia, Russia, or Tajikistan. There is a gap in the literature, therefore, which this thesis in some part attempts to fill.

TABLE 1.1

Summary of the main studies on returns to education in transition economies

Country	Year	Returns to education	Men	Women	Reference
Belarus	1996	0.100***			Pastore and Verashchagina (2006)
Belarus	2001	0.101***			Pastore and Verashchagina (2006)
Bulgaria	1985-1995 pooled data		0.040***	0.057***	Trostel *et al.* (2002)
Bulgaria	1989-1992 pooled data	0.040***			Jones and Simon (2005)
Bulgaria	1995	0.046***			Kovacheva (2011)
Bulgaria	1997	0.048***			Kovacheva (2011)
Bulgaria	2001	0.058***			Kovacheva (2011)
Bulgaria	2003	0.071***			Kovacheva (2011)
Czech Republic	1988	0.044***	0.034***	0.054***	Flanagan (1995)
Czech Republic	1989	0.039***			Münich *et al.* (2004)
Czech Republic	1985-1995 pooled data		0.035***	0.043***	Trostel *et al.* (2002)
Czech and Slovak	1997	0.080***			Filer *et al.* (1999)
Czech Republic	1984		0.025***	0.040***	Chase (1998)
Czech Republic	2002	0.070***			Münich *et al.* (2004)
East Germany	1988	0.077***	0.071***	0.085***	Krueger and Pischke (1992)
East Germany	1989	0.044***			Bird *et al.* (1994)
East Germany	1985-1995 pooled data		0.026***	0.045***	Trostel *et al.* (2002)
Hungary	1988		0.047***	0.061***	Köllő (2003)
Hungary	1986	0.062***			Jolliffe and Campos (2005)
Hungary	1989	0.074***			Jolliffe and Campos (2005)
Hungary	1992	0.095***			Jolliffe and Campos (2005)
Latvia	1985-1995		0.067***	0.078***	Trostel *et al.* (2002)
Poland	1985-1995 pooled data		0.073***	0.100***	Trostel *et al.* (2002)
Russia	1985	0.028***			Gorodnichenko and Sabirianova (2005)
Russia	1985-1995 pooled data		0.044***	0.053***	Trostel *et al.* (2002)
Russia	2002	0.092***			Gorodnichenko and Sabirianova (2005)
Slovakia	1985-1995 pooled data		0.052***	0.064***	Trostel *et al.* (2002)
Slovenia	1985-1995		0.080***	0.101***	Trostel *et al.* (2002)
Ukraine	1986	0.034***			Gorodnichenko and Sabirianova (2005)
Ukraine	1987	0.037***			Gorodnichenko and Sabirianova (2005)
Ukraine	2002	0.045***			Gorodnichenko and Sabirianova (2005)
Kazakhstan	1996	0.070***			Rama and Scott (1999)
Kazakhstan	2001		0.080***	0.115***	Arabsheibani and Mussorov (2007)
Tajikistan	1999	0.097***			Banzragch (2010)
Tajikistan	2003	0.041***			Banzragch (2010)
Tajikistan	2007	0.041***			Banzragch (2010)
Mongolia	2003	0.065***	0.058***	0.076***	Banzragch (2010)

Notes: ***Statistically significant at 1%, ** statistically significant at 5%, * statistically significant at 10%;
The rates of returns are based on the OLS estimates.

2.4. Summary and conclusions

The analysis of returns to education remains a highly discussed topic in labour economics. Interest comes from a number of potentially disruptive biases associated with conventional OLS estimates of the returns to education. In this chapter, we have examined several estimation strategies used to estimate the rates of return to education and the main problems associated with these estimators. The chapter also provides an overview of the empirical literature related to the rate of returns to education with a special focus on the returns to education in transition countries. More specifically, a number of conclusions can be drawn based on this review. They can be briefly summarized as follows:

• The specification is often estimated by OLS, which suffers from a number of potential biases that can either exert downward or upward pressure on OLS estimates.

• In conventional OLS analysis a potential source of bias is the assumption that the number of years an individual spends at school is exogenous (or determined by factors outside the model). However, if schooling results in higher earnings, higher earnings may in turn result in an increase in the amount of education acquired which is more likely the higher the ability of the individual.

• The human capital literature focuses on ability bias as one of the central problems estimating rates of return to education. If people of higher ability have the capacity to earn more (at a given education level) and if they also tend to acquire more education than others, the failure to take ability differences into account has two consequences: (a) OLS estimates overstate the gross contribution of education to earnings, and (b) understate the opportunity cost of foregone earnings to high-ability persons. As most data do not contain information on ability, we have an omitted variable bias problem. One solution to the ability bias problem is the method of instrumental variables (IV).

• Finding an instrument that identifies demand-side variation in education and is uncorrelated with individual earning is difficult. A suitable instrument must meet the relevance and exogeneity conditions. The relevance condition requires the instrument to be strongly correlated with the number of years of education an individual receives. The exogeneity condition requires that the instrument affects earnings only through the channel of education, and therefore the instrument is uncorrelated with the error term in the earnings equation.

• In estimating earnings function there is an additional complication – the problem of sample selectivity. The sample selection bias in the classic Mincerian equation is extensively examined in the literature, especially when analysing womens' return. However, correcting for sample selection bias by identifying satisfactory restrictions, sometimes substantially changes the estimated rates of return to education.

• Human capital theory implicitly recognises that returns to education may be heterogeneous. Inter alia educational returns can vary across schooling levels and even across individuals with the same schooling level. Evidence based on a QR estimates show that returns are higher for those in the top quantiles of the earnings distribution compared to those at the bottom.

• Evidence from transition studies suggests that returns to education increased during the transition process and that the sharpest increase took place during the early transition period.

• While there are many studies that address these issues in estimating returns to education in developed countries, especially in the US and OECD countries, few studies take into account 'ability' and endogeneity bias for the transition countries. Studies that present results for the pre-transition period are also limited.

• The review of existing studies highlights several issues which are important for empirical research on the returns to education. First, in order to obtain consistent estimates we need to allow for both the endogeneity of education and sample selection problem. Second, it is crucial for the analysis, to address the extent to which returns to education vary across the earnings distribution. Finally, it is of interest to draw conclusions about differences in the rates of return to education across the countries analysed.

CHAPTER THREE

OVERALL BACKGROUND AND EDUCATION IN THE REGION

3.1. Introduction

The economic effects of transition have been as diverse as there are countries which have experienced the transition process. Some of these countries have enjoyed the benefits of industrial growth and comparatively high level of initial GDP. Others have enjoyed none of these benefits and have struggled to avoid chronic economic stagnation (Johnes, 2002). With the exception of Serbia, the countries selected for analysis in this thesis are those that commenced their economic transition to a marked-based economy between 1989 and 1992. They adopted a shock-therapy approach to transition, characterized by fast changes which did not necessarily lead to improved economic performance (Hung, 2002).

The current Chapter provides a brief insight of the economic development and educational systems under communism and the early transition period in Bulgaria, Serbia, Russia and Tajikistan. Section 3.2 provides a brief background of the Bulgarian, Serbian, Russian and Tajikistan economies using a number of indicators of macroeconomic and labour market performance. The section sheds some light on labour market adjustment over the transition period, with particular emphasis on the dynamics of employment, including unemployment and labour force participation, and real wage growth. Section 3.3 describes the main features of the education systems and Section 3.4 concludes.

3.2. Overall background

The structure of a planned economy differs significantly from a market economy. At the beginning of transition, the labour markets of the former centrally planned economies were characterized by full employment, job security, centrally set wages and prices, and output targets for state-owned enterprises (Svejnar, 1999). In particular, the economies of the Soviet Union and Eastern Europe were guided by a principle of distribution according to contribution not need (Atkinson and Micklewright, 1992). Full employment, however, was achieved at a cost of low wages. Low wages and limited income differentials failed to motivate workers to improve their performance. The private sector was practically non-existent or played only a limited role in most countries of the Eastern Bloc (e.g. in Hungary or Bulgaria) (Nešporová, 1999). However, most of the countries selected for analysis in this thesis began the transition period with a relatively good level of human capital development[22] (Paci and Reilly, 2004). But labour market conditions changed rapidly after the introduction of reforms between 1989 and 1991.

The transition process in East and Southeast Europe commenced in the early 1990s with the collapse of the Soviet Union. It implied the end of the economic system that had been in force for over four decades: the planned economy. Progress towards a new economic system, a market economy, was initiated. The main objectives were to create a decentralized and incentive based economy, in which the private sector dominated the state sector (Stålhandske, 2007). One of the central aspects of transition was the process of structural reallocation of production among sectors of the economy and industries. The market reforms implemented across the region included wage and price liberalization, trade liberalization and privatization of state-owned enterprises. Transition induced major shifts in demand and supply conditions. The old centralized wage setting system was replaced by new arrangements. These ranged from decentralized plant level negotiations – in Russia – to collective bargaining in Eastern Europe (Paci and Reilly, 2004).

While market reforms intended to raise economic efficiency, the political and economic changes, or what is now called the transition process, involved significant costs in the form of lost output and unemployment, widening income inequality, the growth of poverty and changing demand for jobs. The market reforms implemented across the region contributed to gender wage inequality, increasing unemployment and deterioration in health care, pensions, and in the education system. Economic transition affected the level of earnings, as well as the level of relative wages. As a result of demand and supply shocks that accompanied the reforms, real wages in all

[22] The human capital indicators (for instance, gross and net enrolment rates, average years of schooling and literacy levels) tend to be better in the late 1980s (Beirne and Campos, 2007).

transition economies experienced a substantial fall, along with the GDP (Rutkowski, 1996). In addition, the transition process in Central and Eastern European countries exposed serious weaknesses in educational systems. The direction and pace of market reforms have been different across countries, which also contributed to more diverse labour market institutions in the region (Word Bank, 2001). According to Gimpelson and Lippoldt (2001) one of the reasons for this large-scale disorganisation was the collapse of bureaucratic coordination while market signals were still weak and imperfect. Firms were left to face greater uncertainty about prices and wages.

The speed of transition varied widely among countries, with those that implemented reforms more quickly, generally experiencing more rapid recovery. Table 3.1 shows some of the changes in the Eastern European economies documented by the European Bank for Reconstruction and Development (EBRD). The first column gives the value, in 1997, of the EBRD's 'transition index', which summarises countries' cumulative progress from a planned to a market economy. The index takes into account a variety of different dimensions of transition – price liberalisation, privatization, restructuring, competition policy and reforms in financial institutions (Flemming and Micklewright, 1999).

According to the EBRD's index of progress, the most advanced countries in 1997 were those in the Central Europe and the Baltic region[23]. These countries were more successful in achieving a high private sector share of GDP and recovered more rapidly from the shock of transition. In contrast, countries such as Bulgaria and Albania lagged behind. The least progress was made by the Central Asian republics. The slow reformers experienced the largest fall in output over the period 1989-1997, which was accompanied by a high rate of inflation. The relationship between the pace of transition and labour market developments was complex. While the labour markets in Central and South-Eastern Europe were characterised by high unemployment, in the former Soviet republics unemployment was at lower levels. In Hungary, for example, which is a relatively successful model of transition, the unemployment rate peaked in 1992 because job creation in the

[23] All the countries in these two groups had a value of transition index greater than 3.0 and on average the private sector accounted for over two-thirds of the GDP.

growing sectors was initially too slow to offset job losses in the restructuring sectors, but it has since declined gradually to single digits. However, in the countries of the Commonwealth of Independent States (CIS), which lagged in the reform process, employment remained stable during the early 1990s even when there was a significant collapse in output (Schiff *et al.*, 2006).

TABLE 3.1: Indicators of transition

	EBRD transition index	Private sector share of GDP 1997 (%)	Change in real GDP 1989-97 (%)	Average annual inflation 1991-96 (%)	Average registered unemployment 1991-96 (%)	Government expenditure as share of GDP 1996 (%)
Central Europe	3.4	68	-1	34	10.6	46
SE Europe	2.7	59	-30	179	17.3	37
Baltics	3.2	67	-37	254	4.9	38
Western CIS	2.4	46	-50	775	1.3	38
Caucasus	2.4	50	-63	1926	3.1	19
Central Asia	2.2	41	-42	758	1.6	24

Notes: Central Europe – Czech Republic, Hungary, Poland, Slovakia and Slovenia, South East (SE) Europe – Albania, Bulgaria, FR Yugoslavia, Macedonia and Romania; Baltics – Estonia, Latvia and Lithuania; Western CIS – Belarus, Moldova, Russia and Ukraine; Caucasus – Armenia, Azerbaijan and Georgia; Central Asia – Kazakhstan, Kyrgyzstan, Tajikistan, Turkmenistan and Uzbekistan. The figures in the table are all unweighted averages of the data for each country.

Source: EBRD (1997) for the transition index and private sector share, EBRD (1998) for GDP change and government expenditure shares, and UNICEF (1998) for inflation and unemployment.

Under state planning regime, the school system and curricula were subject to rigid control by central governments. Enrolment targets were set by planners to serve the needs of industry. Secondary and tertiary education, were strongly influenced by industrial ministers. Employment was guaranteed for all recent graduates, with party affiliation more important than academic achievement for many jobs (Laporte and Ringold, 1997). A natural question raised in the literature is whether the transition process has produced a worsening in educational opportunities in the former socialist countries and whether the extent of differences in educational reforms at the end of the 1990s is a cause of major concern (Micklewright and Nagy, 1999).

3.2.1. Bulgaria

Located in the Balkan Peninsula, modern Bulgaria is one of the smaller countries in Europe, with a population of approximately 7.6 million and an area of 43,000 square miles. Bulgaria shares borders with Serbia, Macedonia, Romania, Greece, and Turkey. Its transition to a democracy and market economy after the collapse of communism has not been easy and standards of living have been low. Since the onset of the reforms in 1989 Bulgarian society has undergone profound political, economic and social changes[24].

Bulgaria began transition to a market economy in particularly unfavourable conditions, which combined with a number of serious external shocks – a collapse of trade with the former Soviet Union, on which the country was heavily dependent; a need to import energy at world prices to replace Soviet supplies; a lack of access to commercial credit as a result of the unilateral moratorium on external debt declared in March 1990; problems resulting from the international embargo on trade as a consequence of the Gulf and Yugoslav wars; and an exceptionally high external debt (127% of GDP in 1990).

In early 1991 a package of fiscal and monetary policies were introduced similar to a Polish style 'big bang shock therapy' to open and stabilize the economy (World Bank, 1996). During the 'big bang shock therapy' period, Bulgaria implemented a wide range of policies: prices were liberalized; an independent central bank and commercial banks were created from the old state banking system; the central bank implemented a tight monetary policy and introduced a managed float of the exchange rate by making it partially convertible; most restrictions on imports and exports were removed; and state enterprises were de-concentrated and de-monopolized (Levinson, 1995).

However, the first phase of economic transformation was characterised by the slow implementation of reforms and inconsistent economic policies. The deep recession in the initial phase was followed by a weak and fragile recovery in 1994-95, only to set the stage for a new and deep recession in 1996 (Dobrinsky, 2000). In 1996 and 1997 banking and currency crises increased the country's internal debt (Barlemann and Nenovsky, 2003). Like other countries in transition, Bulgaria experienced a huge one-time price jump as prices were initially liberalized and real output fell. A doubling of consumer price inflation in 1996 was followed by a nine-fold increase

[24] Over the period of 13 years, there were nine government changes.

in 1997, which was one of the highest inflation rates among the transition economies. High inflation rates can be attributed to the price reforms, particularly the freeing of prices and budget deficits. The introduction of a value-added tax in 1994 also contributed to a high inflation rate. Another factor was the devaluation of the Bulgarian currency by nearly 100% in 1994. One important aspect of the exchange rate regime in Bulgaria at that stage of its transition was its instability. The government lost all credibility and, as a result, the Bulgarian lev depreciated by 33% in November 1996, by 39% in December 1996, by 110% in January 1997, and by a further 187% in the first two weeks of February 1997. As a result, for approximately 300 days between April, 1996 and February 1997, the Bulgarian lev depreciated by 3500% against the dollar (Ganev, 2001).

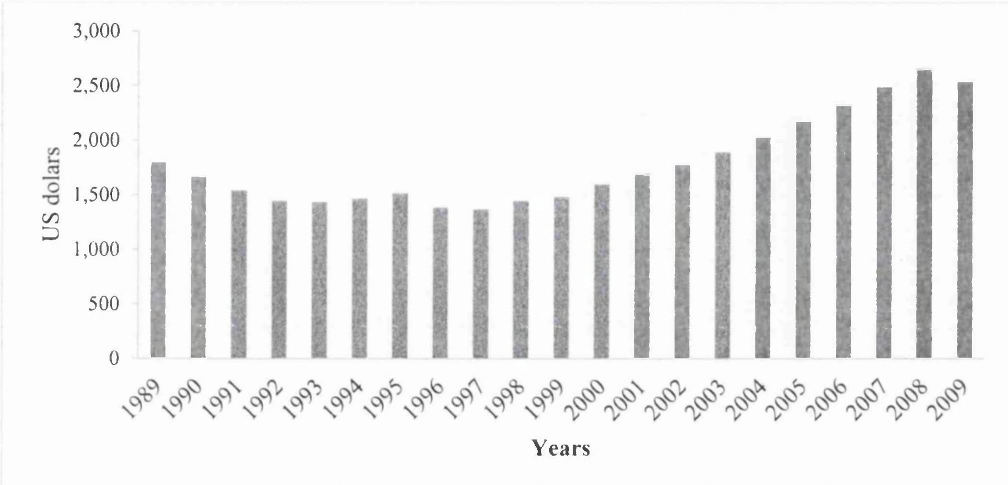
Relative economic stability was achieved with the introduction of a Currency Board arrangement in 1997[25]. This arrangement along with the economic reforms was crucial for stabilizing the economy. The introduction of the Currency Board led to a rapid decline in the rate of inflation, falling to a single digit (Valev, 2004). Relatively stable prices, privatization, and increased competition in financial markets contributed to an increase in investment levels. For the first time since the beginning of transition, Bulgaria experienced positive economic growth after 1997. However, despite the positive performance of some economic indicators, the recovery was weak. The industrial recession was still present during the first months of 1998 and the overall level of privatization was still very slow.

The level of income, using GDP per capita, provides a better understanding of the Bulgarian economy. Figure 3.1 indicates that during the initial stage of the transition to a market economy, GDP per capita declined by 20% in the first three years but after 1997 the trend was reversed due to privatization, fiscal policy and structural reforms. The industrial output showed some improvement in late 1994 and early 1995. The main reason for the decline was that reforms continuously weakened uncompetitive industries. Another reason was the slow growth of the private sector. In 1994 the share of the private sector in the Bulgarian GDP was

[25]The Currency Board is a system that does not allow the Central Bank to implement its own, independent fiscal and money supply policy.

40%, compared to 65% in the Czech Republic, 55% in Poland, Hungary, and Slovakia, and 35% in Macedonia and Romania. The massive privatization process and enterprise liquidation resulted in 75% of GDP being produced in the private sector by 2001. GDP per capita continued to rise thereafter, reaching about 2,661 US dollars in 2008.

FIGURE 3.1: The Bulgarian GDP per capita, 1989-2009

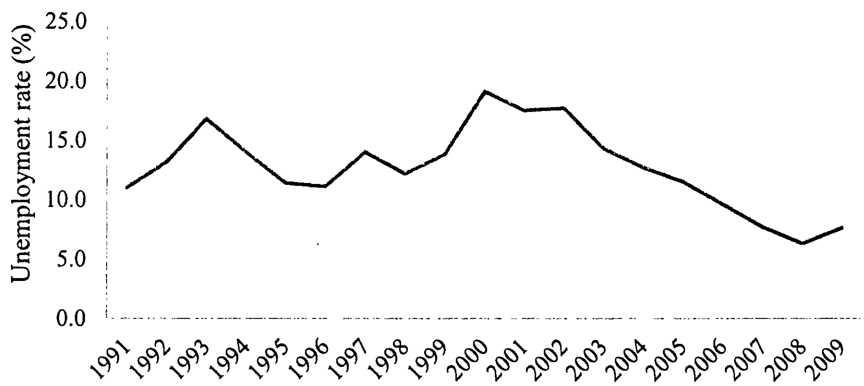


Source: World Bank, World Development Indicators database.

One of the characteristic features of the transition from a planned to a market economy in Central and Eastern European countries (CEEC) was a reduction in employment and the emergence of a high unemployment. This was especially true for Bulgaria, where the number of employed was reduced by more than one-third from the beginning of the reforms. Table 3.2 provides detailed information on the scale of employment losses and employment reallocation in the country. Registered unemployment in Bulgaria was among the highest of any of the Eastern and Central European countries in transition (Beleva *et al.*, 1995). Specifically, the unemployment rate rose from negligible levels pre-1992 to over 11% in 1991, over 15% in 1992, and 19% in 2001 (see Figure 3.2 and Table 3.2). The unemployment rise in 1999 was due to privatizations and restructuring of public enterprises (Beleva and Tzanov, 2001). In addition, long-term unemployment was a major problem. Of the unemployed, the share that has been out of work for more than 1 year increased from 53% in 1993 to nearly 66% in 2002. The unemployment rate among those in the lowest category of educational attainment was nearly 45% (Beleva *et al.*, 1995).

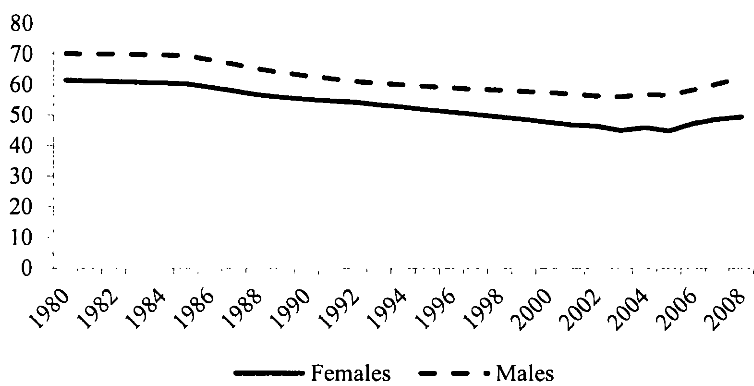
Overall, the economic recession during the 1990s led to a sharp decline in the labour force participation of both men and women. According to the International Labour Organization (ILO) key indicators database, the participation rates for men in Bulgaria decreased from 70% in 1980 to nearly 56% in 2003 and from 61% in 1980 to 45% in 2003 for women (see Figure 3.3).

FIGURE 3.2: Annual registered unemployment rates in Bulgaria, 1991-2009



Source: Bulgaria National Statistics Office

FIGURE 3.3: Labour participation rate in Bulgaria, 1980-2008



Notes: % of population ages 15 and over.
Source: ILO, Key Indicators of the Labour Market database.

The transition period has also been characterized by large shifts in a labour between the public and private sectors. Private sector employment in Bulgaria has grown rapidly, both in absolute terms and as a share of total employment. By 2002, the private sector accounted for 62.5% of total employment (see Table 3.2)[26].

[26] More discussions on the sector employment will be given in the Section 9.2.1.

TABLE 3.2: Employment and unemployment structure in Bulgaria

	1993	1994	1995	1996	1997	1998	1999	2000	2001	2002
Working age population[1]	6,881	6,887	6,899	6,903	6,903	6,924	6,890	6,890	6,779	6,740
Economically active [2]	55.4%	52.9%	51.8%	51.7%	51.8%	51.0%	49.5%	48.8%	49.6%	49.4%
Employed [3]	43.5%	42.2%	43.3%	44.4%	44.3%	43.8%	41.7%	40.6%	39.8%	40.7%
Of which[4]:										
Private sector	22.4%	25.8%	28.2%	31.3%	37.0%	43.3%	39.6%	53.0%	59.5%	62.5%
Youth[5]	9.4%	9.2%	8.8%	8.4%	8.1%	8.7%	8.6%	8.0%	8.0%	7.7%
Females	46.6%	46.7%	46.9%	47.0%	46.8%	46.9%	46.8%	46.8%	47.8%	47.5%
Employees	89.7%	89.4%	88.8%	88.6%	86.2%	86.1%	86.7%	84.9%	84.7%	85.2%
Private enterprises	12.7%	17.2%	19.4%	22.7%	27.3%	34.5%	30.7%	45.2%	52.7%	56.5%
Employers	1.7%	1.6%	1.8%	1.8%	1.9%	2.1%	2.3%	2.4%	3.5%	3.3%
Self-employed	8.1%	7.6%	8.1%	8.1%	9.7%	9.9%	9.3%	10.8%	9.8%	9.5%
Unemployed (in thousands)	815	737	590	505	513	497	534	567	664	592
in % of economically active	21.4%	20.2%	16.5%	14.1%	14.4%	14.1%	15.7%	16.9%	19.7%	17.8%
Of which[6]:										
Youth	30.5%	28.0%	28.9%	26.9%	25.8%	25.1%	23.7%	21.6%	21.2%	20.4%
Females	48.3%	46.8%	48.1%	46.8%	47.0%	45.9%	64.8%	45.9%	45.1%	44.7%
Long term[7]	52.5%	59.0%	64.1%	61.7%	58.4%	57.0%	54.3%	56.8%	61.8%	65.7%
Basic education[8]	47.3%	46.8%	48.1%	45.4%	40.5%	40.5%	40.6%	38.0%	35.7%	35.6%
Secondary education	45.1%	46.0%	45.4%	47.6%	51.7%	50.9%	51.9%	53.2%	53.7%	53.6%
Tertiary education	7.6%	7.2%	6.5%	6.9%	7.9%	8.6%	7.5%	8.8%	10.6%	10.8%
Urban areas	66.0%	64.7%	59.4%	67.2%	69.3%	69.1%	67.6%	67.9%	68.8%	70.5%
Unemployment by reasons										
Lost jobs[9]	55.8%	58.5%	58.7%	57.0%	56.9%	54.6%	55.8%	59.6%	45.1%	43.4%
Temporary job terminated	5.0%	6.4%	7.9%	10.1%	9.7%	12.4%	13.8%	12.1%	7.6%	8.9%
Seeking first job	21.4%	21.0%	21.8%	21.4%	21.6%	20.8%	19.5%	18.8%	23.4%	23.8%
Unemployment rates by age										
Age 15-19	66.9%	64.7%	57.4%	51.5%	49.9%	45.9%	50.4%	52.9%	58.7%	54.1%
Age 20-24	40.5%	38.2%	34.4%	30.5%	31.0%	28.9%	30.6%	31.9%	35.3%	33.1%
Age 25-49	17.7%	17.3%	13.9%	12.2%	12.5%	12.3%	14.0%	15.0%	17.7%	16.2%
Age 50 and over	15.6%	14.7%	11.0%	9.2%	9.6%	10.2%	11.5%	13.8%	16.5%	14.5%
Economically inactive[10]	3,071	3,244	3,325	3,332	3,330	3,392	3,480	3,529	3,416	3,408

Notes: [1] Age 15 and above (in thousands); [2] % of working-age population; [3] % of working-age population; [4]Percentage share in total employed; [5] Age 15-24; [6] Percentage share in total unemployed; [7] One year and over; [8] International Standard Classification of Education (ISCED); [9] From 2001 data refer to unemployed who lost job within past eight years; [10]In thousands;

Source: National Statistics Institute of Bulgaria

The very high inflation level in 1997 further contributed to the decline in real wages caused by a collapse in output. Bulgaria was one of the countries that suffered the largest cumulative fall in real wages among all transition economies of Eastern Europe. Real wages in 1998 reached 40% of their level in 1989, with a cumulative fall of 60% (Garibaldi *et al*, 2001b). Changes in real wages were mainly due to inflationary shocks, in particular in 1991, 1994, and 1996-1997 (see Table 3.3). During these years, nominal wage growth failed to compensate for increase in the inflation rate.

FIGURE 3.4: Real wage growth in Bulgaria (1989=100)

Source: Garibaldi *et al.* (2001b).

TABLE 3.3: Real wages and real gross income per capita, Bulgaria, 1990-1999

Year	Nominal wage (yearly changes)	Inflation annual changes	Real wage[a] yearly change	Real gross income per capita (1989=100)
1990	37.9	23.9	5.3	96.2
1991	167.8	338.5	-39.0	59.0
1992	102.3	79.4	12.8	63.5
1993	57.8	56.1	1.1	60.4
1994	53.5	87.1	-17.9	53.6
1995	53.4	62.1	-5.5	49.0
1996	89.2	123.0	-17.6	33.5
1997	884.1	1087.8	-18.8	28.4
1998	47.0	22.3	22.3	36.9
1999	11.9	2.6	9.1	37.1

Source: NSI, Statistical Yearbook.
Notes: [a] Calculated as the ratio between nominal wages and inflation.

Despite recent labour market improvements, in 2008 the employment rate in Bulgaria was approximately 2% point lower than the EU-27 average and 6% point lower than the Lisbon target (a 70% employment rate) (World Bank, 2007). The country still lags far behind not only the countries of the Western Europe but also the more developed transition economies such as Hungary and Poland.

3.2.2. Serbia

Located between Central and Southern Europe, Serbia is located in the Balkan Peninsula sharing borders with Bulgaria, Romania, Hungary, Bosnia-Herzegovina, Croatia, the Republic of Macedonia and Montenegro. Serbia has a population of 7.5 million which is predominantly Serb, but with a significant and ethnically diverse minority[27]. The largest minority group, according to 2001 Census data, are Hungarians (3.9%), Roma (1.4%), Croats (0.9%) and Albanians (0.8%). Despite fast economic growth attributed to the far-reaching economic reforms introduced after October 2000, Serbia still struggled to overcome the economic and social damage suffered during the 1990s. Serbia began the 1990s relatively well integrated with the world economy, and with a higher standard of living than many other transition economies. However, armed conflicts, international sanctions, and trade shocks caused by the break-up of the Socialist Federal Republic of Yugoslavia (SFRY) during the 1990s, had a detrimental effect. Combined with economic mismanagement, these events resulted in hyperinflation and a virtual collapse of the economy by the end of 1993 (Arandarenko, 2007).

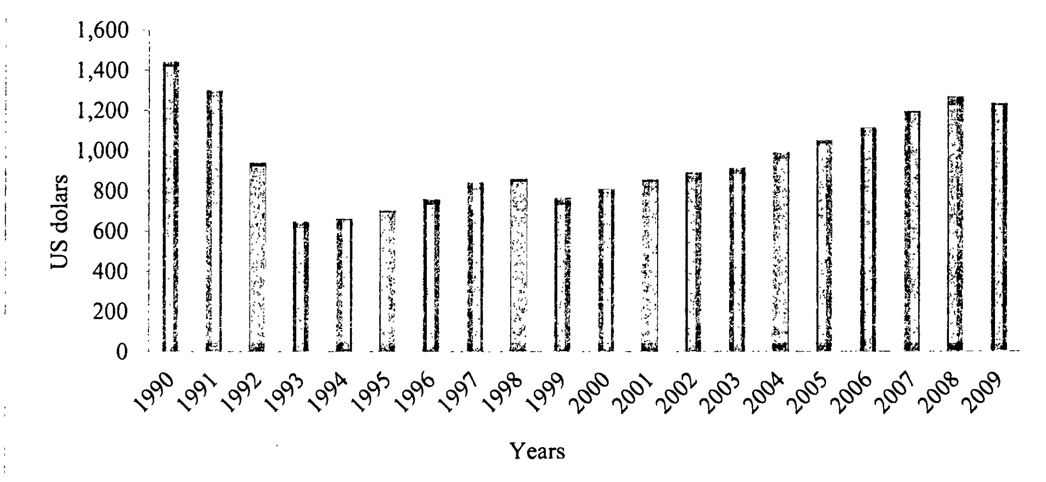
Like other countries in transition, Serbia's economy suffered the loss of markets due to the break-up of the unified Yugoslav market. The economic system in the Milošević-era (1991-2000) was characterized by a war economy, a large informal sector, 'black' markets, an oligarchy of tycoons controlling large state-owned and private business enterprises, state-driven hyperinflation, state-sponsored smuggling, client-patron relationships, state-backed pyramidal schemes, inefficient fiscal system, and lack of investments (Brunhart and Gajić, 2005). The share of the 'grey' economy during the unprecedented hyperinflation of 1993 exceeded 50% of GDP (Reilly, 2003). Reilly and Krstic (2000) attributed the huge rise in informal activity to a combination of declining income pushing individuals into informal activities, and to soft penalties imposed by the government. This means that most of those employed in the informal sector held a formal sector full-time job and used the informal sector to supplement their declining incomes (Kecmanovic, 2009).

In terms of economic policies, the government introduced a number of measures to deal with the recession. In 1992 price controls on many products were introduced, and a law was passed forbidding firms to fire employees. In 1993 a fall in output of more than 30% was experienced and hyperinflation took hold, with the inflation rate at one point measured in billions of percentage points. This was the second highest inflation rate recorded in monetary history. A stabilization package,

[27] 83% of total population are Serbs (European Training Formation, 2006).

introduced in 1994 brought inflation down below 20% (Crnobrnja and Savić, 2008). The dramatic fall in the GDP that occurred during the break-up of Yugoslavia and the imposition of sanctions can be seen in Figure 3.5. The decline in GDP per capita was far deeper than the average for transition countries and reached a low point of 650 US dollars in 1993.

FIGURE 3.5: The Serbian GDP per capita, 1989-2009



Source: World Bank, World Development Indicators database, accessed March 2011.

Since January 2001, Serbia has undertaken an ambitious programme of economic reforms, including macroeconomic stabilisation, liberalisation of foreign trade and domestic prices, and restructuring of the banking system (Paunovic *et al.*, 2005). The structural changes in the economy have put significant pressure on the Serbian labour market. Accurate data on economic activity, employment and unemployment in Serbia is difficult to get because of the extent of the informal sector. The main labour market trends can be identified using data from the Labour Force Survey (LFS). The participation rate (employed and unemployed as percentage of active labour force aged 15 and 64 years old) reached 65.2% in 2005 (European Training Forum (ETF), 2006). The employment rate however, has been constantly declining over the last few years. The employment rate of 50.8% in 2009 was significantly below the EU-15 average, but also below the rates of Croatia and those of the EU members Bulgaria, Romania and Slovenia (see Table 3.4). The 2006 ETF country profile report shows that large differences exist in the employment rates for groups with different educational attainment. For instance, the employment rate is 42.8%

for low-skilled, 58.2% for medium-skilled and 77.8 % for people with higher education.

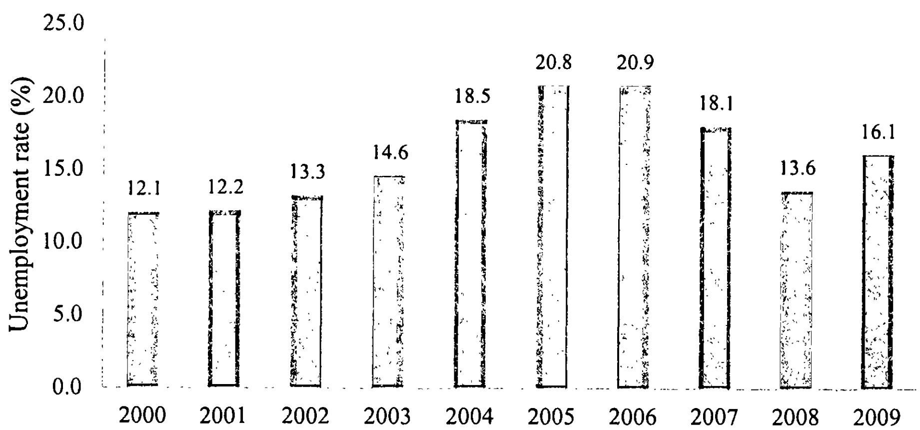
TABLE 3.4: Employment rates in South East Europe

	2000	2001	2002	2003	2004	2005	2006	2007	2008	2009
Serbia	50.1	50.3	48.6	47.6	45.2	-	49.8	51.5	53.3	50.8
Croatia	53.2	51.6	52.9	53.4	54.9	55	55.6	57.1	57.8	56.6
Kosovo	-	-	23.8	25.3	27.7	25.7	25.8	23.7	21.8	-
Bulgaria	-	-	-	53.1	55.1	55.8	58.6	61.7	64	62.6
Romania	64.2	63.3	58.6	58.7	58.7	57.6	58.8	58.8	59	58.6
Slovenia	62.7	63.6	64.3	62.5	65.6	66	66.6	67.8	68.6	67.5
EU-15	63.2	63.9	64.2	64.4	64.6	65.4	66.2	66.9	67.3	65.9

Source: Eurostat; wiiw Database incorporating national statistics.
Notes: Total employed in % of working age population 15-64.

Figure 3.6 shows that the unemployment rate in Serbia remains an important economic and social problem. It increased from 18.5% in 2004 to 20.9% in 2006 and despite economic growth it has remained high.

FIGURE 3.6: Annual registered unemployment rates in Serbia, 2000-2009



Source: World Bank, World Development Indicators database.

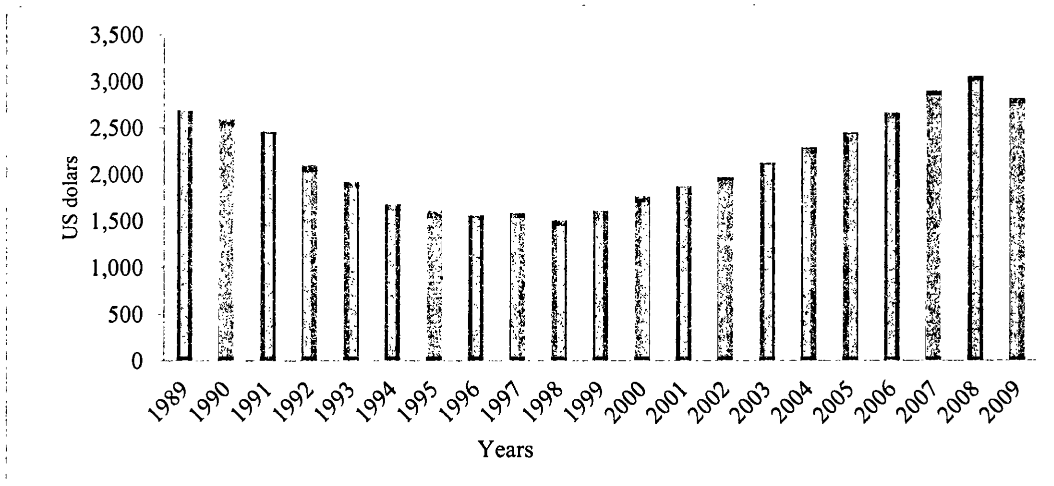
3.2.3. Russia

Russia is one of the most interesting countries that moved from central planning to a market economy not only because it is the largest and most powerful of the group of former communist countries, but because it practiced central planning far longer than any other country (Schwarz *et al.*, 2002). Nowadays Russia has a population of 146 million and a land area of 17 million square kilometres. Over the period 1992 to 2000, Russia's population decreased by 2.8 million, representing over 2% of the total. This decrease was the result of emigration and falling birth rates, which in turn reflected the economic crisis in the country (Bray and Borevskaya, 2001). Prior to the dissolution of the Union of Soviet Socialist Republics (USSR) in December 1991, the population of Russia enjoyed virtual job security and the official rate of unemployment was zero. However, subsequent economic transformations have led to an increase in the number of unemployed as well as under-employed individuals (Foley, 1997). Prices were liberalised, precipitating high inflation, and a mass privatization programme was embarked upon.

The Russian recession was deeper and longer than those in most Central and Eastern European (CEE) countries (Gimpelson and Lippoldt, 2001). The first seven years of Russia's transition from a centrally planned economy (1991-1998), which coinciding with the regime of President Boris Yeltsin, were a time of economic chaos (Cooper, 2009). There was a rapid decline in GDP, by about 45% between 1989 and 1998 (Figure 3.7). The crisis then turned into an economic catastrophe. National income fell by 3.5% in the third quarter of 1990 compared to the same period in 1989 (Khanin, 1990). In this regard Russia did worse than other transitional countries, where the fall in the GDP was about 26% on average. The decline in GDP was accompanied by high rates of inflation and an enormous increase in the 'shadow' economy. The rate of inflation was over 2,000% in 1992 and over 800% in 1993 (Markandya and Averchenkova, 2000). Russia's economy came to a head in the financial crisis of August 1998. The devastating crisis in 1998 resulted in a situation where Russia defaulted on its debts and devalued the rouble by almost 75%.

Kapelyushnikov (2003) indicates that wage determination was the most puzzling aspect of the Russia's transition labour market. Average real wages fell dramatically after price liberalisation: by more than 30% in 1990, 40% in 1991, and nearly 60% in 1992 (ILOCEET, 1996). The collapse in real wages was even more dramatic during the period between 1994 and 1999. During the transition years, a central issue was the movement of labour from the former state sector to a newly emerging private sector, where resources were more efficiently utilised (Huffman and Rizov, 2011).

FIGURE 3.7: The Russian GDP per capita, 1989-2009



Source: World Bank, World Development Indicators database, accessed March 2011.

According to Gimpelson and Kapeliushnikov (2011) a key distinguishing feature of Russian wage formation was its high degree of flexibility. The guaranteed part of pay was very low by international standards, at about 50%-60%. Accordingly, much of any remuneration package was variable and made up of bonuses, benefits, and lump-sum payments. In terms of overall wage inequality, Russia is one of the most unequal countries in the world. Wages are closely related to regions, which is a key differentiating factor because of high migration costs and low levels of labour mobility; to the industry to which the enterprise belongs; to the professional status of the worker; and to the financial and economic performance of the enterprise (Gimpelson and Kapeliushnikov, 2011).

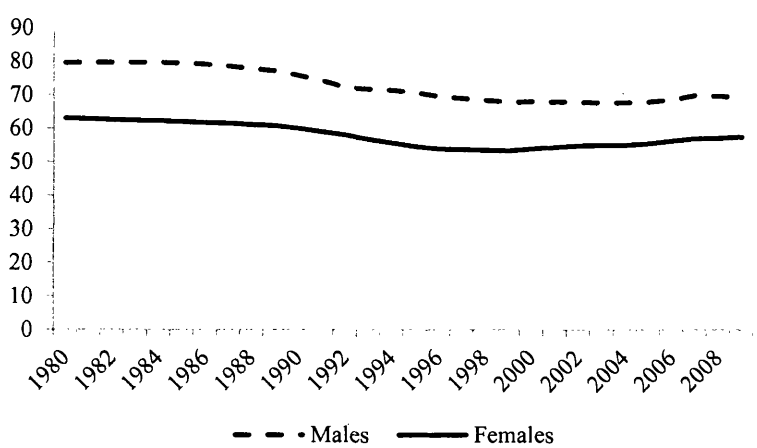
Another interesting characteristic of the Russian labour market was a so called 'wage arrears' effect. Wage arrears is a situation in which firms do not pay their employees in full (or at all) potentially even for years, which affects up to half of all employees and plays an important role in the determination of wages in Russia. Delays in Russian wage payments first became substantial in 1993, and according to the Russian State Statistical Committee (Goskomat, 1998) the aggregate stock of overdue wages grew to a total of 50 trillion rubles (around 8 billion dollars in US) by the beginning of 1998 (Earle and Sabirianova, 2002). According to Ivanova and Wyplosz (1999) total wage arrears, that is arrears to the consolidated government, private and public sectors and suppliers and banks amounted to around 35% of GDP in 1998.

Most studies of the Russian labour market have generally treated wage arrears as a way for firms to reduce their wage costs. Layard and Richter (1995) describe wage arrears as a form of wage flexibility, in which workers are willing to accept pay cuts in order to preserve jobs. The authors identify the lack of binding bankruptcy rules in the Russian economy as an explanation for wage arrears, which are related to the broader patterns of economics and financial decline in Russia. Another explanation suggests the firm's customers fail to pay on time, resulting in no money to pay wages. Alternatively where there is little external finance available, firms may take advantage of the possibility of using interest-free loans from their workers.

The unemployment rate in Russia remained at less than 10% during most of transition period. It increased steadily to 11.5% by 1998, reaching a peak of 12.9% in 1999 after the Russian financial crisis of 1998. Thereafter unemployment gradually decreases to 8.3% in 2004 (IMF, 2005). From Figure 3.8 it is evident that the participation rate of men and women decreased by 10% over the period 1980 to 2009.

From 1999 to mid-2008, Russia experienced strong economic growth (see Figure 3.7). Real GDP per capita increased on average by 6.9% per year and reached 2,805 US dollars in 2009. In addition, over the period 1999 to 2008, average real wages increased by around 10.5% per year (Cooper, 2009).

FIGURE 3.8: Labour participation rate in Russia, 1980-2008



Source: ILO, Key Indicators of the Labour Market database.
Notes: % of population ages 15 and over.

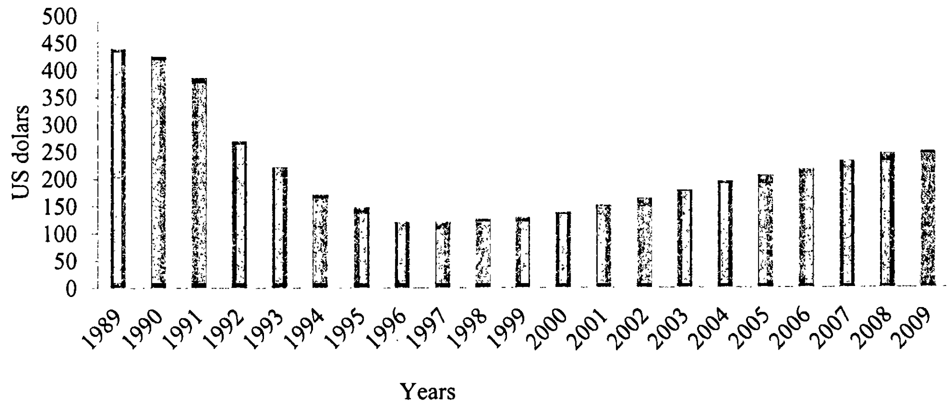
3.2.4. *Tajikistan*

The country is located on the southern edge of the Central Asian group of nations, bordering Afghanistan to the south, China to the east, Kyrgyzstan to the north, and Uzbekistan to the West. The smallest of the five former Soviet republics of Central Asia, Tajikistan has an area of 143,100 square kilometres, of which 400 square kilometres is water (Country profile, 2007). The country, which is one of the new independent States of Central Asia, has one of the lowest per capita GDP among the 15 former Soviet republics[28]. Tajikistan's economy was among the worst affected by the transition, with hyperinflation and the collapse of industrial production aggravated by a five-year civil war (1992 to 1997) (Falkingham, 2000). The financial crisis in Russia in August 1998 additionally altered the economic environment in the country.

The decline in the output in Tajikistan has been much sharper and more sustained than elsewhere in the region. By 1996 real GDP was worth less than 40% of its value in 1989 (see Figure 3.9). The GDP per capita declined from 440 US dollars in 1989 to around 122 US dollars in 1996. The country's economy started to recover in 1997. Even after several years of nearly 7% annual growth real GDP has not returned to the level prevailing at the start of the civil war. The civil war that erupted in 1992, and was due to ethnic and religious differences, resulted in nearly 1 million refugees who fled to Afghanistan and to neighbouring Central Asian countries. This outflow of population was partly responsible for the dramatic economic and social decline in Tajikistan between 1992 and 1997.

[28]Tajikistan became independent in 1991 following the breakup of the Soviet Union, and experienced a civil war between regional factions from 1992-1997, which inflicted widespread physical damage and heavy human losses of up to 50,000 lives.

FIGURE 3.9: The Tajikistan GDP per capita, 1989-2009



Source: World Bank, World Development Indicators database.

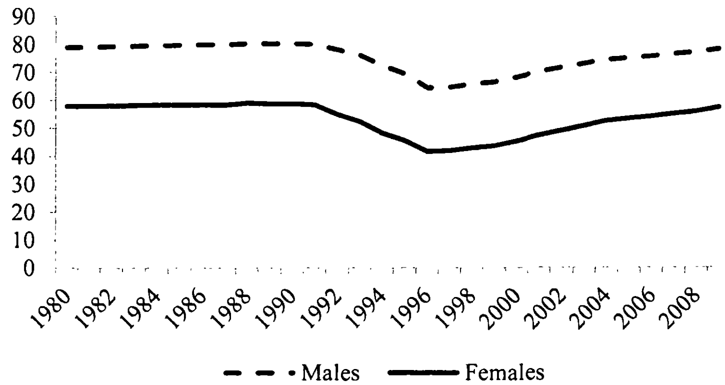
While, the country has experienced steady economic growth since 1997, nearly two-thirds of the population continue to live in poverty. In 2003, 68% of Tajiks lived below the national poverty line (Kohl *et al.*, 2008). Because of a lack of employment opportunities in Tajikistan, nearly half of the labour force works abroad, primarily in Russia and Kazakhstan, supporting their families in Tajikistan through remittances (ETF, 2010). Unemployment is a major push factor of migration. Olimova and Bosc (2003) find that 14.5% of all interviewed migrants had not worked before leaving the country, indicating poor labour prospects[29]. The increase in the working-age population affected the informal sector, where employment grew by almost 300% over the period 1991 to 2005 (ETF, 2010)[30].

As discussed in previous economies, the country's labour market changed dramatically during transition. Since independence in 1991 decreasing production and the closure of state-owned companies has led to massive job losses (Kröger and Meier, 2011). Due to the privatisation policy pursued in the period 1991-2007, the share of people employed in the private sector has increased and the share of people employed in the state-owned sector has decreased (Table 3.5). According to official statistics, in the period 1991-2007, employment distribution by economic sector changed radically. The proportion of the workforce employed in industry declined and the proportion employed in agriculture increased more than 1.6 times (Table

[29] The collapse of the Soviet Union produced a massive increase in migration within the Commonwealth of Independent States (CIS), including internal and cross-border movements within the region.
[30] European Training Forum (2010), Tajikistan. ETF Country information note.

3.6). In 2008, 67.2% of the labour force participation was in agriculture, 7.5% in industry, and 25.3% in services[31]. Figure 3.10, which plots the labour force participation rate in Tajikistan since 1980, indicates that under Soviet rule, participation rates were high with almost 80% of men and 60% of women being employed. After the collapse of the Soviet Union and the civil war between 1992 and 1997, participation of men and women decreased significantly.

FIGURE 3.10: Labour participation rate in Tajikistan, 1980-2008



Source: ILO, Key Indicators of the Labour Market database.
Notes: % of population ages 15 and over.

TABLE 3.5

Employment by ownership and economic sector in Tajikistan, 1991-2007

	1991	1998	2000	2002	2004	2005	2006	2007
Total employment (million)	1,971	1,796	1,745	1,857	2,090	2,112	2,137	2,150
Ownership (%)								
State-owned	59.4	42.7	32	27.8	26.5	25.7	25.1	24.2
Private[1]	19.1	33.1	43.1	44	50	51.2	50.7	51.8
Collective[2]	20.4	23.2	23.8	27.5	22.6	22.2	23.2	22.9
Economic sector (%)								
Industry	13.0	8.2	6.9	6.6	5.6	5.7	5.5	5.5
Construction	7.5	2.9	2.1	1.8	1.6	3.3	2.9	3.0
Agriculture	44.7	60.7	65.0	66.6	67.6	66.6	67.5	67.0

Notes: (1) Personal subsidiary plots and farms, individual labour activities, private enterprises. (2) Collective farms, cooperatives, public associations.
Source: State Statistics Committee.

An analysis of the outcomes of various surveys shows a decline of unemployment rate for the period 1999-2004. According to the Tajikistan Living Standards Measurement Survey (LSMS) conducted in 1999, the unemployment rate was 16%

[31] Central Intelligence Agency World Fact Book, (2010).

and it declined to 11.4% in 2002 based on the Poverty Reduction Monitoring Survey.

TABLE 3.6: Changes in general unemployment in Tajikistan (%), 1991-2004

	1991	1999	2000	2002	2003	2004
Unemployment rate		16.0	9.3	11.4	12.0	7.4
Labour force participation rate	76.5	69.1	63.2	-	64.2	66.5

Source: 1991, official statistics; 1999, Tajikistan Living Standards Measurement Survey; 2000, Census; 2002 Poverty Reduction Monitoring Survey; 2003 Tajikistan Living Standards Survey; 2004 LFS.

3.3. Educational systems

The countries of Eastern Europe and the former Soviet Union are unusual in many ways but compared to developing economies transition economies have high human capital relative to GDP per capita (Spagat, 2006). Education was a top priority of communism, serving both ideological and economic roles: to promote the goals of the Soviet Union and to meet the skills needs of a centrally planned economy (Banzragch, 2010). Under communism, the centrally-planed educational system was designed to fill the skill requirements of state firms (Orazem and Vodapivec, 1997). Overall, the system encouraged students to select vocational curricula and to leave school after completing upper secondary education. In 1989, adult literacy was generally universal and participation and completion rates for both males and females were high at all levels of education. Education through to university level was provided free which meant that the opportunity costs of education were low. Employment after graduation was guaranteed by state enterprises. Beirne and Compos (2007) argue that most human capital indicators (for instance, gross and net enrolment rates, average years of schooling and literacy levels) tend to be better in the late 1980s for these countries than for, say, the members of the Organisation for Economic Co-operation and Development (OECD).

Banzragch (2010) indicates several reasons for high enrolments in all level of education during the early transition period. Firstly, the organization of labour and enforcement of compulsory attendance laws were all under the control of the government authority committed to achieving high enrolment rates for education.

Secondly, during communism when the government was the only employer, the completion of such education led to secure and permanent employment. Thirdly, the necessity of education was one of the key points of the regime. Hence, people assigned value to education not because of its future wage rewards but because of the prestige it brought. Finally, education at all levels was free.

The situation changed abruptly with the transition to a market economy, when the structure of incentives was progressively altered in favour of college and more general education. The expansion of skill-intensive services (such as finance, insurance and information services) dramatically increased the demand for more educated employees. The share of students enrolled in general education between 1989 and 1997, for example, rose from 24% to 28% in Hungary, from 23% to 32% in Poland, from18% to 22% in the Czech Republic and from 18% to 25% in the Slovak Republic (Brunello *et al.*, 2011).

Enrolment rates are the best available measure of learning opportunities in regions being considered. In Table 3.7 we show preschool, basic, upper-secondary and tertiary education enrolment rates for all of the countries of interest during the period of transition. Preschool enrolments varied across countries, being especially low in Tajikistan. Basic education enrolment rate have tended to fall in all four countries. The lower and generally declining enrolment rates at the upper secondary levels in Tajikistan indicate emerging educational inequalities. In 1989 enrolment rates in general and vocational upper secondary education in Tajikistan were significantly lower than enrolment rates for basic education. In Bulgaria the tertiary enrolment rate increased from 16.4% in 1989 to 27.1% in 1997. Tajikistan has the lowest upper-secondary and tertiary enrolment rates.

TABLE 3.7: School enrolment rates in comparative perspective

	1989	1990	1991	1992	1993	1994	1995	1996	1997
Preschool Enrolment rates (net rates, % of relevant population)									
Bulgaria	63.9	63.9	55.9	57.4	56.2	57.5	61.6	62.6	58.8
FR Yugoslav	24.1	23.8	21.9	20.5	21.8	24.6	26.3	28.1	29.2
Russia	69.3	66.4	63.9	56.8	57.4	56.2	55.5	55.0	56.0
Tajikistan[e]	16.7	15.4	14.1	11.5	11.5	11.1	8.2	7.7	-
Basic Education Enrolment rates (gross rates, % of 6/7-14/15 age groups)									
Bulgaria[a]	98.4	98.6	97.3	95.1	94.0	94.3	93.7	93.6	94.0
FR Yugoslav[b,c]	95.3	95.0	94.4	72.7	74.3	72.5	71.6	72.7	71.8
Russia[d]	93.0	93.6	94.4	93.3	91.9	90.7	91.3	91.4	90.8
Tajikistan[d]	94.1	94.0	94.2	89.6	85.1	86.4	86.6	85.0	85.5
General Upper Secondary Enrolment rates (gross rates, % of 15-18 age group)									
Bulgaria	30.9	29.9	29.4	29.3	29.6	31.3	31.7	31.4	30.7
FR Yugoslav	-	4.0	6.1	9.6	12.4	12.8	12.9	13.3	13.7
Russia	24.7	22.5	33.9	22.9	22.6	23.7	24.6	26.0	27.7
Tajikistan	41.5	40.7	37.7	29.7	26.8	25.3	23.6	22.3	22.5
Vocational Upper Secondary Enrolment rates (gross rates, % of 15-18 age group)									
Bulgaria	47.0	47.3	46.2	43.1	41.5	42.6	42.5	42.2	42.4
FR Yugoslav	13.9	10.7	11.2	12.1	10.0	10.0	11.1	10.6	-
Russia	54.1	50.8	49.0	46.4	44.2	41.7	41.8	41.9	41.9
Tajikistan	20.2	19.3	-	16.2	-	-	12.4	-	-
Tertiary Enrolment rates (gross rates, % of 18-22 age group)									
Bulgaria[f]	16.4	18.8	18.7	19.8	20.9	23.0	26.0	27.3	27.1
FR Yugoslav	17.1	16.9	15.8	13.7	14.8	14.5	14.9	16.5	-
Russia	16.7	17.0	17.1	16.9	16.4	16.1	16.9	17.6	18.7
Tajikistan	9.0	9.4	9.4	9.3	8.6	9.2	-	9.4	8.9

Source: UNICEF-ICDC Trans MONEE Database.
Notes: a. 6-14 year olds; b. net rates; c. 7-14 year olds; d. 7-15 year olds; e. 1-6 year olds; f. Includes part and full-time students 19-25 age group;

Although countries differ in how far they have 'travelled down the road' to new economic destinations, they have all had to address similar problems in their education systems (Berryman, 2000). Reforms of the old system were introduced practically in all Central and Eastern European countries and implied a revolution with long-term consequences. General reforms were at the root of a series of more specific educational reforms: reforms of institutional structures, of curriculum, of management, governance and finance of educational systems (Cerych, 1997). One argument made in the literature is that political and economic changes in the 1990s rendered some teaching skills obsolete and added others almost overnight. Examples include the need to retrain Russian teachers in Central Europe and, the need to teach applied as well as theoretical knowledge in subjects such as mathematics and the sciences (UNICEF 1998). Berryman (2000) argues that during the transition the education sector adjusted to budget constraints in two ways. One was to allow the real wages of staff in the sector to decline. Wages for teachers in Tajikistan, for example, dropped from 88% of the average wage in 1990 to 49% in

1996. Rutkowski (1995) finds that being in a social service job in Bulgaria, including education, rather than in manufacturing almost doubled the risk of low wages. The other adjustment used by the sector was to scale back budgets on other inputs in the production of education. Several European and Central Asian (ECA) countries, for example, almost eliminated budgets for textbooks, teaching materials and maintenance[32] (Berryman, 2000).

Greater differences exist between countries in terms of the reform process itself and its outcomes. Economically and educationally, the countries of South-Eastern Europe (Romania, Bulgaria, Albania, and the former Yugoslav republics) developed relatively late, mainly after the Second World War, and they were heavily influenced by the Soviet Union. The next sections outline the main features of the Bulgarian, Serbian, Russian, and Tajik education systems.

3.3.1. The educational system in Bulgaria

Education in Bulgaria, although fundamentally national in character, has significant foreign influences. The Soviet influence was most evident during the period of the national revival in the nineteenth century and reflected the ideas of Slavophilism and pan-Orthodoxy. After the Second World War, the communist regime wanted to establish a system of education similar to that of the Soviet Union, but at the same time they wanted to maintain a national culture. The length of time at elementary secondary level was reduced to 11 years, instead of 12 years, and the first 4 years were compulsory. The primary aim of the communist educational authorities was to increase the number of students at each educational level. No questions were asked about the quality of education since the majority of teachers either needed special training in the new state demands or there simply were no qualified teachers. As in Russia, the Bulgarian education system was very much controlled by the party. After 1945 an intensive effort was made to include more women, both in the economy and in political organizations. Women were required to participate in all events which would have been unthinkable before the communist regime. They were also included in all communist youth organizations. By 1965, the number of

[32] In Russia the percentage of textbooks provided by region budgets varied from 6% in Chelybinsk region to 65% in Irkutsk (Berryman, 2000).

women in schools, in the economy and in the party was almost equal to the number of men (Shimoniak, 1970).

Enrollment rates in Bulgaria were related to the level of family income. Children of poorly educated parents tended to complete fewer years of education. While enrolment rates in basic and secondary education did not vary much by expenditure quintiles in 1995, by 1997 the net enrolment rates for basic education, while stable for the top four quantiles, dropped by about 20% for the lowest quantile. Net enrolment rates at the upper secondary level for the lowest quantile also declined by more than 10 %. Although overall public spending was distributing neutrally, it was pro-poor at the preschool and basic education levels and favored the non-poor at secondary and tertiary levels (World Bank, 1999). The macroeconomic decline in Bulgaria also reduced public funds available for education, which declined from 5% of the GDP in 1990 to 3.2% in 1996 (UNESCO, 2000).

During the period 1990 to 2006, a number of democratic changes affected society as a whole and the educational system in particular. In 1991 a new Public Education Act was adopted, which is still in force in Bulgaria. The principle of competition for electing school management was introduced, and democratic principles of governance and the administration of schools were applied. A serious attempt was made to harmonize Bulgarian educational legislation with the standards of other European countries, while preserving the achievements of the Soviet era (Eurybase, 2005).

In Bulgaria, education is compulsory for children aged between 7 to 16 years old. Children at the age of 6 can also be enrolled as first-grade pupils, if their physical and mental development allows it. Since 2003/2004 pre-school education has become compulsory (Eurybase, 2005). Prior to entry into higher education, the education system in Bulgaria consisted of 12 school grades, organized in two major levels of study: basic and secondary. Basic education (grades one to eight) is divided into two sub-levels: elementary (grades one to four) and pre-secondary (grades five through eight). Secondary education normally encompasses grades eight to twelve and there are two major types of secondary schools: secondary comprehensive, usually called gymnasia (high school) and secondary vocational, most often referred to as technikum (vocational school). University education is provided at universities and specialized higher schools – academies or institutes. It includes three stages: first stage – a course of study of at least four-years, leading to a Bachelor degree upon graduation; second stage – a course of study of at least five-years, or one-year following a bachelor degree, leading a master degree upon graduation; third stage – a three-year course of study following a master degree, leading to a doctoral degree.

Although the majority of European national policies are now encouraging higher education institutions to rely increasingly on private sources of funding, direct public funding continues to represent a substantial share of the higher education budget. In 2003, within the 27 Member States of the European Union, 79.9 % of the funding for higher educational institutuions came from public sources. In Bulgaria this proportion was 55.2%. The amount of tuition fees in the country are determined by the central education authorities and the proportion of expenditure in higher education from households reach about 40% (Eurydice, 2008).

3.3.2. The educational system in Serbia

The progress towards a stable democratic system in Serbia has been slow but despite all of its problems Serbia has begun to rebuild and reform its education system. Before the global crisis began, public investment in education increased from 2.7% of the GDP in 2001 to 3.7% in 2007. Investments into school infrastructure also increased but they were far from sufficient (Mijatovic, 2008).

Latest Labour Force Survey (2009) data confirms the low educational attainment of the Serbian population: 3% of the population older than age of 15 do not have full elementary education; 35% of the population have only elementary education; 48% secondary education; and 14% higher education. For almost half of the population aged 15 or above, secondary education represents the highest educational attainment. The link between poverty and education in Serbia is very strong, with 71% of those living below the poverty line having no education or only primary school education[33]. According to the last Census (2002), 3.45% of the population

[33] *UNICEF* Serbian Annual Report, 2007.

were illiterate and almost a million had not completed primary schooling[34]. Among sub groups of the population, Roma children are the most disadvantaged and excluded group in Serbia's education system. Only 36% of Roma children complete primary school, 9.1% secondary school and 0.9% tertiary education (OSI, 2008).

The main intention of the reforms implemented in Serbia was to make it possible for students to compare their higher education qualifications with those acquired by students at other European higher education institutions, which would encourage student exchange programmes. In 2005 Serbia adopted the Law on Higher Education which is considered with the Bologna Declaration (Ćirić and Đurđić, 2009). To support educational reform the Ministry of Education undertook a major legislative initiative by revising the laws governing pre-schools, primary and secondary schools in Serbia. The key innovations adopted by the Parliament in 2009 set out a framework for an integrated education policy. This included new principles of equity, improved access, quality and efficiency of education that sought to match competencies to the requirements of jobs, standardise learning achievements, increase the autonomy of schools and promote flexible learning programmes (Klenha *et al.*, 2010).

The education system in Serbia includes preschool, primary, secondary, higher, and university education. Compulsory education (pre-school and primary) and secondary education are free and funded from the state budget, but parents and pupils have to cover the expenses for textbooks, stationary and school trips. Preschool covers children from 6 to 7 years old. Primary education starts at age of 7 and is compulsory until the age of 16. Primary education in Serbia is comprehensive, currently organised in 2 cycles, grades 1-4 as class instruction and 5-8 as a subject-based instruction. The 2003 legislation introduced nine years of compulsory education in 3 cycles, but the 2004 amendments reinstated the previous 8 grade system (Cerović, 2007). Secondary education follows primary education and while it is not compulsory it is free for all. Secondary schools are divided into gymnasiums and vocational schools, each of which lasts 3 or 4 years. Tertiary

[34] The population in Serbia was 7.5 million in 2002, excluding Kosovo.

education is offered at 6 state universities, 6 private universities and at numerous tertiary vocational schools that offer 2-year programmes (Cerović, 2007).

Higher education institutions in Serbia acquire funds from various sources and manage these sources independently. These are mainly: the founder; students; donations; projects and contracts related to the carrying out of courses of study, research and consulting services. State-founded and private-founded universities are in a different position in relation to funding. There are two categories students with respect of funding: state funded students who study at state-funded higher education institutions and self-funded students, who pay the tuition fee determined by higher education institutions. All self-funded students at a particular higher education institution pay the same fee regardless of age. Tuition fees are determined by the higher education institutions (TEMPUS, 2010).

3.3.3. The educational system in Russia

Russia has a long-standing tradition in promoting high-quality education and has always shown great concern for education. It probably has also one of the best mass-education systems in the world producing a literacy rate (98%) exceeding most Western European countries. Education was one of the top priorities during the new communist regime and it is highly valued not only because it offers higher wages but because of prestige self-esteem associated with its acquisition (Cheidvasser and Benitez-Silva, 2007). One general characteristic of the education system prior to the transition was a high degree of centralisation. Curricula, personnel, criteria for enrolling pupils and many other dimensions of the school system were controlled by the central authorities, and there was little autonomy at either the school or district level (Bray and Borevskaya, 2001). Free educational establishments provided equal educational opportunities to all. Compulsory education was 10 years in duration. A large proportion of secondary school graduates went to specialized vocational schools. Access to higher education was controlled and the content of higher education emphasized the technical needs of the economy (Banzragch, 2010).

Following the break of the USSR in December 1991 the entire Soviet education system, given the collapse of Soviet ideology, were restructured and transformed. Educational reforms moved much more rapidly than the democratic transformation of society. Major provisions of the education reforms included overhauling the curriculum to make it less complex; adding new courses in social and political science; increasing the vocational component of education by requiring vocational training at all levels; achieving universal computer literacy; and increasing resources for education (Kaufman and Hardt, 1993). In 1992 the Law on Education and in 1993 the Federal Programme of Educational Development were implemented. Article 41 of the Law stipulated that educational institutions would be financed on a per student basis according to federally and locally established criteria for financing all type of educational establishments. Eleven years of secondary education were made universal and compulsory during the 1980s. However, because of qualitative and financial problems, the government decided to reduce the duration of compulsory education to 9 years in 1992. Despite these changes, much of the Soviet education system remains in place in Russia today.

Tuition-based education and private education have been legalised and the non-state education sec-tor has witnessed rapid growth in recent years. By 2004 there were 655 state and about 530 non-state higher education institutions in Russia. Of the 530 non-state higher education institutions, 364 were accredited by 2004. In 2002/2003 there were 5,228,700 students in state institutions and 718,800 in non-state institutions.

If we look at the two widely used indicators of the extent to which governments give priority to education – public expenditure as a percentage of GNP and public expenditure on education as a percentage of total public expenditures, the figures for Russia are lower than for many other countries (see Table 3.8). Moreover, public expenditures on education declined from 3.9% of the GNP in 1975 to 3.1% in 1985, and only increased to 4.2% in 1997 (UNESCO, 2000).

TABLE 3.8: The public expenditures on education as % of GNP, 1985-1997

Country/region	1985	1990	1995	1997
Russia	3.2	3.5	3.5	4.2
North America	5.0	5.4	5.3	5.4
Europe	5.2	5.1	5.3	5.3
Sub-Sarah Africa	4.5	4.6	5.1	5.1
Arab States	5.8	4.9	5.0	5.4
Latin America and Caribbean	4.0	4.0	4.5	4.6
Eastern Asia and Oceania	3.1	3.0	2.9	2.9
Southern Asia	3.4	3.7	3.2	3.3

Source: Berryman (2000, p.123), UNESCO (2000).

Table 3.9 shows the proportion of government total expenditure. The statistics refer to consolidated budgets, e.g. the sum of the national, regional and sub-regional budgets. The situation in Russia was markedly different prior to the market reforms and there was a steady decline in public expenditure on education as a share of total government expenditure until 1995. Even in 1995, the Russian figures were significantly lower compared to the shares reported for both Europe and Asia.

TABLE 3.9

The public expenditures on education as % of total public expenditures in Russia

	1980	1990	1992	1993	1994	1995	1996	1997	1998
Russia	16.7	14.8	11.0	12.7	11.7	12.4	12.8	14.4	12.9
Asia	-	-	-	-	-	15.1	-	-	-
Europe	-	-	-	-	-	13.4	-	-	-

Source: Bray and Borevskaya (2001).

Education in Russia is split into a compulsory basic education, and higher education. The stages of compulsory schooling are: primary education for ages 6-7 to 9-10 inclusive; senior school for ages 10-11 to 12-13 inclusive, and senior school for ages 13-14 to 14-15 inclusive. If a secondary school pupil wishes to go on to higher education, he or she must stay at school for another two years. Primary and secondary schooling together account for 11 years of study, split into elementary (grades 1-4), middle (grades 5-9) and senior (grades 10-11) classes. Eleven-year secondary education in Russia has been compulsory since September 1, 2007. No entrance examination is required for admission to programmes of basic vocational education. Entrance to basic vocational education is after year 9. Students entering basic vocational education after eleven years of general education follow shortened programmes. Higher education is provided by state and non-state higher education institutions. There are three levels of higher education – incomplete higher education (2 years at least), 4 year programmes leading to bachelor degrees and postgraduate study of 1-2 years leading to masters degrees.

3.3.4. The educational system in Tajikistan

With the breakup of the Soviet Union and independence, Tajikistan witnessed formidable changes in the education sector – including the provision of postsecondary school graduates; maintaining free and compulsory secondary education; ensuring an adequate number of teachers; and developing legal frameworks, new standards and curricula. The country saw significant changes in the financing of education. At the beginning of the decade, the central government was virtually the only source of funding. After the civil war, government revenues fell and funding for education dropped from 22.1% of the national budget (10.7% of the GDP) in 1990 to 11.8% in 1999 (2.1% of the GDP) (Asian Development Bank, 2004). The education system in Tajikistan underwent two major changes. First, there was an increased emphasis on the use of the Tajik language, and second, there was a considerable reduction of government funding of the educational system as a direct consequence of the 1992 to 1997 civil war (Banzragch, 2010). Tajikistan was the first country in Central Asia to pass legislation adopting a national language other than Russian. The Law on Language (1989) established Tajik as the official language and Russian as an international language.

Education reforms in Tajikistan were initiated later than in other Central Asian countries. The adoption of the Law of Education in 1993 and its amendments in 1996 and 1999 set up a new legal framework for the reform of education. The Law extended the rights and autonomy of education institutions while promoted alternative sources of finance for education by allowing financial support from other sources, such as parents, sponsors, enterprises, international organisations and foreign investors. In 1995 a Concept of National Education was established, which defined the purpose and future directions of the education system. The issues of inadequate teacher supply and the need to upgrade the qualification of existing teachers were addressed.

There was a significant fall in preschool enrolments and the proportion of total expenditure on education going to preschool education declined from 13.8% in 1995 to 4.5% in 2000 (Asian Development Bank, 2004). However, general education suffered less and broader ranges of programs were introduced, providing students with more options. In addition to secondary schools, gymnasiums and colleges were established and the share of general education in total education expenditure increased from 57.4% to 80.1% between 1995 and 2000. The number of postsecondary institutions also increased from 13 in 1991 to 30 in 2000, and student enrolment increased from 69,300 to 79,200. Fewer girls, however, were enrolled and remained in schools especially at the secondary level. Indeed in some districts girls were not allowed to attend secondary school by their parents because of religious beliefs (Asian Development Bank, 2004). Since 2000, private schools and non-profit training institutions have appeared in Tajikistan, responding to a demand for emerging secondary education. By 2006, the number of gymnasiums had more than doubled and 26% of these were private schools.

In general, the education sector pays low wages, and this is a main cause of the decline in teacher professionalism in secondary schools and vocational education (ETF, 2011). As a result of poor salaries, many teachers left teaching for jobs in other fields, and this caused a severe shortage of qualified staff. In 1998, the average salary of employees in the education sector was 6.9 US dollars a month while the living wage was estimated to be 28.3 US dollars per month. The low quality of education was a serious concern in rural areas (Banzragch, 2010).

The Tajik education system has two major stages: general and professional. The general education stage is comprised of three levels - primary, basic education (also called incomplete secondary education) and secondary education. The second stage, professional education compromises three elements – initial, secondary, and higher professional education. At the age of seven, children begin 9 years of compulsory education consisting of 4 years of primary school and 5 years of lower education with no vocational qualification of any kind. Those leaving the education system at this point therefore enter the labour market as unskilled workers.

Education legislation guarantees each student free access to any school at the upper secondary (general or vocational) level and on a competitive base, to technical (secondary specialist) and higher level schools where they can obtain work-related qualifications. Following completion of compulsory basic education the Tajik education system offers a number of different options, including general upper secondary, vocational and technical education, or combination of these (ETF, 2011). Following completion of general and upper secondary education students may continue to either higher education, technical or vocational education. The state only supported tertiary education students, paying stipends and providing them accommodation, no stipend is envisaged for initial vocational education students and existing professional-technical school premises. After technical education students can continue to higher education and may enter post-secondary and higher-level public schools free of charge on the basis of a competition for limited places. Almost half of all students are admitted into universities on a fee-payment basis. They can also gain access on a fee-paying basis. Students can also complete postgraduate studies in a 3 year program, which combines coursework and dissertation writing (Banzragch, 2010).

The development of higher education is ensured by financing from the state budget and grants as well by creating favourable conditions for self-financing of the process of training specialists and academic staff. Higher education in Tajiksitan is financed from three sources: national budget, private funds (student fees) and other sources, which consist of income received from research and manufacturing contracts and international funds (grants). Course fees are fixed by the higher education institutions in coordination with the Minsitry of Education and the Ministry of Economy and Trade. The fees depend on the speciality and it varies from 880 somoni (150 EUR) per year for a teacher-training courses up to 5,280 somoni (935 EUR) for law. Only full-time students with academic progress have the right to receive a grant (European Commission, 2010).

3.4. Summary and conclusions

In this Chapter we have provided an overview of economic developments during the transition process and briefly summarized the key features of the education systems in Bulgaria, Serbia, Russia, and Tajikistan. The review of educational developments is important as it reflects the diversity of the educational systems in place in the countries examined in this thesis. The overview of the transition process in Central European and Central Asian countries highlights three main facts: a move from the communist political system to a democratic one; overcoming deep structural economic crises by developing free market economies; and modernization and adjustment to global changes.

The evidence reviewed suggests that the selected countries resemble the characteristics of the wage settings and education systems observed in most Central and Eastern European economies under the central planning and thus providing a particularly interesting combination for estimating the returns to education. Specifically, wages in all four countries were paid under highly centralized wage grids and little variation was observed by their level of education or occupation. The Soviet commitment to universal education and employment meant that these transition economies inherited a legacy of high education levels. The educational systems emphasized secondary and vocational training which were relevant to the demands of state run industries. In all four countries, during the transition economic uncertainty increased and job security disappeared, leading to problems in their labor markets. Almost all of these countries suffered severe recessions, rising unemployment, a collapse of economic activity and a sharp deterioration in growth performance. This decline reduced the public funds available for education in all four countries.

CHAPTER FOUR

EVOLUTION OF THE RETURNS TO EDUCATION IN BULGARIA: AN EMPIRICAL INVESTIGATION

4.1. Introduction

Changing conditions in the labour market, and in the education system, suggest that returns to education may not be constant over time. A number of studies have shown that transition process had significant effects on labour market outcomes and the distribution of income[35]. Early empirical evidence suggests that, on average, returns to education rose during the early transition from central planning to a market economy. Andrén *et al.* (2005), for example, find that the returns to education in Romania increased from 3% during the pre-transition period to 8.5% in 2000. A similar pattern is reported by Compos and Jolliffe (2002) for Hungary, where the returns to education rose from 6.4% in 1986 to 11.2% in 1998. Münich *et al.* (2004) find for the Czech Republic that in 1989 women with university education earned on average 45% more than women with secondary education and that this had increased to 85% by 2002. Münich *et al.* (2004) also show that the major increase in the rate of returns to education occurred in the first phase of transition (1989-1996) and no further changes were observed in the later phase (1996-2002). Similarly, study by Fleisher *et al.* (2005) suggests that the sharpest increase in returns to education took place during the early transition (around the early 1990s).

There is an alternative hypothesis, however, namely that the rate of returns to education and experience fell from the pre-transition to the transition period. Education was poorly rewarded in many transition economies in the initial stages of transition and directed toward the productive and ideological goals of the communist regimes. Returns to schooling were low and hence have been limited in

[35] See for example Fleisher *et al.* (2005) analysis of returns to education. They use metadata between 1975 and 2002 collected from 33 studies of 10 transition economies, to assess changes in the returns to education over the transition period.

its ability to respond to changing economic and employment patterns essential for the success of market economy (Laporte and Ringold, 1997). An interesting view among researchers on transition economies is that education obtained under central planning is too narrow and firm specific, and not appropriate for the new market environment (Kertesi and Köllő, 1999; Filer *et al.*, 1999). Some individuals (in particular, low-educated or women) are more likely to become unemployed and to experience more difficulty in moving out of unemployment into a job, whereas for the young educated people it is relatively easy to find new job. Some studies argue that the 'losers' of transition are women and older workers, whose human capital has been relatively devaluated and have low incentives to acquire the new skills relevant to the emerging market economy[36]. There is also evidence of factors making the 'losers' more difficult to move from the group of 'losers' to that of 'winners', such as marked regional imbalances in the allocation of employment opportunities, and the narrow base of many vocational education curricula, which reduce the eligibility of the workforce (Boeri and Terrell, 2001).

Changes in the returns to education typically arise from three sources: changes in the demand for educated workers, changes in the supply of educated workers or from the institutional features of the wage formation. Internationally, a rising return to education has been attributed to an increasing demand for highly educated workers, particularly as a result of technological changes and innovation (Barth and Røed, 2001). For instance, Fleisher *et al.* (2005) find that both the speed of economic transformation and the degree of economic disequilibrium measured by macroeconomic volatility explain differences in the rates of return to schooling over time across Central Eastern European countries and Russia. Moreover, the economic and political transition in the countries of transition has led to considerable reduction in political supervision of the wage structure. Orazem and Vodopivec (1997) categorise the fundamental forces which have influenced the wage structure during transition into three groups. The first group consist of corrections in distortions presented in the labour market during communism, meaning measures that secured the egalitarian wage structure. The second group consists of changes in the final demand for goods and services which indirectly led

[36] See for example Orazem and Vodapivec (1995).

to reduced demand for labour in industry. The final group consists of shifts in demand pertaining to the imbalances which emerged during the process of transition, i.e. the fact that education and entrepreneurship associated with it were not sought in the previous system, so due to the elasticity in the labour supply over the short period, its supply is lower than could be expected given the incentives established. All these forces pushed the wage structure in the same direction, toward growing wage premiums for education (Vujčić and Šošić, 2007).

Despite the large number of empirical studies on returns to education in transition economies, the results for the pre-1989 transition period are limited. There is no evidence for returns to education in Bulgaria and in particular on the evolution of returns to education in the 1980s. In this context, we believe it is interesting to examine whether the transition process has caused changes in economic returns to education in Bulgaria, a country which went through radical economic transformation after the 10th of November 1989.

This chapter attempts to contribute to the literature in several ways. First, we address the question what has happened to returns to human capital in Bulgaria before and during the transition from centrally planned to market economy by examine the evolution of returns to education for the 1986 to 2003 period. Second, the study adds to the previous literature by using econometric techniques, which have not previously been applied to Bulgarian data. We begin by presenting OLS estimates of a Mincerian earning equation. Next we account for a potential selectivity bias. We compare parametric and semi-parametric sample selection results. We also deal with possible endogeneity bias of education by employing an instrumental variable (IV) approach. The analysis is undertaken separately for men and women. Finally, we provide evidence on how returns to schooling in Bulgaria have evolved at various points along the conditional wage distribution. To the best of our knowledge, this is the first study that estimates the evolution of returns to education in Bulgaria.

The empirical analysis covers quite an important period in Bulgarian's transition. The year 1986 falls within the socialist period, whereas in 1995 Bulgaria was into the initial transition phase. As we have discussed in Section 3.2.1, the end of 1996 and the beginning of 1997, the country saw an increase of 827 percentage points in the inflation rate from an already high base of 122.9 percentage points[37]. Real GDP fell by another 10.1% in 1996 and 7% in 1997 (Dimova and Wolff, 2008). At the end of 1996 about one-third of Bulgaria's banking system dissolved. Households had to bear the largest part of the currency devaluation and the GDP contraction as those losses worked their way through the banking sector and the government budget down to the consumer (Ivanov, 2003).

The structure of the Chapter is organised as follows. In Section 4.2 the data sets and the empirical methodological framework are presented. Main results are discussed in Section 4.3, and finally Section 4.4 draws some conclusions.

4.2. Data and methodology

4.2.1. The data

The empirical analysis in this Chapter is based on five household surveys conducted in Bulgaria. One is administrated before transition and the others after the initial economic reforms were held. The 1986 Town and Village Survey, conducted by the Institute of Sociology of the Bulgarian Academy of Science in Sofia, was carried out in conjunction with the national census delivered in the winter of 1985 and contains information on the structure of Bulgarian wages prior to the transition (Giddings, 2002b). This survey is the only existing pre-transition survey in Bulgaria. The sample is representative of the population and contains 10,333 respondents. After eliminating those who reported zero earnings and excluding individuals who were not of working age, the sample consists of 6,451 individuals[38].

The next data sets comes from the nationally representative Bulgarian Integrated Household Surveys (BIHS) collected in 1995, 1997, 2001 and 2003. The Gallup Organization in Sofia managed the survey and the Bulgarian Ministry of Labour, the National Institute of Statistics and the World Bank provided technical assistance. The BIHS questionnaire follows the structure of a typical Living

[37] Inflation reached 4-digit numbers from February until August 1997 (Ivanov, 2003).
[38] We sample individuals aged between 16 and 65 in all surveys.

Standard Measurement Survey (LSMS). The sampling procedures ensured that all the BIHS samples were highly representative. Exclusion of individuals who do not report wages accounts for a substantial reduction in sample sizes. Our final working samples consist of 1,317 individuals in 1995, 1,531 individuals in 1997, 1,438 individuals in 2001 and 2,482 individuals in 2003.

While the LSMS and the 1986 Town and Village Survey were conducted for different purposes, the surveys are comparable and representative at country level. An comparison of 1986 and 2001 data indicates that both have similar gender, age, and regional distributions. For instance, 50% and 49% of females in both samples are in paid employment. The regional variation have been found to be close between surveys with about 4% of the individuals in 1986 sample and 4.4% in 2001 worked in the Pleven region (see Appendix Table A4.17).

The dependent variable used in the analysis is the logarithm of hourly earnings resulting from the primary occupation and excludes earnings from secondary jobs, or from agricultural production, and non-monetary benefits[39]. Wages are net of taxes. The variable also includes all additional payments in cash, in kind and bonuses. Education is represented by binary variables measuring the completion of indicated levels of schooling. These levels consist of primary, secondary, and university education. However, our empirical analysis is based on the total number of years in education, which are also available within the data sets. It comprises of the number of educational grades completed by the individual and takes on 18 different values: from 0 in the individual achieved no grade to 18 if the individual has a postgraduate education. The binary variable Bulgarian takes value 1 if an individual is ethnic Bulgarian and 0 if an individual is a member of an ethnic minority group, of which the main groups are Roma and Turkish. The vector of exogenous control variables used in the estimations also includes potential experience[40] (linear and quadratic terms), dummy variables for individuals years of tenure within the firm (less than 1 year, between 1 and 2 years, between 3 and 5 years, between 6 and 10 years and more than 10 years) and urban settlement. We

[39] Hourly earnings are defined as reported monthly net earnings divided by 4.34 and then divided by reported weekly hours of work.
[40] Potential experience is imputed as follows: age – 7– years of schooling.

have also included a public sector dummy so as to control for the remaining effect of the old public sector wage structure.

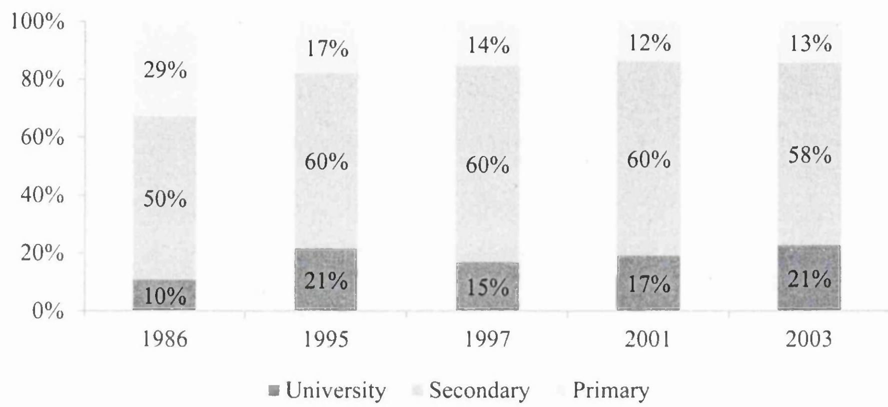
In Table 4.1 we report descriptive statistics of the resulting samples. For all years, the average log hourly wage rate is higher for men. Interestingly, in 1986 in terms of the hourly wages and education, there is no significant difference between the genders mainly explained by the socialist ideology of equality and wage setting structure. Wages under the communist regime were paid under highly centralized wage grids and differentials were kept within certain limits. Moreover, under communism, tuition was free and all students were eligible for textbooks and meals that were subsidised by the state. Our descriptive statistics show dramatic fall in wages in 2001, and the reason for that is the 1999 redenomination of the Bulgarian currency[41].

According to the 1986 data, 9.9% of employed men and 9.5% of employed women have a university degree. Public sector employment in 1986 accounted for around 81% of total employed males and females. Experience, defined as age minus years in school minus 7, is higher for females than males in 1986. However, potential experience might overstate women's actual labour market experience as it does not account for female absence due to childbearing activities. It is also interesting to compare particular social groups with a certain level of education and the completed education levels of their offspring. The 1986 Town and Village Survey provide information of the educational level of the respondent's father and mothers. According to this data, 13% of respondent's mothers had a high degree (university and college education), which indicates that our respondents were slightly better educated than their parents. In contrast to the 1986 data, the 1995, 1997, 2001 and 2003 samples indicate that women earn less than men, which support the thesis that transition has resulted in increased gender inequality and large changes in the distribution of wages. For the whole period, women have more years of schooling, measured as a number of total year's education. Moreover, the data shows that 26.3% of employed women in 1995 had a university degree compared with 17.4 % for men. Women have higher participation rate in the public sector and we can see

[41] As of 5 July 1999 the Bulgarian lev was redenominated whereby 1,000 old levs shall be exchange for 1 new lev (Law of Redenomination of the BG lev adopted by the 38th National Assembly on 19 February 1999).

that for the period 1986-2003 the percentage of public sector workers significantly decreased (from 82% of working women employed in public sector in 1986 to 39% in 2003). The majority of working males and females live in urban areas.

FIGURE 4.1: Educational structures of employed respondents in Bulgaria



Notes: Bulgarian Household Surveys descriptive statistics based on working individuals.

All the cross-section data clearly indicate that in the beginning of the period, the majority of the working sample was poorly educated and only a small proportion had university degrees. For the period 1986-2003 the percentage of working individuals with secondary and university education increased, while the proportion of those with primary decreased from 29% in 1986 to 13% in 2003 (see Figure 4.1). Significant changes in the educational structure of the working population, which were expressed in a 10% increase of the share of secondary and university graduates, took place in 1995. Overall, individuals with secondary education dominated the educational structure. This might be an indication for the transition wage structure, which did not provide adequate incentives to invest in education.

TABLE 4.1

Descriptive statistics – Bulgarian working samples, 1986-2003

Variables	1986		1995		1997		2001		2003	
Description	*Males*	*Females*	*Males*	*Females*	*Males*	*Females*	*Males*	*Females*	*Males*	*Females*
Log of hourly	0.955	0.940	4.914	4.617	5.912	5.664	1.519	1.342	1.622	1.460
wage	(0.461)	(0.460)	(0.594)	(0.615)	(0.855)	(0.775)	(0.668)	(0.539)	(0.558)	(0.519)
Education										
Total years in	10.142	10.193	11.182	11.452	11.675	12.132	11.407	12.099	13.056	13.875
education	(3.361)	(3.456)	(3.077)	(3.038)	(2.838)	(2.862)	(3.202)	(3.304)	(3.420)	(3.424)
1 if primary	0.304	0.275	0.199	0.146	0.150	0.122	0.143	0.102	0.154	0.105
	(0.460)	(0.447)	(0.400)	(0.354)	(0.357)	(0.327)	(0.350)	(0.302)	(0.361)	(0.307)
1 if secondary	0.504	0.491	0.614	0.576	0.640	0.556	0.635	0.557	0.620	0.532
	(0.500)	(0.500)	(0.487)	(0.495)	(0.480)	(0.497)	(0.482)	(0.497)	(0.486)	(0.499)
1 if university	0.099	0.095	0.174	0.263	0.127	0.172	0.137	0.207	0.167	0.256
	(0.298)	(0.293)	(0.379)	(0.440)	(0.333)	(0.377)	(0.344)	(0.406)	(0.374)	(0.437)
potential exp	30.072	31.366	22.468	20.941	22.343	20.251	21.79	20.25	19.972	18.653
	(14.638)	(15.244)	(11.399)	(9.937)	(10.837)	(9.472)	(11.739)	(10.799)	(12.264)	(11.179)
Tenure within the firm										
1 if<1 year	0.033	0.039	0.151	0.126	0.309	0.259	0.136	0.159	0.270	0.226
	(0.178)	(0.194)	(0.358)	(0.333)	(0.463)	(0.438)	(0.343)	(0.366)	(0.444)	(0.418)
1 if 1-2 years	0.030	0.027	0.061	0.051	0.073	0.060	0.071	0.075	0.177	0.163
	(0.169)	(0.162)	(0.239)	(0.220)	(0.260)	(0.237)	(0.258)	(0.263)	(0.382)	(0.369)
1 if 3-5 years	0.095	0.090	0.127	0.123	0.139	0.096	0.148	0.133	0.191	0.206
	(0.293)	(0.286)	(0.333)	(0.329)	(0.346)	(0.294)	(0.355)	(0.339)	(0.393)	(0.404)
1 if 6-10 years	0.158	0.170	0.151	0.192	0.133	0.199	0.255	0.240	0.124	0.142
	(0.365)	(0.375)	(0.358)	(0.394)	(0.340)	(0.400)	(0.436)	(0.427)	(0.330)	(0.349)
1 if>10 years	0.684	0.674	0.510	0.507	0.345	0.387	0.390	0.394	0.238	0.263
	(0.465)	(0.469)	(0.500)	(0.500)	(0.476)	(0.487)	(0.488)	(0.489)	(0.426)	(0.440)
1 if married	0.819	0.856	0.806	0.818	0.820	0.820	0.760	0.726	0.680	0.728
	(0.385)	(0.351)	(0.395)	(0.386)	(0.385)	(0.385)	(0.427)	(0.446)	(0.467)	(0.445)
1 if Bulgarian	0.852	0.866	0.919	0.910	0.906	0.912	0.877	0.918	0.901	0.896
	(0.355)	(0.340)	(0.272)	(0.287)	(0.291)	(0.284)	(0.329)	(0.274)	(0.298)	(0.305)
1 if in public	0.811	0.819	0.814	0.859	0.743	0.821	0.432	0.444	0.306	0.390
	(0.392)	(0.385)	(0.390)	(0.349)	(0.437)	(0.384)	(0.496)	(0.497)	(0.461)	(0.488)
1 if urban	0.526	0.538	0.740	0.767	0.780	0.790	0.748	0.811	0.766	0.808
	(0.499)	(0.499)	(0.439)	(0.423)	(0.415)	(0.408)	(0.435)	(0.392)	(0.423)	(0.394)
1 if in Sofia	0.110	0.140	0.114	0.118	0.092	0.102	0.112	0.128	0.165	0.167
region	(0.313)	(0.347)	(0.319)	(0.323)	(0.289)	(0.303)	(0.316)	(0.335)	(0.371)	(0.373)
Sample size	*3213*	*3238*	*708*	*609*	*727*	*804*	*729*	*709*	*1296*	*1186*

Notes: Standard deviations in brackets.

4.2.2. Endogeneity of education

As we have already discussed, the OLS estimate of the return to education is unbiased only if measured schooling is exogenous. Endogeneity arising from measurement error in schooling and is generally thought to bias the estimate of the schooling coefficient toward zero, although this effect is believed to be small because the reliability of measuring the schooling data is typically quite high. The endogeneity can also arise because of omitted ability bias. That is, the return to schooling coefficient is biased (upward) because chosen schooling levels are positively correlated with omitted ability, and ability is correlated with earnings

(Trostel *et al.,* 2002). To address the problem of endogeneity bias we need to instrument education, using a variable that is correlated with education but not with earnings other than through the variable that is being instrumented. Specifically, our instrument needs to be orthogonal to the unobserved component of the earning equation.

The following two-equation model describing the log of earnings ($\ln Y_i$) and years of schooling (S_i) is applied:

$$\ln Y_i = \beta S_i + Z_{1i}\delta + \xi_i \tag{4.1}$$

$$S_i = Z_i \alpha + v_i \tag{4.2}$$

where Z_{1i} is a vector of exogenous variables that determine individual earnings (Y_i), Z_i is a vector of exogenous variables that influence the educational decision, $E(Z_i, v_i) = 0$, Z_{1i} is a subset of Z_i. The returns to schooling is measured by β and since ξ_i is defined as the residual from a regression of v_i on Z_i, it is uncorrelated with Z_{1i} by construction, and $Z_i\alpha = Z_{1i}\alpha_1 + Z_{2i}\alpha_2, \alpha_2 \neq 0$ is the rank condition needed for identifying equation (4.1), or as discussed by Heckman (1990) and Card (1993), identification is provided by including variables in the vector Z_i that are not contained in Z_{1i}. Equation (4.1) is then estimated by substituting the fitted values from the first-stage regression of S_i on Z_i as defined in equation (4.2). In practise, to model the relationship between education and earnings, we need to compute the predicted values S^* for education S_i and then replace the education variable in the earnings function with this predicted value. Since S^* is not correlated with the error term in the earnings function, the estimated return based on predicted education is unbiased. The addition of S^* is sufficient to eliminate endogeneity bias arising from the possible correlation between education and unobservable influences on earnings.

Identification in the endogeneity adjusted two-step estimations is achieved by making use of a mother's education as an instrument, assuming that there is high correlation between children's educational attainments and those of their parents. Higher parental education is also associated with more substantial family investments in children[42]. Relatively well-educated parents, cognizant of the importance of human capital in determining socioeconomic well-being, are arguably more willing and able to help in the education of their children (Chen, 1998). However, the exclusion restriction that a mother's education has no direct effect on earnings is equally, if not more, problematic. Trostel *et al.* (2002) argue that it is not difficult to think of good reasons why it should not hold. Even if family characteristics are considered to be correlated with earnings, they are still widely used in the literature.

Theoretically if the instrument is valid, IV estimates yields consistent estimates of the impact of schooling on earnings. As we have discussed, a suitable instrument must meet two conditions: relevance and exogeneity. First, relevant instruments are highly correlated with the endogenous regressor even after controlling for other exogenous regressors. The relevance of the instruments can be tested in the first-stage regression. As suggested by Bound *et al.* (1995), we report F-tests on the excluded variables and the partial R^2 from the first stage regression[43]. As shown by Staiger and Stock (1997), the first-stage F-statistic on the excluded instrument served as an indicator for weak identification. Bound *et al.* (1995) show, that instrumental validity exists if there is a strong correlation between the potential instrument and education. Intuitively, the stronger the association, the stronger will be the identification of the model. If the potential instrument is weakly correlated with schooling, it is likely to obtain much less precise estimated results. The exogeneity requirement, however, needs a stronger theoretical argument. As for the exogeneity test, to ensure that the selected instruments have no relationship with the unobserved error term, the over-identifying restrictions test is performed and a second instrument is utilised.

An important potential source of regional divergence, featured in the literature, is a local human capital concentration. A recent article by Jurajda (2011) indicates that

[42] Parent's education was one of the first instruments proposed in the literature to obtain IV estimators. This choice has recently been criticized because parent's education is arguably not exogenous to the wage determination process.

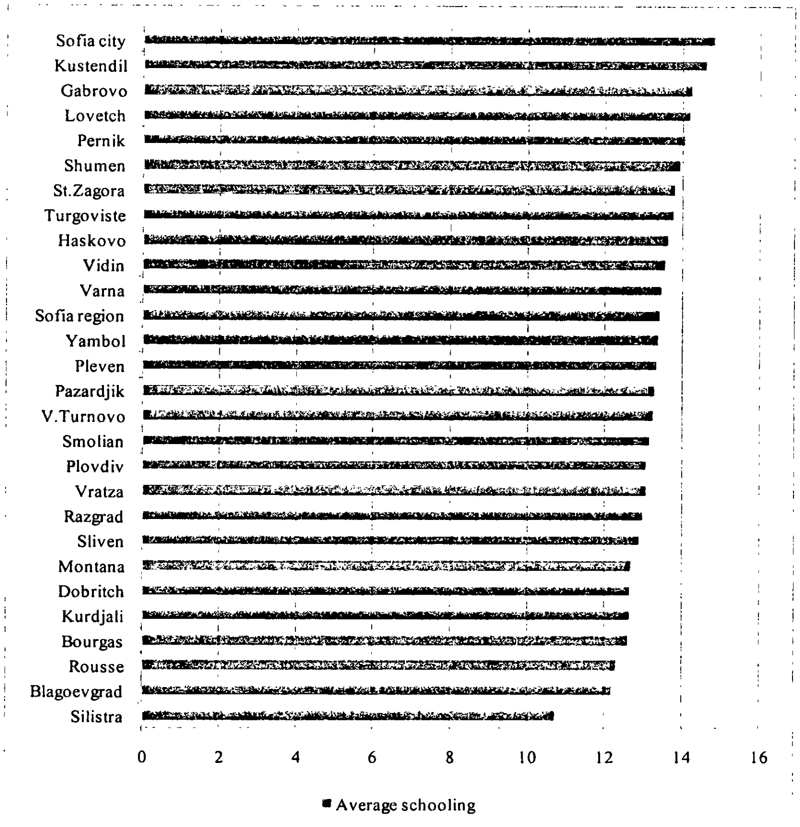
[43] To examine whether the instrument is sufficiently correlated with the amount of education, Bound *et al.* (1995) have proposed researchers to report the partial R^2 and the F-statistics of the identifying instruments in the first-stage estimation. As a rule of thumb, the F-statistic of a joint test whether all additional instruments are significant should be bigger than 10. Partial R^2 relates to the variation in the first stage regression estimates of the year of schooling on the instrument and exogenous variables.

areas with a higher concentration of human capital experience a larger increase in their human capital endowment, and have higher returns to education. There exists a significant positive impact from living in a 'good' neighbourhood for children's educational attainment (Chen, 1998). Based on that theoretical argument, we propose as a second instrument in the analysis the regional average differences in schooling. The average educational attainment may be higher in some districts because of the presence of a college, or some other geographical dimension factors that affect schooling decision. Any significant effects stemming from differences in geographic location point to differences in the macro-level structure of those regions. Specifically, we look for significant effects on educational attainment from different geographic locations. In our data, the district's educational composition provides evidence of regional divergence driven by a concentration of human capital and we can see a unique schooling distribution in each region.

This allows us to generate the average schooling instrumental variable as a continuous variable based on the means of total number years spent in education within 28 different regions in Bulgaria, respectively for 1986, 1995, 1997, 2001 and 2003[44]. Significant difference does exist between educational levels in these 28 regions. We find high concentration of more endowed workers within the capital Sofia. For instance, in 2003 individuals living in district of Silistra have on average 10.7 years spent in education compared to the average of 14.8 years for the respondents in Sofia city (see Figure 4.2).

[44] Since 1999, Bulgaria has been divided into 28 provinces (Bulgarian: области, *oblasti;* also translated as 'region') which correspond approximately to the 28 districts that existed before 1987.

FIGURE 4.2: Regional variations in the average schooling in Bulgaria, 2003



Source: Bulgarian LSMS 2003

Finally, we report the Durbin-Wu-Hausam (DWH) test for endogeneity of schooling, which allows us to reject or accept the null hypothesis that the OLS estimates are consistent (Davidson and MacKinnon, 1993). If the DWH test statistics is significant, we reject the null hypothesis and we show that the endogeneity adjusted results are more robust.

4.2.3. Heckman two-step estimation for sample selection

Sample selection bias arises as individuals have a choice whether to work, and thus earnings are observed only for those in work. This decision is not made randomly and is typically linked to certain characteristics. We address the question of the potential impact of sample selection bias on returns to education by adopting Heckman's (1979) two-step procedure. The sample selection model has the following form:

Wage equation: $\ln Y_i = \alpha + \beta S_i + u_{1i}$ (4.3) where $\ln Y_i$ is the log of earnings, S_i is schooling for individual i;

Selection equation: $z_i\gamma + u_{2i} > 0$ (4.4)

where $u_1 \sim N(0, \sigma)$

$u_2 \sim N(0, 1)$

corr $(u_1, u_2) = \rho$

In the first step, the female or male probability of being employed is estimated from an auxiliary probit regression on a sample of employed and non-employed individuals to obtain estimates of γ.

$$\Pr(y_i \text{ observed} \mid z_j) = \Phi(z_j\gamma)$$

Thus, following Heckman two-step procedure, the selectivity corrected wage equation can be estimated by Ordinary least Squares (OLS) which provide unbiased coefficient estimates for the population as a whole, since the selection variable removes the part of the error term that is correlated with the explanatory variables:

$$\ln Y_i = \alpha + \beta S_i + \sigma\hat{\lambda}_{ij} + \omega_{ij} \quad (4.5)$$

λ_{ij} is estimated from the employment probit $\hat{\lambda}_i = \dfrac{\phi(\hat{\gamma}_j z_j)}{\Phi(\hat{\gamma}_j z_j)}$ and is the inverse of Mills' ratio, where $\phi(.)$ and $\Phi(.)$ denote the probability density and cumulative distribution functions of the standard normal distribution, σ is the covariance between the error term in the wage equation and the employment equation.

The dummy variables that define labour force participation decision of females and males are set to 1 if they are economically active and 0 otherwise. As for the identification the number of young children under 6 years in the family is used, assuming that it is an exogenous variable. The rationale for this is that it should influence the employment participation of women or men in the labour market but not their performance and earnings. We expect lower employment participation especially for females if there are dependent children in the household[45].

[45] The presence of young children is expected to reduce female participation since with fewer children or older children, women will be better positioned to avail themselves for work outside the home. For consistency the same identifying variable has been applied in male's specification, where the assumption is that the presence of younger children is expected to encourage male participation since men are likely to be the main family

4.2.4. Semi-parametric two-step for sample selection

An alternative approach to account for the sample selection bias is based on Newey (1991, 2009) two-step estimator. The estimator is analogous to Heckman (1976) two-step procedure for the Gaussian disturbances case. This approach amounts to replacing a linear regression of the selection term with a series expansion. The method is described in details in Buchinsky (1998). The estimation is semi-parametric in a sense that it does not assume any distributional form for the disturbance term[46]. However, it assumes that the choice probability function depends on the parametrically specified index function[47]. To perform the semi-parametric correction procedure, we use a power series expansion of the inverse Mill's ratio of the normalized estimated index through participation decision in a first step and relax the normality assumption[48]:

$$g_i = \gamma z_i + \xi_i \tag{4.5}$$

where g_i is an index function.

To get unbiased estimates of β for the male and female respondents it is necessary to introduce an extra term:

$$\ln Y_i = \beta S_i + h(g_i) + \xi_{\theta i} \tag{4.6}$$

This new term $h(g_i)$ includes information about the unobservable characteristics of the respondents which affect their labour force participation decision and it plays a similar role as the Mills ratio in the usual Heckman (1979) two-step procedure. The estimated probability function provides the location for the index $g(\hat{\gamma}Z_i)$ and the

supporters. In addition, expectations about men's involvement in the care of children are changing, albeit slowly. In sum, a common trend for families with young children is for women to reduce out of home working hours while men often increase theirs. Some other identifiers were considered such as urban residence which is believed to be associated with various factors that help in reducing fertility and increasing female participation rate. However, the size of Mills ratio does not vary much with the suggested identifying variable.

[46] Semi-parametric models combine components of parametric and nonparametric models. They avoid the assumptions of parametric models and functional form since they do not assume a known link function, and solve the curse of dimensionality of nonparametric models. However, if the distribution of parametric distributional assumption is valid, a semi-parametric estimator will be less efficient than its parametric counterparts (Lei, 2005).

[47] There are many distribution-free estimates that are available for the first-step, including those of Manski (1975) when there is conditional median restriction, Cosslett (1983) when the selection disturbance is independent of regressors, Klein and Spady (1993) and Ichimura (1993) estimators.

[48] We also experimented with power series in the index however the results do not depend on doing so.

values of $g = \hat{\gamma}Z_i$ are now used to expand $h(g_i)$ in a power series by approximating[49]:

$$g(\gamma Z_i) = \sum_{k=1}^{k} \alpha^k (g_i)^{k-1} \tag{4.7}$$

where k is the number of terms in the approximating series and is allowed to grow with the sample size. In the results reported here experimentation with different power series indicated that a second order power series was sufficient in each case[50]. Thus, equation (4.6) can be defined as:

$$\ln Y_i = \beta S_i + \sum_{k=1}^{k} \alpha^k (\hat{\gamma} Z_i)^{k-1} + \xi_i \tag{4.8}$$

4.2.5. Quantile regression approach

Distributional approach is based on the use of quantile regression (QR) method (Koenker and Bassett, 1978)[51] which provides estimates of the effect of education on earnings at different points in the earnings distribution. Estimating the effect of education at conditional quantiles, therefore, allows for heterogeneity in the returns to education. Just as OLS models the conditional mean of the dependent variable Y relative to the covariates X used in the analysis, QR gives estimates of the effect of covariates at different percentiles of the conditional earnings distribution[52]. In a wage equation setting, the QR model can be written as:

$$\ln Y_{\theta i} = X_i \beta_\theta + u_{\theta i} \text{ with } Q_\theta(\ln Y_i \mid X_i) = \beta_\theta X_i \tag{4.9}$$

where as before, the notation $\ln Y_i$ denotes the logarithm of hourly earnings for the sample of individuals $i=1 \ldots n$, $Q_\theta(\ln Y_i \mid X_i)$ denotes the conditional quantile θ of $\ln Y_i$, conditional on the regressor vector of characteristics X_i, β_θ denotes the

[49] Newey (2009) explains that this power series can lead to several different types of sample selection correction and we can have a power series in the index, in the inverse Mills ratio or in the normal cumulative density function (CDF). Newey (2009) also allows the correction term to have an unknown functional form.
[50] According to Buchinsky (1998) addition of more terms to the series expansion resulted, in some cases, in severe multi-collinearity problems.
[51] Among many others, Buchinsky (1994, 1998) and Powell (1986) extend the use of QR to get information about the effect of exogenous explanatory variables on the dependent variable at different parts of the distribution.
[52] Another advantage of QR is that it tends to be less sensitive to the presence of outliers in the dependent variable. This is because in the QR the residuals to be minimized are not squared as in the OLS, and as a result outliers receive less emphasis. Moreover, if the error term of the regression is not distributed normally, the QR may be more efficient than the mean regression (Buchinsky, 1998).

vector of quantile regression coefficients and $u_{\theta i}$ denotes the random error term with an unspecified distribution.

The θ^{th} regression quantile, $0 < \theta < 1$ are defined as a solution to the problem:

$$\min_{\beta \in R^k} \left\{ \sum_{i:\ln Y_i \geq x_{i\beta}} \theta \mid \ln Y_i - \beta_\theta X_i \mid + \sum_{i:\ln Y_i < x_{i\beta}} (1-\theta) \mid \ln Y_i - \beta_\theta X_i \mid \right\} \tag{4.10}$$

This is written as:

$$\min \sum_i \rho_\theta (\ln Y_i - \beta_\theta X_i) \tag{4.11}$$

where $\rho_\theta(\varepsilon)$ is the check function defined as $\rho_\theta(\varepsilon) = \theta_\varepsilon$ if $\varepsilon \geq 0$, or $\rho_\theta(\varepsilon) = (\theta - 1)\varepsilon$ if $\varepsilon < 0$. The model specifies the θ^{th} quantile of the conditional distribution of the $\ln Y_i$ given the covariates X_i as:

$$Q_{\theta i}(\theta \mid X_i) = \beta_\theta X_i, \theta \in (0,\ 1) \tag{4.12}$$

We obtain different quantiles by increasing θ from 0 to 1. As θ is increased, the entire distribution of $\ln Y_i$ is traced conditional on X_i. We assume that both $\ln Y_i$ and X_i are observed with no error and that equation (4.9) is correctly specified. Thus, we can view the model as the best linear predictor for the conditional quantile. However, the distribution of the error term is left unspecified and we only assume that the θ^{th} quantile of the error term is zero, so that $Q_\theta(\varepsilon_{\theta i} \mid X_i) = 0$. The median regression can be defined by minimizing the sums of the absolute errors rather than minimizing the sum of squares as in the OLS framework. This estimator is known as the Least Absolute Deviations (LAD) estimator.

Estimates at different quantiles can be interpreted as showing the response of the dependent variable to the regressors at different points in the conditional wage distribution. The relative positioning of workers in the conditional wage distribution, therefore, can be related to systematic differences in unobservable, which generically may be referred to as 'ability' and include a diverse range of attributes like motivation, labour market connections, family human capital, school quality, etc (Arias *et al.*, 2001). Variation in ability concerns variation in the intercept of the wage equation. One feature of the model is that variation in ability also concerns the slope, or in other words, ability influences the wage-effect of education. If it only influenced the intercept, individuals with higher ability might well invest less in education, since they have a higher opportunity cost of school attendance.

4.3. Empirical results

In Table 4.2 and Figure 4.3 we report the estimated OLS and IV returns to education coefficients for males and females in Bulgaria in five different years. The full set of results applying the OLS, Heckman two-step, semi-parametric, instrumental variable and the QR approaches are given in the Appendix Tables A4.1-A4.15. The probit results are also reported. We have attempted to fit comparable specifications to the samples across years, to put the analysis in a common framework. The vector of control variables includes respondent's potential experience (linear and quadratic terms), variables for years of tenure with the current firm, being married, Bulgarian ethnicity, public sector, urban settlement and living in the capital city. The instrumental variable in the IV estimates is defined as a dichotomous variable that takes value 1 if a respondent's mother reports higher education degree and 0 otherwise. The over-identification restrictions test is performed by combining the two instruments – mother's education and average regional schooling. A test for misspecification of the OLS models is also conducted by using the Ramsey stability test (RESET). The test is obtained by adding powers of the fitted values from the estimated model to the original regression equation. It uses the squares of fitted values to test the assumption that the relationship between the dependent and independent variables is linear and therefore test the correctness of the functional form of the model which is based on the linearity assumption. The diagnostic checks are reported at the bottom panel of Tables A4.1, A4.3, A4.5, A4.7 and A4.9.

Turning to the OLS estimates in Table 4.2 and Figure 4.3, we can see that for both men and women there was a rise in the returns to education in the early 1995, followed by a decline in 1997. Overall, the returns to education in Bulgaria are increasing over time. More specifically, the OLS estimate of the rate of returns to education for men rises from 1.1 % in 1986 to 5.1% in 2003 and from 2.1% to 5.9% for women. This finding is in line with Arandarenko *et al.* (2006) who conclude that in Bulgaria the returns to all types of educational degrees increased substantially between 1995 and 2003. The very low return to education in 1986 suggests that the structure of wage rates in a centrally-planned economy does not create incentives for obtaining more education. The returns of 1.1% for males and 2.1% for females in 1986 are also in line with Hung (2002), who reports return to education of 1.9% for Bulgaria in 1989/1990. Clearly, there is a marked downturn in estimated returns after 1995, which might be due to the recession of 1996/97[53]. Accordingly, there is a recovery after 2001 when the Bulgarian economy has improved. Turning to the market forces, Bulgaria has experienced an increase in the supply of higher education over the period – between 1995 and 2002, the number of university graduates increased from 33,000 to 50,000[54].

The OLS estimates of the return to education are higher for women than for men. For instance, the estimated return to education in 2001 is 4.3% for males and 6.6% for females. By 2003 the estimated return to education for males increases to 5.2%, while females' coefficient fell slightly to 5.9%. The upward trend in the returns to education found here is in line with previous findings. For example, in the Czech Republic the returns to education for men rose from 2.4% to 5.2% between 1983 and 1993 (Chase, 1998). Similarly, increases in returns to schooling also occurred in Poland (Rutkowski, 1996) and Slovenia (Orazem and Vodopivec, 1995).

[53] As we have already discussed during 1996, the Bulgarian economy collapsed due to an unstable and decentralized banking system, a wave of hyperinflation throughout several countries of Eastern Europe, and slow reforms, which led to the collapse of the Bulgarian currency. In 1997 the government set up a currency board to stabilize the currency.

[54] Country Profile (2006) Bulgaria, Library of congress country studies program, p.6.

TABLE 4.2

Estimated returns to education over time in Bulgaria, by gender

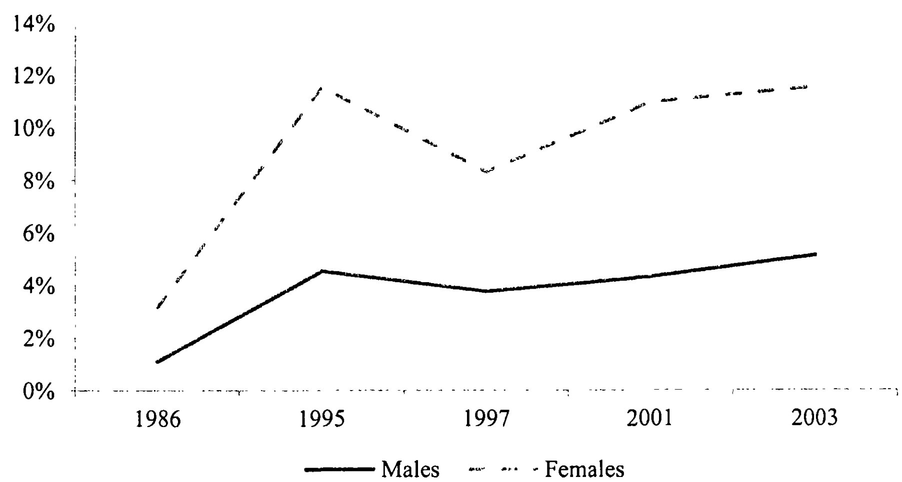
	MALES			*FEMALES*		
	OLS	IV_1	IV_2	*OLS*	IV_1	IV_2
1986	0.0110***	0.0215***	0.0228***	0.0207***	0.0415***	0.0415***
	(0.0030)	(0.0079)	(0.0077)	(0.0029)	(0.0072)	(0.0071)
1995	0.0453***	0.0850**	0.0923***	0.0700***	0.0838*	0.0971**
	(0.0097)	(0.0331)	(0.0332)	(0.0113)	(0.0434)	(0.0433)
1997	0.0397***	0.0745***	0.0774***	0.0454***	0.0595***	0.0635***
	(0.0119)	(0.0198)	(0.0197)	(0.0115)	(0.0179)	(0.0174)
2001	0.0431**	0.1281**	0.1036***	0.0664***	0.1005***	0.0983***
	(0.0117)	(0.0614)	(0.0474)	(0.0094)	(0.0212)	(0.0184)
2003	0.0512***	0.0996**	0.1144***	0.0598***	0.1051***	0.1056***
	(0.0048)	(0.0496)	(0.0310)	(0.0046)	(0.0143)	(0.0132)

Notes: ***, ** and * denote significance at the 1%, 5% and 10% level respectively;
IV_1 – specification where mother's education is used as an instrument; IV_2 – over-identification specification - both mother's education and average regional differences in schooling are used as instruments.

The stability tests used to detect the general functional form misspecification in the OLS, report insignificant χ^2 and F-statistics and thus indicate that we cannot reject the null hypothesis of no model specification problem.

FIGURE 4.3

Returns to education over time in Bulgaria, males and females



Notes: OLS estimates of return to education, Bulgaria.

The corresponding IV results reveal a similar increasing pattern. For instance, by using mother's education as an instrument, the IV coefficient for men rises from 2.2% in 1986 to 9.9% in 2003. Similarly for women, the IV estimates increase from 4.2% in 1986 to 10.5% in 2003. We obtain higher IV estimates for both genders than the corresponding OLS and this is in accordance with the results reported in the literature. For example, our IV estimate for males in 2003 is 42% higher than the corresponding OLS estimate. Similarly, for females in 2003, the IV estimate of the return to education increases by 47%. The coefficients do not change significantly in the over-identified specifications where both instruments – a mother's education and regional average schooling are used. For instance, the IV estimates for males in 1986 are 2.2% when the mother's education is used as an instrument and 2.3% in the over-identification specification by adding average regional schooling as an additional instrument.

We also report the results from the first stage regressions. As might be expected, there is a strong relationship between a mother's education and individual's school attainment. Throughout all the years the instrument does show a strong significant and positive effect on the school outcome of individuals. For both males and females, the *F*-statistics of the excluded instrument indicates the significance of the coefficient in their first stage regression. Moreover, the partial R^2 of the effect of the instrument on the years of schooling is higher than the guidelines given by Bound *et al.* (1995). For instance in 1986 just-identified IV estimates, the *F*-statistics of 511.63 (partial R^2 of 0.138) for men and 586.78 (partial R^2 of 0.154) for women confirm validity of the instrument and reject the null hypothesis that this instrument is weak (the critical value is 16.38). Therefore we can conclude that in the 1986 specification, men and women with highly educated mothers have on average 3.3 more years of schooling compared to their counterparts (see Table A4.2). Slightly lower, but still valid following Bound *et al.* (1995), are the tests for the 1995 just-identified estimates, where the tests of excluding the instrument from the reduced form equations, yield an *F*-statistics of 54.13 (partial R^2 of 0.073) for males and 36.03 (partial R^2 of 0.057) for females (see Table A4.4).

Our main concern however, lies in the direct effect of the mother's education on earnings. In an exactly identified model we cannot test the hypothesis that the instrument is valid. In that case, the assumption that the instrument is valid will essentially have to be taken in faith (Söderbom, 2011)[55]. As we discussed, having more than one potential instrument, will enable us to undertake the Sargan instrument validity test of the over-identifying restriction and to provide support for our empirical approach. The Sargan test results in failure to reject the null hypothesis[56] in all specifications (for instance, in 1986 estimates the p-value of 0.4748 for males and 0.9497 for females). The validity tests were passed for both males and females, which provides some support for the approach adopted here. In addition, we apply the Stock and Yogo (2005) (SY) test. We compare the minimum eigenvalue statistic (MES) with the critical values (we are willing to accept at most a rejection rate of 10% of a nominal 5% Wald test) in the 1986 sample. Since the MES statistics of 271.25 for males and 299.68[57] for females exceed the critical value of 19.93, we passed the SY test in the over-identified model.

The endogeneity test (DWH test[58]) indicates that the hypothesis of endogeneity of education is rejected at a 1% level only in female specifications for 1986 and 2001 and for both male and female over-identified specifications in 2003, and we can conclude that the OLS estimates tend to underestimate the average effect of schooling. The DWH test is also rejected at a 10% level of significance for males in 1997 estimates. The sign of the bias is in accordance with the results reported in the literature (Card, 1999, 2001; Griliches, 1977). In all other specifications we do not reject the null hypothesis (for instance in 1995, DWH p-value=0.2059 for males and p-value=0.7428 for females, see Table A4.4).

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[55] In order to show why we cannot test for validity of this assumption, the following exactly identified model can be considered: $y_1 = \beta_0 + \beta_1 y_2 + u_1$

$$y_2 = \pi_0 + \pi_1 z_1 + v_2$$

Expressing the structural equation as a function of the predicted value of leads:

$$y_1 = \beta_0 + \beta_1(\hat{\pi}_1 z_1) + u_2 = (\beta_0 + \beta_1\hat{\pi}_0) + \beta_1(\hat{\pi}_1 Z_1) + u_1$$

One cannot test the hypothesis $\text{cov}(z_1, u_1) = 0$, simply because u_1 is not observed and without further information, is not possible to obtain an estimate of u_1 unless we assume $\text{cov}(z_1, u_1) = 0$. That is, the estimate of u_1 will be uncorrelated with z_1 by construction.

[56] The null hypothesis of over-identifying restriction test is that all included instrumental variables are jointly exogenous.

[57] These are statistics in the 1986 over-identified specification.

[58] A DWH test is performed by comparing $\hat{\beta}_{OLS}$ and $\hat{\beta}_{IV}$ estimates. The null hypothesis is that there are no endogenous variables or that endogeneity does not affect the OLS estimator. Under the null hypothesis, $\hat{\beta}_{OLS}$ is consistent and efficient and $\hat{\beta}_{IV}$ is consistent but inefficient. Under alternative hypothesis, $\hat{\beta}_{OLS}$ is inconsistent but $\hat{\beta}_{IV}$ remains consistent.

The coefficients on the return to education do not change significantly between Heckman and semi-parametric approaches. For example, the estimated rate of return to education for males in 2003 using the Heckman parametric two-step approach is 4.8%, while the semi-parametric approach gives an estimate of 4.7%. The similarity of the parametric and semi-parametric two-step estimates suggests that inclusion of the two approximation terms is essentially performing the same task as the inclusion of the inverse Mills ratio. A similar finding was provided in Vella (1998). Moreover, Heckman two-step estimates do not confirm selectivity bias as the included Mills ratio is statistically insignificant in all specifications for both males and females. This also suggests that there is no correlation between the error terms in the earnings equation and in the selection equation. Apart from the 1995 male estimates, the coefficients of the two approximation power series correction terms in the semi-parametric approach are jointly not significant in both male and female specifications. Therefore, we may draw a conclusion that the corrected returns to education in the Heckman specification is not significantly different from the one obtained in the straightforward OLS specification.

Importantly, the selection equation (probit equation) has to be properly specified. The dependent variable in our probit specifications is a dummy variable that takes the value 1 if females or males are in employment and 0 otherwise. The possible exclusion variable in our data sets is the number of children up to the age of 6, currently present in the family. The presence of young children is expected to reduce female participation since with fewer children or older children; women will be better positioned to avail themselves for work outside the home. For consistency the same identifying variable has been applied in male's specification, where the assumption is that the presence of younger children encourages male participation since men are likely to be the main family supporters. In addition, expectations about men's involvement in the care of children are changing, albeit slowly. In sum, a common trend for families with young children is for women to reduce out of home working hours while men often increase theirs. In order to eliminate the possibility of reaching a conclusion regarding selectivity bias based on misspecification of the model, we experimented with several specifications. Some other identifiers were considered such as urban residence which is believed to be associated with various factors that help in reducing fertility and increasing female participation rate. The size of Mills ratio does not vary much with each specific choice of identifying variables included in the probit function and all specifications lead to the conclusion that no selectivity bias is present in the data[59]. The probability models indicate that having children in a household does seem to have significant effect in females' participation decision only in 1986, 1995 and 1997 specifications. In addition, married females in the 1986 and 1995 specifications are less likely to be in the labour force.

There are several potential explanations for the increasing trend in returns to education in Bulgaria. In the literature, a lot of emphasis has been placed on the role of skill-based technical changes[60]. By shifting the relative demand of educated labour relative to available supply, the skill-based technical change has increased the returns to education (Denny and Harmon, 2001). The skill-based technical progress increased the relative productivity of skilled labour and generated a continuous upward shift in its demand (see Acemoglu and Autor, 2010, for a recent discussion). Supply and demand factors affecting the returns to education include changes in the relative supply of educated workers, the changing composition of the labour force as retiring cohorts were replaced with younger workers over time, skill-biased technological change and globalization of the economy, i.e. international trade (Zhang *et el.*, 2005). Desjonqueres *et al.* (1999) emphasizes that the rise in skill premium has been accompanied by increases in the ratio of skilled to unskilled employment in all sectors. Moreover, the skill premium has risen in less-developed and newly industrializing countries as well as in OECD countries.

The most obvious explanation for the steady increase in the returns to education in Bulgaria is that as a transition economy, the country moved from a system of government compressed wage scales to a market-oriented system. The downsizing of heavy industry in Bulgaria during the transition period caused a reduction in the demand of unskilled labour. The expansion of skill-insentive services (for example, finance, insurance and real estate, information services) increased the demand for more educated employees[61]. While the current study is unable to examine a

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[59] Since most women in Bulgaria are in work this result is expected.

[60] See Card and Lemieux (1999) for a recent review.

[61] As shown in Appendix Table A4.16 the service sector in tersm of employment expanded significantly. Between 2000 and 2003, for example, employment in the real estate sector increased by 27%. Further

systematic link between returns to education and economic reforms, as in Flabbi *et al.* (2008) we can indicate that heterogeneity in estimated returns to education coefficients suggests that non-competitive forces might have determined wage levels and their distributions in Bulgaria. We draw this conclusion from the QR estimates showing how return to education has evolved at different points of the earning distribution. Table 4.3 shows estimated returns to education over time at the 10th, 25th, 50th, 75th and 90th percentiles for males and females respectively. The full set of estimates is presented in Appendix Tables A4.11 – A4.15.

TABLE 4.3

Evolution of returns to education across distribution, Bulgaria

			Males		
QR	θ=0.10	θ=0.25	θ=0.50	θ=0.75	θ=0.90
1986	0.0281***	0.0205***	0.0098***	-0.0005	0.0000
	(0.0043)	(0.0034)	(0.0033)	(0.0035)	(0.0001)
1995	0.0451***	0.0292***	0.0236**	0.0308**	0.0395*
	(0.0161)	(0.0086)	(0.0094)	(0.0122)	(0.0218)
1997	0.0172	0.0270	0.0352***	0.0450***	0.0678***
	(0.0205)	(0.0167)	(0.0107)	(0.0144)	(0.0161)
2001	0.0344**	0.0310***	0.0427***	0.0305***	0.0166
	(0.0146)	(0.0090)	(0.0093)	(0.0104)	(0.0198)
2003	0.0383***	0.0463***	0.0461***	0.0490***	0.0589***
	(0.0057)	(0.0055)	(0.0046)	(0.0049)	(0.0082)
			Females		
QR	θ=0.10	θ=0.25	θ=0.50	θ=0.75	θ=0.90
1986	0.0524***	0.0305***	0.0171***	0.0082**	-0.0032
	(0.0050)	(0.0027)	(0.0029)	(0.0035)	(0.0021)
1995	0.0557***	0.0526***	0.0644***	0.0448***	0.0458**
	(0.0093)	(0.0058)	(0.0070)	(0.0120)	(0.0192)
1997	0.0571***	0.0383***	0.0492***	0.0604***	0.0436**
	(0.0144)	(0.0116)	(0.0153)	(0.0175)	(0.0213)
2001	0.0591***	0.0672***	0.0566***	0.0434***	0.0533***
	(0.0097)	(0.0069)	(0.0066)	(0.0130)	(0.0139)
2003	0.0478***	0.0505***	0.0625***	0.0699***	0.0673***
	(0.0073)	(0.0034)	(0.0066)	(0.0063)	(0.0080)

Notes: Standard errors in parenthesis; ***, ** and * denote significance at the 1%, 5% and 10% level respectively.

The results demonstrate that the average returns to education in Bulgaria, for both males and females, are driven by an increasing trend in return in the upper end of the conditional wage distribution. In 1986 the point estimates at the top of the distribution are insignificant and even negative. In 2003, the rate of return to education at the 90th quantile rises to 5.9% for males and 6.7% for females. For

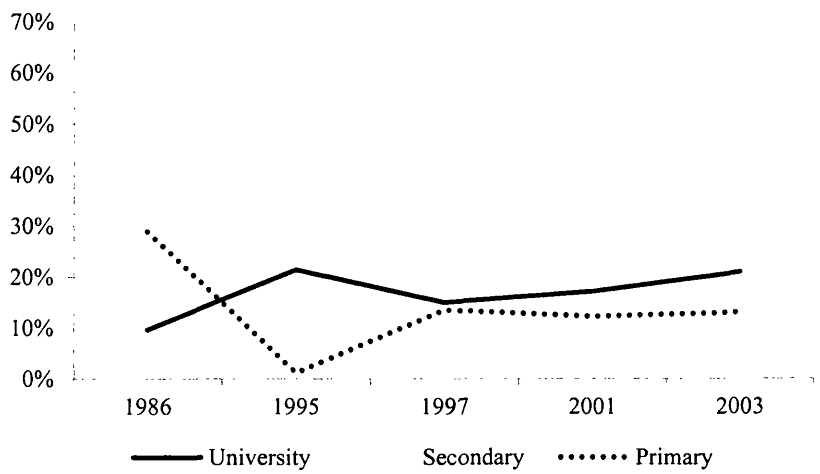
examination of the explanations for the increased return to education refer back to Table 3.5 which shows that the increased rate of return to education is unlikely to be due to increased shortatge of highly educate labour arising from changes in the educational participation rate (since this rate has been rising).

females, the changes in returns to education at the lower part of the distribution are roughly proportional over time. The spread in returns between the 10th and 90th percentile is higher for males as compared with females.

In addition, Figure 4.4 shows the evolution of the employment structure based on qualification in Bulgaria. One can see that employment possibilities for primary educated workers declined considerably during the first years of transition and the share of university workers rapidly increased in 1995, thus we may suggest that adjustment in the labour market could have been made through the change in supply of educated workers.

FIGURE 4.4

Evolution of the structure of employees based on qualifications in Bulgaria



Notes: The figure is based on the descriptive statistics for workers who report wages.

Finally we briefly consider the return to the other covariates included in the models. The estimated coefficients in the potential experience are insignificant in most of the OLS estimates and this may be explained by the structural reforms that have taken place in Bulgarian economy during the transition. Only in 2003 the experience coefficients are statistically significant and earnings tend to increase with tenure. Married women in 1986 report 37% higher earnings when compared to other categories (single and divorced). The results also show that workers who reside in an urban or capital city show positive relationship between years of schooling and the type of settlement and thus earnings increase with urban residence. This is consistent with expectations that individuals living in urban areas have more opportunity to exploit skills acquired through higher education than do those living in rural areas. One would have expected public sector employment to be associated with higher earnings. However, the coefficients on public sector employment tend to be insignificant in most of the OLS specifications and even negative in value for the 1995 OLS estimates. Only for the 1997 and 2003 OLS estimates, the public sector wages are statistically significant standing at 19.7% for males and 5.9% for females in the latest year. Looking at the returns to public sector across the conditional earnings distribution, however, it appears that for both males and females in 1986, returns to public sector employment tend to be significant and positive in the lower part of the distribution, and significant and negative in the top end.

4.4. Conclusions

The empirical analysis in this Chapter provides a broad assessment of the returns to education in Bulgaria over the transition to market economy. The main question addressed in the Chapter concerns the hypothesis whether Bulgarian returns to education rise or fall during the transition. Unlike most studies on returns to education, we have taken into account several empirical methods – the OLS, Heckman two-step, semi-parametric two-step approach, instrumental variable and the QR. Based on the empirical estimations, the following conclusions can be drawn:

• Over the period of transition, which continued until the early 2003, we find an upward trend in the evolution of returns to education in Bulgaria. The average returns to one additional year of education raised from 1.1% to 5.1% for males and from 2.1% to 5.9% for females. Similarly, in the IV just-identified estimates, return to education increased from 2.2% in 1986 to 9.9% in 2003 for males and from 4.2% to 10.5% for females. The results are not surprising, given the documented increase in the rate of returns to education over the transition period. Estimates of earning function using data from Russia and Eastern Europe exhibit the same increasing tendency (Brainerd, 1998; Chase, 1998). These results clearly suggest that market reforms in Bulgaria led to the productivity augmenting factors being rewarded accordingly.

• The OLS estimates show that returns to education in Bulgaria are statistically significant and lower than the international average of 7% to 8% and certainly broadly similar to other transition countries, such as the Ukraine for example[62]. Our findings of a very low return to education in 1986 are in line with previous studies that estimate returns to education in transition economies (Brainerd, 1998; Katz, 1999; Newell and Reilly, 1996).

• In line with Flabbi *et al.* (2008) findings, our results confirm that the largest increase in the rate of returns to education took place in the early transition period. For males, the OLS estimate of the return to education increased from 1.1% in 1986 to 4.5% in 1995 and from 2.1% to 7.0% for females.

• We have been able to control for the endogeneity of education by employing the IV estimation procedure. The IV estimates, when using mother's education as an instrument increased by approximately 42% in all specifications, and this increase is clearly documented in the literature. Mother's education presents a strong correlation with the schooling variable. The instrumental validity was strengthened by the over-identification tests, which result in failure to reject the null hypothesis in all specifications. However, the DWH test rejected the null hypothesis of exogeneity only for females in 1986 and for both males and females in 2003 over-identified specifications.

• Selectivity bias with respect to the participation of men and women in the labour market shows that there are no reasons to suspect the estimates of returns to education are misleading. Only in the 1995 male estimates, the two approximation selectivity terms are statistically significant at 5% level.

• The similarity of the parametric and semi-parametric two-step estimates indicates that inclusion of the two approximation terms perform the same task as the inclusion of the inverse Mills ratio.

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[62] See Gorodnichenko and Sabirianova (2005).

• We found that not all individuals benefited equally from education. Early in transition, individuals at the lower part of the distribution appear to have benefitted more from education compared to their counterparts at the upper part of the distribution, lending support to the notion that education could have been substituted for the low ability. Whereas the return for males in 1986 at the 10th percentile was 2.8%, it tended to be insignificant at the 90th percentile. Similarly for females, the rate of returns to education at the 10th percentile was 5.2% and it turned to negative at the top of the distribution.

• Over the examined period, the most prominent increase in the wage premium occurred at the top of the distribution, where the rate of returns to education, in particular for females, increased from a negative and insignificant sign in 1986 to 7% in 2003.

CHAPTER FOUR APPENDIX

TABLE A4 Classification of the levels of education

Terms used in the study	ISCED classification
Primary education The first stage of education designed to provide a sound of basic education in reading, writing, and mathematics. Entry age: between 6 and 7.	ISCED 1
Lower secondary education Completes provision of basic education, usually in more subject-oriented way. Duration is 3 years.	ISCED 2
Upper secondary education Stronger subject specialisation than at lower secondary level. Students typically expected to have completed 9 years of education or lower secondary schooling before entry and are generally 15 or 16 years old.	ISCED 3
Post-secondary non-tertiary education This level straddles the boundary between upper secondary and post-secondary education. Programme content may not be significantly more advanced than that in upper secondary, but is not as advanced as that in tertiary programmes. Duration usually is the equivalent of between 6 months and 2 years. Students tend to be older that those enrolled in upper secondary education.	ISCED 4
Tertiary education Programmes designed to provide qualifications for entry to advanced research programmes. Duration at least 3 years.	ISCED 5
Advanced research programmes Programmes that lead to the award of an advanced research qualification, e.g. Ph.D. The duration of these programmes is 3 years	ISCED 6

TABLE A4.1

Returns to education in Bulgaria, by gender, 1986

	Males				Females			
	OLS	Heckman	Probit	Semi-parametric	OLS	Heckman	Probit	Semi-parametric
	(1)	*(2)*	*(3)*	*(4)*	*(1)*	*(2)*	*(3)*	*(4)*
school	0.0110***	0.0110***	-0.0113*	0.0109***	0.0207***	0.0208***	0.0120**	0.0205***
	(0.0030)	(0.0031)	(0.0060)	(0.0031)	(0.0029)	(0.0029)	(0.0060)	(0.0030)
exp	0.0000	-0.0000	-0.0069***	-0.0000	-0.0004	-0.0003	0.0069***	-0.0004
	(0.0009)	(0.0010)	(0.0013)	(0.0010)	(0.0007)	(0.0009)	(0.0013)	(0.0010)
expsq	0.0000	-0.0000	-0.0002***	0.0000	0.0000	0.0000	0.0002***	0.0000
	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
1-2 years	0.1222	0.1248	0.1804	0.1243	0.0247	0.0208	-0.1797	0.0239
	(0.0753)	(0.0768)	(0.1248)	(0.0771)	(0.0733)	(0.0746)	(0.1247)	(0.0748)
3-5 years	0.1237*	0.1254*	0.1379	0.1250*	-0.1465**	-0.1483**	-0.1369	-0.1478**
	(0.0644)	(0.0649)	(0.0982)	(0.0651)	(0.0585)	(0.0588)	(0.0982)	(0.0589)
6-10 years	-0.0358	-0.0359	0.0417	-0.0359	-0.2235***	-0.2219***	-0.0394	-0.2242***
	(0.0652)	(0.0652)	(0.0975)	(0.0652)	(0.0575)	(0.0578)	(0.0975)	(0.0580)
>10 years	0.0087	0.0058	-0.1319	0.0064	-0.1642***	-0.1582**	0.1370	-0.1647**
	(0.0699)	(0.0723)	(0.1062)	(0.0730)	(0.0618)	(0.0657)	(0.1062)	(0.0669)
married	0.2166***	0.2133***	0.1033**	0.2139***	0.3754***	0.3797***	-0.1026**	0.3758***
	(0.0258)	(0.0324)	(0.0504)	(0.0335)	(0.0289)	(0.0337)	(0.0504)	(0.0353)
bulgarian	0.1367***	0.1357***	-0.1103**	0.1360***	0.0629**	0.0650**	0.1096**	0.0629**
	(0.0275)	(0.0281)	(0.0492)	(0.0282)	(0.0291)	(0.0298)	(0.0492)	(0.0302)
public	0.0167	0.0171	0.0275	0.0171	-0.0362	-0.0371	-0.0285	-0.0363
	(0.0270)	(0.0270)	(0.0441)	(0.0271)	(0.0254)	(0.0256)	(0.0441)	(0.0257)
urban	0.0596***	0.0582***	-0.1156***	0.0585***	0.0398**	0.0424*	0.1148***	0.0407*
	(0.0187)	(0.0205)	(0.0381)	(0.0210)	(0.0193)	(0.0220)	(0.0381)	(0.0224)
sofia region	0.0572**	0.0532	-0.3110***	0.0546	0.0432*	0.0501	0.3061***	0.0467
	(0.0266)	(0.0344)	(0.0579)	(0.0378)	(0.0262)	(0.0371)	(0.0579)	(0.0382)
mills		0.0217				0.0362		
		(0.1237)				(0.1433)		
children			0.1148***				-0.1151***	
			(0.0204)				(0.0204)	
$(z'\gamma)$				0.0450				0.1432
				(0.1796)				(0.3354)
$(z'\gamma)^2$				-0.0178				-0.0848
				(0.1266)				(0.2224)
constant	0.4777***	0.4739***	1.1026***	0.4647***	0.5042***	0.4525**	-1.1165***	0.4457*
	(0.1021)	(0.1034)	(0.1604)	(0.1226)	(0.0921)	(0.2228)	(0.1605)	(0.2399)
N	3213	3213	6450	3213	3238	3238	6450	3238
R^2	0.0724	0.0724		0.0725	0.1096	0.1096		0.1097
F	16.08	14.86		13.81	24.42	22.53		20.92
Log-likelihood	-1946.76	-1946.75	-4390.15	-1946.74	-1888.39	-1888.34	-4389.75	-1888.22
Ramsey test	χ^2=2.81[0.038]				χ^2=0.32[0.810]			

Notes: Standard errors in parentheses; ***, ** and * denote significance at the 1%, 5% and 10% level respectively.Figures in squared brackets are p-values.

TABLE A4.2

IV estimates of returns to education in Bulgaria, by gender, 1986

	Males				Females			
	First stage	IV_1	First stage	IV_2	First stage	IV_1	First stage	IV_2
	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
pschool		0.0215***		0.0228***		0.0415***		0.0415***
		(0.0079)		(0.0077)		(0.0072)		(0.0071)
exp	0.0153***	-0.0000	0.0156***	-0.0000	0.0373***	-0.0009	0.0377***	-0.0009
	(0.0036)	(0.0006)	(0.0036)	(0.0006)	(0.0034)	(0.0006)	(0.0034)	(0.0006)
expsq	-0.0017***	0.0000	-0.0016***	0.0000	-0.0014***	0.0001***	-0.0014***	0.0001***
	(0.0001)	(0.0000)	(0.0001)	(0.0000)	(0.0001)	(0.0000)	(0.0001)	(0.0000)
1-2 years	-0.1833	0.1270**	-0.2268	0.1276**	-0.4078	0.0312	-0.4294	0.0311
	(0.3535)	(0.0633)	(0.3521)	(0.0633)	(0.3494)	(0.0609)	(0.3489)	(0.0609)
3-5 years	-0.6155**	0.1322***	-0.6293**	0.1332***	-0.2145	-0.1394***	-0.2288	-0.1394***
	(0.2838)	(0.0512)	(0.2827)	(0.0511)	(0.2730)	(0.0476)	(0.2727)	(0.0476)
6-10 years	-1.2236***	-0.0197	-1.2445***	-0.0178	-0.7371***	-0.1993***	-0.7391***	-0.1993***
	(0.2794)	(0.0514)	(0.2783)	(0.0513)	(0.2667)	(0.0471)	(0.2663)	(0.0471)
>10 years	-3.3591***	0.0522	-3.3418***	0.0574	-2.6361***	-0.0900	-2.6132***	-0.0902
	(0.2928)	(0.0614)	(0.2917)	(0.0610)	(0.2871)	(0.0557)	(0.2867)	(0.0556)
married	-0.7533***	0.2210***	-0.7068***	0.2216***	-0.3686***	0.3755***	-0.3429**	0.3755***
	(0.1262)	(0.0227)	(0.1260)	(0.0227)	(0.1341)	(0.0232)	(0.1341)	(0.0232)
bulgarian	0.9950***	0.1246***	0.8758***	0.1232***	1.4311***	0.0327	1.3356***	0.0328
	(0.1297)	(0.0249)	(0.1312)	(0.0248)	(0.1403)	(0.0266)	(0.1431)	(0.0266)
public	1.0041***	0.0054	0.9749***	0.0041	0.9260***	-0.0588**	0.8899***	-0.0587**
	(0.1202)	(0.0231)	(0.1198)	(0.0230)	(0.1245)	(0.0230)	(0.1248)	(0.0230)
urban	1.3298***	0.0430*	1.3073***	0.0410*	1.1953***	0.0092	1.1913***	0.0093
	(0.1018)	(0.0221)	(0.1015)	(0.0219)	(0.1070)	(0.0214)	(0.1069)	(0.0213)
sofia region	2.4348***	0.0233	1.5432***	0.0193	2.0736***	-0.0192	1.5380***	-0.0190
	(0.1616)	(0.0379)	(0.2359)	(0.0375)	(0.1555)	(0.0339)	(0.2251)	(0.0337)
mother_educ	3.2874***		3.2247***		3.2988***		3.2436***	
	(0.1453)		(0.1453)		(0.1362)		(0.1370)	
average_sch			0.3321***				0.2095***	
			(0.0642)				(0.0637)	
constant	11.7082***	0.3449***	8.4913***	0.3291***	9.6087***	0.2733**	7.5809***	0.2740***
	(0.3850)	(0.1211)	(0.7307)	(0.1190)	(0.4119)	(0.1063)	(0.7416)	(0.1057)
N	3212	3212	3212	3212	3237	3237	3237	3237
R^2	0.4537	0.0704	0.4582	0.0708	0.4769	0.1038	0.4787	0.1039
F	221.37	20.18	208.04	20.30	244.94	31.11	227.62	31.16
Log-likelihood	-7476.16	-1950.07	-7462.78	-1949.41	-7553.77	-1898.63	-7548.35	-1898.37
F (excluded)	511.63***		271.25***		586.78***		299.68***	
p-value	0.0000		0.0000		0.0000		0.0000	
Sargan test			0.511				0.004	
p-value			0.4748				0.9497	
Partial R^2	0.1379		0.1450		0.1540		0.1568	
DWH test	1.9807		2.6435		9.8647		10.0199	
p-value	0.1593		0.1039		0.0016		0.0015	

Notes: Standard errors in parentheses; ***, ** and * denote significance at the 1%, 5% and 10% level respectively.
IV_1 – specification where mother education is used as an instrument; IV_2 – over-identification specification - both mother education and average regional differences in schooling are used as instruments

TABLE A4.3

Returns to education in Bulgaria, by gender, 1995

	Males				*Females*			
	OLS	Heckman	Probit	Semi-parametric	OLS	Heckman	Probit	Semi-parametric
	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
school	0.0453***	0.0444***	0.0405***	0.0328***	0.0700***	0.0710***	0.0605***	0.0713***
	(0.0097)	(0.0096)	(0.0185)	(0.0108)	(0.0113)	(0.0115)	(0.0185)	(0.0115)
exp	0.0121	0.0136	0.0230***	0.0285**	-0.0013	-0.0060	-0.0730***	-0.0064
	(0.0122)	(0.0123)	(0.0175)	(0.0135)	(0.0106)	(0.0114)	(0.0175)	(0.0125)
expsq	-0.0003	-0.0003	-0.0015***	-0.0007**	0.0003	0.0004	0.0018***	0.0004
	(0.0003)	(0.0003)	(0.0004)	(0.0003)	(0.0003)	(0.0003)	(0.0004)	(0.0003)
1-2 years	0.0823	0.0811	0.1126	0.0455	0.3543**	0.3496**	0.2125	0.3508**
	(0.1013)	(0.1011)	(0.0085)	(0.1022)	(0.1460)	(0.1450)	(0.2085)	(0.1461)
3-5 years	0.1048	0.1037	-0.1826	0.0770	0.1470	0.1408	0.1826	0.1411
	(0.0829)	(0.0829)	(0.1884)	(0.0835)	(0.1150)	(0.1140)	(0.1884)	(0.1142)
6-10 years	0.2466***	0.2460***	-0.1837	0.2132**	0.1983*	0.1905*	0.1837	0.1910*
	(0.0847)	(0.0846)	(0.1792)	(0.0863)	(0.1119)	(0.1123)	(0.1792)	(0.1130)
>10 years	0.1630**	0.1610**	-0.2733	0.1149	0.1140	0.1103	0.2733	0.1108
	(0.0768)	(0.0767)	(0.1751)	(0.0784)	(0.1095)	(0.1098)	(0.1751)	(0.1104)
married	0.0537	0.0689	0.1737***	0.1801**	0.1429*	0.1252	-0.5737***	0.1221
	(0.0600)	(0.0625)	(0.0199)	(0.0746)	(0.0830)	(0.0793)	(0.1199)	(0.0815)
bulgarian	-0.0298	-0.0274	0.1525	-0.0025	0.0894	0.0979	-0.1525	0.0977
	(0.0800)	(0.0801)	(0.1976)	(0.0794)	(0.0778)	(0.0760)	(0.1976)	(0.0760)
urban	0.1276**	0.1292**	-0.0161	0.1171**	0.0856	0.0811	0.0661	0.0818
	(0.0536)	(0.0536)	(0.0138)	(0.0540)	(0.0654)	(0.0645)	(0.1138)	(0.0644)
sofia_region	0.2209***	0.2252***	0.2915	0.2628***	0.1660**	0.1627**	-0.1915	0.1628**
	(0.0617)	(0.0619)	(0.0344)	(0.0635)	(0.0665)	(0.0665)	(0.1344)	(0.0666)
public	-0.0633	-0.0635	0.1150	-0.0927	0.0925	0.0944	0.1650	0.0948
	(0.0617)	(0.0618)	(0.1304)	(0.0626)	(0.0860)	(0.0860)	(0.1304)	(0.0860)
children			0.1537				-0.9535***	
			(0.0507)				(0.2050)	
mills		0.0596				-0.0692		
		(0.0688)				(0.0505)		
$(z'\gamma)$				0.6597**				-0.0927
				(0.2624)				(0.2065)
$(z'\gamma)^2$				-0.1635**				0.0156
				(0.0755)				(0.1329)
constant	4.1018***	4.0657***	0.8135***	3.8529***	3.1930***	3.2509***	-0.8135***	3.2555***
	(0.1634)	(0.1714)	(0.3062)	(0.1831)	(0.2215)	(0.2104)	(0.3062)	(0.2183)
N	708	708	1379	708	609	609	1379	609
R^2	0.1326	0.1333		0.1361	0.1439	0.1469		0.1469
F	7.3059	6.7918		6.9158	8.2355	7.5914		7.1434
Log-likelihood	-575.58	-575.28	-486.86	-574.14	-514.00	-512.94	-486.86	-512.94
Ramsey test	χ^2=1.24[0.433]				χ^2=0.47[0.701]			

Notes: Standard errors in parentheses; ***, ** and * denote significance at the 1%, 5% and 10% level respectively.
Figures in squared brackets are p-values.

TABLE A4.4

IV estimates of returns to education in Bulgaria, by gender, 1995

	Males				*Females*			
	IV_1	First stage	IV_2	First stage	IV_1	First stage	IV_2	First stage
	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
school	0.0850**		0.0923***		0.0838*		0.0971**	
	(0.0331)		(0.0332)		(0.0434)		(0.0433)	
exp	0.0098	0.0722**	0.0093	0.0721**	-0.0019	0.0642	-0.0026	0.0635
	(0.0086)	(0.0348)	(0.0087)	(0.0348)	(0.0109)	(0.0427)	(0.0109)	(0.0427)
expsq	-0.0002	-0.0036***	-0.0002	-0.0036***	0.0003	-0.0043***	0.0004	-0.0042***
	(0.0002)	(0.0007)	(0.0002)	(0.0007)	(0.0003)	(0.0009)	(0.0003)	(0.0009)
1-2 years	0.0720	0.1048	0.0702	0.1048	0.3498***	-0.0438	0.3455***	-0.0386
	(0.1036)	(0.4822)	(0.1042)	(0.4825)	(0.1250)	(0.4797)	(0.1255)	(0.4783)
3-5 years	0.1190	-0.2059	0.1216	-0.2061	0.1438	-0.1372	0.1407	-0.1606
	(0.0841)	(0.3097)	(0.0846)	(0.3106)	(0.0967)	(0.3433)	(0.0971)	(0.3454)
6-10 years	0.2532***	-0.0803	0.2544***	-0.0804	0.1944**	0.1256	0.1907**	0.1183
	(0.0834)	(0.3457)	(0.0839)	(0.3459)	(0.0938)	(0.3353)	(0.0942)	(0.3353)
>10 years	0.1819**	-0.2658	0.1853**	-0.2657	0.0970	1.1242***	0.0805	1.1118***
	(0.0786)	(0.3408)	(0.0790)	(0.3418)	(0.1042)	(0.3102)	(0.1044)	(0.3085)
married	0.0291	0.9028***	0.0246	0.9028***	0.1536**	-0.3999	0.1638**	-0.4030
	(0.0618)	(0.2214)	(0.0622)	(0.2217)	(0.0743)	(0.2823)	(0.0745)	(0.2819)
bulgarian	-0.1280	2.2660***	-0.1459	2.2659***	0.0529	2.5212***	0.0177	2.5363***
	(0.1138)	(0.3141)	(0.1144)	(0.3147)	(0.1416)	(0.3637)	(0.1416)	(0.3626)
public	-0.0853	0.4490*	-0.0893	0.4493*	0.0970	-0.3182	0.1014	-0.3000
	(0.0626)	(0.2365)	(0.0629)	(0.2376)	(0.0778)	(0.2926)	(0.0781)	(0.2941)
urban	0.0886	0.9170***	0.0815	0.9167***	0.0712	0.8752***	0.0572	0.8391***
	(0.0591)	(0.2003)	(0.0594)	(0.2041)	(0.0730)	(0.2181)	(0.0731)	(0.2198)
sofia_region	0.1523*	1.7025***	0.1398	1.6951**	0.1558*	0.7697**	0.1459*	0.0770
	(0.0887)	(0.3158)	(0.0892)	(0.8200)	(0.0807)	(0.3299)	(0.0810)	(0.7876)
moth_educ		3.8402***		3.8396***		2.5001***		2.4890***
		(0.4355)		(0.4390)		(0.4264)		(0.4319)
average_sch				0.0030				0.2786
				(0.2990)				(0.2788)
constant	3.7726***	7.9400***	3.7127***	7.9109***	3.0612***	9.4479***	2.9340***	6.7341**
	(0.2995)	(0.4870)	(0.3009)	(2.9086)	(0.4365)	(0.5480)	(0.4358)	(2.7447)
N	708	708	708	708	609	609	609	609
R^2	0.1074	0.3556	0.0974	0.3556	0.1414	0.4035	0.1342	0.4044
F	6.94	33.85	6.96	31.31	4.84	35.88	4.91	33.07
Log-likelihood	-585.59	-1626.72	-589.49	-1626.72	-514.88	-1456.32	-517.38	-1455.83
F (excluded)		54.13***		27.10***		36.03***		18.23***
p-value		0.0000		0.0000		0.0000		0.0000
Sargan test				19.055				7.380
p-value				0.0000				0.0066
Partial -R^2		0.0730		0.0732		0.0576		0.0584
DWH test		1.5998		2.2410		0.1076		0.4209
p-value		0.2059		0.1343		0.7428		0.5164

Notes: Standard errors in parentheses; ***, ** and * denote significance at the 1%, 5% and 10% level respectively.
IV_1 – specification where mother education is used as an instrument; IV_2 – over-identification specification - both mother education and average regional differences in schooling are used as instruments

TABLE A4.5

Returns to education in Bulgaria, by gender, 1997

	Males				*Females*			
	OLS	Heckman	Probit	Semi-parametric	OLS	Heckman	Probit	Semi-parametric
	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
school	0.0397***	0.0168**	0.0482	0.0363**	0.0454***	0.0515***	0.0329	0.0586***
	(0.0119)	(0.0084)	(0.0521)	(0.0182)	(0.0115)	(0.0149)	(0.0539)	(0.0141)
exp	0.0119	-0.0034	-0.0355	0.0138	0.0040	-0.0439***	-0.0196	0.0021
	(0.0113)	(0.0069)	(0.0337)	(0.0144)	(0.0121)	(0.0160)	(0.0631)	(0.0136)
expsq	-0.0004	0.0002	0.0003	-0.0004	-0.0001	0.0015***	-0.0005	-0.0002
	(0.0002)	(0.0001)	(0.0006)	(0.0002)	(0.0003)	(0.0004)	(0.0012)	(0.0003)
1-2 years	0.2277	0.1167	0.2701	0.2254	0.3047**	-0.1982	0.1592	0.3512***
	(0.1450)	(0.1201)	(0.3618)	(0.1555)	(0.1224)	(0.2302)	(0.3727)	(0.1258)
3-5 years	0.0769	-0.0713	0.0555	0.0797	0.1462	0.0527	0.3294	0.0010
	(0.0883)	(0.0652)	(0.3345)	(0.0888)	(0.1018)	(0.1066)	(0.2203)	(.0.0834)
6-10 years	0.1276	0.0416	0.6456	0.1004	0.1136	-0.1197	0.2074	0.1532*
	(0.1095)	(0.0713)	(0.4385)	(0.1788)	(0.0830)	(0.1149)	(0.3438)	(0.0875)
>10 years	0.2662***	-0.0174	0.9946***	0.2049	0.2452***	-0.0327	0.2928	0.3056***
	(0.0740)	(0.0671)	(0.3699)	(0.2477)	(0.0745)	(0.0951)	(0.2688)	(0.0869)
married	0.1287	-0.0520	0.3060	0.1107	0.0152	-0.1195	-0.1225	-0.0326
	(0.0857)	(0.0520)	(0.3452)	(0.1073)	(0.0804)	(0.1070)	(0.3732)	(0.0922)
bulgarian	0.0621	-0.1199	0.9171***	0.0719	0.0562	-0.0796	-0.1680	-0.0210
	(0.1338)	(0.0954)	(0.2889)	(0.2744)	(0.1070)	(0.1781)	(0.4322)	(0.1208)
urban	0.3151***	0.0153	-0.3451	0.3343***	0.2378***	-0.0679	-0.1966	0.1782**
	(0.0771)	(0.0574)	(0.2418)	(0.1189)	(0.0715)	(0.1066)	(0.2952)	(0.0837)
sofia region	-0.0532	0.0306	-0.8239**	-0.0173	0.0674	0.1237	0.2996	0.1510
	(0.1189)	(0.1034)	(0.3333)	(0.2406)	(0.1003)	(0.1455)	(0.4739)	(0.1144)
public	0.2445***	0.0247	0.1808	0.2368**	0.1414*	0.2090	0.8261***	0.3238**
	(0.0757)	(0.0552)	(0.2314)	(0.0931)	(0.0787)	(0.1616)	(0.2472)	(0.1413)
mills		0.1886				-0.3081		
		(0.6090)				(0.6842)		
children			-0.3445				-0.5453*	
			(0.3748)				(0.3177)	
$(z'\gamma)$				-0.2837				-0.2400
				(0.4032)				(0.3036)
$(z'\gamma)^2$				0.0721				0.0067
				(0.0706)				(0.0591)
constant	4.6832***	1.5499***	1.1602	4.9508***	4.5390***	0.9434***	2.1390**	4.9588***
	(0.1888)	(0.1738)	(0.7252)	(0.4359)	(0.1900)	(0.2511)	(1.0627)	(0.4215)
N	727	727	895	727	804	804	895	804
R^2	0.1204	0.0206		0.1219	0.0893	0.0513		0.0944
F	9.1314	1.0771		8.0272	6.7998	3.3229		6.1457
Log-likelihood	-870.18	-576.77	-60.45	-869.54	-897.57	-1055.77	-62.45	-807.75
Ramsey test	χ^2=1.57 [0.195]				χ^2=0.69 [0.557]			

Notes: Standard errors in parentheses; ***, ** and * denote significance at the 1%, 5% and 10% level respectively.
Figures in squared brackets are the p-values.

TABLE A4.6

IV estimates of returns to education in Bulgaria, by gender, 1997

	Males				Females			
	IV_1	First stage	IV_2	First stage	IV_1	First stage	IV_2	First stage
	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
school	0.0745***		0.0774***		0.0595***		0.0635***	
	(0.0198)		(0.0197)		(0.0179)		(0.0174)	
exp	0.0110	0.0450	0.0109	0.0426	0.0030	0.0873**	0.0027	0.0922**
	(0.0119)	(0.0302)	(0.0119)	(0.0301)	(0.0125)	(0.0383)	(0.0125)	(0.0376)
expsq	-0.0003	-0.0025***	-0.0003	-0.0024***	-0.0000	-0.0038***	0.0000	-0.0038***
	(0.0003)	(0.0006)	(0.0003)	(0.0006)	(0.0003)	(0.0009)	(0.0003)	(0.0009)
1-2 years	0.2255*	0.2745	0.2253*	0.2473	0.3118**	-0.1651	0.3138***	-0.1352
	(0.1240)	(0.2692)	(0.1241)	(0.2664)	(0.1216)	(0.2494)	(0.1217)	(0.2507)
3-5 years	0.0902	0.1514	0.0913	0.1262	0.1483	-0.0208	0.1488	-0.0224
	(0.0984)	(0.2626)	(0.0985)	(0.2639)	(0.0994)	(0.2910)	(0.0995)	(0.2842)
6-10 years	0.1185	0.2629	0.1177	0.2568	0.1127	-0.1543	0.1124	-0.1281
	(0.1012)	(0.2334)	(0.1013)	(0.2322)	(0.0791)	(0.1984)	(0.0792)	(0.1924)
>10 years	0.2626***	0.1910	0.2623***	0.1768	0.2366***	0.3741*	0.2341***	0.3725**
	(0.0798)	(0.2020)	(0.0799)	(0.2016)	(0.0721)	(0.1907)	(0.0721)	(0.1891)
married	0.1041	0.4198**	0.1020	0.4165**	0.0224	-0.1673	0.0244	-0.1678
	(0.0861)	(0.1964)	(0.0862)	(0.1975)	(0.0722)	(0.1860)	(0.0723)	(0.1862)
bulgarian	-0.0419	2.4116***	-0.0506	2.3313***	0.0135	2.3512***	0.0013	2.3735***
	(0.1230)	(0.2816)	(0.1229)	(0.2885)	(0.1102)	(0.3103)	(0.1096)	(0.3122)
public	0.2377***	0.2552	0.2372***	0.2799	0.1364*	0.0347	0.1350*	-0.0044
	(0.0728)	(0.1825)	(0.0729)	(0.1818)	(0.0739)	(0.1707)	(0.0739)	(0.1688)
urban	0.2783***	0.6083***	0.2753***	0.5556***	0.2257***	0.6303***	0.2222***	0.4306**
	(0.0804)	(0.1995)	(0.0804)	(0.2002)	(0.0727)	(0.1972)	(0.0727)	(0.1963)
sofia_region	-0.0297	-0.3973*	-0.0277	-0.0863	0.0837	-0.7157***	0.0883	-0.3243
	(0.1094)	(0.2308)	(0.1095)	(0.2481)	(0.0951)	(0.2123)	(0.0950)	(0.2071)
moth_educ		4.6623***		4.6029***		4.1143***		3.9815***
		(0.1458)		(0.1502)		(0.1385)		(0.1441)
average_sch				0.3136***				0.5104***
				(0.1184)				(0.1083)
constant	4.3903***	8.3218***	4.3659***	4.6994***	4.4049***	8.9740***	4.3668***	2.9585**
	(0.2296)	(0.3594)	(0.2288)	(1.3905)	(0.2302)	(0.4220)	(0.2271)	(1.2615)
N	727	727	727	727	804	804	804	804
R^2	0.1104	0.5382	0.1087	0.5429	0.0874	0.5629	0.0862	0.5769
F	8.34	147.01	8.43	136.63	5.94	148.60	6.11	152.05
Log-likelihood	-874.27	-1508.56	-874.98	-1504.82	-898.41	-1653.11	-898.95	-1639.97
F (excluded)	435.81***			223.55***	471.88***			256.59***
p-value	0.0000			0.0000	0.0000			0.0000
Sargan test				1.303				0.981
p-value				0.2537				0.3220
Partial-R^2	0.3790			0.3854	0.3737			0.3938
DWH test	5.0225			6.0588	0.9977			1.7906
p-value	0.0250			0.0138	0.3178			0.1808

Notes: Standard errors in parentheses; ***, ** and * denote significance at the 1%, 5% and 10% level respectively.
IV_1 – specification where mother education is used as an instrument; IV_2 – over-identification specification - both mother education and average regional differences in schooling are used as instruments

TABLE A4.7

Returns to education in Bulgaria, by gender, 2001

	Males				Females			
	OLS	Heckman	Probit	Semi-parametric	OLS	Heckman	Probit	Semi-parametric
	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
school	0.0431***	0.0820**	-0.0241*	0.0944***	0.0664***	0.0699***	0.0231*	0.0703***
	(0.0117)	(0.0330)	(0.0128)	(0.0345)	(0.0094)	(0.0088)	(0.0128)	(0.0092)
exp	-0.0002	0.0157	-0.0108	0.0279	0.0040	-0.0076	0.0083	-0.0078
	(0.0107)	(0.0169)	(0.0116)	(0.0196)	(0.0089)	(0.0149)	(0.0116)	(0.0149)
expsq	-0.0001	-0.0006	0.0004*	-0.0010*	-0.0001	0.0004	-0.0004	0.0004
	(0.0002)	(0.0005)	(0.0002)	(0.0006)	(0.0002)	(0.0005)	(0.0002)	(0.0005)
1-2 years	0.1417	0.1272	0.0115	0.1097	0.0099	-0.0254	0.0239	-0.0298
	(0.1055)	(0.1089)	(0.1384)	(0.1073)	(0.0770)	(0.0834)	(0.1385)	(0.0859)
3-5 years	0.2092**	0.0270	0.1188	-0.0358	0.1045	0.2111	-0.0897	0.2200
	(0.1058)	(0.1954)	(0.1121)	(0.2001)	(0.0764)	(0.1324)	(0.1121)	(0.1352)
6-10 years	0.4157***	0.2686	0.0971	0.2070	0.1205	0.1839	-0.0539	0.1881*
	(0.1068)	(0.1680)	(0.1066)	(0.1728)	(0.0868)	(0.1122)	(0.1065)	(0.1120)
>10 years	0.4514***	0.3941***	0.0391	0.3673***	0.1333	0.1298	0.0007	0.1291
	(0.1066)	(0.1238)	(0.1096)	(0.1228)	(0.0812)	(0.0804)	(0.1095)	(0.0806)
married	0.1254*	-0.0391	0.0887	-0.0923	-0.0128	0.1060	-0.0927	0.1182
	(0.0719)	(0.1517)	(0.0786)	(0.1567)	(0.0469)	(0.1243)	(0.0787)	(0.1308)
bulgarian	0.2096**	0.3428***	-0.0878	0.4196***	0.0171	-0.1632	0.1360	-0.1702
	(0.0971)	(0.1139)	(0.1131)	(0.1205)	(0.0813)	(0.1755)	(0.1135)	(0.1786)
public	0.0107	0.0175	-0.0040	0.0209	0.0166	0.0003	0.0168	-0.0023
	(0.0594)	(0.0588)	(0.0664)	(0.0581)	(0.0489)	(0.0497)	(0.0664)	(0.0508)
urban	0.1116*	0.2219**	-0.0730	0.2689**	-0.1083*	-0.1706*	0.0502	-0.1747*
	(0.0591)	(0.1108)	(0.0819)	(0.1181)	(0.0599)	(0.0907)	(0.0819)	(0.0924)
sofia_region	0.0892	0.0249	0.0402	0.0025	0.0727	0.1077	-0.0280	0.1108
	(0.0844)	(0.0991)	(0.0794)	(0.1019)	(0.0670)	(0.0742)	(0.0795)	(0.0742)
children			0.0472				-0.0438	
			(0.0732)				(0.0732)	
mills		-2.4700				-1.9253		
		(0.9544)				(0.8612)		
$(z'\gamma)$				-7.0139*				-3.0135
				(0.8957)				(0.6147)
$(z'\gamma)^2$				2.4249				0.5467
				(0.8648)				(0.4233)
constant	0.3314**	1.7841	0.3242	3.6064**	0.4621***	2.4814	-0.3566*	3.0297
	(0.1542)	(1.2049)	(0.1981)	(1.7769)	(0.1467)	(1.9714)	(0.1981)	(2.5849)
N	729	729	1712	729	709	709	1712	709
R^2	0.1510	0.1531		0.1550	0.1350	0.1365		0.1367
F	8.85	9.83		9.25	8.23	7.83		7.34
Log-likelihood	-664.39	-663.55	-1172.87	-662.80	-468.02	-467.45	-1171.84	-467.40
Ramsey test	χ^2=0.74 [0.526]				χ^2=0.45[0.789]			

Notes: Standard errors in parentheses; ***, ** and * denote significance at the 1%, 5% and 10% level respectively. Figures in squared brackets are p-values.

TABLE A4.8

IV estimates of returns to education in Bulgaria, by gender, 2001

	Males				*Females*			
	IV_1	First stage	IV_2	First stage	IV_1	First stage	IV_2	First stage
	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
school	0.1281**		0.1036**		0.1005***		0.0983***	
	(0.0614)		(0.0474)		(0.0212)		(0.0184)	
exp	-0.0302*	0.2718***	-0.0242*	0.2669***	-0.0174*	0.2600***	-0.0169*	0.2690***
	(0.0174)	(0.0672)	(0.0144)	(0.0670)	(0.0091)	(0.0680)	(0.0087)	(0.0664)
expsq	0.0005	-0.0065***	0.0004	-0.0063***	0.0004*	-0.0074***	0.0004*	-0.0076***
	(0.0004)	(0.0013)	(0.0003)	(0.0013)	(0.0002)	(0.0014)	(0.0002)	(0.0014)
1-2 years	0.2208*	-0.5896	0.2038*	-0.5427	0.0280	0.6271	0.0289	0.6121
	(0.1242)	(0.4786)	(0.1175)	(0.4848)	(0.0894)	(0.5012)	(0.0890)	(0.4870)
3-5 years	0.2964***	-0.4263	0.2884***	-0.3964	0.1492**	-0.3435	0.1484**	-0.3479
	(0.0979)	(0.4504)	(0.0940)	(0.4467)	(0.0759)	(0.5040)	(0.0756)	(0.4982)
6-10 years	0.4540***	-0.5377	0.4423***	-0.5156	0.1816**	-0.3084	0.1813***	-0.3204
	(0.0935)	(0.3854)	(0.0888)	(0.3827)	(0.0705)	(0.4120)	(0.0703)	(0.3936)
>10 years	0.5153***	-0.3214	0.5115***	-0.3166	0.2267***	0.1340	0.2275***	0.1010
	(0.0893)	(0.4097)	(0.0862)	(0.4080)	(0.0717)	(0.3836)	(0.0714)	(0.3706)
married	0.1557**	0.0312	0.1581**	0.0071	0.0177	0.0024	0.0176	-0.0746
	(0.0689)	(0.3243)	(0.0665)	(0.3259)	(0.0482)	(0.2615)	(0.0481)	(0.2612)
bulgarian	-0.0532	3.1841***	0.0161	2.9861***	0.0289	2.3997***	0.0338	2.2425***
	(0.1923)	(0.3670)	(0.1560)	(0.3749)	(0.0890)	(0.4164)	(0.0858)	(0.4146)
public	0.0476	0.7045***	0.0663	0.6937***	0.0285	0.5222**	0.0306	0.5202**
	(0.0710)	(0.2152)	(0.0629)	(0.2139)	(0.0475)	(0.2274)	(0.0463)	(0.2228)
urban	0.0446	1.1395***	0.0715	1.1133***	-0.0478	1.1097***	-0.0453	1.1111***
	(0.0921)	(0.2286)	(0.0799)	(0.2299)	(0.0586)	(0.2791)	(0.0572)	(0.2730)
sofia_region	0.1418*	0.5289	0.1583**	0.0573	0.0838	0.6409*	0.0859	-0.2490
	(0.0846)	(0.3327)	(0.0782)	(0.3546)	(0.0611)	(0.3298)	(0.0600)	(0.3620)
moth_educ		-0.9050***		-0.9313***		-2.0092***		-1.9276***
		(0.2722)		(0.2696)		(0.2379)		(0.2329)
average_sch				0.4513***				0.8518***
				(0.1565)				(0.1699)
constant	-1.3356***	6.3307***	-1.1843***	1.3428	-1.1075***	8.4099***	-1.0910***	-1.3727
	(0.3978)	(0.6602)	(0.3149)	(1.8368)	(0.1900)	(0.8149)	(0.1728)	(2.2226)
N	729	729	729	729	709	709	709	709
R^2	0.0511	0.2903	0.1130	0.2987	0.1159	0.2712	0.1213	0.2977
F	11.10	18.46	11.89	17.79	7.68	22.53	8.20	23.05
Log-likelihood	-747.11	-1757.32	-722.67	-1752.98	-551.14	-1740.64	-548.99	-1727.52
F (excluded)		16.55***		13.13***		77.53***		52.90***
p-value		0.00001		0.0000		0.0000		0.000
Sargan test				0.469				0.045
p-value				0.4933				0.8321
Partial R^2		0.0227		0.0356		0.1014		0.1336
DWH test		2.8615		2.5043		5.4962		6.8024
p-value		0.0907		0.1135		0.0190		0.0091

Notes: Standard errors in parentheses; ***, ** and * denote significance at the 1%, 5% and 10% level respectively. IV_1 – specification where mother education is used as an instrument; IV_2 – over-identification specification - both mother education and average regional differences in schooling are used as instruments

TABLE A4.9

Returns to education in Bulgaria, by gender, 2003

	Males				Females			
	OLS	Heckman	Probit	Semi-parametric	OLS	Heckman	Probit	Semi-parametric
	(1)	(2)	(3)	(4)	(1)	(2)	(3)	(4)
school	0.0512***	0.0484***	0.0199	0.0475***	0.0598***	0.0580***	0.0288	0.0555***
	(0.0048)	(0.0052)	(0.0277)	(0.0060)	(0.0046)	(0.0062)	(0.0285)	(0.0086)
exp	0.0069*	0.0068*	0.0030	0.0067*	0.0179***	0.0192***	-0.0272	0.0217**
	(0.0040)	(0.0040)	(0.0195)	(0.0040)	(0.0050)	(0.0067)	(0.0286)	(0.0094)
expsq	-0.0002**	-0.0002**	0.0001	-0.0002**	-0.0004***	-0.0004**	0.0008	-0.0005*
	(0.0001)	(0.0001)	(0.0004)	(0.0001)	(0.0001)	(0.0002)	(0.0007)	(0.0003)
1-2 years	0.0739*	0.0890**	-0.0985	0.0941**	0.0476	0.0732	-0.5179*	0.1207
	(0.0389)	(0.0402)	(0.2277)	(0.0436)	(0.0422)	(0.0744)	(0.2784)	(0.1304)
3-5 years	0.1251***	0.0995**	0.2206	0.0912*	0.1673***	0.1910***	-0.4878*	0.2370*
	(0.0419)	(0.0447)	(0.2720)	(0.0519)	(0.0414)	(0.0692)	(0.2640)	(0.1226)
6-10 years	0.2546***	0.2609***	-0.0512	0.2632***	0.1697***	0.2049**	-0.6366**	0.2669
	(0.0495)	(0.0497)	(0.2906)	(0.0505)	(0.0500)	(0.0940)	(0.2854)	(0.1662)
>10 years	0.2799***	0.3759***	-0.5279**	0.4055***	0.2485***	0.2829***	-0.6251**	0.3427**
	(0.0435)	(0.0805)	(0.2348)	(0.1234)	(0.0420)	(0.0869)	(0.2787)	(0.1584)
married	0.0216	-0.0133	0.1862	-0.0250	-0.0379	-0.0301	-0.1251	-0.0201
	(0.0308)	(0.0389)	(0.1768)	(0.0540)	(0.0292)	(0.0360)	(0.1805)	(0.0440)
bulgarian	0.1186**	0.1377***	-0.0960	0.1433***	0.0582	0.0527	0.0937	0.0437
	(0.0462)	(0.0478)	(0.2444)	(0.0519)	(0.0447)	(0.0470)	(0.2356)	(0.0518)
public	0.1973***	0.1951***	0.0092	0.1944***	0.0593**	0.0501	0.1175	0.0365
	(0.0316)	(0.0316)	(0.1802)	(0.0317)	(0.0294)	(0.0336)	(0.1716)	(0.0445)
urban	0.1180***	0.0824**	0.2096	0.0712	0.0563*	0.0466	0.1317	0.0334
	(0.0322)	(0.0403)	(0.1681)	(0.0517)	(0.0328)	(0.0398)	(0.1762)	(0.0510)
sofia region	0.2358***	0.4834**	-0.9160***	0.5346**	0.1161**	0.2060	-0.7577***	0.2626
	(0.0711)	(0.1942)	(0.2431)	(0.2425)	(0.0506)	(0.1999)	(0.2350)	(0.2579)
mills		-1.3150				-0.5018		
		(0.9273)				(0.0560)		
children			0.1024				0.0814	
			(0.1989)				(0.1676)	
$(z'\gamma)$				-1.9023				-1.7776
				(2.1079)				(2.7801)
$(z'\gamma)^2$				0.9164				2.6860
				(3.3492)				(4.2924)
constant	0.0952	0.6346***	1.5749***	0.6839***	-0.1615**	0.2828***	1.9665***	0.3514**
	(0.0730)	(0.1179)	(0.4084)	(0.1936)	(0.0756)	(0.1074)	(0.4990)	(0.1785)
N	1296	1296	1418	1296	1186	1186	1609	1186
R^2	0.3209	0.3223		0.3224	0.2970	0.2971		0.2975
F	52.66	48.81		45.55	39.76	36.82		34.32
Log-likelihood	-831.75	-830.37	-158.53	-830.32	-694.87	-694.72	-160.34	-694.45
Ramsey test	χ^2=3.19 [0.023]				χ^2=1.13 [0.336]			

Notes: Standard errors in parentheses; ***, ** and * denote significance at the 1%, 5% and 10% level respectively. Figures in squared brackets are p-values.

TABLE A4.10

IV estimates of returns to education in Bulgaria, by gender, 2003

	Males				Females			
	IV (1)	First stage (2)	IV (3)	First stage (4)	IV (1)	First stage (2)	IV (3)	First stage (4)
school	0.0996**		0.1144***		0.1051***		0.1056***	
	(0.0496)		(0.0310)		(0.0143)		(0.0132)	
exp	0.0116**	-0.0552**	0.0124**	-0.0566**	0.0232***	-0.0970***	0.0232***	-0.0843***
	(0.0052)	(0.0233)	(0.0048)	(0.0228)	(0.0046)	(0.0236)	(0.0045)	(0.0234)
expsq	-0.0002*	-0.0007	-0.0002	-0.0007	-0.0004***	-0.0004	-0.0004***	-0.0007
	(0.0001)	(0.0005)	(0.0001)	(0.0005)	(0.0001)	(0.0006)	(0.0001)	(0.0006)
1-2 years	0.0596	0.4767**	0.0526	0.3872	0.0247	0.4022	0.0245	0.3237
	(0.0492)	(0.2378)	(0.0468)	(0.2369)	(0.0437)	(0.2548)	(0.0437)	(0.2529)
3-5 years	0.0927*	0.6313***	0.0834*	0.5363**	0.1147**	0.9771***	0.1141***	0.9094***
	(0.0544)	(0.2325)	(0.0498)	(0.2305)	(0.0446)	(0.2520)	(0.0442)	(0.2482)
6-10 years	0.2400***	0.8209***	0.2282***	0.6811***	0.1121**	1.1047***	0.1115**	0.9936***
	(0.0657)	(0.2627)	(0.0593)	(0.2595)	(0.0508)	(0.2807)	(0.0503)	(0.2782)
>10 years	0.2306***	1.2590***	0.2118***	1.1611***	0.1712***	1.4816***	0.1703***	1.4128***
	(0.0780)	(0.2325)	(0.0614)	(0.2302)	(0.0517)	(0.2933)	(0.0508)	(0.2877)
married	-0.0485	0.5690***	-0.0572	0.6317***	-0.0554*	1.1598***	-0.0556*	1.1614***
	(0.0454)	(0.1779)	(0.0402)	(0.1782)	(0.0317)	(0.1790)	(0.0317)	(0.1782)
bulgarian	-0.0043	2.2015***	-0.0369	2.1068***	-0.0521	2.3199***	-0.0533	2.2358***
	(0.1194)	(0.3063)	(0.0837)	(0.3068)	(0.0570)	(0.3337)	(0.0555)	(0.3350)
public	0.1437***	0.4538**	0.1369***	0.4260**	-0.0029	1.1625***	-0.0036	1.0894***
	(0.0424)	(0.1874)	(0.0393)	(0.1861)	(0.0370)	(0.1823)	(0.0362)	(0.1816)
urban	0.0362	1.2031***	0.0182	1.0694***	-0.0138	1.3770***	-0.0145	1.2502***
	(0.0697)	(0.1920)	(0.0516)	(0.1924)	(0.0412)	(0.2084)	(0.0404)	(0.2086)
sofia region	0.2464***	-0.0435	0.2473***	-0.0349	0.1436**	-0.5439	0.1439**	-0.5034
	(0.0744)	(0.2928)	(0.0765)	(0.2927)	(0.0666)	(0.3851)	(0.0665)	(0.3850)
moth_educ		4.2242***		3.8407***		3.3675***		3.2341***
		(0.6775)		(0.5797)		(0.2027)		(0.2035)
average_sch				0.4508***				0.4669***
				(0.0911)				(0.0926)
constant	-0.3743	9.7416***	-0.5177*	3.9589***	-0.6600***	10.2563***	-0.6653***	4.1379***
	(0.4872)	(0.3672)	(0.3083)	(1.2103)	(0.1684)	(0.3848)	(0.1572)	(1.2641)
N	1296	1296	1296	1296	1186	1186	1186	1186
R^2	0.1554	0.2969	0.1053	0.3141	0.2390	0.4234	0.2377	0.4359
F	24.28	34.59	23.73	35.93	29.60	101.80	30.37	98.42
ll	-695.40	-2441.21	-725.97	-2428.11	-741.86	-2815.43	-742.83	-2802.45
F (excluded)		14.85***		22.18***		163.96***		102.01***
p-value		0.0000		0.0000		0.0000		0.0000
Sargan test				0.363				0.025
p-value				0.5467				0.8755
Partial R^2		0.0140		0.0406				0.1480
DWH test		1.6596		10.039		12.1658		15.783
p-value		0.1976		0.0015		0.00049		0.00007

Notes: Standard errors in parentheses; ***, ** and * denote significance at the 1%, 5% and 10% level respectively.

TABLE A4.11

QR estimates of returns to education in Bulgaria, by gender, 1986

	Males					Females				
	10th	25th	50th	75th	90th	10th	25th	50th	75th	90th
school	0.0281***	0.0205***	0.0098***	-0.0005	0.0000	0.0524***	0.0305***	0.0171***	0.0082**	-0.0032
	(0.0043)	(0.0034)	(0.0033)	(0.0035)	(0.0001)	(0.0050)	(0.0027)	(0.0029)	(0.0035)	(0.0021)
exp	0.0037***	0.0026***	-0.0015**	-0.0033***	-0.0003***	0.0005	-0.0010*	-0.0004	-0.0020***	-0.0009**
	(0.0010)	(0.0008)	(0.0007)	(0.0007)	(0.0000)	(0.0010)	(0.0006)	(0.0006)	(0.0007)	(0.0004)
expsq	0.0000*	-0.0000	-0.0000	-0.0000	0.0000***	0.0001***	0.0000	-0.0000	-0.0001***	-0.0000***
	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
1-2 years	0.2163**	0.1785**	0.1244*	0.0449	0.0132***	0.1367	-0.0471	-0.0601	0.0292	0.0323
	(0.0958)	(0.0747)	(0.0708)	(0.0753)	(0.0022)	(0.1078)	(0.0588)	(0.0626)	(0.0747)	(0.0425)
3-5 years	0.2393***	0.1488**	-0.0236	0.0998	0.0487***	0.0226	-0.1617***	-0.3656***	-0.1752***	-0.0647*
	(0.0776)	(0.0602)	(0.0570)	(0.0609)	(0.0017)	(0.0866)	(0.0466)	(0.0490)	(0.0587)	(0.0362)
6-10 years	0.2456***	0.0445	-0.1881***	-0.2025***	-0.0664***	0.0603	-0.1753***	-0.4411***	-0.3578***	-0.1486***
	(0.0799)	(0.0610)	(0.0563)	(0.0602)	(0.0017)	(0.0871)	(0.0464)	(0.0479)	(0.0571)	(0.0351)
>10 years	0.4010***	0.1663**	-0.1631***	-0.2564***	-0.0667***	0.2251**	-0.0999**	-0.3758***	-0.3667***	-0.1921***
	(0.0888)	(0.0664)	(0.0600)	(0.0634)	(0.0018)	(0.0965)	(0.0509)	(0.0521)	(0.0614)	(0.0363)
married	0.3134***	0.3673***	0.2583***	0.1692***	0.0469***	0.6130***	0.6105***	0.4394***	0.2239***	0.0728***
	(0.0381)	(0.0281)	(0.0252)	(0.0264)	(0.0008)	(0.0428)	(0.0233)	(0.0239)	(0.0281)	(0.0173)
bulgarian	0.2013***	0.1723***	0.1694***	0.1357***	0.0002	0.1225***	0.1022***	0.0990***	-0.0096	-0.0137
	(0.0355)	(0.0279)	(0.0263)	(0.0277)	(0.0008)	(0.0453)	(0.0243)	(0.0255)	(0.0296)	(0.0172)
public	0.3101***	0.1380***	0.0392	-0.2402***	-0.1743***	0.1171***	0.0716***	-0.0125	-0.1530***	-0.1569***
	(0.0327)	(0.0257)	(0.0244)	(0.0250)	(0.0007)	(0.0392)	(0.0214)	(0.0225)	(0.0255)	(0.0152)
urban	0.2013***	0.1014***	0.0348*	-0.0089	-0.0003	0.1736***	0.0844***	0.0473**	-0.0064	-0.0102
	(0.0285)	(0.0221)	(0.0210)	(0.0215)	(0.0006)	(0.0328)	(0.0182)	(0.0196)	(0.0230)	(0.0139)
sofia_region	0.1914***	0.0765**	0.0501	0.0019	0.0109***	0.1575***	0.0945***	0.0947***	0.0297	0.0047
	(0.0439)	(0.0349)	(0.0334)	(0.0348)	(0.0010)	(0.0487)	(0.0265)	(0.0284)	(0.0335)	(0.0204)
constant	-1.1783***	-0.3555***	0.6351***	1.5524***	1.6740***	-1.2709***	-0.2307***	0.6794***	1.5409***	1.8491***
	(0.1257)	(0.0958)	(0.0874)	(0.0931)	(0.0026)	(0.1423)	(0.0758)	(0.0798)	(0.0955)	(0.0528)
N	3213	3213	3213	3213	3213	3238	3238	3238	3238	3238
Pseudo R^2	0.1629	0.1082	0.0394	0.0398	0.0535	0.1797	0.1404	0.0647	0.0305	0.0426

Notes: Standard errors in parentheses; ***, ** and * denote significance at the 1%, 5% and 10% level respectively.

TABLE A4.12

QR estimates of returns to education in Bulgaria, by gender, 1995

	Males					Females				
	10th	25th	50th	75th	90th	10th	25th	50th	75th	90th
school	0.0451***	0.0292***	0.0236**	0.0308**	0.0395*	0.0557***	0.0526***	0.0644***	0.0448***	0.0458**
	(0.0161)	(0.0086)	(0.0094)	(0.0122)	(0.0218)	(0.0093)	(0.0058)	(0.0070)	(0.0120)	(0.0192)
exp	0.0118	0.0066	0.0047	-0.0069	-0.0166	-0.0159	-0.0049	-0.0101	0.0059	-0.0096
	(0.0162)	(0.0090)	(0.0096)	(0.0114)	(0.0209)	(0.0134)	(0.0070)	(0.0077)	(0.0117)	(0.0170)
expsq	-0.0004	-0.0002	-0.0002	0.0001	0.0002	0.0004	0.0001	0.0003*	-0.0000	0.0005
	(0.0004)	(0.0002)	(0.0002)	(0.0002)	(0.0004)	(0.0003)	(0.0002)	(0.0002)	(0.0003)	(0.0004)
1-2 years	0.0038	0.0447	0.0364	0.1194	0.1074	0.1859	0.3358***	0.2995***	0.3641***	0.2607
	(0.2061)	(0.1179)	(0.1187)	(0.1350)	(0.2246)	(0.1511)	(0.0854)	(0.0933)	(0.1390)	(0.2011)
3-5 years	0.0440	0.0860	0.0467	0.1004	0.3185	0.0289	0.2644***	0.1523**	0.0876	0.0684
	(0.1628)	(0.0965)	(0.0984)	(0.1136)	(0.2065)	(0.1342)	(0.0684)	(0.0730)	(0.1080)	(0.1538)
6-10 years	0.0973	0.1860*	0.2398**	0.3283***	0.3812*	0.0651	0.2449***	0.1817**	0.1071	0.4741***
	(0.1520)	(0.0962)	(0.0989)	(0.1152)	(0.2093)	(0.1234)	(0.0634)	(0.0706)	(0.1075)	(0.1583)
>10 years	0.1221	0.2208**	0.1629*	0.2211**	0.2619	0.1172	0.2695***	0.2013***	0.0276	0.0628
	(0.1387)	(0.0892)	(0.0911)	(0.1081)	(0.2229)	(0.1175)	(0.0602)	(0.0684)	(0.1063)	(0.1605)
married	0.0636	-0.0047	0.0092	0.0622	0.2089	0.1325	0.0699	0.1005**	0.1164	0.1247
	(0.1177)	(0.0673)	(0.0688)	(0.0786)	(0.1429)	(0.0902)	(0.0491)	(0.0502)	(0.0725)	(0.1061)
bulgarian	-0.1565	-0.0076	0.0656	0.0632	-0.1970	0.1179	0.1147*	0.0553	0.1525	0.3127**
	(0.1499)	(0.0920)	(0.0974)	(0.1073)	(0.1973)	(0.1008)	(0.0585)	(0.0653)	(0.0993)	(0.1367)
urban	0.1011	0.0864	0.1177**	0.1344**	0.1431	-0.0039	0.0727*	0.0784*	0.1173*	0.0451
	(0.1015)	(0.0584)	(0.0594)	(0.0678)	(0.1235)	(0.0771)	(0.0411)	(0.0443)	(0.0660)	(0.0923)
sofia_region	0.2762*	0.3448***	0.2284***	0.1004	0.1961	0.2075**	0.1921***	0.1424**	0.1989**	-0.0672
	(0.1470)	(0.0783)	(0.0810)	(0.0922)	(0.1550)	(0.0919)	(0.0535)	(0.0559)	(0.0829)	(0.1184)
public	0.0753	0.0027	-0.0053	-0.0459	-0.2556*	0.2278**	0.0629	0.0352	0.0005	-0.0273
	(0.1168)	(0.0695)	(0.0709)	(0.0811)	(0.1408)	(0.0906)	(0.0509)	(0.0581)	(0.0872)	(0.1248)
constant	3.6313***	4.0199***	4.3159***	4.5669***	5.2127***	3.0536***	3.2455***	3.4560***	3.8488***	4.1302***
	(0.2733)	(0.1466)	(0.1555)	(0.1835)	(0.3220)	(0.1757)	(0.1020)	(0.1138)	(0.1734)	(0.2677)
N	708	708	708	708	708	609	609	609	609	609
Pseudo R^2	0.0723	0.0740	0.0655	0.0632	0.0646	0.1008	0.1218	0.1100	0.0793	0.0844

Notes: Standard errors in parentheses; ***, ** and * denote significance at the 1%, 5% and 10% level respectively.

TABLE A4.13

QR estimates of returns to education in Bulgaria, by gender, 1997

	Males					Females				
	10th	25th	50th	75th	90th	10th	25th	50th	75th	90th
school	0.0172	0.0270	0.0352***	0.0450***	0.0678***	0.0571***	0.0383***	0.0492***	0.0604***	0.0436**
	(0.0205)	(0.0167)	(0.0107)	(0.0144)	(0.0161)	(0.0144)	(0.0116)	(0.0153)	(0.0175)	(0.0213)
exp	-0.0135	0.0017	0.0071	0.0229	0.0166	-0.0067	-0.0063	0.0170	0.0206	-0.0065
	(0.0206)	(0.0157)	(0.0104)	(0.0153)	(0.0192)	(0.0145)	(0.0145)	(0.0175)	(0.0180)	(0.0193)
expsq	0.0000	-0.0001	-0.0002	-0.0006*	-0.0005	0.0001	0.0001	-0.0003	-0.0003	0.0001
	(0.0004)	(0.0003)	(0.0002)	(0.0003)	(0.0004)	(0.0003)	(0.0003)	(0.0004)	(0.0004)	(0.0005)
1-2 years	-0.2178	0.0990	0.2579**	0.4495***	0.5207***	0.3266**	0.2768**	0.2508	0.2627	0.5613***
	(0.2186)	(0.1668)	(0.1082)	(0.1518)	(0.1609)	(0.1545)	(0.1295)	(0.1714)	(0.1824)	(0.1956)
3-5 years	0.1727	0.0962	0.0736	0.0481	0.1689	0.1228	0.1972*	0.2141	0.1500	0.1828
	(0.1616)	(0.1322)	(0.0866)	(0.1214)	(0.1395)	(0.1199)	(0.1048)	(0.1392)	(0.1513)	(0.1739)
6-10 years	-0.0079	-0.1135	0.0033	0.2863**	0.5304***	0.0574	0.1346	0.2375**	0.0468	0.3175**
	(0.1682)	(0.1311)	(0.0891)	(0.1245)	(0.1508)	(0.1049)	(0.0855)	(0.1115)	(0.1227)	(0.1427)
>10 years	0.3336***	0.1787*	0.2379***	0.2707***	0.3417***	0.2458***	0.2805***	0.2743***	0.1784	0.2889**
	(0.1264)	(0.1043)	(0.0700)	(0.0993)	(0.1121)	(0.0930)	(0.0758)	(0.1009)	(0.1122)	(0.1305)
married	0.2733**	0.1315	0.1581**	-0.0287	0.0624	0.2020**	0.0364	-0.0682	0.0270	-0.0703
	(0.1299)	(0.1130)	(0.0754)	(0.1041)	(0.1242)	(0.0890)	(0.0772)	(0.1007)	(0.1105)	(0.1268)
bulgarian	0.5608***	0.1877	0.0140	-0.1910	-0.1943	-0.1204	0.0845	0.0405	-0.2405	0.0869
	(0.1958)	(0.1571)	(0.1001)	(0.1431)	(0.1741)	(0.1417)	(0.1063)	(0.1422)	(0.1567)	(0.1516)
public	0.2385**	0.2869***	0.2184***	0.3215***	0.2516**	0.1592*	0.0819	0.0585	0.0858	0.4040***
	(0.1180)	(0.0949)	(0.0634)	(0.0896)	(0.1029)	(0.0954)	(0.0804)	(0.1038)	(0.1136)	(0.1246)
urban	0.1050	0.3496***	0.4582***	0.3848***	0.2940**	0.2528**	0.2396***	0.2562**	0.2496**	0.2676**
	(0.1390)	(0.1063)	(0.0690)	(0.0964)	(0.1146)	(0.1005)	(0.0776)	(0.0998)	(0.1108)	(0.1221)
sofia_region	-0.2395	-0.0826	-0.1542	0.1409	0.1165	0.0650	-0.1932*	-0.0660	0.1059	0.3123**
	(0.1941)	(0.1436)	(0.0958)	(0.1289)	(0.1489)	(0.1195)	(0.1008)	(0.1296)	(0.1384)	(0.1249)
constant	3.9127***	4.2854***	4.7065***	5.1969***	5.4763***	3.6831***	4.2516***	4.4049***	4.9607***	5.3982***
	(0.3341)	(0.2579)	(0.1657)	(0.2352)	(0.2699)	(0.2369)	(0.1981)	(0.2599)	(0.2861)	(0.3118)
N	727	727	727	727	727	804	804	804	804	804
Pseudo R^2	0.0676	0.0814	0.0910	0.0755	0.0784	0.0610	0.0771	0.0577	0.0447	0.0539

Notes: Standard errors in parentheses; ***, ** and * denote significance at the 1%, 5% and 10% level respectively.

TABLE A4.14

QR estimates of returns to education in Bulgaria, by gender, 2001

	Males					*Females*				
	10^{th}	25^{th}	50^{th}	75^{th}	90^{th}	10^{th}	25^{th}	50^{th}	75^{th}	90^{th}
school	0.0344**	0.0310***	0.0427***	0.0305***	0.0166	0.0591***	0.0672***	0.0566***	0.0434***	0.0533***
	(0.0146)	(0.0090)	(0.0093)	(0.0104)	(0.0198)	(0.0097)	(0.0069)	(0.0066)	(0.0130)	(0.0139)
exp	-0.0119	-0.0021	-0.0073	-0.0083	-0.0183	-0.0286**	-0.0027	-0.0056	0.0010	-0.0022
	(0.0145)	(0.0096)	(0.0093)	(0.0096)	(0.0174)	(0.0120)	(0.0086)	(0.0075)	(0.0125)	(0.0129)
expsq	0.0000	-0.0002	-0.0000	0.0000	0.0004	0.0006**	-0.0000	0.0000	-0.0001	-0.0000
	(0.0003)	(0.0002)	(0.0002)	(0.0002)	(0.0004)	(0.0003)	(0.0002)	(0.0002)	(0.0003)	(0.0003)
1-2 years	0.3264*	0.1302	0.1169	0.0475	0.1444	0.1632	0.1899*	0.1434*	0.0242	-0.2422*
	(0.1826)	(0.1196)	(0.1180)	(0.1133)	(0.2064)	(0.1323)	(0.0971)	(0.0866)	(0.1419)	(0.1434)
3-5 years	0.3811**	0.2595***	0.3368***	0.2226**	0.1954	0.3104***	0.2344***	0.2024***	0.0979	-0.0896
	(0.1584)	(0.0984)	(0.0971)	(0.0948)	(0.1844)	(0.1143)	(0.0815)	(0.0737)	(0.1207)	(0.1231)
6-10 years	0.4175***	0.3386***	0.3563***	0.3078***	0.5136***	0.3324***	0.2404***	0.1735**	0.1313	0.0390
	(0.1428)	(0.0916)	(0.0899)	(0.0903)	(0.1742)	(0.1024)	(0.0760)	(0.0689)	(0.1146)	(0.1145)
>10 years	0.6039***	0.4742***	0.4382***	0.3826***	0.4424**	0.4531***	0.3706***	0.2695***	0.2001*	-0.0254
	(0.1419)	(0.0936)	(0.0897)	(0.0916)	(0.1872)	(0.1018)	(0.0766)	(0.0700)	(0.1190)	(0.1202)
married	0.2105*	0.2160***	0.1899***	0.1805**	0.1565	-0.0005	-0.0677	0.0267	-0.0004	0.0775
	(0.1073)	(0.0673)	(0.0693)	(0.0702)	(0.1331)	(0.0751)	(0.0516)	(0.0471)	(0.0779)	(0.0804)
bulgarian	0.3554***	0.3461***	0.2571***	0.0056	0.1565	0.0881	0.0441	0.1156	0.1649	-0.0025
	(0.1251)	(0.0845)	(0.0874)	(0.0858)	(0.1467)	(0.1091)	(0.0804)	(0.0748)	(0.1246)	(0.1224)
public	0.1637*	0.0902	0.0948*	0.1249**	0.0227	0.1491**	0.0850*	0.0637	0.0705	-0.0090
	(0.0869)	(0.0556)	(0.0543)	(0.0536)	(0.1079)	(0.0632)	(0.0472)	(0.0426)	(0.0702)	(0.0700)
urban	0.1397	0.1091*	0.0670	0.2215***	0.2549**	-0.0007	-0.0146	0.0483	0.0245	0.0825
	(0.1051)	(0.0629)	(0.0640)	(0.0609)	(0.1108)	(0.0744)	(0.0593)	(0.0530)	(0.0845)	(0.0840)
sofia_region	0.3006**	0.2431***	0.1954***	0.0598	0.1631	0.2504***	0.1211*	0.1276**	0.1153	0.1157
	(0.1234)	(0.0778)	(0.0740)	(0.0733)	(0.1363)	(0.0836)	(0.0621)	(0.0567)	(0.0921)	(0.0909)
constant	-1.5307***	-1.1305***	-0.8239***	-0.2106	0.1045	-1.3899***	-1.1972***	-0.8727***	-0.4472**	0.0337
	(0.2163)	(0.1345)	(0.1329)	(0.1442)	(0.2548)	(0.1618)	(0.1206)	(0.1133)	(0.2075)	(0.1976)
N	729	729	729	729	729	709	709	709	709	709
Pseudo R^2	0.1864	0.1696	0.1235	0.1044	0.0634	0.1494	0.1352	0.1173	0.0949	0.0999

Notes: Standard errors in parentheses; ***, ** and * denote significance at the 1%, 5% and 10% level respectively.

TABLE A4.15

QR estimates of returns to education in Bulgaria, by gender, 2003

	Males					Females				
	10th	25th	50th	75th	90th	10th	25th	50th	75th	90th
school	0.0383***	0.0463***	0.0461***	0.0490***	0.0589***	0.0478***	0.0505***	0.0625***	0.0699***	0.0673***
	(0.0057)	(0.0055)	(0.0046)	(0.0049)	(0.0082)	(0.0073)	(0.0034)	(0.0066)	(0.0063)	(0.0080)
exp	0.0090	0.0030	0.0055	0.0076**	0.0038	0.0185**	0.0173***	0.0211***	0.0123**	-0.0009
	(0.0057)	(0.0049)	(0.0039)	(0.0038)	(0.0059)	(0.0075)	(0.0033)	(0.0059)	(0.0050)	(0.0054)
expsq	-0.0003**	-0.0001	-0.0001*	-0.0002***	-0.0001	-0.0005**	-0.0003***	-0.0004***	-0.0001	0.0001
	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0002)	(0.0001)	(0.0001)	(0.0001)	(0.0001)
1-2 years	0.1139**	0.1482***	0.1064***	0.0768*	-0.0148	0.0540	0.0368	0.0399	0.0617	0.0597
	(0.0513)	(0.0486)	(0.0397)	(0.0405)	(0.0654)	(0.0704)	(0.0322)	(0.0598)	(0.0533)	(0.0612)
3-5 years	0.0350	0.1920***	0.1846***	0.1532***	0.0380	0.1607**	0.1518***	0.1577***	0.1382***	0.1696***
	(0.0497)	(0.0495)	(0.0398)	(0.0397)	(0.0631)	(0.0685)	(0.0313)	(0.0578)	(0.0509)	(0.0561)
6-10 years	0.1752***	0.3055***	0.2705***	0.2892***	0.1862**	0.0608	0.2103***	0.1859***	0.1835***	0.2521***
	(0.0588)	(0.0579)	(0.0465)	(0.0476)	(0.0752)	(0.0808)	(0.0363)	(0.0659)	(0.0585)	(0.0647)
>10 years	0.2494***	0.3087***	0.3388***	0.3332***	0.2963***	0.2190***	0.2294***	0.1894***	0.2699***	0.2971***
	(0.0560)	(0.0535)	(0.0418)	(0.0416)	(0.0667)	(0.0758)	(0.0346)	(0.0639)	(0.0563)	(0.0621)
married	-0.0072	0.0278	0.0501	0.0429	0.0396	-0.0083	0.0096	-0.0196	-0.0659*	0.0050
	(0.0416)	(0.0390)	(0.0310)	(0.0304)	(0.0500)	(0.0542)	(0.0238)	(0.0431)	(0.0374)	(0.0427)
bulgarian	0.0464	0.0056	0.1151**	0.1482***	0.1186*	0.0022	-0.0052	0.0118	0.1030*	0.0764
	(0.0615)	(0.0589)	(0.0469)	(0.0460)	(0.0712)	(0.0691)	(0.0348)	(0.0639)	(0.0565)	(0.0653)
public	0.1937***	0.1974***	0.2038***	0.1744***	0.1749***	0.1991***	0.1136***	0.0544	-0.0913**	-0.1116**
	(0.0410)	(0.0396)	(0.0314)	(0.0311)	(0.0489)	(0.0526)	(0.0239)	(0.0442)	(0.0398)	(0.0433)
urban	0.0958**	0.0852**	0.1256***	0.1564***	0.1483***	0.0233	0.0558**	0.0559	0.0322	0.1456***
	(0.0433)	(0.0420)	(0.0330)	(0.0322)	(0.0524)	(0.0629)	(0.0275)	(0.0491)	(0.0437)	(0.0483)
sofia region	0.1839**	0.1822**	0.1518**	0.2394***	0.5894***	0.1252	0.1861***	0.1214	0.1033	0.0511
	(0.0780)	(0.0759)	(0.0635)	(0.0630)	(0.0985)	(0.1091)	(0.0489)	(0.0909)	(0.0812)	(0.0956)
constant	-0.1684*	-0.0502	0.1014	0.3464***	0.5803***	-0.4047***	-0.3123***	-0.2146**	0.0137	0.2967**
	(0.0869)	(0.0833)	(0.0726)	(0.0775)	(0.1266)	(0.1231)	(0.0553)	(0.1085)	(0.1040)	(0.1242)
N	1296	1296	1296	1296	1296	1186	1186	1186	1186	1186
Pseudo R^2	0.0959	0.1415	0.1666	0.1907	0.1805	0.1270	0.1609	0.1876	0.1944	0.1915

Notes: Standard errors in parentheses; ***, ** and * denote significance at the 1%, 5% and 10% level respectively.

TABLE A4.16

Employees by economic sector in Bulgaria

Economic sectors	2000	2001	2002	2003	2004	2005	2006
Total	**1,900,940**	**1,899,874**	**1,927,690**	**2,079,932**	**2,152,301**	**2, 177 226**	**2,267,727**
Agriculture	89,650	79,534	78,539	74,935	72,017	69,983	66,120
Manufacturing	602,714	599,179	607,649	630,703	637,491	636,009	656,067
Electricity & water	59,285	59,234	59,112	58,748	58,316	56,978	55,533
Construction	98,053	97,053	95,336	108,110	119,618	141,829	166,259
Trade	210,678	222,722	239,072	282,492	301,573	312,272	336,810
Distribution & hotels	49,904	55,734	58,177	75,059	79,489	83,201	91,212
Transport & communication	164,248	163,419	162,152	159,742	162,981	160,153	162,115
Finance	27,891	27,486	27,655	29,125	30,934	33,599	37,995
Real estate	89,263	100,063	105,116	113,660	124,630	128,139	141,402
Public admin, Health & education	445,875	430,749	427,129	438,082	445,875	449,367	452,092
Others	63,379	64,701	67,753	109,276	119,377	105,696	102,122

Source: National Statistics Institute, Bulgaria

TABLE A4.17

Descriptive statistic – 1986 Town and Village survey and 2001 LSMS

	1986		2001	
	Mean	*SD*	*Mean*	*SD*
Female	0.5019	(0.500)	0.4930	(0.5001)
Age 16-35	31.763	(2.823)	28.076	(4.674)
Age 35-55	46.167	(5.692)	45.116	(5.643)
Sectors				
Agriculture	0.1176	(0.3222)	0.0514	(0.2210)
Manufacturing	0.3252	(0.4685)	0.2308	(0.4215)
Construction	0.0680	(0.2518)	0.0452	(0.2078)
Trade	0.0561	(0.2301)	0.1133	(0.3171)
Science& education	0.0548	(0.1547)	0.1070	(0.3093)
Regions				
Blagoevgrad	0.0338	(0.1807)	0.0466	(0.2108)
Bourgas	0.0540	(0.2259)	0.0758	(0.2648)
Varna	0.0561	(0.2302)	0.0626	(0.2423)
Veliko Turnovo	0.0416	(0.1996)	0.0355	(0.1850)
Vidin	0.0149	(0.1211)	0.0104	(0.1016)
Vratza	0.0310	(0.1734)	0.0285	(0.1665)
Gabrovo	0.0223	(0.1478)	0.0209	(0.1430)
Kurjali	0.0259	(0.1588)	0.0188	(0.1358)
Kustendil	0.0256	(0.1579)	0.0334	(0.1797)
Lovetch	0.0225	(0.1483)	0.0174	(0.1307)
Montana	0.0242	(0.1536)	0.0348	(0.1833)
Pazardjik	0.0394	(0.1945)	0.0320	(0.1760)
Pernik	0.0223	(0.1478)	0.0160	(0.1255)
Pleven	0.0400	(0.1960)	0.0438	(0.2047)
Plovdiv	0.0839	(0.2772)	0.0987	(0.2984)
Razgrad	0.0219	(0.1462)	0.0167	(0.1282)
Rousse	0.0274	(0.1634)	0.0299	(0.1704)
Silistra	0.0203	(0.1411)	0.0104	(0.1016)
Sliven	0.0223	(0.1478)	0.0216	(0.1453)
Smolian	0.0205	(0.1416)	0.0174	(0.1307)
Blagoevgrad	0.0338	(0.1807)	0.0466	(0.2108)
Bourgas	0.0540	(0.2259)	0.0758	(0.2648)

Notes: Standard deviations in parentheses. Samples relate to working individuals.

CHAPTER FIVE

RETURNS TO EDUCATION IN FOUR TRANSITION COUNTRIES: A QUANTILE REGRESSION APPROACH

5.1. Introduction

Understanding the heterogeneous pattern of returns to education across the conditional earnings distribution requires recognition of the effect that 'ability' and/or 'endogeneity' bias can have on the estimated returns. The underlying hypothesis is that ability and education are two distinct factors in the construction of human capital and that there exist some sort of complementarily between these two factors. The human capital theory implicitly recognises that the return to education may be heterogeneous[63]. Inter alia educational returns can vary across schooling levels and even across individuals with the same schooling level. Typically mean based regression models, like OLS, constitutes only a limited aspect of possibly more extensive changes involving the entire earnings distribution (Bushinsky, 1994). To place this idea into context we can envisage a process in which individuals are likely to differ with respect to not only the perceived benefits of education, but also the cost of education and the choices subsequently made in the labour market. In such circumstances the return to education is unlikely to be a single parameter; instead it is likely to vary systematically according to differences in individual's unmeasured characteristics, which in turn determine where in the overall earnings distribution an individual is placed. More generally any uncontrolled effect that is systematically correlated with an individual's position in the earnings distribution and which is also correlated with education attainment implies that the return to education is likely to vary across the earnings distribution. Accounting for this heterogeneity, therefore, requires an estimation strategy that allows the returns to education to differ at different points in the earnings distribution.

In this Chapter we use the QR model to address three important empirical questions. First, we examine the extent to which the return to education varies across the

[63] See Heckman *et al.,* 1997; Blundell *et al.*, 2000, Blundell *et al.*, 2001.

conditional earning distribution in four transition countries (Bulgaria, Serbia, Russia, and Tajikistan) at the same time period. Second, as male and female employment rates differ substantially, sample selection is potentially an important issue for this type of analysis. We consider the impact sample selection bias has on the returns to education in the QR framework by using Buchinsky's (1998, 2001) power series estimator[64]. Third, we investigate the empirical implications of allowing education to be endogenous (individual self-selection in the education process) across the distribution, using a control function approach proposed by Lee (2007). The importance of controlling for endogeneity lies in the fact that it allows us also to discuss the role that changes in the distribution of unobserved ability play in the observed changes of the earnings distribution. Fourth, in many cases it may be important to correct for both sample selection and endogeneity in education. Finally, having data for each country for the same year allows us to compare returns to human capital across the countries. The results of our studies dealing with these economies may provide important information for the policy makers.

The current Chapter is organized as follows. In Section 5.2, we comment on the data and methodology used in the QR estimates. Section 5.3 and Section 5.4 discuss the main results and conclusions.

5.2. The data

Our results are based on data from Bulgaria, Serbia, Russia and Tajikistan. The Bulgarian data are taken from the Living Standard Measurements Survey (LSMS), conducted in 2003. Since the construction of the Bulgarian data and the definitions of the variables have already been explained in Section 4.2.1, this discussion is not repeated here. Exclusion of individuals who were not of working age and restricting the sample to respondents who reported earnings accounts for a substantial reduction in the sample size. The subset of the Bulgarian data used in the estimation consists of a sample of 1,296 working men and 1,186 working women. The Serbian sample is drawn from the Serbian Living Standard Measurement Survey (LSMS)[65] in 2003.

[64] Buchinsky (2001) applies the sample selection model in a QR framework to estimate women's return to education in the U.S.

[65] The basic survey was carried out in 2002 on a representative sample of households in Serbia (without Kosovo and Metohija). The survey was repeated in 2003 on a panel sample. The 2003 survey was conducted in accordance

The Labour Market module is in line with a simplified version of the Labour Force Survey (LFS), with special additional questions to capture various informal sector activities and provide information on earnings. The sample size in 2003 consists of 2,548 individuals of which 2,450 report wages. The Serbian sample used in the analysis consists of 1,466 working men and 984 working women. The Russian NOBUS Survey (2003) provides our third source of data[66]. The NOBUS was conducted by the Russian Statistical Agency and the World Bank in 2003. It is a cross-sectional survey with large sample of around 45,000 households covering all Russian regions. The survey collects detailed information on household consumption and income, household demographics and labour market participation, access to health, education and social programs, and subjective perceptions of household welfare. The Russian working sample consists of 21,874 men and 24,318 women. Finally, the Tajikistan data are taken from the 2003 Tajikistan Living Standard Survey (TLSS)[67], which was done by the Word Bank and UNICEF in collaboration with the National Committee for Statistics (Goskomstat). The working Tajik sample in 2003 consists of 3,006 men and 1,832 women. The descriptive statistics for the countries and variables used are listed in Table 5.1.

There is an advantage in having data for each country for the same year. We have attempted to fit comparable specifications to the samples across the countries to put the analysis in a common framework[68]. In all four countries, we only retain individuals aged between 15 and 65 who are not currently full-time students and who supply information on their earnings. The dependent variable in the analysis is the logarithm of hourly earnings resulting from the primary occupation of the individuals and excludes earnings from secondary jobs, or from agricultural production, and non-monetary benefits[69]. The schooling variable is defined as total number of school grades completed by the individual, which coincides with the number of years of

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with the same methodological principles used in 2002 survey, with necessary changes referring only to the content of certain modules and the reduction in sample size. The aim of the repeated survey was to obtain panel data to enable monitoring of the change in the living standard within a period of one year, thus indicating whether there had been a decrease or increase in poverty in Serbia in the course of 2003.

[66] The NOBUS is the translated Russian abbreviation for the National Survey of Household Well-being and participation in Social Programs.

[67] The data are freely available at http://www.worldbank.org/lsms.

[68] Only in Russia an additional variable that captures a wage arrears effect is inlcued.

[69] Hourly earnings in all four countries are defined as reported monthly net earnings divided by 4.34 and then divided by reported weekly hours of work.

schooling. It takes on 18 different values: from 0 if the individual achieved no grade to 18 in the individual finished postgraduate activity. The vector of exogenous control variables used in the estimations also includes potential experience[70] (linear and quadratic terms), variables for individuals years of tenure with the current firm, and a set of regional variables in order to control for potential measurement issue that years of schooling may result in different levels of human capital accumulation over different regions if there are differences in schooling quality across regions. The regional dummies may also reflect differences in occupational structure, infrastructure quality, agglomeration economies etc. We have also included government sector variables as to control for the remaining effect of the old public sector wage structure. The Russian specifications are also supplemented with an additional variable that captures a wage arrears effect[71].

There are differences in the characteristics of men and women with respect to education as revealed in the summary statistics of the variables. All countries exhibit average levels of schooling above 10, the highest value being that of Bulgaria (13.9 for females). Moreover, in all four countries, the average length of time spent in education is higher for women than for men. For Bulgarian females for example, the average length of time spent in education is 0.8 years higher than for males. Average experience (which corresponds, in all countries, to Mincer's potential experience, age – schooling – 7) is generally above 19 and below 26 years. The high proportion of Serbian respondents has tenure in excess of 10 years, and this may point to a relatively less mobile labour market. The data also indicates that in all four countries women earn less than men and the gender differential is highest in Tajikistan. The lower is the proportion of Bulgarian respondents employed in public sector (around 30% for men and 39% for women). Not surprisingly, women's employment is more concentrated than men's in the public sector, and as a result women are less represented in private sector where both job opportunities and employment flexibility are less likely to be attractive to female workers. For instance, 69% of female employment in Russia is in the public sector compared to 60% of male employment. Only in Tajikistan the male employment in the public sector is higher.

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[70] Potential experience in all four countries is imputed as follows: age – 7– years of schooling.

[71] Wage arrears or unprecedented delays in the payment of wages have become an endemic feature of the Russian labour market. There are several forms wage arrears can take in Russia: 1) not paid wages 2) delayed but paid in full wages, 3) paid in time but not in full or 4) paid in part and not in time wages.

TABLE 5.1

Descriptive statistics by countries, 2003 working samples

		Bulgaria		Serbia		Russia		Tajikistan	
Variable	*Description*	*Males*	*Females*	*Males*	*Females*	*Males*	*Females*	*Males*	*Females*
lhwage	log of hourly wage	1.622	1.460	4.135	3.974	2.862	2.605	3.555	2.911
		(0.558)	(0.519)	(0.713)	(0.697)	(0.812)	(0.730)	(1.022)	(0.947)
school	total number years	13.056	13.875	11.421	11.428	11.337	11.882	11.596	10.844
	in education	(3.420)	(3.424)	(2.746)	(3.022)	(2.247)	(2.178)	(2.655)	(2.506)
exp	potential exp	19.972	18.653	25.164	24.146	21.587	21.488	19.444	18.456
		(12.264)	(11.179)	(12.970)	(12.485)	(11.463)	(11.002)	(12.156)	(11.576)
edbef91	=1 if graduate before	0.673	0.674	0.779	0.787	0.732	0.750	0.636	0.627
	1991	(0.469)	(0.468)	(0.414)	(0.409)	(0.442)	(0.432)	(0.481)	(0.484)
married	=1 if married	0.680	0.728	0.749	0.709	0.761	0.624	0.814	0.633
		(0.467)	(0.445)	(0.434)	(0.454)	(0.426)	(0.484)	(0.389)	(0.482)
tenure1	=1 if<1 year	0.270	0.226	0.218	0.263	0.150	0.120	0.197	0.178
		(0.444)	(0.418)	(0.413)	(0.441)	(0.357)	(0.325)	(0.398)	(0.383)
tenure3	=1 if 1-2 years	0.177	0.163	0.029	0.032	0.201	0.191	0.155	0.160
		(0.382)	(0.369)	(0.169)	(0.175)	(0.401)	(0.393)	(0.362)	(0.367)
tenure4	=1 if 3-5 years	0.191	0.206	0.080	0.072	0.135	0.120	0.205	0.199
		(0.393)	(0.404)	(0.272)	(0.259)	(0.342)	(0.325)	(0.403)	(0.400)
tenure5	=1 if 6-10 years	0.124	0.142	0.098	0.118	0.172	0.169	0.145	0.138
		(0.330)	(0.350)	(0.297)	(0.323)	(0.377)	(0.374)	(0.352)	(0.345
tenure6	=1 if>10 years	0.238	0.263	0.575	0.515	0.342	0.400	0.298	0.324
		(0.426)	(0.440)	(0.495)	(0.500)	(0.474)	(0.490)	(0.457)	(0.468)
public	=1 if in public sector	0.306	0.390	0.457	0.506	0.595	0.694	0.465	0.443
		(0.461)	(0.488)	(0.498)	(0.500)	(0.491)	(0.461)	(0.499)	(0.497)
arrears	=1 if arrears effect	-	-	-	-	0.189	0.138	-	-
						(0.392)	(0.345)		
region1	=1 if region1	0.100	0.116	0.143	0.191	0.214	0.222	0.085	0.108
		(0.299)	(0.321)	(0.350)	(0.393)	(0.410	(0.415	(0.280)	(0.311)
region2	=1 if region2	0.148	0.130	0.261	0.273	0.139	0.140	0.297	0.266
		(0.355)	(0.336)	(0.439)	(0.446)	(0.346)	(0.347)	(0.457)	(0.442)
region3	=1 if region3	0.122	0.118	0.119	0.092	0.131	0.131	0.368	0.395
		(0.327)	(0.323)	(0.324)	(0.290)	(0.337)	(0.338)	(0.482)	(0.489)
region4	=1 if region4	0.194	0.226	0.104	0.098	0.130	0.127	0.130	0.127
		(0.395)	(0.418)	(0.306)	(0.297)	(0.336)	(0.333)	(0.336)	(0.333)
region5	=1 if region5	0.212	0.185	0.205	0.196	0.134	0.131	0.120	0.103
		(0.409)	(0.389)	(0.404)	(0.397)	(0.341)	(0.337)	(0.325)	(0.304)
region6	=1 if region6	0.225	0.224	0.166	0.149	0.082	0.081		
		(0.417)	(0.417)	(0.373)	(0.357)	(0.275)	(0.272)		
region7	=1 if region7	-	-	-	-	0.169	0.169	-	-
						(0.375)	(0.375)		
N		1296	1186	1466	984	21874	24318	3006	1832

Source: Bulgaria, LSMS (2003), Serbia, LSMS (2003), Russia NOBUS (2003), Tajikistan, TLSS (2003).
Notes: Standard deviations in brackets.
The control regional variables for Bulgaria are: region 1- North West, region 2- North East, region 3- Central, region 4- South West, region 5- South East, region 6- Sofia region;
For Serbia: region 1- Belgrade, region 2- Vojvodina, region 3- West Serbia, region 4- Šumadija and Pomoravlje, region 5- East Serbia, region 6- South-East Serbia;
For Russia: region 1-Central, region 2- North-West, region 3- Siberia, region 4- South, region 5- Far-East, region 6- Urals, and region 7- Volga; Tajikistan is divided into 4 regions: Leninabad (renamed into Sogd) in the northwest of the country, Khatlon in the southwest, Rayons of Republican Subordination (RRS) in the middle and to the west of the country, and Gorno-Badakhshan Autonomous Oblast (GBAO) in the East. The capital, Dushanbe, in the RRS oblast, is a separately administrated area. The regional control variables for Tajikistan are: region 1- GBAO, region 2- Sogdian, region 3- Kahtlon, region 4- Dushanbe, and region 5-RRS;

5.3. Methodology

5.3.1. Sample selection in the QR model

The estimation approach in the present Chapter is based on the quantile regressions. Since the QR methodology has been already explained in Section 4.2.5, this discussion is not repeated here. In this Section we focus on the sample selection bias in a QR framework as the selectivity correction in the QR models is less well developed and little consensus exist on the appropriate sample selection procedure. Importantly, the bivariate normality assumption between the error terms in the earnings and participation equation does not hold in the quantile regression model as the conditional mean of the quantile regression's disturbance term need not be zero. Buchinsky[72] (1998 and 2001) was the first to consider the problem of estimating QR in the presence of sample selection. He suggests an approach using the non-parametric procedure of Newey (1991, 2009) to deal with this problem. Specifically, Buchinsky (1998) assumes conditional independence between the error terms and the regressors given the probability of selection. As in the seminal work of Heckman (1974), the outcome equation and the latent selection function are linear in the covariates. The error terms in both equations are independent of the covariates (conditional on the selection probability), but in contrast to the model of Heckman their joint distribution is completely unrestricted (Huber and Melly, 2011).

The estimation procedure used can be briefly described as follows. In the first step the selection parameter is estimated from a standard Probit model[73]. In the second step, as suggested by Buchinsky (1998), we use a power series expansion of the inverse Mill's ratio of the normalized estimated index which is designed to approximate the unknown quantile functions of the truncated bivariate distribution of the error terms in the wage and participation equations[74]. The key assumption is the single index restriction on the error term. Buchinsky's selection correction technique[75] requires a valid exclusion variable. A range of familiar variables are used as covariates in the

[72] Some recent studies criticize Buchinsky method and prove that his assumptions are very strong. Melly (2006) argues that Buchinsky estimator is consistent only when all quantile regression lines are parallel. Melly (2006) also concludes that rejecting the hypothesis for homogeneity (that all slope coefficients are constant), biases the estimator of Buchinsky.

[73] We follow Buchinsky (1998) modification that the error distribution is not assumed to take any parametric form.

[74] The first order approximation will be sufficient if the error term is normally distributed. According to Buchinsky (1998) addition of more terms to the series expansion resulted, in some cases, in severe multicollinearity problems.

[75] This method is incorporated into the Machado and Mata technique by Albrecht *et al.* (2009).

participation equation, including the presence of dependent children in the household which, as in Buchinsky (1998), is used to identify participation on the assumption that this variable is exogenous[76]. In the analysis, we follow the same identifying restriction.

To get unbiased estimates of β_θ for the male and female respondents the power series approximation term is introduced:

$$\ln Y_{\theta i} = \beta_\theta X_i + h_\theta(g_i) + \xi_{\theta i} \tag{5.1}$$

where:

$$h_\theta(g_i) = Quant_\theta(\xi_{\theta i} | Z_i, g_i > 0) \tag{5.2}$$

The values of $g_i = \hat{\gamma} Z_i$ are used to expand $h_\theta(g_i)$ in a power series approximation. The procedure has been discussed in semi-parametric two-step estimation for sample selection bias at the mean (Section 4.2.4). The term $h_\theta(g_i)$ includes information about the unobservable that affects individual labour force participation decisions and it plays a similar role as the Mills ratio plays in the usual Heckman (1979) procedure, but is quantile-specific and more general so as not to assume normality.

5.3.2. Endogeneity in the QR model

In many empirical regression models, it is quite common to have a regressors that are endogeneous[77]. If schooling is endogenous, the returns to schooling obtained by standard quantile regressions may be misleading. Amemiya (1982) was the first to seriously consider QR in the presence of endogenous regressors. He shows the consistency and asymptotic normality of a class of two-stage median regression estimators. Subsequent work by Powell (1983) and Chen and Portnoy (1996) extend this approach but maintain the focus primarily on the conditional median problem (Melly, 2006). Arias *et al.* (2001) applied IV-QR in the same way as the classical

[76] A commonly made assumption in the literature is that fertility decisions are exogenous to decisions about labour force participation.

[77] Such endogeneity may arise when this regressor and the dependent variable are determined simultaneously or when this regressor is a consequence of self-selection. Education is the result of constrained optimizing decision. Another possibility for endogenety in schooling is that the variable may be measured with errors. IV methodology can overcome the endogeneity bias by making use of natural experiments that enable the identification of variation in education that is exogenous to wages.

Two-Stage Least Squares estimator[78] the so called 'fitted value' approach. Chernozhukov and Hansen (2005), however, show that Arias *et al.* (2001) approach is inconsistent unless treatment effects are homogeneous. Chernozhukov and Hansen (2006, 2008)[79] approach is practical in the sense that it can be computed through a series of quantile regression steps. Abadie *et al.* (2002) considered quantile regression methods for estimating endogenous treatment effects focusing on a binary treatment case. Their estimator applies only to a restrictive case – a binary treatment variable and a single binary instrument and imposes a monotonicity condition. Chen *et al.* (2003) consider a partially linear median regression model with endogenous regressors. Chesher (2003) develops a general nonlinear model which may be viewed as an extension of the recursive causal chain models discussed by Strotz and Wold (1960). Chesher's approach can be viewed as a 'local, non-separable, nonparametric' control function approach (Lee, 2007). Ma and Koenker (2006) propose two estimators assuming a finite-dimensional parametric restriction and integrating over the nonparametric estimates.

To control for endogeneity bias in a QR framework, we adopt the control function approach proposed by Lee (2007). As an alternative to existing methods in the literature, Lee's (2007) methodology extends the control function approach to structural quantile regression model semi-parametrically. He shows that under suitable conditions, the estimator obtained from the control function approach is consistent and asymptotically normally distributed.

Formally Lee (2007) considers the following model, which is a semi-parametric quantile regression version of Newey (1999):

$$Y = X\beta(\tau) + Z_1'\gamma(\tau) + U \qquad (5.3)$$

[78] Other empirical papers have followed a similar strategy; see Levin (2001) and Ribeiro (2001). The consistency and asymptotic normality of the QR estimator with instrumental variable (IV), in a two-stage framework has been obtained by Chen and Portnoy (1996). The author extends the results of Amemia (1982) and Powell (1983) for the last absolute deviation (LAD) estimator in a more general framework under weak hypotheses. Kim and Muller (2004) criticise Arias *et al* (2001), Garcia *et al.* (2001) approach, arguing that their methodology may be delicate for the general type of problem consider since using LS estimation in the first stage combined with Amemiya's reformulation of the dependent variable may produce asymptotic bias and also it can destroy the robustness properties of the quantile regressions.

[79] Their estimator has already been used in empirical implications, e.g. Hausman and Sidak (2004), Chevalier *et al.* (2006), Kondylis (2005), Melly (2005), Autor *et al.* (2006), Eren (2009), Galvao and Rojas (2009).

$$X = \mu(\alpha) + Z'\pi(\alpha) + V \tag{5.4}$$

where Y is the dependent variable, X is real-valued continuously distributed endogenous explanatory variable, $Z \equiv (Z_1, Z_2)$ is a $(d_z \times 1)$ vector of exogenous explanatory variables, U and V are real-valued unobserved random variables, $\beta(\tau)$ and $\gamma(\tau)$ are unknown structural parameter of interest, $\mu(\alpha)$ is an unknown parameter, $\pi(\alpha) \equiv [\pi_1(\alpha), \pi_2(\alpha)]$ vector is a $(d_z \times 1)$ vector of unknown parameters for some τ and α such that $0 < \tau < 1$, and $0 < \alpha < 1$. For identification it is assumed that there is at least one component of Z that is not included in Z_1, and that there is at least one non-zero coefficient for the excluded components of Z. That is, $d_{z1} < d_z$ and $\pi_2(\alpha) \neq 0$, where d_{z1} is the dimension of Z_1. The model is completed with the following assumptions:

$$Q_{U|X,Z}(\tau \mid x, z) = Q_{U|V,Z}(\tau \mid v, z) = Q_{U|V}(\tau \mid v) \equiv \lambda_\tau(v) \tag{5.5}$$

$$Q_{V|Z}(\alpha \mid z) = 0 \tag{5.6}$$

where $\lambda_\tau(.)$ is a real- valued, unknown function of V, $Q_{U|X,Z}(\tau \mid x, z)$ denotes the τ th quantile of U conditional on $X = x$, and $Z = z$. The first equality in (5.5) holds when v is the value of V that satisfies $v = x - \mu(\alpha) - z'\pi(\alpha)$. The second equality in (5.5) suggests a quantile independence of U on Z conditioned on V (Lee, 2007).

The estimation procedure consists of two steps. The first step is construction of estimated residuals $\hat{V}_i = X_i - \hat{\mu} - Z'_i \hat{\pi}$ $(i = 1,...,n)$ by a linear quantile regression of X on (1, Z), where $(\hat{\mu}, \hat{\pi})$ is a solution to

$$\min_{\mu,\pi} n^{-1} \sum_{i=1}^{n} \rho_\alpha (X_i - \mu - Z_i'\pi) \tag{5.7}$$

where ρ_α is the 'check' function such that $\rho_\alpha(u) = |u| + (2\alpha - 1)u$ for $0 < \alpha < 1$. The second step is estimation of a partially linear quantile regression of Y on (X, Z_1, V)

using the estimated residuals $\hat{V}_i$ in place of unobserved V_i's via power series estimation[80] (Lee, 2007).

In the return to education estimates, the reduced-form schooling residuals V are interpreted as 'individual ability' and U is not assumed to be independent of V. The approach corrects for endogeneity by adding a residual power series as additional explanatory variables and therefore is interpreted as a variant of the control function approach (e.g., Newey, 1999; Blundell and Powell, 2003). Under suitable conditions, Lee (2007) shows $n^{-1/2}$ consistency and asymptotic normality of the estimator[81].

Finding an instrument that is orthogonal to the disturbance term in the earning equation has been a topic of great debate. As in chapter four, in the control function approach, we utilise the regional human capital concentration as an instrument. The choice of the instrument is motivated by data availability and by our objective of maintaining the same specification across countries. The relevance of the instrument is tested at the mean of the first-stage regression. Durbin-Hausman Wu test (DWH) (Davidson and MacKinnon, 1993) is used to test the hypothesis of endogeneity of schooling[82].

Figure 5.1 illustrates the cross-sectional district variation in average schooling by countries. In all four countries, we find high concentration of more endowed workers within the capitals. In Bulgaria, for instance, respondents in Sofia region have on average 14.5 years spent in education, which is significantly higher compared to the North-West region. Similarly, in Serbia, individuals in West region appear to be the most disadvantaged.

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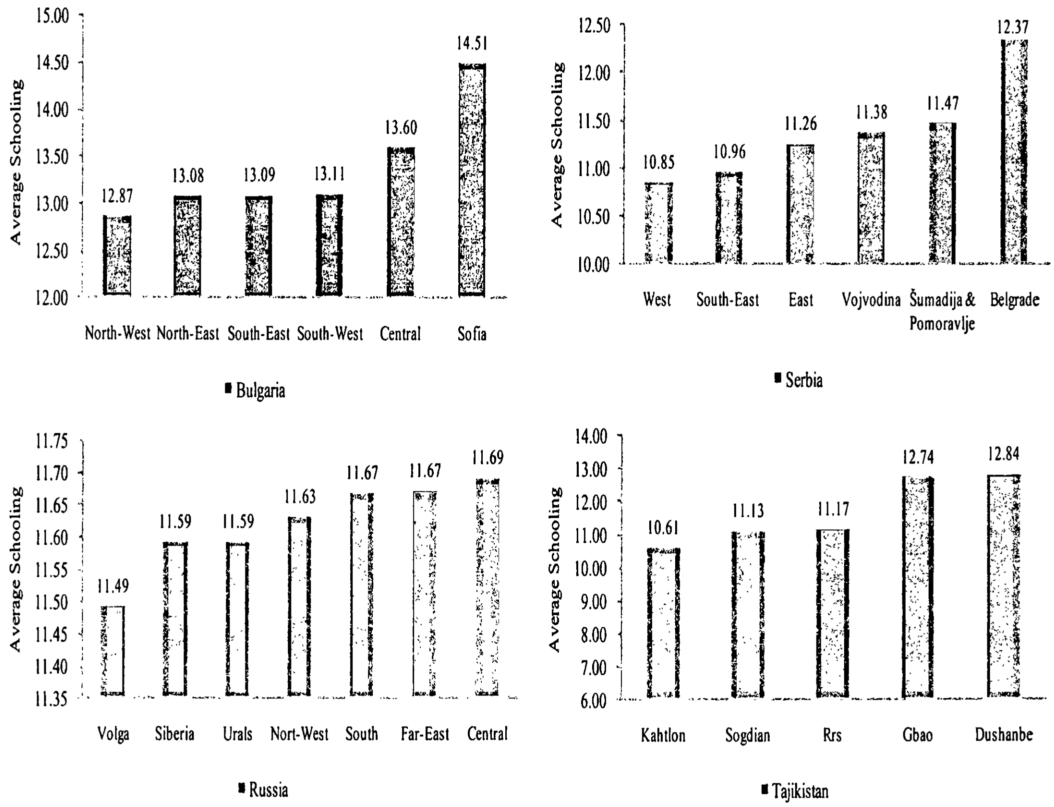
[80] This two-step estimator resembles closely the approach of Buchinsky (1998) which we apply in our quantile specification, where the sample selection bias is corrected for non-parametrically by a two-stage procedure. The author also imposes no functional form restrictions on the stochastic relationship between the reduced - form residual and the disturbance term in the primary equation conditional on observable explanatory variables. This residual term captures the effect of any unobserved ability variables such as ability which influence individual productivity.

[81] See Lee (2007) for details.

[82] The DWH test was conducted at the mean as no equivalent test exists for the QR model. We do not need to go beyond testing for endogeneity in the linear model even when we estimate QR model in the end. The reduced form model in either case is the same and we still want to test the underlying hypothesis that the error terms in the reduced form model are uncorrelated regardless to whether we are estimating a QR or OLS model.

FIGURE 5.1: Regional variations in the average schooling by countries



Source: Bulgaria, LSMS (2003), Serbia, LSMS (2003), Russia NOBUS (2003), Tajikistan, TLSS (2003).

5.4. Empirical results

Tables 5.2 to 5.5 provide the main results concerning the returns to education across the wage distribution. The unadjusted results are presented in the first row followed by the unadjusted estimates where schooling is interacted with an indicator whether individuals have graduated before 1991, sample selection and endogeneity adjusted QR estimates. This allows comparison of the results from correction methods. We report the QR estimates for three values of θ (10th, 50th, and 90th percentiles) for males and females respectively. The full set of estimates is presented in Appendix Tables A5.1-A5.20. The interpretation of the QR coefficients is analogous to the OLS regression, where the coefficients measure the influence of the regressor variables on the conditional mean of the dependent variables. In the quantile case the coefficients represents the influence of the regressors on the conditional θ-quantile of the dependent variable. In Tables 5.6 and 5.7 we also test to see whether differences across the quantiles are statistically significant. The second and the third columns represent the F-test for the equality of coefficients at the 10th and 90th percentile for males and females, followed by a joint test of equality of coefficients at all quantiles. The estimated returns to education are also plotted by countries and by gender for each percentile along with the 95 % confidence interval for each point estimates (Figures 5.2 to 5.5). Superimposed on the plots is a dashed line representing the OLS estimate of the effect of education on earnings. Each side of the OLS estimate is a dotted line which shows the associated 95% confidence interval of the estimate. It is important to be able to contrast the QR results with those for the mean regression technique which allows their comparison.

5.4.1. Unadjusted QR estimates

In the QR estimates where we treat schooling as exogenous (the first rows of Tables 5.2 to 5.5), the effect of education on earnings is positive and statistically different from zero at each of the reported percentiles. Specifically, the unadjusted estimates of the returns to education are higher at higher points of the conditional earning distribution for both males and females in Bulgaria and for females in Tajikistan. For instance, the return to education in Bulgaria increases from 3.9% to 6.2% in the male specification and from 4.7% to 7.2% in the female specification between the 10^{th} and 90^{th} quantiles (see Table 5.2).[83] To put it differently, the earning increments associated with education is higher for those individuals whose unobservable characteristics place them at the top of the conditional earnings distribution. The results for Bulgaria, reported here are consistent with previous estimates reported in the literature. Martins and Pereira (2004) and Flabbi *et al.* (2008), for example, report higher returns to education at the top end of the conditional wage distribution. The more extreme case of this increasing pattern is female Tajikistan case. In Tajikistan the return at the first quanitle for females is no greater than 6% and it reaches 12% at the 90^{th} quantile. Compared to the corresponding returns for Tajik males, where we find little heterogeneity, the returns to education for females tend to increase rapidly, suggesting that inequality is more pronounced for females than males in terms of educational returns.

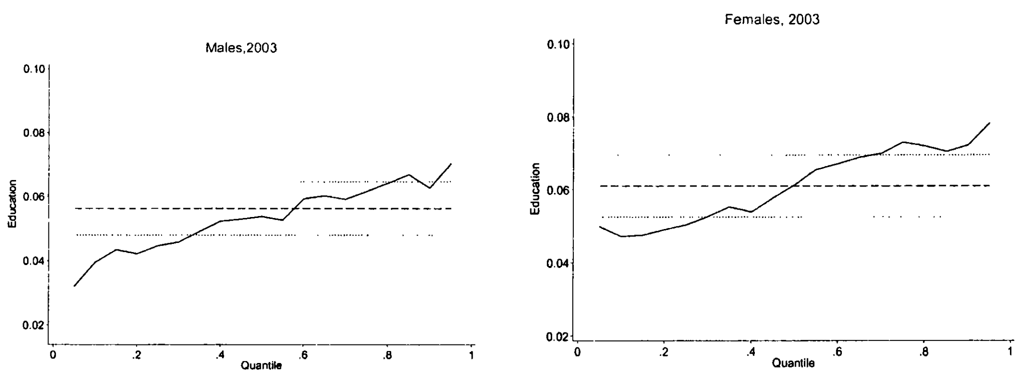
[83] The difference between the returns to education at the 10^{th} and 90^{th} in Bulgaria for males is significant at 5% level (the *F*- statistic for males is 4.65^{**}).

TABLE 5.2: Returns to education by quantiles, Bulgaria, 2003

	Males			Females		
	θ=0.10	θ=0.50	θ=0.90	θ=0.10	θ=0.50	θ=0.90
(1)Unadjusted QR estimates						
Education	0.0395***	0.0537***	0.0626***	0.0472***	0.0612***	0.0722***
	(0.0050)	(0.0044)	(0.0095)	(0.0062)	(0.0057)	(0.0087)
Pseudo R^2	0.1400	0.2071	0.1978	0.1567	0.2037	0.2086
N	1296	1296	1296	1186	1186	1186
(2)Unadjusted QR estimates with interactions						
Education	0.0516***	0.0505***	0.0556***	0.0378***	0.0577***	0.0686***
	(0.0090)	(0.0079)	(0.0164)	(0.0110)	(0.0100)	(0.0148)
Educ*bef1991	-0.0153	0.0054	0.0148	0.0112	0.0061	0.0039
	(0.0105)	(0.0092)	(0.0185)	(0.0122)	(0.0118)	(0.0175)
Before 1991	0.1553	-0.0685	-0.2446	-0.0125	-0.1181	0.0071
	(0.1578)	(0.1307)	(0.2477)	(0.1897)	(0.1807)	(0.2652)
Pseudo R^2	0.1420	0.2073	0.1996	0.1588	0.2043	0.2094
N	1296	1296	1296	1186	1186	1186
(3 Sample selection adjusted QR estimates[a]						
Education	0.0336***	0.0502***	0.0627***	0.0481***	0.0574***	0.0879***
	(0.0109)	(0.0081)	(0.0115)	(0.0171)	(0.0130)	(0.0145)
Pseudo R^2	0.1408	0.2073	0.1997	0.1572	0.2040	0.2093
N	1296	1296	1296	1186	1186	1186
(4)Endogeneity adjusted QR estimates[b]						
Education	0.0465**	0.0511***	0.1070***	0.0451**	0.0499***	0.1071***
	(0.0184)	(0.0131)	(0.0313)	(0.0216)	(0.0095)	(0.0202)
Pseudo R^2	0.1608	0.2203	0.2173	0.1581	0.2162	0.2140
N	1296	1296	1296	1186	1186	1186

Notes: [a] QR estimates with the second order power series correction terms; [b] QR estimates with the fifth order polynomials of the reduced form residuals; (i) Bootstrapped errors in parentheses; ***, ** and * denote significance at the 1%, 5% and 10% level respectively ; (ii) Full results are given in Appendix 5.1; (iii) A tests on excluding our potential instrument (average regional schooling) from the reduced form equations conducted at the mean, yield an F-statistics of 19.39***and partial R^2 of 0.0149 for males, and F= 17.54***and partial R^2 of 0.0147 for females; (iv) The DWH test conducted at the mean in males specification is 24.966, p-value=0.0000 and 17.985, p-value=0.00002 in females, so we do reject the null hypothesis in both specifications; (v) Sargan test, when combining mothers education and regional average schooling results in a p-value= 0.152 for males and p-value= 0.696 in female specification, which is failure to reject the null hypothesis.

FIGURE 5.2: Returns to education by quantiles, Bulgaria 2003



Notes: The y-axis measures the return to education coefficients; the x-axis depicts the selected quantiles of the conditional wage distribution for male and females. The horizontal line plots the OLS estimate and its 95% confidence interval.

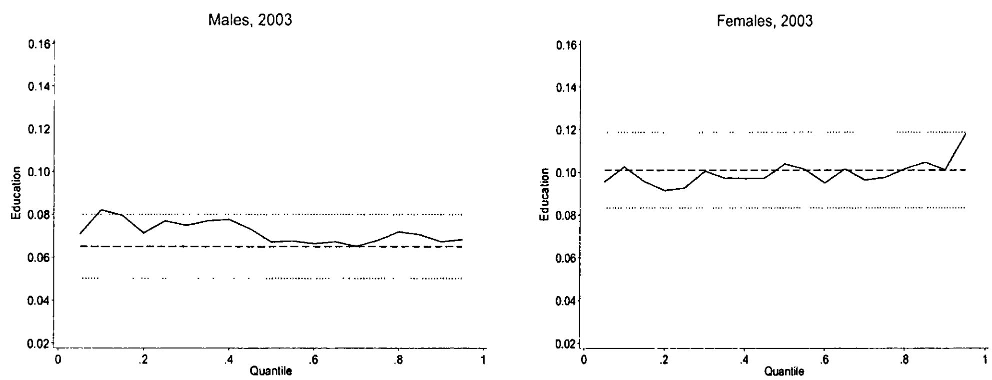
TABLE 5.3

Returns to education by quantiles, Serbia, 2003

	Males			Females		
	θ=0.10	θ=0.50	θ=0.90	θ=0.10	θ=0.50	θ=0.90
(1)Unadjusted QR estimates						
Education	0.0821***	0.0672***	0.0674***	0.1025***	0.1039***	0.1012***
	(0.0133)	(0.0064)	(0.0213)	(0.0157)	(0.0082)	(0.0202)
Pseudo R^2	0.1003	0.0740	0.0684	0.1683	0.1443	0.1282
N	1466	1466	1466	984	984	984
(2)Unadjusted QR estimates with interactions						
Education	0.0437	0.0376**	0.0975**	0.0673**	0.1284***	0.1596***
	(0.0365)	(0.0191)	(0.0489)	(0.0278)	(0.0254)	(0.0552)
Educ*bef1991	0.0342	0.0337*	-0.0389	0.0396	-0.0361	-0.0628
	(0.0384)	(0.0203)	(0.0536)	(0.0316)	(0.0274)	(0.0577)
Before 1991	-0.5609	-0.4137	0.5694	-0.4002	0.4102	0.8379
	(0.4833)	(0.2523)	(0.6339)	(0.4049)	(0.3575)	(0.7259)
Pseudo R^2	0.1023	0.0750	0.0700	0.1705	0.1459	0.1320
N	1466	1466	1466	984	984	984
(3) Sample selection adjusted QR estimates[a]						
Education	0.0832***	0.0676***	0.0633**	0.1034***	0.1083***	0.1124***
	(0.0126)	(0.0064)	(0.0208)	(0.0155)	(0.0094)	(0.0226)
Pseudo R^2	0.1022	0.0758	0.0718	0.1687	0.1467	0.1320
N	1466	1466	1466	984	984	984
(4)Endogeneity adjusted QR estimates[b]						
Education	0.0182	0.0879***	0.1631**	0.0261	0.1107***	0.2096***
	(0.0306)	(0.0090)	(0.0335)	(0.0334)	(0.0128)	(0.0426)
Pseudo R^2	0.1058	0.0809	0.0816	0.1818	0.1598	0.1431
N	1466	1466	1466	984	984	984

Notes: [a] QR estimates with the second order power series correction terms; [b] QR estimates with the fifth order polynomials of the reduced form residuals; (i) Bootstrapped errors in parentheses; ***, ** and * denote significance at the 1%, 5% and 10% level; (ii) Full results are given in Appendix; (iii) A tests on excluding our potential instrument (average regional schooling) from the reduced form equations conducted at the mean, yield an F-statistics of 37.56*** and partial R^2 of 0.0251 for males, and F=14.71*** and partial R^2 of 0.0149 for females; (iv)The DWH test conducted at the mean is 51.154, p-value=0.000 in males specification and 27.105, p-value 0.000, and we do reject the null hypothesis in both specifications.

FIGURE 5.3: Returns to education by quantiles, Serbia 2003



Notes: The y-axis measures the return to education coefficients; the x-axis depicts the selected quantiles of the conditional wage distribution for male and females. The horizontal line plots the OLS estimate and its 95% confidence interval.

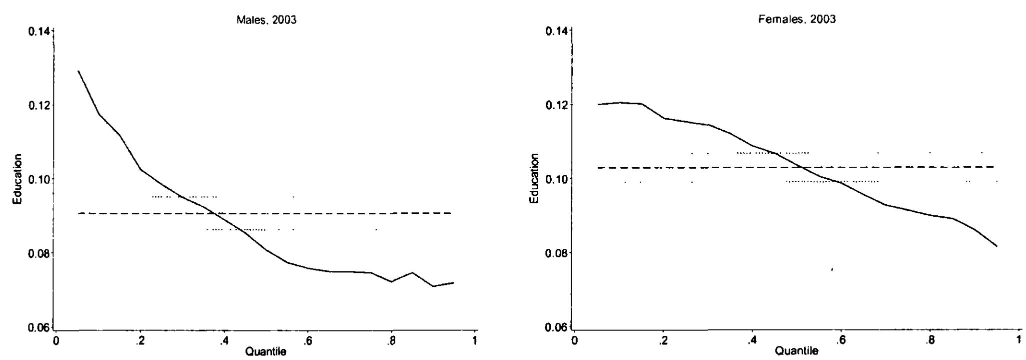
TABLE 5.4

Returns to education by quantiles, Russia, 2003

	Males			Females		
	θ=0.10	θ=0.50	θ=0.90	θ=0.10	θ=0.50	θ=0.90
(1) Unadjusted QR estimates						
Education	0.1174***	0.0808***	0.0709***	0.1205***	0.1035***	0.0859***
	(0.0045)	(0.0023)	(0.0035)	(0.0033)	(0.0025)	(0.0031)
Pseudo R^2	0.1514	0.1387	0.1324	0.1321	0.1449	0.1343
N	21874	21874	21874	24318	24318	24318
(2) Unadjusted QR estimates with interactions						
Education	0.1244***	0.0834***	0.0769***	0.1079***	0.0978***	0.0850***
	(0.0092)	(0.0045)	(0.0068)	(0.0067)	(0.0042)	(0.0058)
Educ*bef1991	-0.0074	-0.0037	-0.0094	0.0186**	0.0079*	0.0009
	(0.0108)	(0.0052)	(0.0078)	(0.0077)	(0.0048)	(0.0067)
Before 1991	-0.0391	-0.0580	0.1029	-0.2406**	-0.1367**	-0.0809
	(0.1274)	(0.0629)	(0.0942)	(0.0980)	(0.0606)	(0.0843)
Pseudo R^2	0.1518	0.1392	0.1325	0.1324	0.1450	0.1345
N	21874	21874	21874	24318	24318	24318
(3) Sample selection adjusted QR estimates[a]						
Education	0.1463***	0.0854***	0.0677***	0.1940***	0.1519***	0.1216***
	(0.0069)	(0.0037)	(0.0055)	(0.0283)	(0.0157)	(0.0212)
Pseudo R^2	0.1565	0.1412	0.1336	0.1324	0.1451	0.1344
N	21874	21874	21874	24318	24318	24318
(4)Endogeneity adjusted QR estimates[b]						
Education	0.1169***	0.0829***	0.1287***	0.0858***	0.1030***	0.1268***
	(0.0143)	(0.0053)	(0.0210)	(0.0109)	(0.0062)	(0.0170)
Pseudo R^2	0.1517	0.1391	0.1334	0.1338	0.1464	0.1368
N	21874	21874	21874	24318	24318	24318
(5)Endogeneity and selection adjusted QR estimates[c]						
Education	0.1341***	0.0894***	0.1373***	0.1560***	0.1663***	0.1834***
	(0.0150)	(0.0066)	(0.0199)	(0.0274)	(0.0190)	(0.0294)
Pseudo R^2	0.1568	0.1416	0.1348	0.1343	0.1468	0.1370
N	21874	21874	21874	24318	24318	24318

Notes: [a] QR estimates with the second order power series correction terms; [b] QR estimates with the fifth order polynomials of the reduced form residuals; [c] QR estimates with the second order power series correction terms and fifth order polynomials of the reduced form residuals;
(i) Bootstrapped errors in parentheses; ***, ** and * denote significance at the 1%, 5% and 10% level; (ii) Full results are given in Appendix; (iii) A tests on excluding our potential instrument (average regional schooling) from the reduced form equations conducted at the mean, yield an F-statistics of 28.91*** and partial R^2 of 0.0013 for males, and F= 22.65*** and partial R^2 of 0.0009 for females; (iv) The DWH test conducted at the mean is 85.734, p-value =0.0000 in males specification and 140.062, p-value 0.0000 for females, and we do reject the null hypothesis.

FIGURE 5.4: Return to education by quantiles, Russia 2003



Notes: The y-axis measures the return to education coefficients; the x-axis depicts the selected quantiles of the conditional wage distribution for male and females. The horizontal line plots the OLS estimate and its 95% confidence interval.

TABLE 5.5

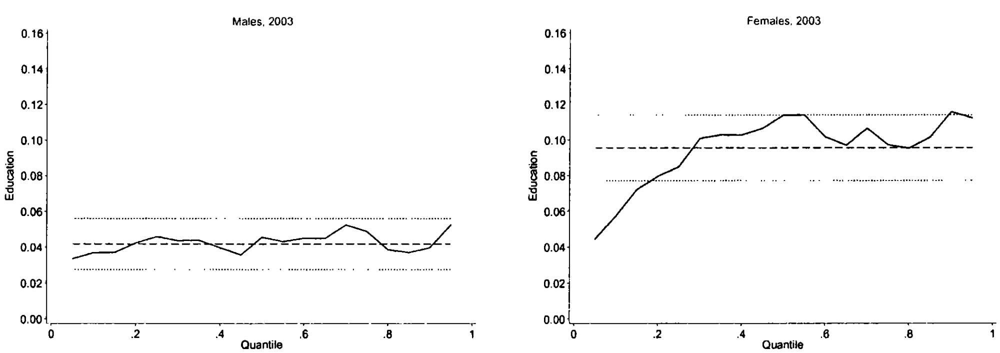
Returns to education by quantiles, Tajikistan, 2003

	Males			Females		
	θ=0.10	θ=0.50	θ=0.90	θ=0.10	θ=0.50	θ=0.90
(1)Unadjusted QR estimates						
Education	0.0370**	0.0457***	0.0398**	0.0575***	0.1139***	0.1157***
	(0.0151)	(0.0088)	(0.0165)	(0.0061)	(0.0124)	(0.0213)
Pseudo R^2	0.0802	0.0974	0.0815	0.0765	0.1353	0.1870
N	3006	3006	3006	1832	1832	1832
(2)Unadjusted QR estimates with interactions						
Education	0.0466**	0.0729***	0.1007***	0.0547***	0.1246***	0.1511***
	(0.0229)	(0.0127)	(0.0247)	(0.0093)	(0.0156)	(0.0278)
Educ*bef1991	-0.0145	-0.0502***	-0.0657**	0.0128	-0.0259	-0.0648*
	(0.0269)	(0.0148)	(0.0307)	(0.0121)	(0.0197)	(0.0347)
Before 1991	-0.0666	0.4963***	0.7528*	-0.0970	0.4074*	0.8615**
	(0.3333)	(0.1864)	(0.3858)	(0.1516)	(0.2408)	(0.4219)
Pseudo R^2	0.0815	0.0994	0.0827	0.0771	0.1364	0.1904
N	3006	3006	3006	1832	1832	1832
(3) Sample selection adjusted QR estimates[a]						
Education	0.0741***	0.0794***	0.0554***	0.0591***	0.1148***	0.1253***
	(0.0162)	(0.0111)	(0.0185)	(0.0079)	(0.0141)	(0.0196)
Pseudo R^2	0.0896	0.1048	0.0863	0.0766	0.1360	0.1903
N	3006	3006	3006	1832	1832	1832
(4)Endogeneity adjusted QR estimates[b]						
Education	0.0346	0.0469***	0.0784**	0.0332**	0.0832***	0.1660***
	(0.0268)	(0.0129)	(0.0290)	(0.0143)	(0.0197)	(0.0356)
Pseudo R^2	0.0801	0.0991	0.0835	0.0676	0.1374	0.1897
N	3006	3006	3006	1832	1832	1832
(5)Endogeneity and selection adjusted QR estimates[c]						
Education	0.0421*	0.0791***	0.1059***	0.0506***	0.0870***	0.1752***
	(0.0234)	(0.0123)	(0.0298)	(0.0191)	(0.0213)	(0.0384)
Pseudo R^2	0.0942	0.0959	0.0906	0.0687	0.1388	0.1940
N	3006	3006	3006	1832	1832	1832

Notes: [a] QR estimates with the second order power series correction terms; [b] QR estimates with the fifth order polynomials of the reduced form residuals; [c] QR estimates with the second order power series correction terms and fifth order polynomials of the reduced form residuals;

(i) Bootstrapped errors in parentheses; ***, ** and * denote significance at the 1%, 5% and 10% level; (ii) Full results are given in Appendix. (iv) A tests on excluding our potential instrument (average regional schooling) from the reduced form equations conducted at the mean, yield an F-statistics of 168.61*** and partial R^2 of 0.0533 for males, and F= 336.02 *** and partial R^2 of 0.1558 for females; (iv) The DWH test conducted at the mean in male specification is 158.921, p-value= 0.0000 and 83.235, p-value= 0.0000 in female, so we do reject the null hypothesis.

FIGURE 5.5: Returns to education by quantiles, Tajikistan 2003



Notes: The y-axis measures the return to education coefficients; the x-axis depicts the selected quantiles of the conditional wage distribution for male and females. The horizontal line plots the OLS estimate and its 95% confidence interval.

One can also see that the horizontal lines in the figures, which plots the OLS estimate and its 95% confidence interval, indicates that the estimated mean returns to schooling for Bulgaria, Russia and Tajikistan are not representative of the effect education has on earnings at all points in the earnings distribution. One possible explanation for the high premium estimated at the top of the distribution in Bulgaria and for females in Tajikistan is that the degree holders at the top of the distribution would be expected to be of higher ability, thereby biasing the estimated educational premium. Arias *et al.* (2001) have interpreted a positive ability-returns relationship as evidence that education and ability are complements in the human capital generation process, which if true suggests that more able individuals in these countries benefit most from educational investment. However, there might be other explanations for this pattern. Because personal abilities and skills (cognitive and non-cognitive) are unobserved by researchers, it is difficult to isolate the effect that drives the heterogeneous pattern of returns to education across the wage distribution. For example, workers with identical education do not necessarily have to have the same level of productivity because of the influence of unobserved variables that are systematically correlated with both measured education and an individual's place in the earnings distribution.

In the set of countries considered here, the Russian unadjusted estimates do not increase as we move up to the distribution. For both males and females, the return to schooling is higher in the lower part of the earnings distribution than at the top end of the distribution. For instance, the returns to education for males fall from 11.7% to 7.1% between the 10^{th} and 90^{th} quantile and for females the equivalent fall is from 12.1% to 8.6% (see Table 5.4). These differences are also significant, as an *F*-test decisively rejects the equality of the estimates at the 10^{th} and 90^{th} percentiles for both male and female workers (see Table 5.6). At the top end of the earnings distribution, the estimated returns to education for both males and females are below than for the average earnings. However, we should note that the endogeneity and selectivity adjusted point estimates in Russia are quite different and show considerable homogeneity across the distribution for women (see the next Sections 5.4.3 and 5.4.4).

Mwabu and Schultz (1996) and Arias *et al.* (2001) interpret the negative ability-returns relationship as evidence of education and ability being substitutes, which implies that maximising the returns to education may require increasing educational opportunities for less able individuals in Russia. Flabbi *et al.* (2008) also find evidence for a higher return to education in the lower part of the earnings distribution in Russian in the early (1991-1996) and late transition (1997-2002) periods. Similarly, Gorodnichenko and Sabirianova (2005) find that the university wage premium in Russia is higher in the lower part of the earnings distribution than in top part of the earnings distribution. There are, however, a number of alternative explanations for this pattern in the literature. First, a demand-side effect could drive down the return to education at different points in the earnings distribution because of an oversupply of well-educated workers in the economy (the supply effect dominates the demand effect at higher points in the earnings distribution). Second, a negative relationship between 'ability' and the return to schooling could also reflect differences in the educational attainment of the labour force (Herrnstein and Murray, 1994). Similarly, lower returns to education at the higher end of the earnings distribution suggests there are factors leading to high-paying employment that act independently of education-generating human capital process. It is also possible to interpret the results in terms of a "state" or "foreign" ownership effect. State ownership is much more relevant to the lower tail of the wage distribution and relatively low paid workers earn more in state owned firms. This state ownership effect tends to die away as there is movement up through the earnings distribution (Machado and Mata, 2001). Finally, as Table 5.1 shows the public sector is much bigger in Russia than in the other three countries and this might suggest that redistributive policies may play a greater role there than elsewhere in supporting those with low human capital endowments.

The QR results also provide evidence of heterogeneity in returns to education across countries as well as within countries. Whereas the return for females at the 90th percentile is 7.2% in Bulgaria, it reaches 11.6% in Tajikistan. For males, the estimated unadjusted returns at the 10th percentile vary from 3.9% in Bulgaria, to 8.2% in Serbia, 11.7% in Russia and 3.7% in Tajikistan.

Finally, the unadjusted results indicate that, in general, educational returns are higher for females than for males. For instance, the point estimates at the 90th percentile in Serbia show return to education of 6.7% for males and 10.1% for females. The differences are also significant in Tajikistan, where at the 90th percentile the estimated returns to education is 3.9% for males and 11.6% for females.

We use testing procedures based on the QR statistics to formally test for the presence of heterogeneity in the returns to education and to test that the difference between quantiles is statistically significant[84]. The test of whether the estimated returns to education differ across each of these quantile levels is significantly different for female workers in Bulgaria, Tajikistan and Russia (see Table 5.6). Whereas for females in Tajikistan, the impact of education at the 90th quantile is two times higher than the 5.7% return found at the lower end of the distribution, we find little heterogeneity in how the quantiles of wages vary with schooling in Serbia. The F-test for equality at the 10th and 90th percentiles is insignificant for male's specifications in Serbia and Tajikistan. The confidence intervals for both males and females estimates in Serbia tend to overlap at most points in the earnings distribution the OLS estimates (see Figure 5.3).

TABLE 5.6

Inter-quantile hypothesis testing of unadjusted QR estimates

Countries	**90th equal to 10th**		**All quantiles equal**	
	Males	*Females*	*Males*	*Females*
	Unadjusted QR estimates			
Bulgaria	F(1,1281) = 4.65**	F(1,1171) = 5.71***	F(4,1281) =1.76	F(4,1171) =3.58***
Serbia	F(1,1451) = 0.49	F(1,969) = 0.00	F(4,1451) = 0.40	F(4,969) = 0.38
Russia	F(1, 21857) =71.10***	F(1, 24302) =58.43***	F(4,21857) =27.53***	F(4, 24302) =21.71***
Tajikistan	F(1,2991) = 0.02	F(1,1817) = 8.10***	F(4,2991) = 0.28	F(4,1817) = 4.44***

Notes: The F-test for the equality of unadjusted QR coefficients at the 90th and 10th and the F-test for the equality of all quantiles in unadjusted QR. ***, ** and * denote significance at the 1%, 5% and 10% level;

5.4.2. Unadjusted results with interactions for education

We next adopt the same specifications for testing the presence of an interaction effect between schooling and an indicator of whether individuals have graduated prior 1991. During the transition toward a market oriented wage system, new wage regimes were more likely to be applied to newly graduated workers (those who obtain their education after 1991) and consequently, we should expect higher returns to education for them. The dummy variable that captures whether individuals received their education prior 1991 is included along with the interaction terms. Even if we place the start of the transition period in 1991, as done for instance by Flabbi *et al.,* (2008), these individuals have been exposed to some extent to new education curricula and systems. In most of the countries, however, coefficients on the interaction terms

[84] Following Buchinsky (1995), this test is performed after an inter-quantile regression, which re-estimates the model taking the difference between the coefficients across the wage distribution $\beta_{\theta 1} - \beta_{\theta 2} = 0$.

between schooling and pre-1991 are insignificant and only for males in Tajikistan at the upper part of the distribution, we find significant and negative interaction term, which basically indicates lower returns for workers graduated before 1991, than for the newly graduated respondents. For females in Russia, the interaction terms are positive and significant at 5% lever at the bottom of the distribution.

5.4.3. Sample selection adjusted QR results

There are at least two indications why sample selection may be a concern for our samples. First, about 52% of females in the Russian sample, 60% of females in Bulgarian sample and only 38% of females in Serbian and 28% of females in Tajikistan samples are in employment. Second, from the descriptive statistics, it appears that working and the non-working women differ in their educational attainment. For instance, in Bulgaria working women report on average 13.9 years in education compared to the 12.6 years for their non-working counterparts.

However, sample selection adjusted QR estimates indicate that the two power series correction terms are not significant for either males or females workers in Bulgaria and for males in Serbia. Moreover, comparing returns to education coefficients to the corresponding unadjusted quantile estimates in these two countries, the differences are not statistically significant. Selectivity terms in the series estimator are significant at the conventional levels of significance in the female specification in Serbia at the 25th, 50th and 75th percentiles. By way of contrast, in Russia there is compelling evidence of strong sample selection effect across the distribution. The coefficient at the 10th percentile for males increases from 11.7% in the unadjusted estimates to 14.6% once selectivity is taken into account and from 12.1% to 19.4% for females. We should note that the sample selection estimates may depend on the first stage selection equation obtained for a parametric probit model[85]. Correcting for selection has a dramatic effect on the returns to schooling for males in Tajikistan, increasing the estimates from 3.7% to 7.4% at the 10th percentile. If true this finding would suggest that for males in Tajikistan, education is important for determining participation in the labour force but thereafter has little effect on the earnings of individuals.

[85] Buchinsky (2001) shows that semi-parametric estimates for the first stage selection equation are considerably different than those obtained from a parametric model. Since this is not focus of the current study and considering that semi-parametric method is highly computer intensive (for example, the estimation using Klein and Spady (1993) estimator on a first step takes between 15 and 17 hours), we apply the probit model on a first-stage.

Overall, the selectivity adjusted point estimates across the various quantiles in Bulgaria and Serbia are quite close, suggesting little heterogeneity, and this is support by the inter-quantile tests. In the case of Serbia, we do not reject that returns to education, based on selectivity adjusted estimates, are constant over the earnings distribution. The test of whether the selectivity corrected returns to education differ across each of these quantile levels is significantly different only for female workers in Bulgaria and Tajikistan and for males in Russia (see Table 5.7).

TABLE 5.7

Inter-quantile hypothesis testing of selectivity adjusted QR estimates

Countries	90th equal to 10th		All quantiles equal	
	Males	*Females*	*Males*	*Females*
	Selectivity adjusted QR estimates			
Bulgaria	$F(1,1279) = 4.82^{**}$	$F(1,1169) = 2.53$	$F(4,1279) = 2.10^{**}$	$F(4,1169) = 1.55$
Serbia	$F(1,1449) = 1.03$	$F(1, 967) = 0.16$	$F(4,1449) = 0.43$	$F(4,967) = 0.32$
Russia	$F(1, 21856) = 70.73^{***}$	$F(1,24300) = 31.46^{***}$	$F(4,21856) = 21.94^{***}$	$F(4, 24300) = 1.24$
Tajikistan	$F(1,2989) = 0.71$	$F(1,1815) = 8.08^{***}$	$F(4,2989) = 2.18$	$F(4,1815) = 3.75^{***}$

Notes: The F-test for the equality of unadjusted QR coefficients at the 90th and 10th and the F-test for the equality of all quantiles in unadjusted QR. ***, ** and * denote significance at the 1%, 5% and 10% level.

5.4.4. Endogeneity adjusted QR results

The fourth set of results pertains to endogeneity adjusted estimates. A fifth order polynomial of the reduced form residuals is applied to estimate the rate of returns to education at different values of θ[86]. The results are sensitive to the choice of the order of residual polynomial used in the analysis. Before comparing the results of these estimates, we test whether our instrument meets the two conditions: relevance and exogeneity. The relevance of the instrument is tested at the mean in the first-stage regression. There is significant and positive relationship between regional average schooling and individual's level of schooling. The *F*-statistics on the excluded instrument in the reduced-form schooling estimates indicate that schooling is strongly correlated with the instrument for both males and females and in all four countries. For instance, the test on excluding our potential instrument from the reduced form equation in Bulgaria, yields an *F*-statistics of 19.39 and partial R^2 of 0.0149 for males and *F*= 17.54 and partial R^2 of 0.0147 for females. Similarly, for Serbia the *F*-statistics of 37.56 for males and 14.71 for females compare favourably with those

[86] Lee (2007) provides a condition which restricts the growth rate of the power series k. The necessary smoothness condition is that $r \geq 5$.

reported in Bound *et al.* (1995).

The validity of our endogeneity adjusted estimates however, depends on the exclusion restriction that average regional schooling has no direct effect on individual's current earnings. Having more than one potential instrument will enable us to undertake a Sargan instrument validity test and to provide support for our empirical approach. Unfortunately in most of the countries dealt with in this Chapter we only have one identifying instrument and are, therefore, unable to undertake a similar test. In the case of Bulgaria only we are able to identify mothers within the household having higher degree. The over-identifying test was passed for both male and female Bulgarian samples, which provides some support for the approach adopted here.

For all four countries, the DWH test undertaken at the mean leads to a strong rejection of the null hypothesis of exogeneity of schooling, and we conclude that the endogeneity adjusted QR results differ from the unadjusted QR estimates. An examination of the results in Tables 5.2 to 5.5 suggests that the effects of adjusting for the endogeneity of schooling is most marked at the top end of the earnings distribution. For instance, in Bulgaria, the estimated returns for both males and females rise to nearly 11%. We observe a similar trend for females in Tajikistan, where correcting for endogeneity increases estimated returns to education at the 90th percentile to 16.6%. For males in Tajikistan, we do find insignificant wage premium at the bottom of the distribution. In Russia, the returns to schooling are quite different compared to the unadjusted results. Again, this effect tends to be more pronounced in the top end of the distribution than in the bottom end of the distribution. We can see at the highest percentile (90th) the return to schooling in Russia is 12.7% for females and 12.8% for males, while the equivalent comparison at the 10th percentile is a return of 8.6% for females and 11.7% for males. Moreover, for males the return decreases sharply as we go up the distribution, though with an increase at the 90th centile (see Table A5.12). The endogeneity adjusted estimates for females in Russia show a rising return (see Table A5.13). For males and females in Serbia, returns to education tend to be insignificant at the bottom of the distribution, once the endogeneity is accounted for.

5.4.5. Sample selection and endogeneity adjusted QR results

In many cases it may be important to correct for both sample selection and endogeneity in education. We do control for both biases in the case of Russia and Tajikistan, as sample selectivity were found to be significant in those countries. After controlling for sample selection due to labour force participation as well as endogeneity of schooling decision, the estimated coefficients are overall larger than the unadjusted ones, thus allowing us to conclude that the differences between estimates is attributed to selection and endogeneity biases. For instance, the estimated returns to education at the 90^{th} percentile for females in Russia rises from 8.6% in unadjusted QR estimates to 18.3% once selectivity and endogeneity are considered. Similarly, in Tajikistan the effect is more pronounced in the top end of the distribution, where the estimated coefficients rise from 3.9% to 10.6% for males and from 11.6% to 17.5% for females.

5.5. Conclusions

In this Chapter we analyse the dispersion of the returns to education at four different transition countries, attempting to test for evidence of individual heterogeneity in returns to education. Based on the empirical estimations, the following main conclusions can be drawn:

The unadjusted estimates suggest that the return to education varies as we move across the earnings distribution. There is a tendency for returns to increase and to remain higher in the upper tail of the distribution in Bulgaria and for females in Tajikistan. We interpret these increasing returns as an indication that 'ability', when broadly defined, and education complement each other. The reason for heterogeneity in the returns to education is likely to be due to the fact that differences in 'ability' translate into higher pay differentials between high-ability and low-ability workers. A little heterogeneity was found in the case of Serbia. While in three of the countries the rate of return to education either rises or stays constant as we move up the distribution, the unadjusted estimates for Russia show that it falls sharply.

Overall, specifications with the interaction terms between schooling and indicators for education obtained prior 1991 indicate insignificant coefficients on the interaction terms. Significant and negative interaction term are found in Tajikistan for males at the upper part of distribution and for females at the 90^{th} percentile which suggest significantly lower return to education for those graduated before 1991.

Comparison of the unadjusted QR estimates and estimates corrected for selectivity suggests that estimates of the returns to education for both males and females in Russia and for males in Tajikistan were sensitive to the inclusion of the power series approximation terms. The significance of the power series approximation terms in these two countries indicates that it was necessary to perform the sample selection correction. In Russia there was compelling evidence of strong sample selection effect at the bottom of the distribution where the estimated coefficients increase to 12.8% for males and 12.7% for females. The two power-series correction terms were not significant for either males or females in Bulgaria.

Overall, the endogeneity adjusted estimates were higher as compared to unadjusted returns to education[87] and we conclude that failure to account for endogeneity seems to create slight downward bias estimates of the returns to education. For all four countries the DWH test, undertaken at the mean, leads to a strong rejection of the null hypothesis of exogeneity in education, therefore the endogeneity adjusted results differ from the unadjusted QR estimates.

The resulst for Russia are sensitive to specification. The negative wage premium, found in unadjusted results was no more evident when both endogeneity and selectivity issues are taken into account.

[87] It is common results in the empirical literature that estimated return to education increase as a result of treating education as an endogenous variable.

CHAPTER FIVE APPENDIX

TABLE A5.1: Unadjusted QR estimates of returns to education, Bulgaria 2003

	MALES					FEMALES				
	θ=0.10	θ=0.25	θ=0.50	θ=0.75	θ=0.90	θ=0.10	θ=0.25	θ=0.50	θ=0.75	θ=0.90
school	0.0395***	0.0445***	0.0537***	0.0615***	0.0626***	0.0472***	0.0503***	0.0612***	0.0729***	0.0722***
	(0.0050)	(0.0037)	(0.0044)	(0.0073)	(0.0095)	(0.0062)	(0.0046)	(0.0057)	(0.0061)	(0.0087)
exp	0.0103**	0.0014	0.0052	0.0078	-0.0022	0.0213***	0.0203***	0.0225***	0.0107**	0.0022
	(0.0049)	(0.0035)	(0.0040)	(0.0059)	(0.0077)	(0.0058)	(0.0046)	(0.0054)	(0.0052)	(0.0070)
expsq	-0.0003***	-0.0001	-0.0001	-0.0002	-0.0001	-0.0005***	-0.0004***	-0.0004***	-0.0001	0.0000
	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0002)	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0002)
married	-0.0195	0.0159	0.0441	0.0360	0.0880	-0.0130	0.0092	-0.0488	-0.0421	-0.0408
	(0.0391)	(0.0287)	(0.0320)	(0.0478)	(0.0626)	(0.0435)	(0.0338)	(0.0399)	(0.0397)	(0.0527)
1-2 years	0.1029**	0.1369***	0.1081***	0.0442	-0.0149	0.0491	0.0183	0.0463	0.0765	0.0814
	(0.0479)	(0.0358)	(0.0409)	(0.0628)	(0.0792)	(0.0603)	(0.0452)	(0.0551)	(0.0558)	(0.0776)
3-5 years	0.0623	0.1704***	0.1333***	0.1607***	0.0908	0.1639***	0.1401***	0.1171**	0.1729***	0.2845***
	(0.0487)	(0.0360)	(0.0409)	(0.0623)	(0.0787)	(0.0579)	(0.0441)	(0.0534)	(0.0532)	(0.0741)
6-10 years	0.1813***	0.3013***	0.2723***	0.2918***	0.1433	0.0301	0.1629***	0.1885***	0.2715***	0.2831***
	(0.0547)	(0.0413)	(0.0480)	(0.0737)	(0.0915)	(0.0661)	(0.0505)	(0.0612)	(0.0611)	(0.0834)
>10 years	0.2277***	0.3412***	0.3259***	0.3214***	0.3534***	0.2065***	0.2288***	0.2193***	0.3342***	0.3609***
	(0.0519)	(0.0383)	(0.0431)	(0.0655)	(0.0858)	(0.0626)	(0.0480)	(0.0590)	(0.0592)	(0.0821)
public	0.2207***	0.1949***	0.1963***	0.1952***	0.1805***	0.2159***	0.1190***	0.0424	-0.0860**	-0.0601
	(0.0399)	(0.0288)	(0.0325)	(0.0487)	(0.0635)	(0.0450)	(0.0331)	(0.0407)	(0.0413)	(0.0572)
North West	-0.0803	-0.0841*	-0.1427***	-0.1416*	-0.1426	-0.0813	-0.0924*	-0.1667***	-0.1762***	-0.1662**
	(0.0595)	(0.0445)	(0.0516)	(0.0793)	(0.0983)	(0.0669)	(0.0499)	(0.0608)	(0.0612)	(0.0839)
North East	0.0329	-0.0131	-0.0827*	-0.0610	-0.0913	0.0253	-0.0180	-0.0636	-0.0303	-0.0423
	(0.0533)	(0.0397)	(0.0459)	(0.0707)	(0.0889)	(0.0629)	(0.0489)	(0.0588)	(0.0592)	(0.0814)
Central	-0.1733***	-0.1325***	-0.1599***	-0.1336*	-0.0990	-0.0529	-0.0766	-0.1420**	-0.2217***	-0.1311
	(0.0558)	(0.0424)	(0.0486)	(0.0741)	(0.0925)	(0.0661)	(0.0504)	(0.0605)	(0.0615)	(0.0837)
South East	-0.0099	-0.0095	0.0094	0.0618	0.0025	-0.1201**	-0.0843*	-0.1289**	-0.1345**	-0.0255
	(0.0483)	(0.0366)	(0.0418)	(0.0640)	(0.0794)	(0.0580)	(0.0440)	(0.0530)	(0.0538)	(0.0742)
Sofia region	0.1154**	0.1867***	0.1648***	0.1514**	0.1555**	0.0443	0.0548	0.0612	0.0511	0.1060
	(0.0497)	(0.0368)	(0.0417)	(0.0624)	(0.0765)	(0.0556)	(0.0425)	(0.0508)	(0.0508)	(0.0678)
constant	0.3325***	0.4852***	0.6302***	0.8196***	1.2174***	0.0231	0.1943**	0.3552***	0.5169***	0.7746***
	(0.0873)	(0.0646)	(0.0762)	(0.1276)	(0.1681)	(0.1054)	(0.0788)	(0.1017)	(0.1095)	(0.1602)
N	1296	1296	1296	1296	1296	1186	1186	1186	1186	1186
Pseudo R^2	0.1400	0.1859	0.2071	0.2150	0.1978	0.1567	0.1743	0.2037	0.2059	0.2086

Notes: Standard errors in parentheses * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

TABLE A5.2: Unadjusted QR estimates of returns to education with interaction terms, Bulgaria 2003

	Males					*Females*				
	θ=0.10	θ=0.25	θ=0.50	θ=0.75	θ=0.90	θ=0.10	θ=0.25	θ=0.50	θ=0.75	θ=0.90
school	0.0516***	0.0477***	0.0505***	0.0528***	0.0556***	0.0378***	0.0418***	0.0577***	0.0726***	0.0686***
	(0.0090)	(0.0063)	(0.0079)	(0.0125)	(0.0164)	(0.0110)	(0.0098)	(0.0100)	(0.0105)	(0.0148)
educ*bef91	-0.0153	-0.0058	0.0054	0.0142	0.0148	0.0112	0.0138	0.0061	0.0008	0.0039
	(0.0105)	(0.0074)	(0.0092)	(0.0144)	(0.0185)	(0.0122)	(0.0111)	(0.0118)	(0.0126)	(0.0175)
bef1991	0.1553	0.0362	-0.0685	-0.3557*	-0.2446	-0.0125	-0.1896	-0.1181	-0.0602	0.0071
	(0.1578)	(0.1058)	(0.1307)	(0.2044)	(0.2477)	(0.1897)	(0.1719)	(0.1807)	(0.1913)	(0.2652)
exp	0.0113	0.0045	0.0044	0.0187**	0.0010	0.0074	0.0189**	0.0223***	0.0158**	-0.0047
	(0.0081)	(0.0056)	(0.0062)	(0.0088)	(0.0123)	(0.0089)	(0.0080)	(0.0086)	(0.0078)	(0.0105)
expsq	-0.0003*	-0.0001	-0.0001	-0.0003**	-0.0001	-0.0003	-0.0004**	-0.0004**	-0.0002	0.0002
	(0.0001)	(0.0001)	(0.0001)	(0.0002)	(0.0002)	(0.0002)	(0.0002)	(0.0002)	(0.0002)	(0.0002)
married	-0.0172	0.0118	0.0496	0.0547	0.0839	0.0184	0.0075	-0.0526	-0.0422	-0.0343
	(0.0439)	(0.0302)	(0.0328)	(0.0457)	(0.0616)	(0.0430)	(0.0406)	(0.0425)	(0.0401)	(0.0533)
1-2 years	0.1191**	0.1396***	0.1072**	0.0608	-0.0450	0.0532	0.0151	0.0403	0.0783	0.0531
	(0.0532)	(0.0368)	(0.0417)	(0.0595)	(0.0764)	(0.0584)	(0.0547)	(0.0584)	(0.0564)	(0.0811)
3-5 years	0.0793	0.1665***	0.1398***	0.1328**	0.0706	0.2079***	0.1443***	0.1230**	0.1699***	0.2608***
	(0.0546)	(0.0374)	(0.0417)	(0.0595)	(0.0778)	(0.0568)	(0.0542)	(0.0567)	(0.0544)	(0.0761)
6-10 years	0.1656***	0.2978***	0.2691***	0.2725***	0.1378	0.0642	0.1790***	0.1925***	0.2718***	0.2841***
	(0.0608)	(0.0431)	(0.0488)	(0.0696)	(0.0943)	(0.0670)	(0.0610)	(0.0648)	(0.0627)	(0.0862)
>10 years	0.2457***	0.3321***	0.3279***	0.3458***	0.3325***	0.2313***	0.2214***	0.2101***	0.3293***	0.3545***
	(0.0584)	(0.0397)	(0.0439)	(0.0618)	(0.0832)	(0.0612)	(0.0585)	(0.0622)	(0.0596)	(0.0837)
public	0.2250***	0.2077***	0.1954***	0.2049***	0.1618***	0.1851***	0.1221***	0.0545	-0.0788*	-0.0561
	(0.0442)	(0.0300)	(0.0331)	(0.0462)	(0.0619)	(0.0443)	(0.0398)	(0.0432)	(0.0417)	(0.0583)
North West	-0.0652	-0.0791*	-0.1424***	-0.1089	-0.1075	-0.1068	-0.0835	-0.1645**	-0.1810***	-0.1744**
	(0.0664)	(0.0461)	(0.0526)	(0.0750)	(0.0963)	(0.0654)	(0.0609)	(0.0643)	(0.0620)	(0.0851)
North East	0.0418	-0.0042	-0.0854*	-0.0674	-0.1034	0.0307	0.0046	-0.0660	-0.0313	-0.0705
	(0.0587)	(0.0412)	(0.0468)	(0.0664)	(0.0875)	(0.0625)	(0.0590)	(0.0622)	(0.0595)	(0.0828)
Central	-0.1723***	-0.1268***	-0.1566***	-0.1300*	-0.1129	-0.0633	-0.0932	-0.1405**	-0.2029***	-0.1484*
	(0.0621)	(0.0437)	(0.0495)	(0.0705)	(0.0907)	(0.0648)	(0.0611)	(0.0637)	(0.0619)	(0.0849)
South East	-0.0065	-0.0051	0.0136	0.0499	0.0302	-0.0940	-0.0839	-0.1349**	-0.1342**	-0.0262
	(0.0541)	(0.0379)	(0.0426)	(0.0605)	(0.0785)	(0.0574)	(0.0531)	(0.0561)	(0.0545)	(0.0756)
Sofia region	0.0931*	0.1794***	0.1741***	0.1698***	0.1472*	0.0493	0.0496	0.0556	0.0562	0.1034
	(0.0547)	(0.0380)	(0.0424)	(0.0593)	(0.0757)	(0.0555)	(0.0512)	(0.0536)	(0.0515)	(0.0693)
constant	0.1754	0.4224***	0.6696***	0.9034***	1.3081***	0.1737	0.3165**	0.4240***	0.4927***	0.8627***
	(0.1354)	(0.0998)	(0.1236)	(0.1933)	(0.2575)	(0.1804)	(0.1555)	(0.1629)	(0.1717)	(0.2385)
N	1296	1296	1296	1296	1296	1186	1186	1186	1186	1186
Pseudo R^2	0.1420	0.1865	0.2073	0.2190	0.1996	0.1588	0.1755	0.2043	0.2062	0.2094

Notes: Standard errors in parentheses * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

TABLE A5.3: Selectivity adjusted QR estimates of returns to education, Bulgaria 2003

	MALES					FEMALES				
	θ=0.10	θ=0.25	θ=0.50	θ=0.75	θ=0.90	θ=0.10	θ=0.25	θ=0.50	θ=0.75	θ=0.90
school	0.0336***	0.0348***	0.0502***	0.0519***	0.0627***	0.0481***	0.0383***	0.0574***	0.0660***	0.0879***
	(0.0109)	(0.0070)	(0.0081)	(0.0109)	(0.0115)	(0.0171)	(0.0118)	(0.0130)	(0.0116)	(0.0145)
exp	0.0112*	0.0020	0.0050	0.0088	-0.0014	0.0200	0.0251***	0.0243**	0.0167**	-0.0094
	(0.0060)	(0.0037)	(0.0043)	(0.0055)	(0.0063)	(0.0127)	(0.0088)	(0.0097)	(0.0080)	(0.0100)
expsq	-0.0003**	-0.0001	-0.0001	-0.0003**	-0.0001	-0.0005	-0.0006**	-0.0005*	-0.0003	0.0004
	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0004)	(0.0002)	(0.0003)	(0.0002)	(0.0003)
married	-0.0764	-0.0832	0.0176	-0.0695	0.0901	-0.0136	0.0453	-0.0354	-0.0332	-0.0706
	(0.1025)	(0.0640)	(0.0709)	(0.0914)	(0.0948)	(0.0612)	(0.0448)	(0.0491)	(0.0422)	(0.0544)
1-2 years	0.1400*	0.1891***	0.1283**	0.0958	0.0379	0.0258	0.1639	0.1197	0.1849	-0.1106
	(0.0758)	(0.0473)	(0.0554)	(0.0721)	(0.0726)	(0.2330)	(0.1615)	(0.1750)	(0.1517)	(0.1871)
3-5 years	0.0458	0.1223***	0.1285**	0.0746	0.1217	0.1401	0.2933*	0.1926	0.2738*	0.1364
	(0.0762)	(0.0462)	(0.0529)	(0.0713)	(0.0767)	(0.2261)	(0.1580)	(0.1714)	(0.1478)	(0.1873)
6-10 years	0.1709**	0.2867***	0.2636***	0.2884***	0.1671**	0.0046	0.3563*	0.2793	0.3888**	0.0667
	(0.0666)	(0.0438)	(0.0517)	(0.0688)	(0.0763)	(0.2659)	(0.1844)	(0.2006)	(0.1715)	(0.2152)
>10 years	0.3712*	0.5353***	0.4153***	0.5528***	0.4096**	0.1840	0.4240**	0.3032	0.4565**	0.1156
	(0.2062)	(0.1322)	(0.1498)	(0.1973)	(0.1949)	(0.2855)	(0.1985)	(0.2142)	(0.1835)	(0.2302)
public	0.2195***	0.1920***	0.1885***	0.1575***	0.1372***	0.2109***	0.0848*	0.0398	-0.1040**	-0.0280
	(0.0503)	(0.0314)	(0.0358)	(0.0467)	(0.0502)	(0.0612)	(0.0470)	(0.0540)	(0.0459)	(0.0642)
North West	-0.0831	-0.0852*	-0.1446***	-0.1387*	-0.1246	-0.0787	-0.0697	-0.1428**	-0.1685***	-0.2326***
	(0.0732)	(0.0473)	(0.0559)	(0.0745)	(0.0787)	(0.0739)	(0.0591)	(0.0664)	(0.0571)	(0.0763)
North East	0.0071	-0.0851	-0.0990	-0.1476	-0.0645	0.0361	-0.0668	-0.0689	-0.0516	0.0089
	(0.0889)	(0.0568)	(0.0668)	(0.0904)	(0.0959)	(0.0816)	(0.0639)	(0.0726)	(0.0647)	(0.0861)
Central	-0.2006**	-0.2178***	-0.1953***	-0.2091**	-0.1079	-0.0542	-0.1198*	-0.1433*	-0.2374***	-0.0960
	(0.1020)	(0.0615)	(0.0706)	(0.0931)	(0.1062)	(0.0866)	(0.0668)	(0.0738)	(0.0641)	(0.0851)
South East	-0.0644	-0.1149	-0.0289	-0.0560	0.0250	-0.1188*	-0.1125**	-0.1287**	-0.1488***	0.0071
	(0.1155)	(0.0708)	(0.0823)	(0.1105)	(0.1163)	(0.0654)	(0.0534)	(0.0605)	(0.0532)	(0.0720)
Sofia region	0.1142	0.1378***	0.1535***	0.0974	0.1372*	0.0513	-0.0006	0.0502	0.0195	0.1726**
	(0.0745)	(0.0466)	(0.0533)	(0.0711)	(0.0757)	(0.0870)	(0.0660)	(0.0743)	(0.0656)	(0.0850)
$(z'\gamma)$	0.3947	0.2890	0.7303	-0.3266	-1.2042	1.1431	-2.8415	-2.6008	-3.1774	3.5430
	(1.2016)	(0.7817)	(0.9501)	(1.2953)	(1.2700)	(0.3313)	(0.5061)	(0.8762)	(0.2075)	(0.2730)
$(z'\gamma)^2$	-0.0286	0.0391	-0.1467	0.2151	0.3310	-5.3932	4.0489	9.6140	10.4755	-6.1515
	(0.2801)	(0.1790)	(0.2200)	(0.2997)	(0.2776)	(0.3356)	(0.0414)	(0.3659)	(0.9743)	(0.7455)
constant	-0.2244	-0.0180	-0.1564	0.8644	2.2517*	-0.0170	0.3978*	0.4559*	0.6617***	0.5430*
	(1.2570)	(0.8234)	(0.9932)	(1.3554)	(1.3579)	(0.3365)	(0.2391)	(0.2622)	(0.2349)	(0.3031)
N	1296	1296	1296	1296	1296	1186	1186	1186	1186	1186
Pseudo R^2	0.1408	0.1868	0.2073	0.2167	0.1997	0.1572	0.1750	0.2040	0.2062	0.2093

Notes: Standard errors in parentheses * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

TABLE A5.4: Endogeneity adjusted QR estimates of returns to education, Bulgaria 2003

	MALES					FEMALES				
	θ=0.10	θ=0.25	θ=0.50	θ=0.75	θ=0.90	θ=0.10	θ=0.25	θ=0.50	θ=0.75	θ=0.90
school	0.0465**	0.0348***	0.0511***	0.0932***	0.1070***	0.0451**	0.0219**	0.0499***	0.1047***	0.1071***
	(0.0184)	(0.0071)	(0.0131)	(0.0153)	(0.0313)	(0.0216)	(0.0102)	(0.0095)	(0.0136)	(0.0202)
exp	0.0121**	-0.0007	0.0034	0.0107*	0.0096	0.0217***	0.0183***	0.0190***	0.0141***	-0.0021
	(0.0061)	(0.0033)	(0.0054)	(0.0063)	(0.0072)	(0.0067)	(0.0046)	(0.0044)	(0.0053)	(0.0052)
expsq	-0.0003***	-0.0001	-0.0001	-0.0002	-0.0002	-0.0005***	-0.0005***	-0.0004***	-0.0001	0.0001
	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0002)	(0.0001)	(0.0001)	(0.0001)	(0.0001)
married	-0.0291	0.0153	0.0525	-0.0170	0.0419	-0.0088	-0.0085	-0.0268	-0.0381	0.0143
	(0.0406)	(0.0256)	(0.0449)	(0.0547)	(0.0658)	(0.0505)	(0.0332)	(0.0323)	(0.0382)	(0.0417)
1-2 years	0.0973*	0.1232***	0.1550***	0.0563	0.0199	0.0397	0.0341	0.0358	0.0612	0.0594
	(0.0548)	(0.0313)	(0.0547)	(0.0653)	(0.0824)	(0.0806)	(0.0444)	(0.0435)	(0.0542)	(0.0607)
3-5 years	0.0501	0.1886***	0.1792***	0.1446**	0.0577	0.1682*	0.1960***	0.1335***	0.1306**	0.2488***
	(0.0655)	(0.0312)	(0.0547)	(0.0653)	(0.0827)	(0.0892)	(0.0449)	(0.0428)	(0.0542)	(0.0574)
6-10 years	0.1520**	0.3126***	0.3250***	0.2513***	0.1187	0.0089	0.2369***	0.2080***	0.2117***	0.3221***
	(0.0743)	(0.0369)	(0.0650)	(0.0775)	(0.0978)	(0.1128)	(0.0515)	(0.0496)	(0.0623)	(0.0649)
>10 years	0.2098**	0.3530***	0.3759***	0.3090***	0.2435***	0.1943*	0.3190***	0.2163***	0.2701***	0.3534***
	(0.0923)	(0.0372)	(0.0608)	(0.0695)	(0.0898)	(0.1054)	(0.0507)	(0.0482)	(0.0618)	(0.0631)
public	0.2188***	0.1992***	0.1833***	0.1418**	0.1666**	0.2083***	0.1464***	0.0748*	-0.1288***	-0.0652
	(0.0406)	(0.0251)	(0.0443)	(0.0607)	(0.0781)	(0.0513)	(0.0365)	(0.0386)	(0.0429)	(0.0458)
North West	-0.0710	-0.0640	-0.1241*	-0.1570*	-0.1126	-0.0737	-0.1051**	-0.1905***	-0.1863***	-0.1452**
	(0.0602)	(0.0391)	(0.0693)	(0.0824)	(0.1057)	(0.0757)	(0.0494)	(0.0479)	(0.0618)	(0.0651)
North East	0.0398	0.0230	-0.0547	-0.0892	-0.0613	0.0178	-0.0508	-0.0817*	-0.0903	-0.1497**
	(0.0550)	(0.0343)	(0.0607)	(0.0742)	(0.0935)	(0.0711)	(0.0480)	(0.0464)	(0.0573)	(0.0633)
Central	-0.1578***	-0.0951***	-0.1671***	-0.1567**	-0.1284	-0.0580	-0.0645	-0.1660***	-0.2161***	-0.1645**
	(0.0569)	(0.0367)	(0.0647)	(0.0773)	(0.0953)	(0.0793)	(0.0493)	(0.0481)	(0.0593)	(0.0651)
South East	0.0025	0.0091	0.0258	0.0470	0.0919	-0.1270*	-0.0914**	-0.1424***	-0.1508***	-0.0847
	(0.0487)	(0.0313)	(0.0555)	(0.0667)	(0.0841)	(0.0664)	(0.0431)	(0.0416)	(0.0521)	(0.0586)
Sofia region	0.1250**	0.1826***	0.1880***	0.1202*	0.1309	0.0044	0.0875**	0.0537	0.0371	0.0762
	(0.0528)	(0.0317)	(0.0560)	(0.0680)	(0.0824)	(0.0626)	(0.0420)	(0.0415)	(0.0505)	(0.0540)
r^1	-0.0005	0.0032	0.0068**	0.0106*	-0.0160*	-0.0052	0.0004	0.0061**	0.0110**	-0.0072
	(0.0038)	(0.0020)	(0.0031)	(0.0056)	(0.0096)	(0.0059)	(0.0027)	(0.0027)	(0.0044)	(0.0062)
r^2	-0.0005	0.0004	0.0007	-0.0005	-0.0040*	0.0005	0.0018***	0.0010*	-0.0002	-0.0023
	(0.0007)	(0.0003)	(0.0007)	(0.0006)	(0.0024)	(0.0003)	(0.0005)	(0.0006)	(0.0006)	(0.0016)
r^3	0.0001	-0.0000	-0.0001**	-0.0003***	-0.0002	0.0001	-0.0000	-0.0001	-0.0002*	-0.0001
	(0.0001)	(0.0000)	(0.0001)	(0.0001)	(0.0003)	(0.0001)	(0.0000)	(0.0001)	(0.0001)	(0.0002)
r^4	-0.0000	-0.0000*	-0.0002*	-0.0001**	-0.0000	-0.0000*	-0.0000***	-0.0000	-0.0000*	-0.0000
	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0010)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
r^5	0.0000	0.0000**	-0.0000*	-0.0000**	0.0001	0.0001*	0.0000*	0.0002	0.0001*	0.0001
	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
N	1296	1296	1296	1296	1296	1186	1186	1186	1186	1186
Pseudo R^2	0.1608	0.1967	0.2203	0.2324	0.2173	0.1581	0.1875	0.2162	0.2065	0.2140

TABLE A5.5: Unadjusted QR estimates of returns to education, Serbia 2003

	MALES					FEMALES				
	θ=0.10	θ=0.25	θ=0.50	θ=0.75	θ=0.90	θ=0.10	θ=0.25	θ=0.50	θ=0.75	θ=0.90
school	0.0821***	0.0770***	0.0672***	0.0679***	0.0674***	0.1025***	0.0928***	0.1039***	0.0976***	0.1012***
	(0.0133)	(0.0086)	(0.0064)	(0.0110)	(0.0213)	(0.0157)	(0.0090)	(0.0082)	(0.0115)	(0.0202)
exp	-0.0063	0.0109	0.0062	0.0147*	0.0140	0.0291**	0.0224***	0.0145**	0.0184**	0.0091
	(0.0120)	(0.0073)	(0.0052)	(0.0087)	(0.0168)	(0.0122)	(0.0066)	(0.0057)	(0.0073)	(0.0112)
expsq	0.0000	-0.0002*	-0.0001	-0.0002	-0.0001	-0.0004**	-0.0003***	-0.0001	-0.0001	0.0001
	(0.0002)	(0.0001)	(0.0001)	(0.0001)	(0.0003)	(0.0002)	(0.0001)	(0.0001)	(0.0001)	(0.0002)
married	0.0838	0.0962	0.0382	-0.0027	0.0180	0.0189	0.0003	-0.0040	-0.0502	-0.1193
	(0.0968)	(0.0588)	(0.0414)	(0.0665)	(0.1170)	(0.0918)	(0.0517)	(0.0432)	(0.0546)	(0.0876)
1-2 years	0.0791	0.0605	-0.0225	-0.1404	-0.0432	0.0791	0.0052	-0.0104	0.0493	-0.1996
	(0.2168)	(0.1317)	(0.0933)	(0.1484)	(0.2667)	(0.2498)	(0.1291)	(0.1090)	(0.1325)	(0.2306)
3-5 years	0.1857	0.1149	-0.0227	-0.0752	-0.2101	0.1967	0.1263	0.1165	0.0718	-0.0186
	(0.1412)	(0.0889)	(0.0627)	(0.0994)	(0.1825)	(0.1843)	(0.0951)	(0.0787)	(0.0967)	(0.1653)
6-10 years	0.1532	0.0108	0.0155	-0.1088	-0.2130	0.0687	0.0867	0.1146*	0.1891**	0.0787
	(0.1382)	(0.0857)	(0.0600)	(0.0965)	(0.1689)	(0.1593)	(0.0827)	(0.0676)	(0.0839)	(0.1356)
>10 years	0.1309	0.0072	-0.0118	-0.1238	-0.1542	-0.1726	-0.0138	-0.0072	0.0021	-0.1087
	(0.1117)	(0.0676)	(0.0482)	(0.0777)	(0.1477)	(0.1294)	(0.0683)	(0.0565)	(0.0709)	(0.1097)
public	-0.0297	0.0032	-0.0538*	-0.1783***	-0.2971***	0.1593*	0.2389***	0.0574	-0.0422	-0.0851
	(0.0716)	(0.0454)	(0.0318)	(0.0508)	(0.0919)	(0.0924)	(0.0484)	(0.0410)	(0.0516)	(0.0835)
Belgrade	0.4331***	0.4376***	0.2952***	0.3645***	0.3560*	0.4641***	0.2609***	0.2201***	0.1615*	0.3348**
	(0.1419)	(0.0871)	(0.0611)	(0.0990)	(0.1818)	(0.1681)	(0.0874)	(0.0717)	(0.0879)	(0.1484)
Vojvodina	0.0944	0.1529**	0.0287	0.1056	0.0923	0.3304**	0.1302	0.0784	0.1095	0.2570*
	(0.1240)	(0.0778)	(0.0545)	(0.0879)	(0.1593)	(0.1576)	(0.0821)	(0.0671)	(0.0821)	(0.1389)
West Serbia	-0.0095	0.0665	-0.0235	0.0818	-0.0425	0.1289	-0.0685	-0.1169	-0.1658*	-0.1238
	(0.1427)	(0.0904)	(0.0633)	(0.1019)	(0.1814)	(0.1931)	(0.1001)	(0.0825)	(0.1004)	(0.1726)
Šumadija & Pomoravlje	0.0953	0.1260	0.1456**	0.2556***	0.2097	0.2683	0.1232	0.0486	0.0513	0.1591
	(0.1309)	(0.0812)	(0.0566)	(0.0907)	(0.1626)	(0.1655)	(0.0875)	(0.0711)	(0.0878)	(0.1476)
South East Serbia	-0.1755	-0.0184	-0.0256	0.0031	0.0192	0.1148	-0.0427	-0.0023	0.0497	0.2183
	(0.1364)	(0.0844)	(0.0589)	(0.0944)	(0.1684)	(0.1730)	(0.0903)	(0.0746)	(0.0919)	(0.1577)
Constant	2.3787***	2.5403***	3.1683***	3.4935***	3.9424***	1.4464***	2.0379***	2.3644***	2.7463***	3.1996***
	(0.2187)	(0.1378)	(0.1009)	(0.1685)	(0.3012)	(0.2806)	(0.1506)	(0.1324)	(0.1742)	(0.3075)
N	1466	1466	1466	1466	1466	984	984	984	984	984
Pseudo R^2	0.1003	0.0926	0.0740	0.0688	0.0684	0.1683	0.1572	0.1443	0.1194	0.1282

Notes: Standard errors in parentheses * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

TABLE A5.6: Unadjusted QR estimates of returns to education with interaction terms, Serbia 2003

	MALES					FEMALES				
	θ=0.10	θ=0.25	θ=0.50	θ=0.75	θ=0.90	θ=0.10	θ=0.25	θ=0.50	θ=0.75	θ=0.90
school	0.0437	0.0613**	0.0376**	0.0556**	0.0975**	0.0673**	0.1153***	0.1284***	0.1323***	0.1596***
	(0.0365)	(0.0257)	(0.0191)	(0.0266)	(0.0489)	(0.0278)	(0.0195)	(0.0254)	(0.0262)	(0.0552)
educ*bef91	0.0342	0.0151	0.0337*	0.0112	-0.0389	0.0396	-0.0212	-0.0361	-0.0475*	-0.0628
	(0.0384)	(0.0272)	(0.0203)	(0.0285)	(0.0536)	(0.0316)	(0.0214)	(0.0274)	(0.0282)	(0.0577)
bef1991	-0.5609	-0.2071	-0.4137	-0.1892	0.5694	-0.4002	0.3158	0.4102	0.5851	0.8379
	(0.4833)	(0.3418)	(0.2523)	(0.3552)	(0.6339)	(0.4049)	(0.2811)	(0.3575)	(0.3640)	(0.7259)
exp	0.0045	0.0091	0.0057	0.0181*	-0.0014	0.0204	0.0196**	0.0154	0.0196**	0.0111
	(0.0145)	(0.0105)	(0.0075)	(0.0107)	(0.0208)	(0.0163)	(0.0092)	(0.0099)	(0.0093)	(0.0143)
expsq	-0.0001	-0.0002	-0.0001	-0.0002	0.0001	-0.0003	-0.0003**	-0.0001	-0.0001	0.0001
	(0.0002)	(0.0002)	(0.0001)	(0.0002)	(0.0003)	(0.0002)	(0.0001)	(0.0001)	(0.0001)	(0.0002)
married	0.1177	0.1155*	0.0350	-0.0105	0.0225	0.0188	-0.0010	0.0226	-0.0084	-0.0780
	(0.0897)	(0.0637)	(0.0448)	(0.0613)	(0.1073)	(0.0840)	(0.0520)	(0.0561)	(0.0525)	(0.0879)
1-2 years	0.1171	0.0036	-0.0308	-0.1148	-0.1812	0.1197	-0.0178	0.0056	0.0511	-0.0361
	(0.2019)	(0.1439)	(0.0998)	(0.1313)	(0.2511)	(0.2026)	(0.1315)	(0.1411)	(0.1282)	(0.2262)
3-5 years	0.1745	0.1178	-0.0232	-0.0682	-0.2441	0.1749	0.1415	0.1145	0.0949	0.1554
	(0.1333)	(0.0982)	(0.0687)	(0.0937)	(0.1765)	(0.1560)	(0.0969)	(0.1020)	(0.0928)	(0.1648)
6-10 years	0.1523	0.0235	0.0251	-0.0827	-0.1522	0.0829	0.0453	0.1202	0.1998**	0.1237
	(0.1239)	(0.0931)	(0.0648)	(0.0883)	(0.1547)	(0.1432)	(0.0860)	(0.0885)	(0.0799)	(0.1312)
>10 years	0.1542	0.0267	-0.0068	-0.1120	-0.1297	-0.1959	-0.0357	0.0377	0.0265	-0.0858
	(0.1005)	(0.0732)	(0.0519)	(0.0708)	(0.1359)	(0.1224)	(0.0707)	(0.0739)	(0.0668)	(0.1104)
public	-0.0228	-0.0011	-0.0485	-0.1676***	-0.3665***	0.2142**	0.2375***	0.0634	-0.0456	-0.1173
	(0.0672)	(0.0493)	(0.0342)	(0.0464)	(0.0858)	(0.0849)	(0.0501)	(0.0531)	(0.0486)	(0.0805)
Belgrade	0.4510***	0.4322***	0.3100***	0.3607***	0.3281*	0.4619***	0.2614***	0.2360**	0.2100**	0.3243**
	(0.1306)	(0.0952)	(0.0657)	(0.0908)	(0.1682)	(0.1534)	(0.0872)	(0.0925)	(0.0836)	(0.1421)
Vojvodina	0.1305	0.1494*	0.0346	0.1061	0.0212	0.3472**	0.1273	0.0830	0.1133	0.2565**
	(0.1151)	(0.0845)	(0.0585)	(0.0808)	(0.1457)	(0.1444)	(0.0818)	(0.0866)	(0.0782)	(0.1282)
West Serbia	-0.0070	0.0669	-0.0163	0.0918	-0.0551	0.1720	-0.0647	-0.0955	-0.1530	-0.1014
	(0.1327)	(0.0983)	(0.0678)	(0.0935)	(0.1632)	(0.1736)	(0.0994)	(0.1060)	(0.0970)	(0.1664)
Šumadija & Pomoravlje	0.1307	0.1155	0.1433**	0.2475***	0.1500	0.2989**	0.1291	0.0879	0.0959	0.1380
	(0.1214)	(0.0886)	(0.0608)	(0.0832)	(0.1498)	(0.1522)	(0.0882)	(0.0917)	(0.0839)	(0.1407)
South East Serbia	-0.1813	-0.0289	-0.0224	0.0115	-0.0539	0.1642	-0.0532	0.0117	0.0619	0.2314
	(0.1262)	(0.0915)	(0.0632)	(0.0864)	(0.1549)	(0.1581)	(0.0903)	(0.0964)	(0.0882)	(0.1476)
Constant	2.7375***	2.7573***	3.5199***	3.6234***	3.8027***	1.8929***	1.7807***	2.0295***	2.2340***	2.3484***
	(0.4650)	(0.3203)	(0.2388)	(0.3367)	(0.6246)	(0.4007)	(0.2647)	(0.3341)	(0.3406)	(0.7060)
N	1466	1466	1466	1466	1466	984	984	984	984	984
Pseudo R^2	0.1023	0.0928	0.0750	0.0690	0.0700	0.1705	0.1579	0.1459	0.1227	0.1320

Notes: Standard errors in parentheses * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

TABLE A5.7: Selectivity adjusted QR estimates of returns to education, Serbia 2003

	Males					Females				
	θ=0.10	θ=0.25	θ=0.50	θ=0.75	θ=0.90	θ=0.10	θ=0.25	θ=0.50	θ=0.75	θ=0.90
school	0.0832***	0.0770***	0.0676***	0.0718***	0.0633***	0.1034***	0.1015***	0.1083***	0.1010***	0.1124***
	(0.0126)	(0.0089)	(0.0064)	(0.0095)	(0.0208)	(0.0155)	(0.0082)	(0.0094)	(0.0123)	(0.0226)
exp	-0.0138	0.0080	0.0126*	0.0257***	0.0282	0.0344**	0.0114	0.0062	0.0076	0.0053
	(0.0155)	(0.0095)	(0.0065)	(0.0091)	(0.0198)	(0.0152)	(0.0072)	(0.0077)	(0.0090)	(0.0147)
expsq	0.0002	-0.0002	-0.0002*	-0.0004**	-0.0004	-0.0005**	-0.0002	0.0001	0.0001	0.0001
	(0.0003)	(0.0002)	(0.0001)	(0.0002)	(0.0004)	(0.0003)	(0.0001)	(0.0001)	(0.0002)	(0.0003)
married	0.0638	0.1042*	0.0260	-0.0123	0.0349	0.0131	0.0375	0.0297	0.0256	0.0143
	(0.0982)	(0.0625)	(0.0423)	(0.0568)	(0.1156)	(0.0996)	(0.0491)	(0.0505)	(0.0593)	(0.1033)
1-2 years	-0.1787	0.0867	0.0375	0.0049	0.2387	0.1198	0.0541	0.0681	0.2000	0.0191
	(0.2627)	(0.1699)	(0.1153)	(0.1543)	(0.3272)	(0.2447)	(0.1239)	(0.1297)	(0.1517)	(0.2659)
3-5 years	0.0443	0.1287	0.0408	0.0036	0.0819	0.2472	0.2161**	0.2273**	0.2812**	0.1908
	(0.1973)	(0.1258)	(0.0862)	(0.1144)	(0.2402)	(0.2226)	(0.1007)	(0.1048)	(0.1266)	(0.2231)
6-10 years	-0.0364	0.0333	0.0568	-0.0239	0.0478	0.0999	0.2038**	0.2038**	0.3527***	0.2515
	(0.1809)	(0.1218)	(0.0838)	(0.1134)	(0.2453)	(0.1985)	(0.0903)	(0.0938)	(0.1109)	(0.1827)
>10 years	-0.0152	0.0207	0.0533	-0.0209	0.0864	-0.1262	0.1474	0.1203	0.1896*	0.0468
	(0.1740)	(0.1148)	(0.0779)	(0.1054)	(0.2396)	(0.2031)	(0.0912)	(0.0923)	(0.1095)	(0.1874)
public	-0.1207	0.0240	-0.0459	-0.1252*	-0.1756	0.1667	0.3391***	0.1702**	0.1118	0.0632
	(0.1059)	(0.0705)	(0.0484)	(0.0655)	(0.1349)	(0.1382)	(0.0643)	(0.0671)	(0.0808)	(0.1409)
Belgrade	0.4525***	0.4288***	0.3143***	0.3273***	0.3610**	0.4240**	0.2408***	0.1946**	0.1558	0.3290**
	(0.1407)	(0.0933)	(0.0615)	(0.0851)	(0.1758)	(0.1736)	(0.0785)	(0.0820)	(0.0951)	(0.1656)
Vojvodina	0.2033	0.1399*	0.0188	0.0494	0.0451	0.2901*	0.1160	0.0467	0.1017	0.2742*
	(0.1246)	(0.0834)	(0.0557)	(0.0771)	(0.1555)	(0.1626)	(0.0740)	(0.0779)	(0.0917)	(0.1595)
West Serbia	-0.0132	0.0412	-0.0159	0.0304	-0.0625	0.0890	-0.0567	-0.1321	-0.1597	-0.0932
	(0.1460)	(0.0960)	(0.0637)	(0.0878)	(0.1769)	(0.2006)	(0.0900)	(0.0929)	(0.1077)	(0.1925)
Šumadija & Pomoravlje	0.1376	0.1233	0.1614***	0.1903**	0.1829	0.2403	0.1315*	0.0482	0.0905	0.1287
	(0.1286)	(0.0865)	(0.0568)	(0.0782)	(0.1577)	(0.1704)	(0.0791)	(0.0804)	(0.0926)	(0.1631)
South East Serbia	-0.1358	-0.0251	-0.0250	-0.0570	-0.0139	0.0601	-0.0311	-0.0009	0.0706	0.2118
	(0.1342)	(0.0893)	(0.0593)	(0.0819)	(0.1641)	(0.1745)	(0.0816)	(0.0848)	(0.0979)	(0.1688)
$(z'\gamma)$	-0.6251	0.1391	0.0477	0.2785	0.8198	0.1577	0.7343***	0.6157**	0.8599***	0.7908
	(0.4702)	(0.3019)	(0.2072)	(0.2704)	(0.5516)	(0.5581)	(0.2615)	(0.2752)	(0.3211)	(0.5412)
$(z'\gamma)^2$	0.1755	-0.0569	0.0702	0.0232	-0.1145	-0.0182	-0.2545**	-0.3492***	-0.3330***	-0.2805
	(0.1186)	(0.0779)	(0.0571)	(0.0700)	(0.1395)	(0.2055)	(0.1005)	(0.1094)	(0.1170)	(0.1908)
constant	2.6276***	2.5359***	3.0474***	3.2749***	3.5124***	1.3686***	1.8226***	2.2003***	2.4776***	2.7204***
	(0.3511)	(0.2205)	(0.1467)	(0.1967)	(0.4003)	(0.3328)	(0.1639)	(0.1788)	(0.2138)	(0.3911)
N	1466	1466	1466	1466	1466	984	984	984	984	984
Pseudo R^2	0.1022	0.0928	0.0758	0.0715	0.0718	0.1687	0.1608	0.1467	0.1253	0.1320

Notes: Standard errors in parentheses * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

TABLE A5.8: Endogeneity adjusted QR estimates of returns to education, Serbia 2003

	Males					*Females*				
	θ=0.10	θ=0.25	θ=0.50	θ=0.75	θ=0.90	θ=0.10	θ=0.25	θ=0.50	θ=0.75	θ=0.90
school	0.0182	0.0693***	0.0879***	0.0991***	0.1631***	0.0261	0.0888***	0.1107***	0.1419***	0.2095***
	(0.0306)	(0.0126)	(0.0090)	(0.0151)	(0.0335)	(0.0334)	(0.0161)	(0.0128)	(0.0199)	(0.0426)
exp	-0.0105	0.0067	0.0038	0.0192**	0.0225	0.0101	0.0182***	0.0094	0.0101	-0.0024
	(0.0121)	(0.0071)	(0.0049)	(0.0080)	(0.0142)	(0.0119)	(0.0062)	(0.0064)	(0.0084)	(0.0118)
expsq	-0.0001	-0.0002	-0.0000	-0.0003**	-0.0002	-0.0003**	-0.0003***	0.0000	0.0002	0.0005**
	(0.0002)	(0.0001)	(0.0001)	(0.0001)	(0.0002)	(0.0002)	(0.0001)	(0.0001)	(0.0001)	(0.0002)
married	0.1424	0.1030*	0.0451	-0.0329	0.0705	-0.0181	-0.0155	0.0214	-0.0147	-0.0689
	(0.0953)	(0.0570)	(0.0391)	(0.0573)	(0.0998)	(0.0781)	(0.0464)	(0.0480)	(0.0613)	(0.0853)
1-2 years	0.1764	0.0270	0.0026	-0.0659	-0.0276	0.1272	0.0931	0.0501	0.0474	-0.0628
	(0.2140)	(0.1282)	(0.0880)	(0.1268)	(0.2197)	(0.2176)	(0.1165)	(0.1223)	(0.1471)	(0.1907)
3-5 years	0.1900	0.1145	-0.0148	-0.0459	-0.4129**	0.2215	0.2240***	0.1359	0.1251	0.0780
	(0.1420)	(0.0869)	(0.0594)	(0.0903)	(0.1629)	(0.1437)	(0.0859)	(0.0875)	(0.1074)	(0.1567)
6-10 years	0.1793	0.0211	0.0029	-0.0743	-0.4265***	0.1677	0.1349*	0.1302*	0.1563*	0.0979
	(0.1366)	(0.0830)	(0.0563)	(0.0857)	(0.1491)	(0.1355)	(0.0787)	(0.0757)	(0.0940)	(0.1317)
>10 years	0.2626**	0.0201	-0.0073	-0.0848	-0.3923***	0.0324	0.0478	0.0268	0.0179	-0.1347
	(0.1240)	(0.0659)	(0.0456)	(0.0706)	(0.1404)	(0.1280)	(0.0711)	(0.0644)	(0.0800)	(0.1160)
public	-0.0200	0.0226	-0.0610**	-0.1923***	-0.3500***	0.2195***	0.1899***	0.0876*	-0.1175*	-0.2547***
	(0.0700)	(0.0438)	(0.0299)	(0.0451)	(0.0742)	(0.0799)	(0.0437)	(0.0457)	(0.0623)	(0.0805)
Belgrade	0.5264***	0.4239***	0.2911***	0.2590***	0.4720***	0.6317***	0.3226***	0.2264***	0.2044**	0.2542*
	(0.1553)	(0.0845)	(0.0575)	(0.0926)	(0.1505)	(0.1474)	(0.0788)	(0.0800)	(0.0979)	(0.1414)
Vojvodina	0.0852	0.1207	0.0067	0.0538	0.1456	0.4394***	0.1990***	0.0729	0.1378	0.1392
	(0.1250)	(0.0755)	(0.0512)	(0.0783)	(0.1285)	(0.1318)	(0.0732)	(0.0747)	(0.0912)	(0.1293)
West Serbia	-0.0159	0.0341	-0.0521	0.0616	0.0616	0.2291	0.0245	-0.0980	-0.1789	-0.0318
	(0.1441)	(0.0880)	(0.0595)	(0.0903)	(0.1466)	(0.1611)	(0.0899)	(0.0919)	(0.1139)	(0.1621)
Šumadija & Pomoravlje	0.1217	0.0961	0.1057**	0.1659**	0.2806**	0.3750***	0.1715**	0.0157	0.1108	0.1050
	(0.1303)	(0.0790)	(0.0532)	(0.0807)	(0.1323)	(0.1402)	(0.0780)	(0.0790)	(0.0971)	(0.1385)
South East Serbia	-0.0655	-0.0431	-0.0388	-0.0610	0.0281	0.1807	0.0098	-0.0089	0.0861	0.1344
	(0.1370)	(0.0820)	(0.0554)	(0.0838)	(0.1378)	(0.1448)	(0.0811)	(0.0828)	(0.1020)	(0.1460)
r^1	0.0108***	0.0066**	0.0083***	0.0110***	0.0131***	0.0336***	0.0131***	0.0140***	0.0046	0.0004
	(0.0040)	(0.0033)	(0.0026)	(0.0033)	(0.0043)	(0.0105)	(0.0033)	(0.0044)	(0.0054)	(0.0058)
r^2	0.0013	0.0000	-0.0005**	0.0002	-0.0008	-0.0042**	-0.0013*	-0.0002	-0.0009***	-0.0024**
	(0.0009)	(0.0003)	(0.0002)	(0.0003)	(0.0011)	(0.0019)	(0.0008)	(0.0004)	(0.0003)	(0.0010)
r^3	-0.0002***	-0.0001	-0.0001	-0.0000	-0.0001	0.0002**	0.0001	-0.0000	0.0000	-0.0001
	(0.0001)	(0.0000)	(0.0000)	(0.0000)	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)
r^4	0.0001***	0.0001	0.0001	0.0000	0.0004	0.0003**	0.0002	-0.0003	0.0001	0.0005
	(0.0002)	(0.0000)	(0.0000)	(0.0000)	(0.0001)	(0.0001)	(0.0000)	(0.0002)	(0.0004)	(0.0001)
Constant	2.9931***	2.6756***	2.9248***	3.0822***	2.6600***	2.2675***	2.0195***	2.2385***	2.1853***	1.9278***
	(0.3762)	(0.1754)	(0.1246)	(0.2146)	(0.4342)	(0.4512)	(0.2173)	(0.1800)	(0.2605)	(0.5414)
N	1466	1466	1466	1466	1466	984	984	984	984	984
Pseudo R^2	0.1058	0.0945	0.0809	0.0821	0.0816	0.1818	0.1683	0.1598	0.1327	0.1431

Notes: Standard errors in parentheses * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$; the residuals of the fifth order are insignificant and for a reason of space not reported;

TABLE A5.9: Unadjusted QR estimates of returns to education, Russia 2003

	Males					*Females*				
	θ=0.10	θ=0.25	θ=0.50	θ=0.75	θ=0.90	θ=0.10	θ=0.25	θ=0.50	θ=0.75	θ=0.90
school	0.1174***	0.0984***	0.0808***	0.0747***	0.0709***	0.1205***	0.1153***	0.1035***	0.0913***	0.0859***
	(0.0045)	(0.0028)	(0.0023)	(0.0026)	(0.0035)	(0.0033)	(0.0024)	(0.0025)	(0.0027)	(0.0031)
exp	-0.0086**	-0.0050**	0.0013	0.0070***	0.0070**	0.0034	0.0044**	0.0069***	0.0078***	0.0088***
	(0.0036)	(0.0022)	(0.0018)	(0.0021)	(0.0027)	(0.0026)	(0.0019)	(0.0019)	(0.0020)	(0.0021)
expsq	0.0001	0.0000	-0.0001***	-0.0002***	-0.0002***	-0.0001	-0.0001*	-0.0001***	-0.0001***	-0.0001***
	(0.0001)	(0.0000)	(0.0000)	(0.0000)	(0.0001)	(0.0001)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
married	0.0930***	0.1340***	0.1214***	0.1197***	0.1260***	-0.0387**	-0.0462***	-0.0403***	-0.0235**	-0.0118
	(0.0260)	(0.0161)	(0.0132)	(0.0151)	(0.0196)	(0.0152)	(0.0108)	(0.0110)	(0.0117)	(0.0126)
1-2 years	0.1222***	0.1667***	0.1447***	0.0991***	0.1023***	0.0688***	0.0986***	0.0876***	0.0688***	0.0893***
	(0.0325)	(0.0203)	(0.0167)	(0.0190)	(0.0249)	(0.0263)	(0.0187)	(0.0191)	(0.0202)	(0.0217)
3-5 years	0.2751***	0.2895***	0.2400***	0.1771***	0.1740***	0.1810***	0.1768***	0.1630***	0.1407***	0.1460***
	(0.0361)	(0.0226)	(0.0185)	(0.0211)	(0.0278)	(0.0295)	(0.0210)	(0.0214)	(0.0226)	(0.0244)
6-10 years	0.3738***	0.3652***	0.3083***	0.2365***	0.2075***	0.2630***	0.2568***	0.2285***	0.1898***	0.1872***
	(0.0351)	(0.0218)	(0.0179)	(0.0204)	(0.0268)	(0.0283)	(0.0201)	(0.0204)	(0.0216)	(0.0233)
>10 years	0.3129***	0.3412***	0.3345***	0.2705***	0.2416***	0.3109***	0.3121***	0.2640***	0.2107***	0.1893***
	(0.0329)	(0.0205)	(0.0167)	(0.0190)	(0.0250)	(0.0265)	(0.0187)	(0.0191)	(0.0202)	(0.0219)
arrears	-0.6292***	-0.4873***	-0.3520***	-0.2994***	-0.2861***	-0.2464***	-0.2354***	-0.2149***	-0.2142***	-0.2363***
	(0.0243)	(0.0152)	(0.0126)	(0.0145)	(0.0189)	(0.0205)	(0.0147)	(0.0151)	(0.0160)	(0.0173)
public	0.0948***	0.0450***	-0.0113	-0.0457***	-0.0834***	0.0441***	-0.0432***	-0.0619***	-0.0553***	-0.0716***
	(0.0202)	(0.0124)	(0.0102)	(0.0117)	(0.0155)	(0.0162)	(0.0115)	(0.0117)	(0.0124)	(0.0135)
Central	0.2597***	0.2722***	0.2148***	0.2356***	0.2911***	0.1312***	0.1300***	0.1775***	0.2204***	0.2772***
	(0.0337)	(0.0209)	(0.0171)	(0.0195)	(0.0256)	(0.0252)	(0.0179)	(0.0182)	(0.0192)	(0.0208)
North-West	0.6195***	0.6636***	0.6072***	0.6061***	0.6036***	0.4762***	0.5481***	0.5752***	0.5834***	0.5923***
	(0.0369)	(0.0230)	(0.0188)	(0.0214)	(0.0281)	(0.0276)	(0.0197)	(0.0200)	(0.0212)	(0.0228)
Siberia	0.2870***	0.3355***	0.3376***	0.3588***	0.4353***	0.2313***	0.2922***	0.3100***	0.3446***	0.4140***
	(0.0374)	(0.0233)	(0.0191)	(0.0218)	(0.0285)	(0.0281)	(0.0200)	(0.0203)	(0.0215)	(0.0232)
Far-East	0.7402***	0.7597***	0.7343***	0.7271***	0.7378***	0.6866***	0.7258***	0.7563***	0.7974***	0.8387***
	(0.0372)	(0.0232)	(0.0190)	(0.0216)	(0.0283)	(0.0281)	(0.0200)	(0.0203)	(0.0215)	(0.0232)
Urals	0.4911***	0.4950***	0.5178***	0.6115***	0.7924***	0.2328***	0.3312***	0.4122***	0.5277***	0.7236***
	(0.0427)	(0.0265)	(0.0217)	(0.0247)	(0.0324)	(0.0322)	(0.0229)	(0.0232)	(0.0246)	(0.0265)
Volga	0.1153***	0.1888***	0.1753***	0.1655***	0.1324***	0.0752***	0.1011***	0.1166***	0.1129***	0.1454***
	(0.0353)	(0.0220)	(0.0180)	(0.0205)	(0.0269)	(0.0265)	(0.0189)	(0.0192)	(0.0203)	(0.0219)
Constant	0.2035***	0.7822***	1.4398***	1.8945***	2.2900***	-0.0542	0.3945***	0.9198***	1.4236***	1.7938***
	(0.0674)	(0.0417)	(0.0344)	(0.0393)	(0.0513)	(0.0510)	(0.0373)	(0.0389)	(0.0424)	(0.0473)
N	21874	21874	21874	21874	21874	24318	24318	24318	24318	24318
Pseudo R^2	0.1514	0.1410	0.1387	0.1330	0.1324	0.1321	0.1425	0.1449	0.1407	0.1343

Notes: Standard errors in parentheses * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

TABLE A5.10: Unadjusted QR estimates of returns to education with interaction terms, Russia 2003

	Males					Females				
	θ=0.10	θ=0.25	θ=0.50	θ=0.75	θ=0.90	θ=0.10	θ=0.25	θ=0.50	θ=0.75	θ=0.90
school	0.1244***	0.0942***	0.0834***	0.0816***	0.0769***	0.1079***	0.1073***	0.0978***	0.0908***	0.0850***
	(0.0092)	(0.0057)	(0.0045)	(0.0053)	(0.0068)	(0.0067)	(0.0053)	(0.0042)	(0.0054)	(0.0058)
educ*bef91	-0.0074	0.0068	-0.0037	-0.0084	-0.0094	0.0186**	0.0111*	0.0079*	0.0011	0.0009
	(0.0108)	(0.0067)	(0.0052)	(0.0061)	(0.0078)	(0.0077)	(0.0061)	(0.0048)	(0.0062)	(0.0067)
bef1991	-0.0391	-0.1880**	-0.0580	0.0446	0.1029	-0.2406**	-0.1587**	-0.1367**	-0.0472	-0.0809
	(0.1274)	(0.0799)	(0.0629)	(0.0737)	(0.0942)	(0.0980)	(0.0780)	(0.0606)	(0.0782)	(0.0843)
exp	0.0042	0.0037	0.0116***	0.0128***	0.0085*	0.0030	0.0059*	0.0104***	0.0112***	0.0165***
	(0.0060)	(0.0038)	(0.0030)	(0.0035)	(0.0045)	(0.0043)	(0.0033)	(0.0026)	(0.0032)	(0.0035)
expsq	-0.0001	-0.0001*	-0.0003***	-0.0003***	-0.0002***	-0.0000	-0.0001	-0.0002***	-0.0002***	-0.0003***
	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0000)	(0.0001)	(0.0001)
married	0.0908***	0.1385***	0.1190***	0.1161***	0.1218***	-0.0366**	-0.0462***	-0.0380***	-0.0234**	-0.0093
	(0.0279)	(0.0174)	(0.0136)	(0.0158)	(0.0197)	(0.0158)	(0.0121)	(0.0093)	(0.0118)	(0.0125)
1-2 years	0.1230***	0.1635***	0.1373***	0.0960***	0.1054***	0.0711***	0.0995***	0.0877***	0.0666***	0.0833***
	(0.0350)	(0.0220)	(0.0172)	(0.0199)	(0.0251)	(0.0272)	(0.0209)	(0.0160)	(0.0203)	(0.0216)
3-5 years	0.2669***	0.2820***	0.2335***	0.1721***	0.1706***	0.1904***	0.1831***	0.1564***	0.1378***	0.1404***
	(0.0390)	(0.0245)	(0.0192)	(0.0222)	(0.0280)	(0.0304)	(0.0235)	(0.0180)	(0.0228)	(0.0242)
6-10 years	0.3622***	0.3648***	0.2998***	0.2316***	0.2079***	0.2639***	0.2637***	0.2257***	0.1852***	0.1831***
	(0.0378)	(0.0237)	(0.0185)	(0.0215)	(0.0270)	(0.0293)	(0.0225)	(0.0172)	(0.0218)	(0.0232)
>10 years	0.2982***	0.3431***	0.3301***	0.2712***	0.2466***	0.3144***	0.3166***	0.2642***	0.2077***	0.1846***
	(0.0353)	(0.0222)	(0.0172)	(0.0199)	(0.0252)	(0.0272)	(0.0208)	(0.0160)	(0.0204)	(0.0218)
arrears	-0.6176***	-0.4835***	-0.3530***	-0.2999***	-0.2860***	-0.2415***	-0.2355***	-0.2162***	-0.2147***	-0.2342***
	(0.0260)	(0.0165)	(0.0130)	(0.0151)	(0.0190)	(0.0211)	(0.0164)	(0.0127)	(0.0162)	(0.0172)
public	0.0926***	0.0399***	-0.0136	-0.0475***	-0.0820***	0.0420**	-0.0445***	-0.0631***	-0.0571***	-0.0675***
	(0.0216)	(0.0135)	(0.0105)	(0.0123)	(0.0156)	(0.0167)	(0.0128)	(0.0098)	(0.0125)	(0.0134)
Central	0.2635***	0.2761***	0.2144***	0.2374***	0.2913***	0.1281***	0.1314***	0.1774***	0.2192***	0.2729***
	(0.0361)	(0.0227)	(0.0177)	(0.0205)	(0.0257)	(0.0259)	(0.0200)	(0.0153)	(0.0194)	(0.0206)
North-West	0.6194***	0.6626***	0.6101***	0.6072***	0.6055***	0.4747***	0.5455***	0.5765***	0.5824***	0.5860***
	(0.0394)	(0.0249)	(0.0194)	(0.0224)	(0.0282)	(0.0285)	(0.0220)	(0.0168)	(0.0213)	(0.0226)
Siberia	0.3000***	0.3388***	0.3340***	0.3644***	0.4344***	0.2295***	0.2906***	0.3108***	0.3435***	0.4137***
	(0.0400)	(0.0252)	(0.0197)	(0.0228)	(0.0286)	(0.0289)	(0.0223)	(0.0170)	(0.0216)	(0.0230)
Far-East	0.7372***	0.7634***	0.7356***	0.7317***	0.7374***	0.6827***	0.7222***	0.7577***	0.7961***	0.8299***
	(0.0398)	(0.0251)	(0.0196)	(0.0226)	(0.0284)	(0.0290)	(0.0223)	(0.0171)	(0.0216)	(0.0229)
Urals	0.4936***	0.5032***	0.5231***	0.6159***	0.7955***	0.2392***	0.3272***	0.4140***	0.5257***	0.7215***
	(0.0457)	(0.0287)	(0.0224)	(0.0259)	(0.0325)	(0.0331)	(0.0255)	(0.0195)	(0.0247)	(0.0263)
Volga	0.1202***	0.1826***	0.1727***	0.1683***	0.1354***	0.0809***	0.1041***	0.1152***	0.1135***	0.1393***
	(0.0377)	(0.0238)	(0.0186)	(0.0215)	(0.0270)	(0.0273)	(0.0211)	(0.0161)	(0.0204)	(0.0218)
constant	0.0580	0.8065***	1.3660***	1.7912***	2.2120***	0.1068	0.4848***	0.9748***	1.4167***	1.7715***
	(0.1227)	(0.0757)	(0.0588)	(0.0688)	(0.0869)	(0.0908)	(0.0716)	(0.0557)	(0.0717)	(0.0769)
N	21874	21874	21874	21874	21874	24318	24318	24318	24318	24318
Pseudo R^2	0.1518	0.1415	0.1392	0.1332	0.1325	0.1324	0.1427	0.1450	0.1407	0.1345

TABLE A5.11: Selectivity adjusted QR estimates of returns to education, Russia, 2003

	Males					Females				
	10th	25th	50th	75th	90th	10th	25th	50th	75th	90th
school	0.1463***	0.1090***	0.0854***	0.0770***	0.0677***	0.1940***	0.1766***	0.1519***	0.1389***	0.1216***
	(0.0069)	(0.0047)	(0.0037)	(0.0037)	(0.0055)	(0.0283)	(0.0188)	(0.0157)	(0.0185)	(0.0212)
exp	-0.0182***	-0.0120***	-0.0018	0.0047**	0.0081**	0.0381***	0.0333***	0.0297***	0.0303***	0.0258**
	(0.0042)	(0.0029)	(0.0023)	(0.0023)	(0.0034)	(0.0137)	(0.0091)	(0.0075)	(0.0088)	(0.0101)
expsq	0.0004***	0.0002***	-0.0000	-0.0001***	-0.0002***	-0.0008***	-0.0007***	-0.0006***	-0.0006***	-0.0005**
	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0003)	(0.0002)	(0.0002)	(0.0002)	(0.0002)
married	0.1614***	0.1676***	0.1372***	0.1358***	0.1085***	-0.6583***	-0.5617***	-0.4508***	-0.4248***	-0.3082*
	(0.0341)	(0.0234)	(0.0183)	(0.0185)	(0.0269)	(0.2400)	(0.1593)	(0.1321)	(0.1563)	(0.1785)
1-2 years	0.2231***	0.1949***	0.1393***	0.1003***	0.0713**	0.1695***	0.1907***	0.1536***	0.1348***	0.1388***
	(0.0442)	(0.0304)	(0.0235)	(0.0234)	(0.0340)	(0.0467)	(0.0315)	(0.0263)	(0.0311)	(0.0361)
3-5 years	0.5419***	0.4359***	0.3214***	0.2350***	0.1711***	0.2113***	0.2075***	0.1798***	0.1687***	0.1576***
	(0.0658)	(0.0458)	(0.0354)	(0.0351)	(0.0517)	(0.0331)	(0.0225)	(0.0188)	(0.0223)	(0.0263)
6-10 years	0.3746***	0.3590***	0.2989***	0.2257***	0.1965***	0.3423***	0.3245***	0.2796***	0.2425***	0.2235***
	(0.0365)	(0.0254)	(0.0198)	(0.0198)	(0.0288)	(0.0420)	(0.0284)	(0.0237)	(0.0280)	(0.0324)
>10 years	0.2690***	0.3296***	0.3127***	0.2573***	0.2278***	0.5385***	0.5088***	0.4185***	0.3618***	0.2954***
	(0.0334)	(0.0231)	(0.0180)	(0.0181)	(0.0264)	(0.0915)	(0.0610)	(0.0506)	(0.0598)	(0.0683)
arrears	-0.6793***	-0.4813***	-0.3355***	-0.2933***	-0.2346***	-0.4792***	-0.4241***	-0.3731***	-0.3693***	-0.3499***
	(0.0471)	(0.0324)	(0.0251)	(0.0252)	(0.0370)	(0.0922)	(0.0611)	(0.0508)	(0.0600)	(0.0686)
public	0.1393***	0.0529***	-0.0077	-0.0401***	-0.0864***	0.3093***	0.1836***	0.1153*	0.1251*	0.0601
	(0.0240)	(0.0165)	(0.0128)	(0.0130)	(0.0191)	(0.1075)	(0.0710)	(0.0588)	(0.0697)	(0.0799)
Central	0.1906***	0.2343***	0.2130***	0.2222***	0.2950***	0.1627***	0.1613***	0.1983***	0.2417***	0.2999***
	(0.0384)	(0.0266)	(0.0208)	(0.0206)	(0.0303)	(0.0302)	(0.0204)	(0.0171)	(0.0203)	(0.0241)
North-West	0.6865***	0.6928***	0.6311***	0.6135***	0.6084***	0.4870***	0.5485***	0.5843***	0.5897***	0.6027***
	(0.0385)	(0.0268)	(0.0210)	(0.0212)	(0.0309)	(0.0298)	(0.0201)	(0.0168)	(0.0199)	(0.0236)
Siberia	0.2028***	0.3100***	0.3201***	0.3493***	0.4532***	0.2909***	0.3361***	0.3487***	0.3771***	0.4457***
	(0.0430)	(0.0300)	(0.0234)	(0.0234)	(0.0346)	(0.0369)	(0.0248)	(0.0207)	(0.0245)	(0.0287)
Far-East	0.7471***	0.7607***	0.7334***	0.7258***	0.7329***	0.6894***	0.7232***	0.7587***	0.7938***	0.8419***
	(0.0373)	(0.0260)	(0.0204)	(0.0206)	(0.0300)	(0.0301)	(0.0204)	(0.0170)	(0.0201)	(0.0238)
Urals	0.4224***	0.4750***	0.5095***	0.6072***	0.7928***	0.2740***	0.3536***	0.4459***	0.5491***	0.7457***
	(0.0448)	(0.0310)	(0.0243)	(0.0244)	(0.0357)	(0.0373)	(0.0251)	(0.0210)	(0.0248)	(0.0293)
Volga	0.0738*	0.1543***	0.1704***	0.1515***	0.1346***	0.1359***	0.1459***	0.1582***	0.1497***	0.1810***
	(0.0381)	(0.0265)	(0.0207)	(0.0207)	(0.0305)	(0.0360)	(0.0242)	(0.0202)	(0.0239)	(0.0281)
$(z'\gamma)$	1.5365***	1.6077***	1.1921***	0.8478***	0.8290***	3.0602***	2.4595***	1.9098***	1.6201**	1.1910
	(0.2429)	(0.1688)	(0.1318)	(0.1331)	(0.1997)	(1.1340)	(0.7541)	(0.6298)	(0.7519)	(0.8601)
$(z'\gamma)^2$	-0.2972***	-0.2790***	-0.1995***	-0.1436***	-0.1275***	-0.7373***	-0.5514***	-0.4006**	-0.2204	-0.1610
	(0.0352)	(0.0247)	(0.0195)	(0.0199)	(0.0300)	(0.2799)	(0.1884)	(0.1604)	(0.1942)	(0.2237)
Constant	-2.0349***	-1.5807***	-0.3517*	0.6493***	1.0242***	-2.9927***	-2.0250***	-0.9844	-0.3645	0.4641
	(0.3901)	(0.2722)	(0.2127)	(0.2149)	(0.3226)	(1.1101)	(0.7377)	(0.6135)	(0.7278)	(0.8315)
N	21874	21874	21874	21874	21874	24318	24318	24318	24318	24318
Pseudo R^2	0.1565	0.1460	0.1412	0.1346	0.1336	0.1324	0.1324	0.1451	0.1410	0.1344

TABLE A5.12

Endogeneity adjusted QR estimates of returns to education for males, Russia 2003

	Males				
	θ=0.10	θ=0.25	θ=0.50	θ=0.75	θ=0.90
school	0.1169***	0.1170***	0.0829***	0.0500***	0.1287***
	(0.0143)	(0.0068)	(0.0053)	(0.0075)	(0.0210)
exp	-0.0068*	-0.0048*	0.0013	0.0072***	0.0085***
	(0.0037)	(0.0025)	(0.0019)	(0.0021)	(0.0029)
expsq	0.0001	0.0000	-0.0001***	-0.0002***	-0.0002***
	(0.0001)	(0.0001)	(0.0000)	(0.0000)	(0.0001)
married	0.0848***	0.1320***	0.1208***	0.1185***	0.1143***
	(0.0277)	(0.0178)	(0.0139)	(0.0154)	(0.0209)
1-2 years	0.1301***	0.1672***	0.1474***	0.1016***	0.0936***
	(0.0325)	(0.0225)	(0.0174)	(0.0195)	(0.0267)
3-5 years	0.2782***	0.2800***	0.2422***	0.1848***	0.1684***
	(0.0367)	(0.0250)	(0.0194)	(0.0216)	(0.0297)
6-10 years	0.3763***	0.3515***	0.3103***	0.2439***	0.2033***
	(0.0358)	(0.0241)	(0.0189)	(0.0209)	(0.0286)
>10 years	0.3044***	0.3372***	0.3377***	0.2753***	0.2336***
	(0.0345)	(0.0227)	(0.0179)	(0.0195)	(0.0267)
arrears	-0.6309***	-0.4853***	-0.3502***	-0.2957***	-0.2830***
	(0.0248)	(0.0169)	(0.0137)	(0.0148)	(0.0203)
public	0.1051***	0.0448***	-0.0117	-0.0448***	-0.0802***
	(0.0204)	(0.0138)	(0.0106)	(0.0120)	(0.0166)
Central	0.2703***	0.2691***	0.2175***	0.2383***	0.2840***
	(0.0333)	(0.0232)	(0.0177)	(0.0200)	(0.0274)
North-West	0.6176***	0.6587***	0.6124***	0.6119***	0.5980***
	(0.0364)	(0.0254)	(0.0194)	(0.0219)	(0.0300)
Siberia	0.2949***	0.3317***	0.3405***	0.3723***	0.4281***
	(0.0369)	(0.0258)	(0.0197)	(0.0223)	(0.0305)
Far-East	0.7340***	0.7587***	0.7392***	0.7379***	0.7395***
	(0.0368)	(0.0256)	(0.0196)	(0.0221)	(0.0303)
Urals	0.4924***	0.5023***	0.5210***	0.6187***	0.8020***
	(0.0423)	(0.0293)	(0.0224)	(0.0253)	(0.0346)
Volga	0.1158***	0.1831***	0.1776***	0.1737***	0.1289***
	(0.0349)	(0.0243)	(0.0186)	(0.0210)	(0.0288)
r^1	0.0080	-0.0179***	-0.0012	0.0035*	0.0067
	(0.0052)	(0.0049)	(0.0015)	(0.0021)	(0.0075)
r^2	-0.0006	0.0017**	0.0000	0.0036***	-0.0060**
	(0.0007)	(0.0008)	(0.0006)	(0.0009)	(0.0025)
r^3	-0.0002	0.0006***	0.0001*	-0.0002**	-0.0014***
	(0.0002)	(0.0001)	(0.0000)	(0.0001)	(0.0004)
r^4	0.0000	-0.0001***	-0.0000	-0.0001***	-0.0001***
	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
r^5	0.0003***	-0.0002***	-0.0001***	-0.0002***	-0.0003**
	(0.0000)	(0.0001)	(0.0009)	(0.0002)	(0.0001)
Constant	0.1613	0.6178***	1.4116***	2.1676***	1.4471***
	(0.1306)	(0.0761)	(0.0627)	(0.0900)	(0.3111)
N	21874	21874	21874	21874	21874
Pseudo R^2	0.1517	0.1421	0.1391	0.1339	0.1334

Notes: Standard errors in parentheses * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

TABLE A5.13

Endogeneity adjusted QR estimates of returns to education for females, Russia 2003

	Females				
	θ=0.10	θ=0.25	θ=0.50	θ=0.75	θ=0.90
school	0.0858***	0.1074***	0.1030***	0.1242***	0.1268***
	(0.0109)	(0.0057)	(0.0062)	(0.0051)	(0.0170)
exp	0.0055**	0.0054***	0.0071***	0.0093***	0.0113***
	(0.0028)	(0.0020)	(0.0019)	(0.0018)	(0.0023)
expsq	-0.0002**	-0.0001**	-0.0001***	-0.0001***	-0.0002***
	(0.0001)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
married	-0.0380**	-0.0473***	-0.0400***	-0.0214**	-0.0173
	(0.0153)	(0.0114)	(0.0111)	(0.0108)	(0.0134)
1-2 years	0.0822***	0.1008***	0.0883***	0.0599***	0.0834***
	(0.0265)	(0.0197)	(0.0192)	(0.0187)	(0.0230)
3-5 years	0.1842***	0.1830***	0.1672***	0.1250***	0.1422***
	(0.0297)	(0.0220)	(0.0216)	(0.0212)	(0.0258)
6-10 years	0.2653***	0.2594***	0.2298***	0.1614***	0.1778***
	(0.0286)	(0.0215)	(0.0208)	(0.0205)	(0.0247)
>10 years	0.3305***	0.3200***	0.2664***	0.1689***	0.1811***
	(0.0273)	(0.0205)	(0.0198)	(0.0198)	(0.0232)
arrears	-0.2612***	-0.2347***	-0.2075***	-0.1852***	-0.2352***
	(0.0208)	(0.0157)	(0.0156)	(0.0155)	(0.0183)
public	0.0395**	-0.0420***	-0.0672***	-0.1045***	-0.0803***
	(0.0164)	(0.0124)	(0.0118)	(0.0132)	(0.0143)
Central	0.1234***	0.1358***	0.1836***	0.2134***	0.2715***
	(0.0253)	(0.0188)	(0.0183)	(0.0178)	(0.0220)
North-West	0.4758***	0.5485***	0.5875***	0.5792***	0.5860***
	(0.0278)	(0.0207)	(0.0201)	(0.0196)	(0.0241)
Siberia	0.2295***	0.2963***	0.3133***	0.3410***	0.4188***
	(0.0281)	(0.0210)	(0.0204)	(0.0199)	(0.0245)
Far-East	0.6841***	0.7181***	0.7606***	0.7956***	0.8323***
	(0.0282)	(0.0210)	(0.0204)	(0.0199)	(0.0245)
Urals	0.2312***	0.3302***	0.4092***	0.5307***	0.7075***
	(0.0322)	(0.0240)	(0.0233)	(0.0227)	(0.0280)
Volga	0.0784***	0.1108***	0.1130***	0.1249***	0.1377***
	(0.0266)	(0.0198)	(0.0193)	(0.0188)	(0.0232)
r^1	0.0211***	0.0033	0.0083***	0.0060**	-0.0014
	(0.0051)	(0.0027)	(0.0018)	(0.0023)	(0.0064)
r^2	-0.0012**	0.0003	-0.0000	-0.0017***	-0.0042**
	(0.0006)	(0.0006)	(0.0007)	(0.0005)	(0.0021)
r^3	-0.0004*	0.0001	-0.0002**	-0.0003**	-0.0007*
	(0.0002)	(0.0001)	(0.0001)	(0.0001)	(0.0004)
r^4	0.0000*	-0.0000**	-0.0000	-0.0000	-0.0000*
	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
r^5	-0.0004	0.0012**	0.0000	0.0026***	-0.0003**
	(0.0007)	(0.0008)	(0.0006)	(0.0001)	(0.0025)
constant	0.2505**	0.4537***	0.8938***	0.9852***	1.2191***
	(0.1098)	(0.0655)	(0.0755)	(0.0705)	(0.2538)
N	24318	24318	24318	24318	24318
Pseudo R^2	0.1338	0.1434	0.1464	0.1429	0.1368

Notes: Standard errors in parentheses * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

TABLE A5.14: Endogeneity and selectivity adjusted QR estimates of returns to education for males, Russia 2003

	Males				
	θ=0.10	θ=0.25	θ=0.50	θ=0.75	θ=0.90
school	0.1341***	0.1182***	0.0894***	0.0518***	0.1373***
	(0.0150)	(0.0071)	(0.0066)	(0.0068)	(0.0199)
exp	-0.0174***	-0.0118***	-0.0019	0.0056***	0.0081**
	(0.0042)	(0.0027)	(0.0025)	(0.0021)	(0.0032)
expsq	0.0004***	0.0002**	-0.0000	-0.0002***	-0.0002**
	(0.0001)	(0.0001)	(0.0001)	(0.0000)	(0.0001)
married	0.1696***	0.1688***	0.1404***	0.1303***	0.1046***
	(0.0348)	(0.0216)	(0.0201)	(0.0168)	(0.0251)
1-2 years	0.2359***	0.1916***	0.1394***	0.0929***	0.0671*
	(0.0433)	(0.0280)	(0.0256)	(0.0213)	(0.0321)
3-5 years	0.5543***	0.4331***	0.3266***	0.2330***	0.1692***
	(0.0650)	(0.0421)	(0.0385)	(0.0319)	(0.0486)
6-10 years	0.3866***	0.3499***	0.2951***	0.2216***	0.1835***
	(0.0366)	(0.0233)	(0.0219)	(0.0180)	(0.0272)
>10 years	0.2809***	0.3288***	0.3084***	0.2581***	0.2220***
	(0.0342)	(0.0212)	(0.0203)	(0.0165)	(0.0249)
arrears	-0.6834***	-0.4742***	-0.3416***	-0.2959***	-0.2450***
	(0.0464)	(0.0298)	(0.0276)	(0.0229)	(0.0348)
public	0.1403***	0.0499***	-0.0047	-0.0397***	-0.0906***
	(0.0238)	(0.0151)	(0.0140)	(0.0118)	(0.0180)
Central	0.1864***	0.2383***	0.2134***	0.2315***	0.2811***
	(0.0374)	(0.0245)	(0.0225)	(0.0188)	(0.0286)
North-West	0.6852***	0.6903***	0.6343***	0.6231***	0.5948***
	(0.0375)	(0.0246)	(0.0227)	(0.0193)	(0.0291)
Siberia	0.2005***	0.3058***	0.3185***	0.3641***	0.4363***
	(0.0420)	(0.0275)	(0.0254)	(0.0213)	(0.0326)
Far-East	0.7424***	0.7632***	0.7347***	0.7304***	0.7350***
	(0.0364)	(0.0239)	(0.0221)	(0.0187)	(0.0284)
Urals	0.4165***	0.4775***	0.5098***	0.6040***	0.7990***
	(0.0437)	(0.0284)	(0.0263)	(0.0222)	(0.0336)
Volga	0.0701	0.1551***	0.1659***	0.1651***	0.1273***
	(0.0372)	(0.0243)	(0.0224)	(0.0189)	(0.0288)
$(z'\gamma)$	1.5644***	1.6206***	1.2421***	0.8481***	0.8020***
	(0.2394)	(0.1566)	(0.1432)	(0.1220)	(0.1893)
$(z'\gamma)^2$	-0.3017***	-0.2804***	-0.2095***	-0.1429***	-0.1255***
	(0.0348)	(0.0230)	(0.0212)	(0.0183)	(0.0285)
r^1	0.0066	-0.0150***	-0.0006	0.0042*	0.0056
	(0.0051)	(0.0045)	(0.0018)	(0.0018)	(0.0070)
r^2	-0.0007	0.0016**	-0.0001	0.0036***	-0.0070***
	(0.0007)	(0.0008)	(0.0006)	(0.0007)	(0.0024)
r^3	-0.0000	0.0005***	0.0001	-0.0002**	-0.0015***
	(0.0002)	(0.0001)	(0.0000)	(0.0001)	(0.0004)
r^4	0.0000	-0.0001***	-0.0000	-0.0001***	-0.0001***
	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
r^5	0.0001	0.0007*	0.0000	0.0006**	-0.0020**
	(0.0007)	(0.0002)	(0.0003)	(0.0009)	(0.0005)
constant	-1.9858***	-1.6817***	-0.4601	0.9224***	0.1017
	(0.3909)	(0.2552)	(0.2388)	(0.2089)	(0.4338)
N	21874	21874	21874	21874	21874
Pseudo R^2	0.1568	0.1469	0.1416	0.1357	0.1348

Notes: Standard errors in parentheses * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

TABLE A5.15: Endogeneity and selectivity adjusted QR estimates of returns to education for females, Russia 2003

	Females				
	θ=0.10	θ=0.25	θ=0.50	θ=0.75	θ=0.90
school	0.1560***	0.1753***	0.1663***	0.1722***	0.1834***
	(0.0274)	(0.0173)	(0.0190)	(0.0182)	(0.0294)
exp	0.0405***	0.0376***	0.0371***	0.0313***	0.0360**
	(0.0128)	(0.0081)	(0.0086)	(0.0084)	(0.0112)
expsq	-0.0009**	-0.0008***	-0.0007***	-0.0005***	-0.0007***
	(0.0003)	(0.0002)	(0.0002)	(0.0002)	(0.0002)
married	-0.6523**	-0.6336***	-0.5760***	-0.4187**	-0.4593*
	(0.2228)	(0.1423)	(0.1517)	(0.1485)	(0.1985)
1-2 years	0.1892***	0.1984***	0.1692***	0.1215***	0.1523***
	(0.0436)	(0.0282)	(0.0301)	(0.0297)	(0.0403)
3-5 years	0.2313***	0.2063***	0.1823***	0.1442***	0.1518***
	(0.0309)	(0.0201)	(0.0217)	(0.0215)	(0.0293)
6-10 years	0.3533***	0.3330***	0.2912***	0.2091***	0.2235***
	(0.0393)	(0.0256)	(0.0273)	(0.0269)	(0.0359)
>10 years	0.5666***	0.5354***	0.4567***	0.3131***	0.3369***
	(0.0856)	(0.0548)	(0.0581)	(0.0571)	(0.0759)
arrears	-0.4930***	-0.4509***	-0.4040***	-0.3326***	-0.3959***
	(0.0858)	(0.0547)	(0.0582)	(0.0571)	(0.0762)
public	0.3046**	0.2145***	0.1643*	0.0690	0.1124
	(0.0998)	(0.0635)	(0.0674)	(0.0664)	(0.0888)
Central	0.1692***	0.1709***	0.2148***	0.2350***	0.2896***
	(0.0281)	(0.0182)	(0.0196)	(0.0193)	(0.0267)
North-West	0.4975***	0.5554***	0.5942***	0.5830***	0.5924***
	(0.0277)	(0.0180)	(0.0193)	(0.0190)	(0.0261)
Siberia	0.2922***	0.3439***	0.3613***	0.3790***	0.4567***
	(0.0344)	(0.0222)	(0.0237)	(0.0233)	(0.0319)
Far-East	0.6853***	0.7168***	0.7673***	0.7952***	0.8311***
	(0.0279)	(0.0182)	(0.0195)	(0.0191)	(0.0263)
Urals	0.2811***	0.3612***	0.4469***	0.5542***	0.7304***
	(0.0346)	(0.0225)	(0.0240)	(0.0236)	(0.0325)
Volga	0.1371***	0.1592***	0.1647***	0.1610***	0.1791***
	(0.0335)	(0.0216)	(0.0232)	(0.0228)	(0.0311)
$(z'\gamma)$	3.2952***	3.0091***	2.7943***	1.7889**	2.1967**
	(1.0601)	(0.6773)	(0.7258)	(0.7155)	(0.9618)
$(z'\gamma)^2$	-0.9155***	-0.7796***	-0.7526***	-0.3456	-0.5439**
	(0.2675)	(0.1716)	(0.1871)	(0.1853)	(0.2534)
r^1	0.0202***	0.0046**	0.0092***	0.0062***	-0.0009
	(0.0051)	(0.0023)	(0.0016)	(0.0022)	(0.0069)
r^2	-0.0011	0.0000	-0.0002	-0.0018***	-0.0046**
	(0.0006)	(0.0005)	(0.0007)	(0.0004)	(0.0022)
r^3	-0.0003	0.0001	-0.0002***	-0.0003***	-0.0008
	(0.0002)	(0.0001)	(0.0001)	(0.0001)	(0.0004)
r^4	0.0000	-0.0000	-0.0000	-0.0000	-0.0000
	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
r^5	-0.0006	0.0007	0.0000	0.0006*	-0.0010*
	(0.0000)	(0.0008)	(0.0006)	(0.0001)	(0.0025)
constant	-2.7314***	-2.3333***	-1.6836**	-0.8510	-0.9281
	(1.0277)	(0.6607)	(0.7107)	(0.6967)	(0.9608)
N	24318	24318	24318	24318	24318
Pseudo R^2	0.1343	0.1440	0.1468	0.1432	0.1370

Notes: Standard errors in parentheses * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

TABLE A5.16: Unadjusted QR estimates of returns to education, Tajikistan 2003

	Males					Females				
	10th	25th	50th	75th	90th	10th	25th	50th	75th	90th
school	0.0370**	0.0460***	0.0457***	0.0490***	0.0398**	0.0575***	0.0850***	0.1139***	0.0972***	0.1157***
	(0.0151)	(0.0095)	(0.0088)	(0.0117)	(0.0165)	(0.0061)	(0.0118)	(0.0124)	(0.0155)	(0.0213)
exp	0.0506***	0.0469***	0.0482***	0.0385***	0.0317**	0.0117***	0.0264***	0.0323***	0.0400***	0.0337***
	(0.0113)	(0.0075)	(0.0069)	(0.0095)	(0.0129)	(0.0040)	(0.0080)	(0.0079)	(0.0096)	(0.0117)
expsq	-0.0009***	-0.0009***	-0.0009***	-0.0007***	-0.0005**	-0.0001	-0.0004**	-0.0004***	-0.0005**	-0.0003
	(0.0002)	(0.0001)	(0.0001)	(0.0002)	(0.0003)	(0.0001)	(0.0002)	(0.0002)	(0.0002)	(0.0002)
married	0.0320	0.0535	0.0434	0.1178	-0.2443**	-0.1457***	-0.1568**	-0.1189*	-0.2100***	-0.1603*
	(0.1230)	(0.0774)	(0.0700)	(0.0943)	(0.1220)	(0.0335)	(0.0626)	(0.0608)	(0.0717)	(0.0945)
1-2 years	0.0262	0.0390	-0.1570*	-0.0701	-0.2368*	0.0275	0.1516	0.1293	0.1510	0.1228
	(0.1312)	(0.0885)	(0.0817)	(0.1088)	(0.1422)	(0.0565)	(0.1027)	(0.1014)	(0.1205)	(0.1561)
3-5 years	0.0227	0.0052	-0.1784**	-0.1792*	-0.1860	-0.0037	-0.0005	0.0497	0.0386	0.1668
	(0.1256)	(0.0844)	(0.0779)	(0.1037)	(0.1374)	(0.0553)	(0.0997)	(0.0982)	(0.1169)	(0.1538)
6-10 years	-0.1781	-0.1513	-0.3862***	-0.3461***	-0.3011**	-0.0917	-0.0053	-0.0321	-0.0459	-0.2171
	(0.1383)	(0.0921)	(0.0848)	(0.1126)	(0.1525)	(0.0600)	(0.1095)	(0.1074)	(0.1271)	(0.1628)
>10 years	-0.2346*	-0.2934***	-0.5445***	-0.5659***	-0.6227***	-0.1081*	-0.2425**	-0.2286**	-0.2444**	-0.3777***
	(0.1344)	(0.0883)	(0.0793)	(0.1030)	(0.1396)	(0.0568)	(0.1009)	(0.0983)	(0.1154)	(0.1451)
public	0.0706	-0.0514	-0.1977***	-0.4138***	-0.4800***	-0.0420	-0.1239**	-0.2517***	-0.3200***	-0.4820***
	(0.0741)	(0.0495)	(0.0460)	(0.0608)	(0.0815)	(0.0308)	(0.0556)	(0.0576)	(0.0736)	(0.0971)
GBAO	-0.6765***	-0.5300***	-0.4820***	-0.5025***	-0.2869*	-0.5173***	-0.8782***	-0.9289***	-0.7733***	-0.7913***
	(0.1499)	(0.0998)	(0.0919)	(0.1223)	(0.1610)	(0.0606)	(0.1101)	(0.1069)	(0.1265)	(0.1685)
Sogdian	-0.7761***	-0.6258***	-0.5483***	-0.4535***	-0.4392***	-0.5529***	-0.6081***	-0.5692***	-0.6173***	-0.4861***
	(0.1108)	(0.0757)	(0.0716)	(0.0947)	(0.1223)	(0.0502)	(0.0916)	(0.0916)	(0.1085)	(0.1396)
Kahtlon	-0.9684***	-0.8168***	-0.8674***	-0.8210***	-0.7988***	-0.6523***	-0.9358***	-0.9151***	-1.0412***	-0.9288***
	(0.1037)	(0.0728)	(0.0700)	(0.0948)	(0.1229)	(0.0510)	(0.0908)	(0.0935)	(0.1159)	(0.1551)
RRS	-0.5904***	-0.3451***	-0.2081**	-0.1563	-0.0072	-0.7351***	-0.7440***	-0.8556***	-0.6804***	-0.5156***
	(0.1338)	(0.0911)	(0.0850)	(0.1120)	(0.1476)	(0.0625)	(0.1142)	(0.1125)	(0.1357)	(0.1776)
constant	2.1704***	2.6079***	3.3967***	4.0474***	5.1721***	1.7481***	2.0098***	2.2086***	2.8892***	3.3279***
	(0.2068)	(0.1475)	(0.1419)	(0.1921)	(0.2570)	(0.0970)	(0.1832)	(0.1943)	(0.2419)	(0.3171)
N	3006	3006	3006	3006	3006	1832	1832	1832	1832	1832
Pseudo R^2	0.0802	0.0887	0.0974	0.1023	0.0815	0.0765	0.0944	0.1353	0.1721	0.1870

Notes: Standard errors in parentheses * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

TABLE A5.17: Unadjusted QR estimates of returns to education with interactions terms, Tajikistan 2003

	Males					Females				
	10th	25th	50th	75th	90th	10th	25th	50th	75th	90th
school	0.0466**	0.0669***	0.0729***	0.0917***	0.1007***	0.0547***	0.0765***	0.1246***	0.1186***	0.1511***
	(0.0229)	(0.0158)	(0.0127)	(0.0182)	(0.0247)	(0.0093)	(0.0159)	(0.0156)	(0.0213)	(0.0278)
educ*bef91	-0.0145	-0.0287	-0.0502***	-0.0554**	-0.0657**	0.0128	0.0376*	-0.0259	-0.0398	-0.0648*
	(0.0269)	(0.0184)	(0.0148)	(0.0216)	(0.0307)	(0.0121)	(0.0203)	(0.0197)	(0.0261)	(0.0347)
bef1991	-0.0666	0.3220	0.4963***	0.5995**	0.7528*	-0.0970	-0.3826	0.4074*	0.5733*	0.8615**
	(0.3333)	(0.2292)	(0.1864)	(0.2714)	(0.3858)	(0.1516)	(0.2533)	(0.2408)	(0.3181)	(0.4219)
exp	0.0691***	0.0513***	0.0593***	0.0465***	0.0394*	0.0098	0.0236**	0.0221**	0.0346**	0.0264
	(0.0147)	(0.0103)	(0.0085)	(0.0131)	(0.0203)	(0.0066)	(0.0116)	(0.0109)	(0.0144)	(0.0190)
expsq	-0.0012***	-0.0010***	-0.0011***	-0.0008***	-0.0007*	-0.0000	-0.0003	-0.0003	-0.0004*	-0.0003
	(0.0002)	(0.0002)	(0.0001)	(0.0002)	(0.0003)	(0.0001)	(0.0002)	(0.0002)	(0.0003)	(0.0003)
married	0.0142	0.0314	-0.0192	0.0339	-0.3466**	-0.1544***	-0.1431**	-0.1004*	-0.2481***	-0.1618*
	(0.1162)	(0.0763)	(0.0611)	(0.0933)	(0.1353)	(0.0357)	(0.0613)	(0.0564)	(0.0741)	(0.0948)
1-2 years	-0.0582	0.0235	-0.1387**	-0.0486	-0.2745*	0.0610	0.1667*	0.1222	0.2025	0.0843
	(0.1156)	(0.0827)	(0.0691)	(0.1051)	(0.1526)	(0.0601)	(0.0996)	(0.0933)	(0.1240)	(0.1621)
3-5 years	-0.0427	-0.0196	-0.1880***	-0.1949*	-0.2312	0.0134	0.0126	0.0165	0.0467	0.1921
	(0.1119)	(0.0791)	(0.0658)	(0.1002)	(0.1492)	(0.0588)	(0.0965)	(0.0902)	(0.1208)	(0.1601)
6-10 years	-0.2083*	-0.1898**	-0.4088***	-0.3484***	-0.3377**	-0.0727	0.0125	-0.0609	-0.0324	-0.2146
	(0.1229)	(0.0861)	(0.0716)	(0.1085)	(0.1654)	(0.0646)	(0.1064)	(0.0990)	(0.1315)	(0.1666)
>10 years	-0.2781**	-0.3268***	-0.5255***	-0.5544***	-0.6367***	-0.1030*	-0.2509**	-0.2511***	-0.2366**	-0.3615**
	(0.1185)	(0.0825)	(0.0670)	(0.0994)	(0.1503)	(0.0604)	(0.0978)	(0.0902)	(0.1192)	(0.1470)
public	0.0555	-0.0446	-0.1947***	-0.3892***	-0.4811***	-0.0485	-0.1391***	-0.2174***	-0.2981***	-0.4671***
	(0.0647)	(0.0463)	(0.0389)	(0.0587)	(0.0877)	(0.0330)	(0.0537)	(0.0531)	(0.0749)	(0.0986)
GBAO	-0.7127***	-0.5450***	-0.5019***	-0.4743***	-0.3512**	-0.5163***	-0.8796***	-0.9509***	-0.7667***	-0.7702***
	(0.1308)	(0.0933)	(0.0778)	(0.1178)	(0.1743)	(0.0652)	(0.1061)	(0.0989)	(0.1304)	(0.1726)
Sogdian	-0.7894***	-0.5876***	-0.5615***	-0.4267***	-0.4223***	-0.5537***	-0.5997***	-0.5631***	-0.6346***	-0.4601***
	(0.0970)	(0.0711)	(0.0607)	(0.0915)	(0.1331)	(0.0533)	(0.0891)	(0.0841)	(0.1110)	(0.1424)
Kahtlon	-0.9907***	-0.7661***	-0.8747***	-0.8211***	-0.7886***	-0.6410***	-0.9248***	-0.9244***	-1.0357***	-0.9234***
	(0.0906)	(0.0682)	(0.0593)	(0.0917)	(0.1326)	(0.0536)	(0.0878)	(0.0859)	(0.1178)	(0.1601)
RRS	-0.5474***	-0.3151***	-0.2409***	-0.1542	0.0314	-0.7216***	-0.7267***	-0.8476***	-0.6762***	-0.4296**
	(0.1169)	(0.0858)	(0.0719)	(0.1079)	(0.1594)	(0.0661)	(0.1105)	(0.1033)	(0.1393)	(0.1837)
constant	2.0876***	2.3332***	3.0837***	3.5385***	4.4894***	1.7607***	2.1032***	2.1234***	2.6594***	2.9165***
	(0.2785)	(0.2060)	(0.1720)	(0.2531)	(0.3396)	(0.1251)	(0.2207)	(0.2177)	(0.2972)	(0.3826)
N	3006	3006	3006	3006	3006	1832	1832	1832	1832	1832
Pseudo R^2	0.0815	0.0895	0.0994	0.1042	0.0827	0.0771	0.0954	0.1364	0.1742	0.1904

Notes: Standard errors in parentheses * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

TABLE A5.18: Selectivity adjusted QR estimates of returns to education, Tajikistan 2003

	Males					Females				
	10th	25th	10th	25th	10th	25th	10th	25th	10th	25th
school	0.0741***	0.0801***	0.0794***	0.0954***	0.0554***	0.0591***	0.0863***	0.1148***	0.0997***	0.1253***
	(0.0162)	(0.0098)	(0.0111)	(0.0111)	(0.0185)	(0.0079)	(0.0145)	(0.0141)	(0.0178)	(0.0196)
exp	0.0276**	0.0245***	0.0317***	0.0142*	0.0146	0.0134***	0.0256***	0.0300***	0.0415***	0.0334***
	(0.0118)	(0.0069)	(0.0080)	(0.0081)	(0.0139)	(0.0050)	(0.0092)	(0.0085)	(0.0103)	(0.0099)
expsq	-0.0004*	-0.0004***	-0.0005***	-0.0001	-0.0002	-0.0001	-0.0003*	-0.0004*	-0.0005**	-0.0003*
	(0.0002)	(0.0001)	(0.0002)	(0.0002)	(0.0003)	(0.0001)	(0.0002)	(0.0002)	(0.0002)	(0.0002)
married	0.5149***	0.4954***	0.4617***	0.7076***	-0.0347	-0.1591***	-0.1433	-0.1021	-0.1891*	-0.0225
	(0.1775)	(0.1022)	(0.1164)	(0.1089)	(0.1752)	(0.0545)	(0.0976)	(0.0867)	(0.1003)	(0.0997)
1-2 years	-0.0761	-0.0105	-0.1594*	-0.1162	-0.3222**	0.0421	0.1384	0.0745	0.1328	0.0948
	(0.1224)	(0.0736)	(0.0856)	(0.0865)	(0.1420)	(0.0686)	(0.1121)	(0.1045)	(0.1263)	(0.1284)
3-5 years	-0.0322	-0.0343	-0.1522*	-0.2039**	-0.2562*	-0.0104	-0.0004	-0.0030	0.0211	0.1350
	(0.1160)	(0.0698)	(0.0811)	(0.0819)	(0.1352)	(0.0667)	(0.1081)	(0.1005)	(0.1226)	(0.1263)
6-10 years	-0.1014	-0.1766**	-0.3515***	-0.3785***	-0.3515**	-0.0868	-0.0074	-0.0979	-0.0563	-0.2357*
	(0.1283)	(0.0762)	(0.0882)	(0.0892)	(0.1512)	(0.0750)	(0.1214)	(0.1119)	(0.1349)	(0.1362)
>10 years	-0.2261*	-0.2765***	-0.4845***	-0.5455***	-0.5801***	-0.1227*	-0.2504**	-0.3068***	-0.2787**	-0.4548***
	(0.1232)	(0.0736)	(0.0827)	(0.0818)	(0.1370)	(0.0744)	(0.1200)	(0.1102)	(0.1311)	(0.1268)
GBAO	-0.6646***	-0.5681***	-0.5482***	-0.6496***	-0.4146**	-0.5375***	-0.8843***	-0.9509***	-0.7893***	-0.8103***
	(0.1398)	(0.0841)	(0.0969)	(0.0979)	(0.1619)	(0.0742)	(0.1205)	(0.1110)	(0.1326)	(0.1373)
Sogdian	-0.7044***	-0.5376***	-0.4755***	-0.3531***	-0.4164***	-0.5837***	-0.6120***	-0.5962***	-0.6289***	-0.4573***
	(0.1048)	(0.0634)	(0.0754)	(0.0756)	(0.1203)	(0.0610)	(0.0996)	(0.0933)	(0.1127)	(0.1121)
Kahtlon	-0.8489***	-0.7140***	-0.7997***	-0.7628***	-0.7473***	-0.6720***	-0.9356***	-0.9419***	-1.0529***	-0.9579***
	(0.0986)	(0.0607)	(0.0736)	(0.0758)	(0.1219)	(0.0616)	(0.0990)	(0.0956)	(0.1206)	(0.1249)
RRS	-0.4857***	-0.2982***	-0.1880**	-0.1307	-0.0109	-0.7380***	-0.7413***	-0.8553***	-0.6650***	-0.4842***
	(0.1239)	(0.0759)	(0.0888)	(0.0894)	(0.1452)	(0.0756)	(0.1272)	(0.1174)	(0.1444)	(0.1459)
$(z'\gamma)$	1.6036**	1.6743***	1.6664***	2.1722***	1.9240***	0.0940	0.0529	0.1936	-0.2369	-1.1727***
	(0.6890)	(0.4218)	(0.4646)	(0.4613)	(0.7397)	(0.2784)	(0.4925)	(0.4225)	(0.4726)	(0.4510)
$(z'\gamma)^2$	-0.2769	-0.3580*	-0.3875*	-0.4405**	-0.8253**	-0.0512	-0.0349	-0.1175	0.0619	0.3840***
	(0.3124)	(0.1865)	(0.2128)	(0.2164)	(0.3587)	(0.0917)	(0.1774)	(0.1495)	(0.1622)	(0.1477)
constant	0.2142	0.7077*	1.5329***	1.4859***	3.9481***	1.7087***	1.9853***	2.2272***	3.0682***	3.9949***
	(0.5796)	(0.3658)	(0.4023)	(0.3813)	(0.5714)	(0.2008)	(0.3406)	(0.3076)	(0.3531)	(0.3548)
N	3006	3006	3006	3006	3006	1832	1832	1832	1832	1832
Pseudo R^2	0.0896	0.0964	0.1048	0.1110	0.0863	0.0766	0.0944	0.1360	0.1723	0.1903

Notes: Standard errors in parentheses * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

TABLE A5.19: Endogeneity adjusted QR estimates of the returns to education, Tajikistan 2003

	Males					Females				
	10th	25th	10th	25th	10th	25th	10th	25th	10th	25th
school	0.0346	0.0507***	0.0469***	0.0690***	0.0784***	0.0332**	0.0889***	0.0832***	0.0577**	0.1660***
	(0.0268)	(0.0151)	(0.0129)	(0.0213)	(0.0290)	(0.0143)	(0.0241)	(0.0197)	(0.0276)	(0.0356)
exp	0.0533***	0.0516***	0.0528***	0.0448***	0.0128	0.0054	0.0226***	0.0286***	0.0263***	0.0276***
	(0.0097)	(0.0055)	(0.0052)	(0.0077)	(0.0107)	(0.0043)	(0.0080)	(0.0068)	(0.0087)	(0.0107)
expsq	-0.0010***	-0.0010***	-0.0010***	-0.0008***	-0.0002	-0.0000	-0.0003	-0.0004***	-0.0002	-0.0002
	(0.0002)	(0.0001)	(0.0001)	(0.0002)	(0.0002)	(0.0001)	(0.0002)	(0.0001)	(0.0002)	(0.0002)
1-2 years	-0.0309	0.0408	-0.1312*	-0.0220	-0.2932**	0.0057	0.1368	0.1633*	0.0966	-0.0135
	(0.1234)	(0.0779)	(0.0727)	(0.1051)	(0.1438)	(0.0625)	(0.1082)	(0.0937)	(0.1186)	(0.1482)
3-5 years	0.0698	0.0023	-0.1485**	-0.1328	-0.2565*	0.0220	-0.0068	0.0328	0.0346	0.0926
	(0.1186)	(0.0741)	(0.0690)	(0.0999)	(0.1384)	(0.0614)	(0.1050)	(0.0907)	(0.1153)	(0.1487)
6-10 years	-0.1301	-0.1476*	-0.3683***	-0.2841***	-0.3116**	-0.0674	-0.0444	-0.0239	-0.1155	-0.2927*
	(0.1300)	(0.0808)	(0.0752)	(0.1086)	(0.1528)	(0.0661)	(0.1161)	(0.0990)	(0.1244)	(0.1541)
>10 years	-0.2213*	-0.3099***	-0.5411***	-0.5350***	-0.6526***	-0.0670	-0.2835***	-0.2185**	-0.2931**	-0.4870***
	(0.1241)	(0.0774)	(0.0706)	(0.0991)	(0.1385)	(0.0622)	(0.1062)	(0.0909)	(0.1137)	(0.1397)
public	0.0515	-0.0607	-0.2153***	-0.4861***	-0.5565***	-0.0007	-0.1457**	-0.2068***	-0.2429***	-0.6783***
	(0.0682)	(0.0435)	(0.0423)	(0.0848)	(0.0990)	(0.0342)	(0.0594)	(0.0538)	(0.0802)	(0.1331)
GBAO	-0.7006***	-0.5504***	-0.4395***	-0.4741***	-0.3426**	-0.5003***	-0.8276***	-0.9542***	-0.7893***	-0.7313***
	(0.1428)	(0.0887)	(0.0821)	(0.1186)	(0.1638)	(0.0659)	(0.1157)	(0.0994)	(0.1244)	(0.1606)
Sogdian	-0.7904***	-0.6084***	-0.5163***	-0.4051***	-0.3883***	-0.5548***	-0.5877***	-0.6589***	-0.7460***	-0.3317**
	(0.1057)	(0.0663)	(0.0663)	(0.0932)	(0.1249)	(0.0534)	(0.0964)	(0.0891)	(0.1240)	(0.1550)
Kahtlon	-0.9678***	-0.7792***	-0.8315***	-0.7832***	-0.7093***	-0.6780***	-0.9172***	-0.9998***	-1.2260***	-0.7794***
	(0.0999)	(0.0640)	(0.0668)	(0.0958)	(0.1318)	(0.0547)	(0.0967)	(0.0937)	(0.1378)	(0.1782)
RRS	-0.5720***	-0.3196***	-0.1987**	-0.1477	-0.0271	-0.7311***	-0.7130***	-0.9285***	-0.7565***	-0.2993
	(0.1283)	(0.0802)	(0.0774)	(0.1096)	(0.1510)	(0.0687)	(0.1206)	(0.1084)	(0.1468)	(0.1832)
r^1	0.0056	0.0076**	0.0063***	0.0045*	0.0101***	0.0016	0.0049	0.0058**	0.0156***	0.0145***
	(0.0053)	(0.0034)	(0.0020)	(0.0025)	(0.0023)	(0.0026)	(0.0036)	(0.0027)	(0.0031)	(0.0030)
r^2	0.0005	-0.0011*	-0.0004	-0.0002	0.0006	0.0009*	-0.0001	0.0017*	0.0036**	-0.0003
	(0.0010)	(0.0006)	(0.0006)	(0.0011)	(0.0008)	(0.0005)	(0.0009)	(0.0009)	(0.0015)	(0.0012)
r^3	-0.0001**	0.0000	-0.0000	-0.0000	-0.0000	-0.0000***	-0.0000	-0.0000**	-0.0000	0.0000
	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
r^4	-0.0001	0.0000	-0.0002	-0.0005	-0.0000	-0.0000**	-0.0000	0.0000**	0.0001	0.0008
	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
r^5	0.0000*	0.0000**	-0.0000**	-0.0010	-0.0011	0.0001**	-0.0000	-0.0000**	-0.0000	0.0000
	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
constant	2.1560***	2.5143***	3.3112***	3.7713***	4.6271***	1.9222***	1.8772***	2.5195***	3.4301***	2.6138***
	(0.2709)	(0.1781)	(0.1777)	(0.2831)	(0.4374)	(0.1540)	(0.2826)	(0.2654)	(0.3957)	(0.4983)
N	3006	3006	3006	3006	3006	1832	1832	1832	1832	1832
Pseudo R^2	0.0801	0.0850	0.0991	0.1044	0.0835	0.0676	0.1104	0.1374	0.1737	0.1897

TABLE A5.20: Endogeneity and selection adjusted QR estimates of the returns to education, Tajikistan 2003

	Males					Females				
	10th	25th	10th	25th	10th	25th	10th	25th	10th	25th
school	0.0421*	0.0659***	0.0791***	0.0982***	0.1059***	0.0506***	0.0882***	0.0870***	0.0577**	0.1752***
	(0.0234)	(0.0156)	(0.0123)	(0.0211)	(0.0298)	(0.0191)	(0.0220)	(0.0213)	(0.0276)	(0.0384)
exp	0.0492***	0.0490***	0.0532***	0.0363***	0.0140	0.0076	0.0213***	0.0271***	0.0263***	0.0293***
	(0.0080)	(0.0055)	(0.0048)	(0.0074)	(0.0107)	(0.0058)	(0.0071)	(0.0071)	(0.0087)	(0.0110)
expsq	-0.0009***	-0.0009***	-0.0009***	-0.0006***	-0.0002	0.0000	-0.0002	-0.0003**	-0.0002	-0.0002
	(0.0002)	(0.0001)	(0.0001)	(0.0002)	(0.0002)	(0.0001)	(0.0002)	(0.0002)	(0.0002)	(0.0002)
1-2 years	-0.1268	0.0378	-0.1759***	-0.0583	-0.3007**	0.0034	0.0965	0.1111	0.0966	-0.0238
	(0.1026)	(0.0752)	(0.0647)	(0.0997)	(0.1346)	(0.0837)	(0.0967)	(0.0989)	(0.1186)	(0.1542)
3-5 years	-0.0136	0.0157	-0.1972***	-0.1280	-0.2029	0.0250	-0.0330	-0.0218	0.0346	0.0368
	(0.0978)	(0.0713)	(0.0614)	(0.0942)	(0.1303)	(0.0818)	(0.0928)	(0.0955)	(0.1153)	(0.1529)
6-10 years	-0.1519	-0.1239	-0.4016***	-0.3324***	-0.3379**	-0.0933	-0.0515	-0.0858	-0.1155	-0.3414**
	(0.1082)	(0.0779)	(0.0669)	(0.1021)	(0.1428)	(0.0876)	(0.1035)	(0.1053)	(0.1244)	(0.1616)
>10 years	-0.2428**	-0.2799***	-0.5430***	-0.5081***	-0.5594***	-0.1439*	-0.3518***	-0.3054***	-0.2931**	-0.5718***
	(0.1039)	(0.0746)	(0.0629)	(0.0936)	(0.1288)	(0.0845)	(0.0980)	(0.1003)	(0.1137)	(0.1539)
public	0.2097**	0.0764	-0.0751	-0.2412***	-0.3687***	-0.0626	-0.2230***	-0.2555***	-0.2429***	-0.6947***
	(0.0821)	(0.0595)	(0.0494)	(0.0930)	(0.1109)	(0.0500)	(0.0605)	(0.0657)	(0.0802)	(0.1438)
GBAO	-0.6712***	-0.5808***	-0.5086***	-0.6109***	-0.3981**	-0.5405***	-0.8387***	-0.9620***	-0.7893***	-0.7367***
	(0.1185)	(0.0864)	(0.0736)	(0.1133)	(0.1558)	(0.0859)	(0.1021)	(0.1042)	(0.1244)	(0.1654)
Sogdian	-0.7166***	-0.5756***	-0.4661***	-0.3612***	-0.3256***	-0.5819***	-0.5463***	-0.6441***	-0.7460***	-0.3178*
	(0.0892)	(0.0653)	(0.0594)	(0.0886)	(0.1173)	(0.0718)	(0.0857)	(0.0944)	(0.1240)	(0.1633)
Kahtlon	-0.9172***	-0.7464***	-0.7897***	-0.7625***	-0.6618***	-0.6800***	-0.9127***	-1.0046***	-1.2260***	-0.8442***
	(0.0829)	(0.0624)	(0.0596)	(0.0913)	(0.1242)	(0.0732)	(0.0853)	(0.0985)	(0.1378)	(0.1899)
RRS	-0.4862***	-0.2971***	-0.1741**	-0.1690	0.0665	-0.7298***	-0.7126***	-0.9237***	-0.7565***	-0.3058
	(0.1064)	(0.0774)	(0.0690)	(0.1040)	(0.1421)	(0.0932)	(0.1073)	(0.1163)	(0.1468)	(0.1954)
$(z'\gamma)$	0.8193	1.1913***	1.4644***	1.8575***	1.8665***	-0.1744	-0.2575	0.0131	0.0900	-1.1081**
	(0.5897)	(0.4264)	(0.3398)	(0.5086)	(0.7224)	(0.2785)	(0.3540)	(0.3361)	(0.1161)	(0.4474)
$(z'\gamma)^2$	-0.2133	-0.4654**	-0.5517***	-0.7239***	-0.7595**	0.0307	0.0023	-0.0670	0.0170	0.3718**
	(0.2734)	(0.1982)	(0.1616)	(0.2467)	(0.3488)	(0.0996)	(0.1411)	(0.1317)	(0.1422)	(0.1559)
r^1	0.0058	0.0092***	0.0065***	0.0066***	0.0102***	0.0022	0.0041	0.0056*	0.0156***	0.0153***
	(0.0045)	(0.0034)	(0.0018)	(0.0023)	(0.0023)	(0.0035)	(0.0030)	(0.0029)	(0.0031)	(0.0031)
r^2	0.0008	-0.0014**	-0.0008	0.0003	0.0006	0.0002	0.0007	0.0015	0.0036**	-0.0001
	(0.0008)	(0.0006)	(0.0005)	(0.0010)	(0.0008)	(0.0007)	(0.0008)	(0.0010)	(0.0015)	(0.0012)
r^3	-0.0001***	0.0000	0.0000	-0.0000	-0.0000	-0.0000	-0.0000***	-0.0000**	-0.0000	0.0000
	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
constant	1.4948***	1.6114***	2.0626***	2.3652***	3.1485***	1.9311***	2.2249***	2.6267***	3.4301***	3.2692***
	(0.4426)	(0.3282)	(0.2767)	(0.4266)	(0.6326)	(0.2709)	(0.3038)	(0.3259)	(0.3957)	(0.6118)
N	3006	3006	3006	3006	3006	1832	1832	1832	1832	1832
Pseudo R^2	0.0942	0.0925	0.0959	0.1077	0.0906	0.0687	0.1245	0.1388	0.1737	0.1940

Notes: Standard errors in parentheses * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$; The residuals of the fourth and fifth order are insignificant and for a reason of space not reported.

ESSAY TWO

THE GENDER WAGE GAP – EVIDENCE FROM TRANSITION COUNTRIES

CHAPTER SIX

GENDER WAGE DIFFERENTIAL - A REVIEW OF DECOMPOSITION METHODS AND EMPIRICAL LITERATURE

6.1. Introduction

The labour economics literature exhibits a long interest and tradition in examination of the gender wage differential. Since the introduction of the Oaxaca-Blinder decomposition method in the 1970s different decomposition methods have been developed that try to shed light on the components of the wage differential. This Chapter provides an overview of the literature related to the different decomposition techniques and associated estimation problems. As there is large literature on the gender wage differential, we concentrate on the commonly used methods. It also introduces the literature that analyses gender wage differential in transition economies.

The Chapter is organised as follows: Section 6.2 discusses what we consider to be evidence on gender wage differential and reviews the difficulties and the problems associated with the decomposition methodology. In Section 6.3 we pay attention to the main decomposition techniques applied in the literature. Our emphasis is on the recent methodological developments of quantile regression techniques. In Section 6.4 we discuss the literature on gender wage differentials in transition economies. Finally, we conclude with a few comments on a research agenda in Section 6.5.

6.2. Gender wage gap and estimation problems

Labour market discrimination exists whenever workers with identical productive characteristics receive different rewards for their attributes because of the population group to which they belong (Valmori, 2008). Economic models of discrimination may be divided into two main classes – competitive models in which agents act individually and collective models in which one group acts collectively against another. Almost all of the theoretical works by economists has been within a competitive framework. These models emphasize two types of discrimination. The first is prejudice, which Becker formalizes as a "taste" by least some members of the majority group against interacting with members of the minority group. The second is statistical discrimination by employers in the presence of imperfect information about the skills or behaviour of members of the minority group (Altonji and Blank, 1999). Wage discrimination means that one group, in our case women, are paid systematically less than another group – men, when both have equal productively related characteristics. Employment discrimination arises when probability of employment for any given level of productivity related characteristics, varies across groups. Occupational discrimination occurs when women are restricted from entering certain occupations and are crowded into others for reasons other than their personal preferences (Paci and Reilly, 2004). The focus of the current research is on the wage discrimination between men and women.

A reason why economists study labour market discrimination is based on an efficiency argument. According to D'Amico (1987) one consequence of labour market discrimination is that it generates clear losses of efficiency, since scarce resources are over allocated to relatively unproductive members of the favoured group (men) and under allocated to more-productive members of the minority group (women). This implies that the society's aggregate output will fall below its potential size. In addition to society loss of efficiency, market discrimination also imposes costs on the members of the discriminated group. These personal loses include a lower per capita real income, poorer living conditions and lower social status relative to the situation that would have existed in the absence of discrimination (Guimarães, 2001).

The question that most researchers are interested in is why do women earn less than men? Several economic theories address the above question such as the human capital theory, the compensating theory and the job matching theory. The traditional approach in analysing the determinants of the wage gap is to consider both the role of gender differences in human capital endowments and labour market discrimination. The human capital theory (Mincer and Polachek, 1974) posits that the earnings of individual workers are a function of their past investment in human capital. The characteristics of men and women in the labour market differ with respect to work experience, the level of education and skills, occupational status and sector of employment. For example gender differences in how long individuals participate in the labour market tend to reflect the household responsibilities of women. Equally, women may be expected to invest less in education because of intended discontinuities in labour market participation. Mincer and Polachek (1974) argue that it is optimal for women to invest less in education than men since women specialize in household production and the rearing of children and therefore envisage labour force participation as discontinuous. Given that each of these characteristics have some association with the level of earnings, it is necessary to distinguish what proportion of the overall pay gap is due to differences in individual characteristics and what proportion is due to sex discrimination within the labour market.

In addition to differences in human capital endowments, labour market discrimination is often identified as a source of gender wage differences. Blau and Ferber (1986) consider that labour market discrimination exists when "two equally qualified individuals are treated differently on the basis of their sex". Thus, in accordance with the theory of labour market discrimination, gender disparities in earnings arise from the unequal treatment of equally productive men and women. Some scholars have focused on gender segregations across occupations as an important factor underlying gender differences in earnings. This type of segregation exists when men and women are employed in different types of occupations: those typically occupied by men and those typically occupied by women (Preston, 1999). The segregation theory suggests that gender differences in pay stem from the fact that female dominated occupations generally pay more poorly than those dominated by males (Boraas and Rodgers, 2003). To explain occupational segregation, Terrell (1992) distinguishes between labour supply and labour demand factors. On the labour supply side, the human capital approach views occupational distribution as a function of occupational choice. The basic idea is that, given their traditional role within the family, women tend to invest less in human capital than men and select occupations on the basis of this role and these investments. On the labour demand side, occupational segregation is explained by employer's discrimination of women in their hiring practices in certain occupations.

Most explanations focus on one of these two themes. The first is Smith's theory of equalizing differences. Women might select themselves into less stressful occupations or pick different career paths than men (Glinskaya and Mroz, 2000).

The starting point of discrimination analyses is to set up a simple wage regression model and to obtain consistent estimates of the parameter of main interest. The underlying economic model is a human capital one. The empirical model is based on the Mincer's earnings function in which logarithmic wages are regressed on measures for individual characteristics: work experience, or time out of work periods, education (pre-labour market schooling) and other background variables. The ways in which different factors contribute to the gender wage gap has been most clearly addressed by Oaxaca-Blinder decomposition technique (Blinder, 1973; Oaxaca, 1973). The Oaxaca-Blinder decomposition uses the coefficients from human capital wage specification to decompose the gender wage gap into a component accounted for by differences in mean characteristics and the component accounted for by differences in the returns to those characteristics (see Section 6.3.1). These two components have been referred to as the 'explained' and 'unexplained 'components, also classified respectively as the 'endowment' and the 'treatment' effects. Researchers take the 'unexplained gap' – the difference in wages after controlling for a host of personal and job characteristics in wage regressions, as evidence of discrimination. While the presence of unexplained differences in male/female is certainly consistent with the presence of discrimination, it does not provide a very direct test of the hypothesis. If discrimination affects human capital investments and personal choices that individuals make or if it is affecting job choice, then the 'unexplained gap' will understate discrimination because some of the control variables themselves reflect the impact of discrimination. On the other hand, the specifications used in many wage regressions are limited and researchers typically have only very crude proxies to measure skills and ability (such as year of education) or experience (such as age minus education). If there are omitted variables that are missing from regressions that relate to the human capital and personal tastes of the individual and those are correlated with individual variables, then the 'unexplained gap' will overstate the impact of discrimination; since it will reflect both the impact of omitted and unmeasured productivity variables as well as any effects of discrimination (Altonji and Blank, 1999).

Many studies examine the role of differences in years of education on the gender gap using standard regression techniques. Among younger workers, there is no longer any difference in average years of education between men and women, although older women continue to have lower average education levels (Blau and Kahn, 1997). As male/female education levels have converged, this has narrowed the wage gap, as confirmed in Blau and Kahn (1997) and O'Neill and Polachek (1993). Some papers have shown that women's gain in education is a major factor behind decrease in relative male/female wage differential. According to Altonjii and Blank (1999), even when women have been improving their relative skills in the labour market, certain aggregate labour market trends have been moving against them. In particular, changes in returns to skill have favoured more skilled workers and lowered the wages of less skilled workers. Since women on average are in less-skilled jobs, these shifts should have lowered the wages of women relative to men. Blau and Kahn (1997) find that among more educated workers, the return to skill have increased more among men than women. On the other hand, among less educated women, the returns to skill have declined less than among less educated men. These changes suggest that it is important to differentiate labour market experience by skill level.

The accumulation of work experience is perhaps the most important factor in the distribution of earnings across workers. The increase of wages is the combined effect of the accumulation of general skills, the return to job seniority that may reflect both worker investments in job specific skills, and incentive devices used by firms. There are a number of reasons to expect gender differences in both the accumulation of and returns to experience. Historically, women have had quite different patterns of labour force participation and job mobility than men. The standard model of human capital investment predicts that investments in general training will be lower for individuals with fewer working hours and less working experience. As women have increased their labour force participation over time, however, women's accumulated labour force experience has also increased.

Many datasets have no information on actual experience and hence researchers use potential experience as a proxy for actual experience. Potential experience is likely to overstate actual experience for women because of the amount of time that women spend out of work force. The major problem with the potential measure is that it correlates poorly with the actual female labour force experience. This may be particularly relevant in the case of estimation of wage regressions for samples of females as well as young workers. These two groups may have in fact working life cycles characteristics by more frequent interruptions (Heather and Bedard, 2004). Given that women are likely to acquire less labour force experience than men through the effects of family formation, a failure to account for these factors may have important consequences. It has been found in the literature that the use of potential experience measure has the effect of increasing or exaggerating the unexplained or discriminatory component. For example, the discrimination effect in Wright and Ermisch (1991), using Oaxaca-type decomposition, falls from 0.35 to 0.23 in moving from the use of potential to actual female labour force experience.

Other studies have attempted to improve upon the Mincer proxy by incorporating some adjustment for the intermittency of female labour participation. Specifically, Zabalza and Arrufat (1985) use the probit coefficients of the structural model of employment status to predict for each female in the sample the probability of being employed in each year from the survey date back to the year she first left full-time education. The approach assumes that the relationship determining an individual's employment status is stable over time (Kidd and Shannon, 1997). Wright and Ermisch (1991) find that the imputed experience measure is highly correlated with the actual labour market experience and the Oaxaca decomposition of the gender wage gap are similar when either actual experience or imputed experience is used.

There are two main themes in the recent research on the role of experience, tenure and job mobility on the gender wage gap. First, a number of studies examine the effect of using more complete measures of actual experience (as opposed to potential) and estimate how much of the narrowing of the gender gap is due to convergence in the actual experience levels of male and female workers. Second, some other papers examine differences in job mobility between men and women. These differences in mobility patterns have been related to differences in job training between men and women.

There is also debate in the literature on the explanatory variables to be included in the earning equation[88]. The general evidence is that the greater number of control variables is, the greater the explained portion and the smaller the unexplained portion or the average adjusted pay gap should be. With the addition of more variables, the pay gap

[88] See Cain (1986) and Blau and Ferber (1986).

associated with discrimination diminishes (Gunderson, 1989). Any supply side variable that may result in differences in wages should enter the earning equation. This includes personal productivity related characteristics, such as ability, education, experience, also employee's preferences for non-pecuniary aspects of the job. In practice choice of control variables is likely to be constrained by their availability in the chosen source of data. For example, direct measures of employee's preferences and supply elasticity's are rarely available. Job characteristics and occupations can be used as proxies for individual differences in tastes, but concentrating women into some jobs and occupations may reflect discrimination, so some studies postulate that such variables may be endogenous and should not be included in the vector of characteristics from which the explained component is calculated. Blau and Ferber (1986) show that including occupational and job related controls could underestimate gender discrimination by masking any discriminatory origins of gender differences in occupation and job-related outcomes. In general decomposition methods can be quite sensitive to the choice of explanatory variables and the inclusion of occupation and sector control variables adds considerably to the explanatory power of the model, generating a much reduced unexplained pay gap. However, several reviews of the approach argue that correcting the gender pay gap for differences in occupational activity using detailed occupational classifications will produce an underestimation of discrimination (Cain, 1986). For instance, Kidd and Shannon (1996) conclude that using 36 occupational categories instead of 17 increases the fraction of the wage gap attributed to occupational choice from 17.8% to 27.3%. For these reasons most studies prefer to use relatively broad classifications of around 6 to 12 categories. The sensitivity of these estimates, however, raises some concerns about the implications of the occupational control variables.

A potential problem with most of these studies is the assumption that occupational attainment is exogenous. As argued by Macpherson and Hirsch (1995), if men and women with higher unmeasured skills are more likely to be sorted into male jobs and those with lower skills into female jobs, then the exogenity assumption is violated. Moreover, there are reasons that unobserved individual characteristics may not only influence worker's decision about which job to take, but they may also affect earnings. Men place greater emphasis on earnings and less on job attributes than do women. This difference arises because of the gender division of labour in the home (Sorensen, 1989).

The main source of endogeneity that the literature on gender wage gap has addressed is the correlation of the unobserved individual specific effects and the regressors in the model. Kim and Polachek (1994) show that the problem remains even after using detailed controls for the differences in human capital and background characteristics. One way to deal with the endogeneity problem is to include instrumental variables in the estimation, and the fixed effects estimators are applied to deal with the unobserved heterogeneity.

Many researchers have further disaggregated the two components of the decomposition into separate contributions of the variables. Jones (1983) demonstrates that Blinder's (1973) method for separating out the contribution of the constant term to discrimination is flawed in the presence of a set of dummy variables, as the magnitude of the estimated constant term depends on the left-out reference group. Conventional decomposition methodology cannot identify the separate contributions of dummy variables to the wage decomposition because it is only possible to estimate the relative effects of a dummy variable and the discrimination component is not invariant to the choice of the "left-out" reference group (Oaxaca and Ransom, 1999). Horrace and Oaxaca (2001) show that the method for estimating gender wage gap by industry proposed by Fields and Wolff (1995) suffers from an identification problem, as its results vary according to the choice of the left out reference group. According to Gardeazabel and Ugidos (2005) failure to identify the contribution of individual dummy variables to the wage discrimination raises one additional problem – it is not possible to compare results from different studies as they may use different left out groups. Trying to overcome this issue, a recent study by Yun (2004) provides a way to apply the Blinder-Oaxaca (Blinder, 1973; Oaxaca 1973) decomposition to a non-linear function for both aggregate and detailed decompositions. The proposed detailed decomposition method does not depend on the functional form as long as the dependent variable is a function of a linear combination of independent variables and the function is once differentiable.

6.3. A review of decomposition techniques

In this section we present a selective review of decomposition techniques and different approaches in the literature. We focus on the methodological aspects of estimating the gender wage gap and do not aim to review the respective estimation results. Our emphasis is also on more recent decomposition techniques of the gender wage gap. The section is organised in three parts. In the first subsection, different decomposition methods concerned with the gender wage gap that looked at the mean earnings are introduced. In the second subsection, decomposition techniques which take into account differences across wage distribution are presented. Semi-parametric and non-parametric decomposition methods are discussed in subsection 6.3.3. The sample selection problem in the gender wage gap is discussed in subsection 6.3.4. Finally, our motivation for the chosen methodology is outlined in subsection 6.3.5.

6.3.1. Decomposition methods at the means of the data

The standard econometric method to account for the gender wage gap is to estimate separate earning equations for males and females, taking the logarithm of their earnings. Few studies explicitly test whether the data supports separation. Male and female equations take the following form:

$$\ln W_m = X'_m \beta_m + e_m \qquad (6.1)$$

$$\ln W_f = X'_f \beta_f + e_f \qquad (6.2)$$

where m stands for male and f for female, $\ln W_m$ and $\ln W_f$ are the levels of natural logarithm of earnings, X'_m and X'_f are vectors of observable characteristics of men and women, β_m and β_f are vectors of parameters.

The differential is then decomposed according to the most standard approach in the literature known as the Blinder-Oaxaca procedure (Blinder 1973; Oaxaca 1973). They express the average gender mark-up in the following form:

$$\ln(\overline{W}_m) - \ln(\overline{W}_f) =$$

$$= \underbrace{\ln \overline{W_m} - \ln \overline{W_f}}_{raw\ wage\ gap} = \underbrace{(\overline{X}_m - \overline{X}_f)\hat{\beta}_m}_{explained\ part} + \underbrace{\overline{X}_f(\hat{\beta}_m - \hat{\beta}_f)}_{unexplained\ part} \qquad (6.3)$$

where the bar indicates mean values. This allows the overall average differential in wages between two groups to be decomposed into a part attributable to differences in measured characteristics and a part attributable to differences in the estimated relationship between men and women (i.e., the gender differences in returns), evaluated at the mean set of female characteristics. The first term on the right hand side of equation (6.3) is a measure of the explained part of the raw wage gap and it is non-zero if the two groups do not have equal characteristics. This part can also be interpreted as the wage gain women would experience if they had the same characteristics, on average, as men. The portion due to differences in coefficients, that is the second term on the right hand side of equation (6.3), is the unexplained part of the average wage gap. It is the wage gain women would experience, given their mean characteristics, if they were remunerated like men. This portion of the differential is defined in Oaxaca (1973) as a measure of wage discrimination. It refers to the difference between how the male equation would value the characteristics of females and how the female wage equation values them in reality. Initially, the Oaxaca-Blinder decomposition method was developed for cross section wage models. However, assuming time constant parameters, application to longitudinal wage models follows straightforwardly.

The wage decomposition suggested by Blinder (1973) and Oaxaca (1973) has been subject to criticism on at least two points. Their method is based on the endowment prices of one of the sexes (the male in most applications), thereby introducing a potential asymmetry in the effects depending on which gender is considered as the reference group, the so called 'index' number problem. In most studies the wage structure of males is assumed as non-discrimination with the arguments that in economy men come from the largest group of workers and therefore face virtually no discrimination. The choice of the male wage structure as a non-discriminatory reference is equivalent to considering discrimination as the disadvantage of any group with respect to the most favoured group.

Although the statistical technique appears to separate out an 'explained' from an 'unexplained' component of the gender pay gap, there are good reasons to believe that the distinction is in fact rather more blurred. This is referred to as the 'feedback effect' problem in the literature. The assumption of the model is that controlling for gender differences in characteristics means that one can produce an estimate of how similar workers would potentially be treated in the labour market and thereby claim that any remaining pay difference is attributable to discrimination.

Depending on which wage structure would prevail in the absence of discrimination, researchers have subsequently introduced different weighting schemes to estimate the discrimination component. Cotton (1988) argues that the non-discriminatory structure should be determined by the larger group, so the weighting factor should reflect the proportion of the largest gender group in the sample. The idea, that discrimination not only lowers the wages of females, but also leads to higher wages for the males, forms the starting point of such methods. Cotton (1988) identifies the separate effects of discrimination of different groups. Derivation of the method builds on Becker's (1957) assumption that in absence of discrimination, the wage differences arise due to differences in productivity characteristics. Neumark (1988) develops an alternative procedure from a particular Beckerian discrimination model. Both the Cotton and Neumark papers generalize methods suggested by Oaxaca (1973).

Oaxaca and Ransom (1994), following Cotton (1988) and Neumark (1988) provide a unifying framework for all previous decompositions. They analyse the empirical consequences of using the female, the male or the Cotton and Neumark wage structure as the non-discriminatory structure. The adoption of the female wage structure as a competitive standard leads to a larger discriminatory and smaller productivity differential than the use of male coefficient estimates. The Cotton and Neumark (pooled) method also yields quite different estimates of the productivity differential. These differences imply a division of the discriminatory wage differential into a male overpayment and female underpayment.

Specifically, the Oaxaca and Ransom (1994) propose using the cross product matrices of the sample characteristics as weights for the estimated parameters $\hat{\beta}_m$ and $\hat{\beta}_f$ as follows:

$$\ln \overline{W}_m - \ln \overline{W}_f = (\overline{X}_m - \overline{X}_f)\hat{\beta}^* + \overline{X}_m(\hat{\beta}_m - \hat{\beta}^*) + \overline{X}_f(\hat{\beta}^* - \hat{\beta}_f) \tag{6.4}$$

where $\ln \overline{W}_m$ and $\ln \overline{W}_f$ are the logarithm of earnings for males and females, $\overline{X}_m$ and $\overline{X}_f$ are vectors with the average characteristics for both genders, $\hat{\beta}_m$ and $\hat{\beta}_f$ are the estimates of relevant coefficients, $\hat{\beta}^*$ is the non-discriminatory coefficient structure obtained from the pooled regression of males and females. The non-discriminatory wage structure is given as $\hat{\beta}^* = \Omega\hat{\beta}_m + (1-\Omega)\hat{\beta}_f$. Oaxaca and Ransom (1994) show that Neumark's solution is the same as their own in the case when the weighing matrix, Ω, is defined as

$\Omega = (X'X)^{-1}(X'_m X_m)$, where X and X_m are the matrices of observed productivity characteristics for the pooled sample and for the male sample.

A number of studies have recently employed a more innovative method for decomposing the gender wage gap suggested by Juhn-Murphy-Pierce (1993). Interest in this approach comes from the claim that gender inequality in pay is due to three factors: differences in productivity-related characteristics, differences in reward to workers with equal characteristics and cross-national differences in the overall shape of the wage structure. Juhn-Murphy-Pierce (1993) (JMP) have proposed a simple extension to the Oaxaca decomposition by taking into account the distribution of the residuals. As in the Oaxaca-Blinder approach the estimated coefficients from the male wage regression are taken as the competitive price vector. The decomposition of the raw wage gap then includes three components related to differences in endowments, in estimated coefficients and in residual wage distribution. The wage differentials can be expressed as:

$$\Delta \ln W = \underbrace{\hat{\beta}_m(\overline{X_\theta^m} - \overline{X_\theta^f})}_{\substack{endowment \\ effect}} + \underbrace{(\hat{\beta}_m - \hat{\beta}_f)\overline{X_\theta^f}}_{\substack{renumeration \\ effect}} + \underbrace{(\overline{e}_m - \overline{e}_f)}_{\substack{unobservable \\ effect}} \tag{6.5}$$

where $\overline{X}_\theta$ represents the mean characteristics for each quantile θ.

The JMP method was applied by various researchers to examine wage structure effects on the difference in wages between two groups of workers over time or by countries. In the case of comparison of gender earnings differential between two countries, the decomposition technique distinguish the effect of gender specific factors from those associated with the underlying wage structure of both economies. The wage equation for a male worker from country j can be defined as:

$$\ln W_j^m = \beta_j^m X_j^m + \sigma_j^m \theta_j^m \tag{6.6}$$

where $\varepsilon_j^m = \sigma_j^m \theta_j^m$ and θ_j^m is country j residual standard deviation of wages. The vector θ_j^m can be interpreted as the standardised unobservable component of productivity. Non-discriminatory returns are assumed for both genders so that: $\beta_j^m = \beta_j^f = \beta$. The male-female log wage gap for country j is given by:

$$D \equiv \Delta \ln W = \beta_j^m (X_j^m - X_j^f) + \sigma_j^m (\theta_j^m - \theta_j^f) = \beta_j \Delta X_j + \sigma_j \Delta \theta_j \qquad (6.7)$$

The gender wage differential between two countries j and k can be defined as:

$$D_j - D_k = \underbrace{\beta_j(\Delta X_j - \Delta X_k)}_{observed\ effect} + \underbrace{\Delta X_k(\beta_j - \beta_k)}_{price\ effect} + \underbrace{(\Delta\theta_j - \Delta\theta_k)\sigma_k}_{gap\ effect} + \underbrace{\Delta\theta_j(\sigma_j - \sigma_k)}_{unobserved\ effect} \qquad (6.8)$$

where D_k is defined in the same way as D_j.

According to the decomposition in equation (6.8) a difference in the gender wage differential D between two countries k and j is decomposed into four components. The first component reflects the contribution of inter-country differences in observed human capital endowments to the gender gap. The second term measures the impact of different prices across countries for observed productivity characteristics. The third term, the gap effect, captures the effect of the international difference in the relative wage position of males and females after controlling for observed human capital endowments. Finally, the unobserved effect measures differences in the return to unobservable skills.

However, there are some potential drawbacks of the JMP decomposition. The first is the strong interpretation of changes in the distribution of male wage residuals. Changes could be interpreted as changes of prices. They may as well capture measurement error, sample composition, equation misspecification and the distribution of unmeasured male productivity characteristics. Second, the use of prices derived from the male wage regression implies that the same set of prices applies to females. Hence, it is assumed that inequality affects men and women equally and the wage structure is, therefore measurable for both men and women by the prices derived from the male sample (Kunze, 2000). Third, Suen (1997) states that the use of the standard deviation of residual wages to measure the price of skills and the average percentile rank to measure the quantity is subject to bias. He argues that the decomposition is unbiased only when percentile ranks are independent of the standard deviation. Fourth, Fortin and Lemieux (1998) show that residual improvements in the relative position of women and the estimated wage structure effects critically depend on which distribution is assumed to be the distribution of reference (hence whether the male, pooled or female sample distribution).

Brown *et al.* (1980) decompose the wage gap across the entire distribution of occupations and allows for endogeneity of the distribution of women across occupations[89]. Their technique goes one step beyond the Oaxaca decomposition by treating average wages as a weighted average of wages across occupations. Thus, while the Oaxaca technique decomposes the aggregate wage differential into two elements (coefficients and endowments), Brown *et al.* (1980) do this but also identify the impact of occupational distribution. The authors argue that differences in the occupational distribution of men and women are an important source of the gender wage differential. Indeed individuals with similar characteristics who have attained different occupations often earns different wages. The main idea is to measure how much of the gap is explained by gender differences in wages within occupations and how much is explained by occupational gender segregation. Brown *et al.* (1980) suggest that by including dummy variables for occupations in a wage regression, previous decomposition methods take differences between the occupational distributions of men and women as given and thereby ignore the inherent potential discrimination.

In this case the gender wage gap can be rewritten as the difference in the weighted average log-wages taken across K occupations as follows:

$$\ln \overline{W}^{M} - \ln \overline{W}^{F} = \sum_{j=1}^{K} P_j^M \ln \overline{W}_j^{M} - \sum_{j=1}^{K} P_j^F \ln \overline{W}_j^{F} \tag{6.9}$$

where P_j^M, P_j^F are the portions of men and women in occupation j, $j = 1,\ldots, K$,

and $\ln \overline{W}^M, \ln \overline{W}^F$ are log mean wages within occupation j;

Equation (6.9) is extended by adding and subtracting the term $P_j^F \ln \overline{W}_j^M$

[89] Various empirical studies have applied Brown *et al.* (1980) approach, providing estimates of within occupation and across occupation wage disparities relative to different economies (e.g. Dolton *et al.*, 1994; Kidd and Shannon, 1996; Meng and Miller, 1995).

$$\ln\overline{W}^{M}-\ln\overline{W}^{F}=\underbrace{\sum_{j=1}^{K}(P_j^M-P_j^F)\ln\overline{W}_j^M}_{Inter-occupational}-\underbrace{\sum_{j=1}^{K}P_j^F(\ln\overline{W}_j^M-\ln\overline{W}_j^F)}_{Intra-occupational} \tag{6.10}$$

The first term of the right hand side of equation (6.10) measures the part of the gap due to differences in occupational distribution between male and females. The second term indicates the part of the gap that is due to differences in mean wages within occupations. If the portion of men and women were the same in each occupation, which is the non-segregation case, the first term would be equal to zero.

Let the wage W_{ij} of individual i in occupation j be expressed as:

$$\ln W_{ij}=\alpha_j+X_{ij}\beta_j+\varepsilon_{ij} \tag{6.11}$$

where W_{ij} represents the characteristics of the individual, α_j and β_j are parameters to be estimated, and ε_{ij} is a random error term. If estimates of α_j and β_j are obtained using OLS, then the mean wages in occupation j are: $\ln\overline{W}_j=\hat{\alpha}_j+\overline{X}_j\hat{\beta}_j$

and the mean wages across all occupations are: $\ln\overline{W}=\sum_j(P_j\hat{\alpha}_j+P_j\overline{X}_j\hat{\beta}_j)$

where P_j is the fraction of employees in occupation j.

Then, the wage decomposition proposed by Brown *et al.* (1980) can be written as:

$$\begin{aligned}\ln\overline{W}^{M}-\ln\overline{W}^{F}=&\underbrace{\sum_j P_j^F(\hat{\alpha}_j^M-\hat{\alpha}_j^F)}_{I}+\underbrace{\sum_j P_j^F\overline{X}_j^F(\hat{\beta}_j^M-\hat{\beta}_j^F)}_{WD}+\underbrace{\sum_j P_j^F\hat{\beta}_j^M(\overline{X}_j^M-\overline{X}_j^F)}_{PD}\\&\underbrace{+\sum_j\overline{W}_j^M(P_j^M-\hat{P}_j^F)}_{QD}+\underbrace{\sum_j\overline{W}_j^M(\hat{P}_j^F-P_j^F)}_{OD}\end{aligned} \tag{6.12}$$

where $\hat{P}_j^F$ represents the portion of women in the sample who would be in occupation j if women faced the same occupational structure as men. The components *I* and *WD* are defined as unjustified differences in within-occupation wages, *PD* is the justifiable within occupation wage differential and *QD* and *OD* represents the justifiable and unjustifiable portions of occupational segregation (Brown *et al.*, 1980). The terms that are justifiable capture the wage differentials due to differences in characteristics between men and women, while those which are unjustifiable capture the wage differential which is unexplained and is normally attributed to discrimination.

Decomposing the term due to the inter-occupational effect, Brown *et al.* (1980) use the predicted distribution of women across occupations in the absence of discrimination based on the non-discriminatory occupational structure for women ($\hat{P}_j^F$). This requires a model of occupational attainment[90] to be estimated. Brown *et al.* (1980) specify a reduced form multinomial logit model to capture how variables affect the probability of individual i working in occupation j. This probability may be defined as:

$$P_{ij} = prob(y_i = oc_j) = \frac{e^{x_i\beta_j}}{\sum_{k=1}^{K} e^{x_i\beta_j}} \qquad i = 1,...,N, \quad j = 1,...,K \tag{6.13}$$

where N is sample size, J is number of occupational groups, and x_i is a vector of exogenous variables affecting supply and demand factors. Estimates of the parameters of this model are obtained for male observations and the female data and are then substituted into the estimated model producing for each woman a vector of predicted probabilities of belonging to each occupation. These predicted probabilities of being in each occupation are summed over observations to produce the predicted occupational distribution of women $\hat{P}_j^F$.

This approach supposes that in a world without discrimination women would be distributed across occupations according to the male occupational mechanism. Both terms could be decomposed further into an explained and an unexplained component as follows:

$$\ln \overline{W}^M - \ln \overline{W}^F = \overbrace{\underbrace{\sum_{j=1}^{K}(P_j^M - \hat{P}_j^F)\overline{X}_j^M \hat{\beta}_j^M}_{Inter-occupational} + \underbrace{\sum_{j=1}^{K} P_j^F(\overline{X}_j^M - \overline{X}_j^F)\hat{\beta}_j^M}_{Intra-occupational}}^{\substack{Explained \\ part}} \tag{6.14}$$

[90]The definition of occupational attainment refers to a model for explaining how individuals achieve certain occupational levels. This differs from the sociological models of status attainment. Brown *et al.* (1980) define an individual occupational attainment as a function of employer's willingness to hire that person and of individuals desire to work in specific occupation. Willingness of employer to hire an individual depends on personal qualifications such as education, training and experience.

$$\overbrace{+\underbrace{\sum_{j=1}^{K}(\hat{P}_j^F - P_j^F)\bar{X}_j^M\hat{\beta}_j^M}_{Inter-occupational} + \underbrace{\sum_{j=1}^{K}P_j^F(\hat{\beta}_j^M - \hat{\beta}_j^F)\bar{X}_j^F}_{Intra-occupational}}^{\substack{Unexplained\\part}}$$

where $\hat{P}^F$ is the vector of predicted proportions of women in occupations j in the male occupation outcomes model; $\bar{X}_j$ are the mean values of characteristics for occupation j; $\hat{\beta}_j$ are the estimated wage equation coefficients for occupation j;

An alternative to the multinomial logit model discussed above is the ordered probit model (see McKelvey and Zavoina, 1975). The advantage of a multinomial logit model is that it fits the data better than the ordered probit model, when fit is measured in terms of the success at predicting membership of occupations (Meng and Miller, 1995). On the other hand, the ordered probit model has fewer parameters to be estimated. In this model the ordering among outcomes of the categorical variable is significant, whereas the multinomial logit model ignores this information. However, Meng and Miller (1995) compare decomposition based on the multinomial logit and ordered probit models and show that their findings are not sensitive to the chosen model. Miller and Volker (1985) further argue that neither the multinomial logit model nor ordered probit model is necessarily superior.

6.3.2. Accounting for distributional effect

Most empirical studies concerned with the gender wage gap analyse mean wage differences which fails to take into account differences across the wage distribution. Earning differences are very complex and vary over the wage distribution. As discussed in Chapter five, we have to consider that the wage structure is not constant over the range of wages (e.g. Buchinsky, 1994) and education may have different effect on wages of individuals at the top of the wage distribution than on wages of individuals at the bottom of the distribution. In other words more educated individuals may experience more unequal wage distribution. If more educated workers experience greater wage spreads, increased educational levels may also contribute to an increase in wage inequality. A much richer story about the role of gender in the labour market emerges once we move away from an exclusive focus on outcomes for "average" men and women. In particular, the presence of a 'glass ceiling' leads us to expect the gender wage gap to be larger amongst workers earning relatively high wages, while the existence of 'sticky floors' might suggest the opposite[91]. The extent to which disparity in men's and women's productivity-related characteristics accounts for the gender wage gap also appears to differ between high- and low-wage workers (Barón and Clark, 2008).

The QR technique introduced by Koenker and Bassett (1978) extends the notation of OLS in a location model to a more general class of linear models in which the conditional quantiles have a linear form. Kuhn (1987) points out that conventional mean regression has its limitations in measuring discrimination. The distributional approach provides more comprehensive information. He shows empirically that U.S. women at higher wage levels are more likely to be discriminated against. Buchinsky (1994, 1998) further advances the application of QR in the U.S. labour market in the context of wage estimation and returns to education[92]. In contrast to the OLS approach, the QR procedure is less sensitive to outliers and provides a more robust estimator in the face of departures from normality (Koenker and Bassett, 1978). QR models may also have better properties than OLS in the presence of heteroscedasticity[93].

Useful features of the QR model can be summarized as follows: 1) the model can be used to characterize the entire conditional distribution of a dependent variable given a set of regressors; 2) the QR model has a linear programming representation which makes estimation easy; 3) the QR objective function is a weighted sum of absolute deviations, which gives a robust measure of location, so that the estimated coefficient vector is not sensitive to outlier observations of the dependent variable; 4) when the error-term is not normal, the QR estimates may be more efficient than least squares estimates; 5) potentially different solutions at distinct quantiles may be interpreted as differences in the response of the dependent variable to changes in the regressors at various points in the conditional distribution of the dependent variable (Buchinsky, 1998); 6) the method allows to deal explicitly with the heterogeneity problem, which is common in estimating wage functions (for more detail see Bushinsky (1994) and Chamberlain (1994)). For instance, if some coefficients are changing with the quantiles then this is indicative of some form of heteroskedasticity. Therefore in the quantile

[91] See Booth *et al.*, 2003; Aralampalam *et al.*, 2007.
[92] See our discussions in essay one.
[93] See Deaton (1997).

regression model we can address the heterogeneity of the unobservable effects in an informative and constructive way. A disadvantage of quantile regression is its possible lack of consistency and monotonicity of estimated conditional quantile functions, i.e. the predicted p^{th} conditional quantile can be smaller than the predicted $(p+1)^{th}$ quantile (Buchinsky, 1998).

Many authors have developed alternative methods to decompose the entire wage distribution and to describe differentials at different quantiles taking into account differences in human capital and other observed characteristics. Several popular ones include a reweighting methods, which essentially generates a counterfactual wage distribution (DiNardo *et al.*, 1996); another approach is based on conditional quantile regression and re-sampling (Machado and Mata, 2005; Autor *et al.*, 2006; Melly, 2005, 2006); and an alternative approach uses a semi-parametric hazard functions to obtain the conditional densities of wages (Donald *et al.*, 2000).

DiNardo *et al.* (1996) extend the Oaxaca (1973) analysis in order to estimate counterfactual wage densities by using kernel methods applied to weighted samples. The authors analyse the impact of workers attributes and labour market institutions at different points of the wage densities for different periods. Their counterfactual analysis is based on the question 'what density would have prevailed if individual attributes had remained at their 1986 level and workers had been paid according to the wage schedule observed in 1995'? By comparing this counterfactual density of wages with the actual one, DiNardo *et al.* (1996) show the role of individual attributes on the observed changes in the wage distribution over the period. A drawback of the DiNardo *et al.* (1996) approach is that the use of kernel density estimation to the weighted sample does not allow one to isolate the effect of changes in individuals attributes from the returns to those attributes. Guimarães (2001) argue that DiNardo *et al.* (1996) apply different weights to the distribution of the covariates in order to estimate the counterfactual densities, but no return to individual covariates is estimated.

Albrecht *et al.* (2003) find that while the average gender wage gap is relatively small in Sweden, the gap increases throughout the wage distribution and is larger in the upper tail. They also show an increasing impact of education across the wage distribution and conclude that the 'glass ceiling' effects limit the earnings potential of women in the upper part of the wage distribution. Moreover, the wider unexplained gap at the top of distribution mainly holds for highly educated workers (De la Rica *et al.*, 2008). Conversely, the wage gap is sometimes larger at the bottom of the distribution for low educated workers. An analysis of mean differences between male and female earnings is limited because it could lead to the conclusion that the gender wage gap is of minor importance and that the size of the wage gap is constant throughout the wage distribution. Furthermore, the traditional approach is based on the assumption that the importance of explaining factors do not vary with the wage rate (Albrecht *et al.*, 2003).

There are good reasons to believe that male and female wages along the whole wage distributions are not equally affected by human capital and institutional settings. Garcia *et al.* (1998) and Mueller (1998) decompose the wage difference in the Oaxaca-Blinder manner using the QR estimates at average characteristics. Garcia *et al.* (1998) consider decomposition of the Spanish gender wage gap at a given conditional quanitle evaluated at the unconditional means of a vector of explanatory variables. However, it is not clear whether it is the right strategy to use average characteristics to decompose the wage gap at different quantiles given that characteristics also vary over the distribution. In this context, Gardeazabal and Ugidos (2005) state that it is more appropriate to measure discrimination at unconditional quantiles. They show that it might be considered to weight the difference in return to a certain characteristic (for example primary education) at a given quantile according to the proportion of individuals with this characteristic at that quantile. Based on this methodological approach, their findings for the Spanish wage gap contradict the results of Garcia *et al.* (2001). Bonjour and Gerfin (2001) use hazard function approach proposed by Donald *et al.* (2000) to decompose the wage gap in Switzerland. They find that the gender wage gap is not constant throughout the wage distribution and is larger at the bottom of the distribution.

Several recent papers use the Machado and Mata (2005) QR decomposition technique to analyze the gender gap in log wages across the distribution. This technique allows one to decompose the difference between male and female log wage distributions into a component due to the difference in the distribution of observable characteristics between genders and a component due to the difference in the distribution of rewards to these characteristics between genders. Such studies include Albrecht *et al.* (2003) for

Sweden, De la Rica *et al.* (2008) for Spain, Arulampalam *et al.* (2007) for several European countries.

The Machado and Mata (2005) (MM) method can be viewed as a generalization of the Oaxaca-Blinder decomposition which decompose differences between groups in average outcomes into differences in average characteristics and differences in rewards to those characteristics. The MM method is designed to simulate counterfactual distributions. For example, what would the distribution of full-time log wages for women have been if working women had the same distribution of labour market characteristics as men or if their human capital characteristics were rewarded the same as men? The general idea of the MM method is parametric model estimation for the quantiles of the conditional distribution. Specifically, the idea is to generate two counterfactual densities. The first is the female log wage density that would arise if women were given men's labour market characteristics, but were paid according to the female wage structure. The proposed method is based on the estimation of marginal wage densities implied by counterfactual distributions for some or all workers' measured characteristics. The counterfactual decomposition of changes in the wage density is obtained by means of QR and re-sampling methods. The MM approach can be seen as an extension of DiNardo *et al.* (1996) and Blau and Kahn (2006).

Recently Melly (2006) proposes a numerically identical to the MM estimator based on the construction of a counterfactual distribution of female wages that would have prevailed if women had been endowed with their own characteristics but were paid like men. The conditional distribution is integrated over the range of covariates to obtain an estimate of the unconditional distribution. Melly's approach can be classified as semi-parametric as he assumes that the conditional quantiles satisfy a parametric restriction but no distributional assumption is needed and the covariates are allowed to influence the whole conditional wage distribution.

There are some limitations and problems in measuring discrimination from a distributive point of view. In some cases, problems arise from the conceptual confusion of the distributive aspects of measurement with the distributive effect of discrimination. Also, there are problems that arise in using counterfactual distribution functions in wage discrimination estimation. For instance, Del Rio *et al.* (2008) provide an illustration of the problem when counterfactuals are used, and show that when moving from the observed wage distribution to the wage distribution without discrimination some female workers may change their relative position. This could imply that any earnings differential, evaluated at each quantile, would not show the true differences in discriminatory experiences of female workers.

The decomposition methods suggested by Fortin and Lemieux (1998) and Donald *et al.* (2000) share the same shortcoming as Machado and Mata (2005) in that they do not provide a way of dividing up the composition effect into the contribution of each individual covariate. Firpo *et al.* (2007) propose a two-stage procedure to perform the Oaxaca-Blinder type decomposition on any distributional measure and not only the mean. The first-stage in their approach consists of decomposing the distributional statistic of interest into a wage structure and a composition component using a reweighting approach, where the weights are parametrically or non-parametrically estimated. This first-step of the Firpo *et al.* (2007) decomposition is semi-parametric because it does not assume any functional form for the earnings distribution. The idea of the first-stage is similar to the DiNardo *et al.* (1996). Firpo *et al.* (2007) also provide the assumptions required for identification in the case of other distributional statistics besides the mean. In the second stage, they divide the wage structure and composition effects into the contribution of each covariate using the recentered influence function (RIF) method. The recentered influence function is a linear approximation to the nonlinear function of distributional statistics of interest such as a quantile. This allows identifying the specific characteristics, differentiated across men and women, which lead to widening of the gender earnings gap. The RIF method generates unconditional quantile estimates, while the commonly used QR gives conditional quantile estimates. The advantages of the RIF method are twofold: first unconditional quantiles are usually of real interest in economic applications; second, this approach allows one to estimate the marginal effects of explanatory variables on the targeted unconditional quantiles (Chi and Li, 2008). Central to the RIF's unconditional quantile method is an influence function. The influence function represents the influence of an individual observation on a distributional statistic of interest such as quantile (see Firpo *et al.*, 2007).

6.3.3. Semi- and non-parametric estimation methods

The standard methods for estimating empirical models often assume that the functions of interest and the distribution of unobserved random variables are normally distributed. Such assumption simplifies estimation and statistical inference, however, such inference based on convenient but incorrect assumptions about functional forms and distributions can be highly misleading (Horowitz, 2009). Recently, semi-parametric and non-parametric techniques have been extensively used in a number of empirical applications[94]. One merit of this procedure is the perception that the conditional mean may not provide an adequate account of the impact of gender on the log wage. In particular if large outliers exist, the median might provide a more accurate reflection of central tendency in the data than the mean[95].

Breunig and Rospabe (2007) find that the semi-parametric analysis illuminates several features of the male-female wage gap in France, which are not evident from a parametric analysis. They use the methodology developed by DiNardo *et al.* (1996) to examine the density of wages. Applying nonparametric kernel density estimation they construct counterfactual density that would prevail for women if they had men's distribution of characteristics. They find that there are important differences in the shape of the densities of male and female wages. Proportion of females in the low part of the wage distribution is more than twice as great as for males. The authors conclude that the effect of characteristics in explaining the wage gap is strikingly different at different points in the distribution.

A recent paper by Barsky *et al.* (2002) proposes a non-parametric alternative to the Blinder-Oaxaca method that reweights the empirical distribution of the outcome variable using weights equalizing the distribution of explanatory variables. The technique is similar to the method developed by DiNardo *et al.* (1996). The non-

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[94] See Buchinsky (1994) and Gosling *et al.* (2000).

[95] Koenker and Basset (1978, 1982) and Powell (1983) have demonstrated the consistency and asymptotic normality of the estimated coefficients, providing the statistical basis for the use of conventional test statistics. One important point is that precision of the parameter estimates in the QR model is dependent on the density at each quantile. Specifically, at quantiles located either at the bottom or at the top end of the distribution, where the density of data points is low, the coefficients are more difficult to compute. In addition, the corresponding test statistics have less statistical power, and the null hypothesis when false may not be rejected as often as it should. The asymptotic formula for the computation of the variance-covariance matrix as developed by Koenker and Basset (1982) is known to understate the true variance-covariance matrix in the presence of heteroskedastisity, and the more conventional approach adopted to compute the matrix is through a bootstrapping method.

parametric approach permits investigation of the role of earnings in accounting for the wage gap at multiple points along distribution. The advantage of their method is that it provides a robust alternative to the standard parametric decomposition, avoiding problems of specification error and uncertainty of functional form.

6.3.4. Selectivity bias in the decomposition analysis

As argued by Olivetti and Petrongolo (2006), if selection into employment is non-random, then it makes sense to worry about the way in which selection may affect the resulting gender wage gap. In particular, if women who are employed tend to have relatively high-wage characteristics, low female employment rates may become consistent with low gender wage gaps simply because low-wage women would not feature in the observed wage distribution (Dolton *et al.*, 1989). Many empirical studies that control for selectivity bias in the gender wage differential follow the Heckman parametric procedure[96], which assumes bivariate normality in the relationship between the error terms in the earnings and participation equations and which depends crucially on this distributional assumption.

However, two substantial problems exist in the Heckman procedure. The first problem is the requirement for an adequate set of instruments to identify the selection effects. We need a variable that affects the participation decision but not the level of wages. This is a recurring problem in the literature and many cross-sectional data sets may be poorly endowed with necessary instruments. The type of instruments used in the literature relate to dependent children (their number or age) and other household-level measures. Nevertheless, the correction mechanism has been the subject of criticism, given the sensitivity of the selection estimates to the identifying restrictions used (Manski, 1989). The second problem relates to the parametric assumption of normality in the construction of the selection term. A violation of normality assumption introduces a miss-specified selection term and this has implications for any inferences offered on the estimated selection effects.

[96] See Arabsheibani and Lau (1999) and Dolton and Makepeace (1987).

The literature that deals with correction for selectivity bias in QR models is less well developed than for the mean regression model. Moreover, it is surprising that, apart from Albrecht *et al.* (2009) and Picchio and Mussida (2010), Badel and Pena (2010) and Chzhen and Mumfrod (2011), there are no studies that correct the gender gap across the wage distribution for the selection due to workforce participation. Buchinsky (1998) for instance, exploits the work of Newey (1992) on the mean regression in approximating the selection term by using a power series expansion of the inverse Mill's ratio of the normalized index. This is designed to treat the selection effect as an omitted variable that can be proxied by a polynomial expressed in terms of the conventional inverse Mills ratio. These high order terms are then included in the QR models and the sampling variance of the estimates are obtained through bootstrapping. Thus the first order approximation will be sufficient if the error term is normally distributed. In any case the estimator is consistent since the order of approximation increases with the sample size.

Albrecht *et al.* (2009) estimate gender wage gap in the Netherland having corrected for selection of women into market work following Buchinsky's (1998) semi-parametric approach. They conclude that were all Dutch women working full-time, the gender wage gap would be much higher. The selection corrected gender wage gap is found to be predominantly related to women receiving lower rewards for their characteristics than men. Similarly, Beblo *et al.* (2003) show selection-corrected wage gaps for Germany and conclude that selectivity correction may have a significant impact both on wage estimates and on pay gap decompositions. Nicodemo (2009) decomposes the wage gap between husbands and wives across the earning distribution after allowing for self-selection of married women into employment in five Mediterranean countries (Spain, France, Greece, Italy and Portugal). She finds a substantial wage gap in each country, with the greatest portion being due to differences in rewards, and that sticky floors are more predominant than glass ceilings. Badel and Pena (2010) find that self-selection into the labour force in Columbia is crucial for the wage gap estimates – if all women participate in the labour force the observed gap would be roughly 50% larger at all quantiles. After allowing for selection into full-time employment by British women, Chzhen and Mumford (2011) find a substantially larger gender earnings gap – the selection corrected gap is found to be twice the raw gap and is related to women receiving lower rewards for their characteristics than men. Recent study by Picchio and

Mussida (2010) propose a new semi-parametric estimator of densities in the presence of sample selection and finds that when sample selection is taken into account the gender wage gap in Italy widens, especially at the bottom of the distribution. By contrast, Breunig and Rospabe (2007) conclude that correcting for sample selection does not change their results. Selection only contributes to narrowing the observed gender wage gap in France by about 1.5% and this change is not significant.

Huber and Melly (2011) criticize Buchinsky series estimator. Their preliminary results show that the series estimator is not consistent when there is heteroskedasticity. One implication of the Buchinsky (1998) independence assumption is that all quantile regression curves are parallel, which restricts the usefulness of considering several quanitle regressions. In the majority of the applications of quantile methods, researchers are particularly interested in the heterogeneity of the coefficients across the distribution. In addition, under the independence assumption the quanitle slope coefficients will be identical to the mean slope coefficients. Huber and Melly (2011) show that the quantile coefficients are useful to test the independence assumption, although they are inconsistent when this assumption is not satisfied. They also show that Buchinsky's assumptions were wrong and the estimator is inconsistent. However the authors do not suggest any alternative estimator. They only propose a test for the independence assumption for the identification of sample selection models.

6.3.5. Choosing decomposition methodology

Given our data, we are able to cope with some but not all of the methodological issues discussed above. In line with the more recent work on gender wage differentials, we will adopt the QR method as it allows us to estimate the marginal effect of the covariates on the wages at various points of the distribution, instead of only at the mean. We will decompose the gender wage gap across the distribution, as proposed by Machado and Mata (2005) and taking into account the 'index' number problem suggested in Neumark (1988) and Oaxaca and Ransom (1988 and 1994). Their empirical findings show that the use of female wage structure as the competitive standard produces a larger discriminatory differential than the use of male coefficient estimates. From our point of view, the most important is that the estimation method chosen allows us to take account of differences across the wage distribution and the non-discriminatory wage structure from the pooled sample of males and females. Although, we found Brown *et al.* (1980) decomposition method appealing because of its focus on occupational segregation and ability to provide within-occupation wage differentials, it has been difficult to find conclusive examples in the literature that the method is applicable across the earnings distribution.

Finally, in the quantile regression framework little consensus exists on the appropriate sample selection procedure and as discussed above the evidences of the effect of selection on the observed wage gap are mixed. Some studies claim that selection is crucial for the gender wage gap. Others conclude that the male/female wage disparities in the selection process are not large enough to explain the gender wage gap. Generally, researchers who find no obvious selection biases in the mean regression tend not to consider the issue in the quantile regression framework. As no significant selection effect was detected in the case of Bulgaria and Serbia (see Essay one) and to be consistent in our cross-country analysis, we leave the selectivity issues to future research.

6.4. Gender wage gap in transition – review of the literature

In accordance with the official policy of gender equality, pay differentiation based on gender was restricted under central planning in ex-communist economies. As we have already discussed in Chapter three, wages were set according to industry specific wage grids varying only with the difficulty of the job and with the worker's education and experience and not with gender (Münich *et al.*, 2005). The fall of communism in transition countries ended the wage regulation, which increased returns to education but also caused an increase in wage dispersion (Svejnar 1999). In most countries, the transition process leads to a decrease in the mean wage and an asymmetric change in the tails of the wage distribution. Real wages at the bottom decile of the wage distribution in each transition economy decreased substantially while real wages at the top decile of the wage distribution decreased relatively less[97] (Skoufias, 2003). Today, wage grids and restraining gender earnings differential are used only in the 'budgetary' sector (public administration, health, education) and room for pay discrimination is more open in the unregulated non-budgetary sector (Jurajda 2003). Svejnar (1999) provides an

[97] In some countries such as Hungary and Croatia, the real wages of highly paid workers increased.

overview of labour markets in Central and East European countries during transition. According to all these studies, transition economies show significant gender wage differential and one important determination of the differences is the degree of gender segregation by job and occupation. Further reasons have been provided to explain why the gender wage gap exists in transition countries and why it widens at the bottom and/or at the top of the wage distribution. It has been argued that women, especially those at the bottom of the wage distribution, might be perceived to have or might have a smaller work force attachment (Booth and Francesconi, 2003). Evidence from early transition periods suggests that women in some transition economies have actually improved their economic position relative to men despite the rise in overall earnings inequality (Rutkowski, 1996). Others point out these structural shifts toward a market economy, involving the shift from heavy to light industry and expansion of the service sector, were expected to simulate the demand for female labour (Einhorn, 1991).

In terms of the gender pay gap, the Soviet Union was one of the most heavily researched. Empirical studies from the post-Soviet era shows that, on average women continued to earn less than two-thirds of the wage of their male colleagues. A study by Ofer and Vinokur (1992) uses emigration data from the 1970s and finds that 51% of the wage differential is due to discrimination. Similarly, Katz (1999) uses data from the Soviet city of Taganrog for 1989 and 1992 and finds that the ratio of female to male hourly wages is 75% and that 84.6% of the difference on the hourly wage gap is due to discrimination. Newell and Reilly (1996) use the first round of the Russian Longitudinal Monitoring Survey (RLMS) and report that the women's hourly wages are 30% less than men's wage and that 88% of total hourly wage gap between men and women is due to differences in the estimated parameters. The authors note that almost all the overall gender pay gap is unexplained by the model and they ascribed most of the unexplained gap to unequal gender treatment within a set of one-digit occupational groups.

Compared to the Newell and Reilly's (1996) results, Arabsheibani and Lau (1999) find lower levels of discrimination against women. Their empirical findings propose that 59% of the difference in the wage gap is due to discrimination. The difference in their results can be explained by their use of the 1994 wave of RLMS and sample selection correction. Similarly, Ogloblin (1999) examines the Russian gender pay gap using rounds 1994 to 1996 of the RLMS. It is difficult to compare their results with the earlier work of Newell and Reilly (1996), as the later study includes industry-level information. Ogloblin (1999) also adjusts the earnings measure for employment status and corrects for selectivity bias associated with the wage arrears. He computes the monthly gender pay differential at 39 % and less than one-fifth is interpreted as an unexplained part. However, the author concludes that over 90% of the corrected gender wage gap is attributable to gender segregation.

Brainerd (1998) explore the gender pay gap in early transition Russia. The author reports that in 1991 the earnings of men are 24.8% higher than the earnings of women with similar characteristics and that the differential increases to 44.6% in 1994. According to her findings, this is due to the increase in overall wage inequality and that gender-specific factors appear to explain little of the poor labour market outcomes of women in Russia's transition. Glinskaya and Mroz (2000) use the RLMS to re-examine the gender pay gap in urban areas. The raw estimates for 1993 and 1994 are comparable to those reported in Brainerd (1998) and the authors concludes that observable differences in characteristics between the gender groups explains almost none of the differential and the changes across time. They argue that most of the changes are attributable to changes in upper tail of the male wage distribution. Newell and Reilly (2001) find similar results and indicate that most of the hourly gender pay gap is not attributable to gender differences in characteristics. Using quantile regression methods, Newell and Reilly (2001) find higher gender pay gap at the top end of the conditional wage distribution than at the bottom end.

In another study, Gerry *et al.* (2002) examine the impact of wage arrears on the gender pay gap using RLMS over the period 1994 to 1998. The authors pool all rounds to enhance efficiency of their estimates. They report 29% average hourly wage differential between men and women, which they attribute to treatment differentials rather than gender differences in endowments. According to their results, wage arrears and payments in kind attenuated the gender pay gap and this effect is particularly pronounced among the lowest paid. Ganguli and Terrel (2005) test the hypothesis that the gender gap is lower at the beginning of transition process than in today's Ukrainian economy, because of the egalitarian principles of a communist state. The authors find that the gender pay gap is higher in the top half of the distribution than at the bottom half and that a glass ceiling is persistent across all three points in time. Using the

Machado and Mata (2005) decomposition method and data from the ULMS, they conclude that differences in rewards rather than differences in productive characteristics explain most of the gender wage gap throughout the distribution.

Further evidence of the gender wage differential in transition economies was found by Krugman (1998) and Arabsheibani and Madirimov (2002) for Uzbekistan, and Namazie (2002) for Kyrgyzstan. According to their results discrimination explains a large proportion of the gender wage gap in Central Asian Republics. Newell and Reilly (2001) report an increase of the wage gap between the 90th and 10th quantiles in 1996 for Kazakhstan. The authors also show that the wage differential is higher at the top end of the conditional wage distribution than at the bottom end in Kazakhstan, Ukraine and Russia.

In the case of Bulgaria, Brainerd (2000) notes that a relatively decentralized wage-setting system has evolved despite national agreements to establish wage floors and ceilings. According to Jones (1991) the large increase in the dispersion of wages in 1989 was due to new legal agreements, i.e. the new system of basic wages that gives managers more control over determining the wages of their employees. Moreover, Jolliffe (2002) uses Bulgarian Integrated Household Survey from 1995 and indicates that males' wages are 24% higher than females' wages. An Oaxaca decomposition of this differential shows that differences in characteristics, such as education, experience and sector of employment, explain very little of the wage differential. Using 1995, 1997 and 2001 Integrated Household surveys, Newell and Reilly (2001) find that discrimination effects in Bulgaria are significantly higher at the 90th percentile. Giddings (2002b) analyses the gender differential in Bulgaria's early transition, using survey data from 1986 and 1993. Her results indicate that the gender earnings gap decreased in the early years of the Bulgaria's transition because women were better educated and were disproportionately represented in growing industries such as commerce and services. Using the Integrated Household Survey from 2001, Kolev (2002) also provides gender wage gap estimate for Bulgaria of 26 % at the 50th quantile of the conditional wage distribution.

The Croatian gender wage gap is examined by Nestic (2010). He utilises data from the Labour Force Survey in 1998 and 2005 and applies both OLS and QR techniques to assess gender wage gap across the wage distribution. The gender wage gap is found to be relatively low at the lower part of the wage distribution and gets larger as one move towards the top of the distribution. Finally, Kecmanovic (2009) analyses the gender wage gap in Serbia and finds a decrease in the wage differential after 2001, accompanied by an increase in the 'unexplained' component of the gap. He also find that the gender wage gap between 2001 and 2005 falls at each of the quantiles of the wage distribution, but discrimination that women face in the labour market has worsened during the examined period.

The review of the literature that examines changes in the wage position of women over the transition period has been inconclusive[98]. There is no consensus within the literature that gender wage gap either remained stable or revealed a downward trend over transition period. In Table 6.1 we attempt to summarise the main findings regarding the changes in the gender wage gap over time in transition countries. For example, Brainerd (2000) finds that during the first years of transition, the gender wage gap in Hungary, Poland, Czech Republic and Slovak Republic decreased between 0.05 and 0.12 log points. The author also finds an increase in the gap of 0.15 and 0.27 log points for Russia and Ukraine. In contrast to Brainerd (2000), Ganguli and Terrel (2005) find that the raw gender wage gap in Ukraine declined from 0.40 log points in 1986 to 0.34 log points in 2003. As we have discussed above, Giddings (2002b) finds that gender wage gap decreases in Bulgaria, despite an overall increase in the earnings inequality. Furthermore, Hunt (2002) documents a 10% point decrease in the gender wage gap in East Germany over the period 1990-1994. Similarly, Jolliffe and Campos (2005) find that the gap in Hungary decreased from 0.31 in 1986 to 0.19 in 1998. Grajek (2003) shows that gender wage gap in Poland decreased over 1987-1996. Nestic (2010) finds declining gender wage gap in Croatia between 1998 and 2008. However, the counterfactual gender wage gap in Croatia is found to be almost two times higher than the raw gap and it even increased during the period examined. In contrast, Newell and Reilly (2001) provide an overview of the evolution of the gender wage gap in eleven countries of East Europe and Former Soviet Union (Bulgaria, Czech Republic, Hungary, Poland, Slovakia, Federal Republic of Yugoslavia, Latvia, Russia, Ukraine,

[98] See Brainerd (2000) and Newell and Reilly (2001) for a selection of Central and Eastern European countries and the countries in the Commonwealth of Independent States (CIS); Orazem and Vodopivec (1995) for Slovenia, Jurada (2003) for the Czech and the Slovak Republic; Ogloblin (1999) for Russia.

Kazakhstan and Uzbekistan) and find that gender earnings gap remained relatively stable during the 1990s. Rutkowski (2001) shows that in the later stages of transition, the gender wage gap increased, although it remained rather modest by the international standards. Paci and Reilly (2004) suggest that the relatively small gender wage gap in most transition countries is connected with the higher human capital endowments of women compared to men, such as education.

TABLE 6.1

Changes in gender wage gap in transition countries- summary of results

Reference	Country	Change in the raw gender wage gap
Brainerd (2000)	Hungary, Poland, Czech and Slovak Republics (1988-1994)	Decreased by 0.05-0.12 log points
Brainerd (2000)	Russia and Ukraine (1992-1994)	Increased by 0.15-0.27 log points
Giddings (2002b)	Bulgaria (1986-2003)	Decreased by 0.09 log points
Newell and Reilly (2001)	11 countries (1984-1996)	Stable
Hunt (2002)	East Germany (1990-1994)	Decreased by 0.10 log points
Grajek (2003)	Poland (1987-1996)	Decreased by 0.10 log points
Ganguli and Terrell (2005)	Ukraine (1986-2003)	Decreased by 0.06 log points
Jolliffe and Campos (2005)	Hungary (1986-1998)	Decreased by 0.12 log points
Kecmanovic (2009)	Serbia (2001-2005)	Decreased by 0.08 log points
Nestic (2010)	Croatia (1998-2008)	Decreased by 0.04 log points

Source: Authors' literature review.

Most studies report narrowing of the gender wage differentials for transition countries except Russia and Ukraine. The literature review also supports the fact that there are very few studies, which exclusively investigate the gender wage gap in Bulgaria, Serbia, and Russia. It has been difficult to find a study that empirically investigates the gender wage gap across conditional distribution in Tajikistan.

6.5. Conclusions and comments on a research agenda

The aim of the present Chapter has been to review the empirical literature and to discuss the main decomposition methods used in the gender wage gap estimates. This Chapter reviews different methodological techniques that have been used to estimate the wage differential at the mean and across the conditional wage distribution. An overview of the empirical literature related to the gender wage gap in transition countries has been provided.

The ways in which the gender wage gap is decomposed are diverse. We have concentrated on the most commonly used decomposition methods. An important finding in the literature is that the estimates at the mean provide an incomplete summary of the pattern of earnings across the entire wage distribution. The Oaxaca-Blinder decomposition, based on the ordinary least squares, allows us to obtain an exact decomposition of the average wage gap between males and females. However, in the context of the QR the unconditional quantile wage is equal to its quantile wage conditional on the vector of individual characteristics at that percentile level plus some individual error terms, which are not zero. Moreover, the evidence of distributional approach literature points out that in most of these studies returns to human capital characteristics and the raw gender differential increase as one moves along the wage distribution. One of the shortcomings of the Machado and Mata (2005) methodology is that it uses the Blinder-Oaxaca type decomposition approach, which has been subject to criticism. In most applications the coefficient estimates from the male wage regressions are used as a reference non-discriminatory wage structure. Hence a pooled approach, obtained from the male and female wage structure is required.

Finally, most of the studies of the East European countries show a narrowing of the gender wage differential after the introduction of market reforms. Gender differences in productive characteristics explain only a small portion of the wage gap. The consensus is that the gender wage gap has either remained stable or has actually decreased in Central Europe while it has increased in the countries of the former Soviet Union. The majority of the empirical studies in the case of Russia indicate that differences in the endowment component between males and females weakly contribute to the wage differential. Some other scholars suggest that the pay gap between men and women is relatively low by international standards and that women have actually benefited throughout the systematic change from communist to market economy.

We find that previous studies on gender wage gap at different points of the distribution in the case of Bulgaria, Russia, Serbia and Tajikistan are limited and thus motivate our empirical analysis. The most challenging point for us is the application of the appropriate decomposition method. In the next Chapter we will apply a newly developed econometric technique which allows us to take into account differences across the wage distribution and the 'index' number problem as suggested by Neumark

(1988) and Oaxaca and Ransom (1988 and 1994). We believe that our estimates of the relative positions of women and men overall distribution of earnings in these particular countries might have potentially important policy implications.

CHAPTER SEVEN

GENDER WAGE GAP ACROSS THE DISTRIBUTION: EMPIRICAL EVINDENCE FROM FOUR TRANSITION COUNTRIES

7.1. Introduction

Explaining the persistence in gender earnings differential is essential to the advancement of women's interests in all economies. It is interesting to see how Bulgaria, Russia, Tajikistan and Serbia are compare to the European benchmarks of gender equality in the labour market now, after almost two decades of transition reforms. In order to obtain a more detailed insight into the gender earnings differential, we employ the method of QR, which allows us to control for individual characteristics over the entire conditional wage distribution. A special emphasis has been placed on the findings of the relatively higher gender differential amongst higher earners (the so called 'glass ceiling effect') and amongst low earners (the 'sticky floors effect'). By a 'glass ceiling' effect we refer to a greater earnings gap at the top end of the distribution, suggesting that female workers in the upper-income brackets have lower relative pay than their male counterparts. In contrast, a 'sticky floor' refers to the scenario where females at the bottom of the distribution are at a greater disadvantage and the gap is wider at the bottom (Booth, 2009; Arulampalam *et al.*, 2007). In order to measure the gender wage differential, the observed wage gap is split into two parts – a part due to differences in characteristics and part due to differences in returns to these characteristics.

This Chapter contributes to the literature in several ways. Firstly, the study considers the case of four formerly communist economies – Bulgaria, Russia, Tajikistan and Serbia, countries for which not considerable amounts of data and research on gender wage gap are available. Secondly, it extends the popular Machado and Mata (2005) decomposition procedure to address the 'index' number problem suggested by Neumark (1988) and Oaxaca and Ransom (1988 and 1994). We use the coefficient estimates from the pooled wage regressions as the reference remuneration that is the non-discriminatory wage structure. To the best of our knowledge, the proposed procedure has not been used to estimate the gender wage gap across the conditional distribution in the countries of interest. Thirdly, our focus is on two time periods 2003 and 2007 (2001 and 2003 for Bulgaria), thus allowing us to observe changes in the size of the wage gap over time. The more recent data utilised in the Chapter allows us to broaden understanding of the way gender differences evolved during the entire process of the economic transition from a planned to a market economy and to a democratic society. Finally, we have attempted to fit comparable specifications to the samples across countries in order to compare the results.

This Chapter is organized in the following way. Section 7.2 outlines the labour market institutional background and labour market trends in the countries of interest. We introduce some stylized facts with which to begin investigating whether transition has increased or reduced gender differences in the labour market. In section 7.3, we comment on the data used in the estimations. Section 7.4 outlines our method for decomposing gender wage gap. Finally, sections 7.5 and 7.6 discuss the main empirical results and conclusions.

7.2. Institutional background

In all four transition countries, gender equality was a proclaimed policy goal during the communist era and evidence shows that the difference in wages between women and men were low at that time (Brainerd, 2000). The prevailing view in the literature is that on the eve of transition, occupational segregation in Eastern Europe was lower than in developed market economies (Maltseva and Roshchin, 2006). As discussed in Chapter two, wages in these economies were set according to industry specific wage grids varying only with the difficulty of the job and with worker's education and experience but not with gender (Münich *et al.*, 2005). Moreover, wages were biased in favour of blue collar industries, so that workers in manufacturing were paid better than workers in health and education (Orazem and Vodopivec, 1997). In most countries, the transition process led to a decrease in mean wages and an asymmetric change in the tails of the wage distribution. Skoufias (2003) argues that real wages at the bottom decile of the wage distribution in each transition economy decreased substantially while real wages at the top decile of the wage distribution decreased relatively less[99]. Another important feature of the Soviet-type economies was an excess demand for labour driven by economic plans targeting rapid industrialization and extensive growth of the economy. During the communist era all able individuals of working age were, by law, obligated to be employed. A well-known feature of the communist systems was their ability to integrate women almost fully into the economy (Atkinson and Micklewright, 1992). Female labour was needed to satisfy the excess demand for labour and official policies encouraged women to enter the labour force in many ways. As a result, women in the region had higher rates of economic activity than in any other part of the world (United Nations, 1991). However, the male-female relative labour market has changed and changes in earnings and in employment structure hit males and females differently (Grajek, 2003).

The very high female labour force participation that had been previously attained dropped considerably just after the start of transition (Malysheva and Verashchagina, 2008). Both men and women experienced a decline in economic activity rate and the reduction was larger in Central Eastern European countries than in those that constituted the Former Soviet Union (FSU). The reason was that many state-owned enterprises in the latter group of countries resorted to the practise of so-called labour hoarding in order to prevent mass unemployment and withdrawal from the labour market (Namazie, 2003). Despite men's higher rates, in all countries where women's activity rate was low, men's rate was also low (see Table 7.1). This suggests a sluggish labour market rather than merely overt gender discrimination (UNDP, 2006).

TABLE 7.1: Economic activity rates for women and men by countries

	1990		1995		2000		2004		2007	
	Women	Men	Women	Men	Women	Men	Women	Men	Women	Men
Bulgaria[1]	50.3	59.7	47.9	56.8	44.7	56.2	45.4	56.1	47.5	58.2
Serbia	-	-	51.8	67.8	48.9	65.8	47.0	64.6	42.8	59.7
Russia[2]	61.0	77.4	48.4	63.3	48.3	61.1	46.6	55.8	48.5	57.9
Tajikistan	-	-	-	-	-	-	52.2	74.0	-	-

Source: UNECE Gender Statistics Database.
Notes: Economically active includes all residents who are either employed or unemployed (aged +15) [1] Data for 1990 refer to 1993. [2] Data for 1990, 1995 and 2000 refer to persons aged 15-72.

Women in the East also fared comparatively worse than men in terms of unemployment, which, as discussed in Essay one, rose at the beginning of the 1990s

[99] In some countries such as Hungary and Croatia, real wage of highly paid workers increased.

from the practically non-existent level experienced before transition. Stefanova and Terrel (2007) provide empirical evidence that women had lower job-finding rates once they become unemployed and were therefore more subject to a long-term unemployment. At the same time, educational attainment of women in the post-communist countries was very high. The fact that education was financed by the state provided women with high chances of access to it. This tendency still holds and female outnumber males in university attendance in countries like Russia, Belarus and Kazakhstan (Paci, 2002).

The dominant view in the literature, as we have summarized in Chapter six, is that the gender wage gap has been stable or declining over the transition period. The gap is documented as standing about 20%, primarily based on evidence from Central European countries (Malysheva and Verashchagina, 2008). The UNECE data on monthly earnings indicate that the wage gap in 2003 ranges from 18% in Bulgaria to 36% in Russia.

Today, wage grids and restrained gender earnings differential are used only in the budgetary sector (public administration, health, education) and room for pay discrimination is relatively open in the unregulated non-budgetary sector (Jurajda, 2003). Having in mind these changes, we can group the factors that influence the trends in the gender wage differential into the following: those related to changes in the overall wage structure, changes in labour market skills and productivity and changes in gender specific factors, such as remunerations of skills. The changes of labour market institutions experienced over the period also have potentially had an impact.

One of the most negative aspects of Bulgarian's transition has been a great decrease in real income level. In particular, in areas where more women than men are employed, wages are, as a rule, significantly lower than average pay levels for the country as a whole. The wages of Bulgarian women in industries and services in 2002 were about 70% of those of men in the same sectors. In spite of proclaimed gender equality[100]

[100]The Labour code (1986, amended 2001) prohibits all forms of discrimination, privileges, and limitations based on nationality, origin, gender and race. The amendment to the Code in 2001 introduced the principle of equal pay for women and men (Article 243). By the end of 2003 certain amendments to the Labour Code introduced the definitions of indirect discrimination. Moreover, when advertising job vacancies employers do not have the right to set requirements relating to gender and age.

women experienced specific treatment by the authorities which resembled occupational segregation and promotional discrimination (Grajek, 2003). Jolliffe (2002) suggests that in the case of Bulgaria, changes in the wage structure have disproportionately harmed the position of women. Possible reason suggested for this is that lower paid professions were being increasingly feminized. Another reason is that women were unable to fully realise their professional potential, even when they share the same professions as men, because they have to take care of their children. In addition there are discriminatory practices in the labour market with regard to some groups of women (for example, young women with short professional experience, pregnant women, women with small children, and women over 45 years)[101].

Under the Soviet system, one of the major political goals was to achieve equality between men and women. Women enjoyed substantial rights at the workplace, but occupational segregation based on gender remained a pronounced feature of the Russian labour market. The state assumed that women were different kind of workers from men, marked out by their biological capacity to be mothers. This assumption found its strongest expression in the banning of women from certain professions because participation in them might impair their ability to produce healthy children (Attwood, 1999). The definition of Soviet women as 'worker-mothers', who were expected to bear primary responsibility for running the household, meant that women tended to be seen as 'second-class' workers (Ashwin, 2000). Despite high levels of education and professional training, women in Russia also faced the phenomenon of the 'glass ceiling' in which they were unable to advance to leadership positions (Central European and Eurasian Law Initiative, 2006). As a whole women were typically located in the lower part of the wage distribution and a compression at the bottom of the distribution was therefore likely to raise their average earning relative to men (Kazakova, 2007). This is consistent with Brainerd (2000) who shows the widening wage structure in the early Russian transition was the major determinant of the increase in the gender wage gap. Studies of the wage gap during the Soviet transition suggest that it has remained more or less constant, with women continuing to earn between 60–70% of men's wages[102].

[101] Sziráczki and Windell (1992) suggest that the deteriorating position of women in Bulgaria may be results of discriminating practices. They report results from a survey of managers who were asked whether they had a preference for men or women when hiring for production or professional work. While 25% reported a preference for hiring women for skilled production work, 54% reported preference for hiring men.
[102] See Lapidus (1993).

The Russian Constitution[103], which sets the principles of equal rights and equal opportunity, does not prohibit the possibility of legal limitations on human rights based on sex. Russian law contains provisions that provide for differential treatment of men and women, primarily in the area of employment. For example, the Labour Code both includes the principle of equality and suggests restrictions on women of childbearing age from holding certain jobs that are considered physically difficult, harmful or dangerous and a prohibition on work that requires heavy lifting[104]. The Labour Code refers to an *Index of Heavy Work and Work in Harmful or Dangerous Labour Conditions which are Forbidden to be Fulfilled by Women*[105] that presently lists 456 specific types of work to which women's access is limited or prohibited. This list, which has existed since Soviet times, was renewed and reduced in 2000 but since then it has remained unchanged (American Bar Association and Central European and Eurasian Law Initiative, 2006). The 2006 American Bar Association Report further indicates that many employers in Russia are reluctant to hire women who are perceived to have family obligations. Women report that they have been specifically asked during interviews if they are married, have children, or intend to have children (Lokshina and Lukashevski, 2003). In addition, wage differential was especially great for female employees aged 20-40, which is precisely the age when women have to take their child care responsibilities. These policies and attitudes resulted in a sustaining gender wage gap in the country. According to recent figures, women's salaries are on average 64% of men's[106].

Similarly, women's capacity in Tajikistan to earn higher incomes was severely limited because of inequality in access to land and financial resources, good jobs and education. Access to education in Tajikistan was also hampered by traditional perceptions of the role of women in the society. According to the World Bank (2005) report women's wages in 2003 were only 45% of those of men. The Constitution of the Republic of Tajikistan (Article 35) declares equal rights for women and men in the field of employment. The Labour Code prohibits discrimination in labour remuneration. The new Gender Equality law was adopted in December 2004 to address discrimination

[103] Constitution of the Russian Federation, Article 19, part 1, Article 33 and Article 46.
[104] Labour Code, Article 253. Prior to 2001, the Labour Code prohibited women from undertaking certain types of work.
[105] Index of Heavy Work and Work in Harmful or Dangerous Labour Conditions which are Forbidden to be Fulfilled by Women, adopted by Russian Federation Government Resolution No.162, from February 25, 2000.
[106] Women and men in Russia, Federal Service of Government Statistics 108 (2004).

against women. Despite these policies, there still exists gender segregation – prevalence of women in certain spheres of economy, mostly with low wages, such as agriculture, public health services and education. Thus, 85.5% of women are employed in these branches with 75.1% in agriculture. Salaries in these activities were approximately 4-7 times lower than in industrial branches (industry, construction, transport, and communication)[107].

Serbia signed the Convention on the Elimination of all forms of discrimination against women and has taken steps to achieve the Convention's objectives[108]. The situation in the country, however, remained complex because the population comprises of several diverse ethnic origins, languages and religions. One of the characteristics of Serbian labour market was the large wage disparities among workers with the same or similar qualifications in different sectors. Some scholars indicate that women in Serbia mostly work in poorly paid jobs and they are frequently employed in the informal economy. Those sectors that suffer discrimination, like textiles, have lower wages relative to the wages in male privileged industries (see Krstic and Reilly, 2000). Evidence shows an increasing gender wage gap in Serbia: while in 1996 employed women earned 15 % less than men, the gender wage gap increased by 26 % in 2000 (Krstic and Reilly, 2000).

7.3. The data

The results in this Chapter are based on data from Bulgaria, Russia, Tajikistan and Serbia. The Bulgarian data are taken from the Living Standard Measurements Survey (LSMS), conducted in 2001 and 2003. The Russian data are taken from rounds 12 and 16 of the Russian Longitudinal Monitoring Survey (RLMS) in 2003 and 2007. The Tajikistan Living Standard Survey (TLSS) of 2003 and 2007 provide our third source of data. Finally, Serbian data are taken from the Serbian Living Standard Measurement Survey (LSMS) in 2003 and 2007. In all four countries, we sample individuals aged between 15 and 65. Since the construction of the data and the definitions of the variables have already been explained in Section 3.3.1, this discussion is not repeated here. However, for Russia some discussion is required as the data used in this chapter

[107]Tajikistan - Shadow Report on the realization of the convention on the elimination of all forms of discrimination against women, Dushanbe, 2006.

[108] CEDAW (Committee on the Elimination of Discrimination against Women), Consideration of Reports submitted by States parties under Article 18 of the Convention on the elimination of all forms of discrimination against women: Serbia, Initial Report of States Parties, CEDAW, New York, NY

are different. Samples are tested to see whether the data sets support separation. The reported *t*-statistics in the descriptive tables indicate that the separation of the data between male and female samples is statistically significant for a number of individual characteristics (see Appendix Tables A7.1 to A7.4).

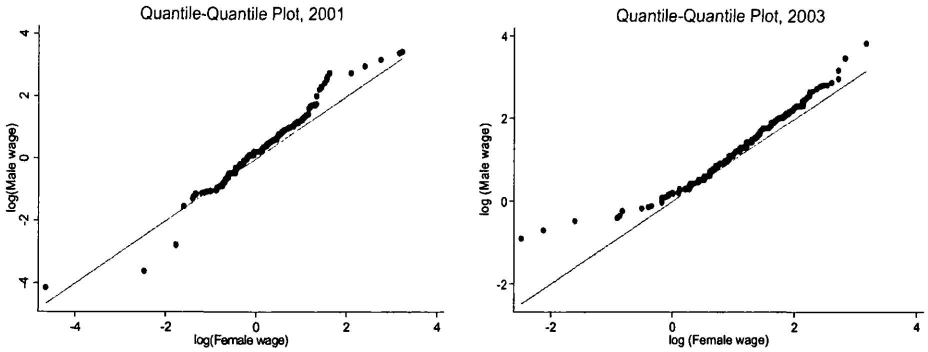
We have fitted comparable specifications to the samples across four countries to place the analysis in a common framework. The dependent variable is hourly earnings. The control variables used in the estimations include education (variables for university, secondary and primary school), potential experience (linear and quadratic terms), variables for individuals years of tenure with the current firm, controls for marital status, public sector employment, and a set of regional variables to pick up regional effects.

The Bulgarian data used in the estimation consists of a sample of 729 men and 709 women in 2001 and 1,296 men and 1,186 women in 2003. Appendix Table A7.1 reports a complete list of variables used in the analysis. For both years, the average log hourly wage rate is higher for men. Education is represented by variables measuring the completion of indicated levels of schooling. The levels consist of primary, secondary and university education. In 2001, 13.7% of employed men and 20.7% of employed women have a university degree. This increases to 16.7% for males and 25.6% for females in 2003. The earnings density functions, estimated using an Epanechnikov kernel estimator, show that the male wage distribution in 2001 is characterized by much lower dispersion than the one in 2003 (see Figure A7.1).

The quanitle-quantile plots (Q-Q plot)[109] of Bulgarian male/female wage distributions in 2001 and 2003, presented in Figure 7.1, facilitates a comparison of the empirical wage distributions for both male and female samples.

[109] The Q-Q plot is a graphical method for comparing two distributions to see how well they match or where differences occur. The graph relates quantiles of the variable on the vertical axis to quantiles of the variable on the horizontal axis. A point on the symmetry line indicates that quantile of one distribution has the same value as the corresponding quantile of the other distribution.

FIGURE 7.1: Comparing empirical wage distributions, Q-Q plot, Bulgaria

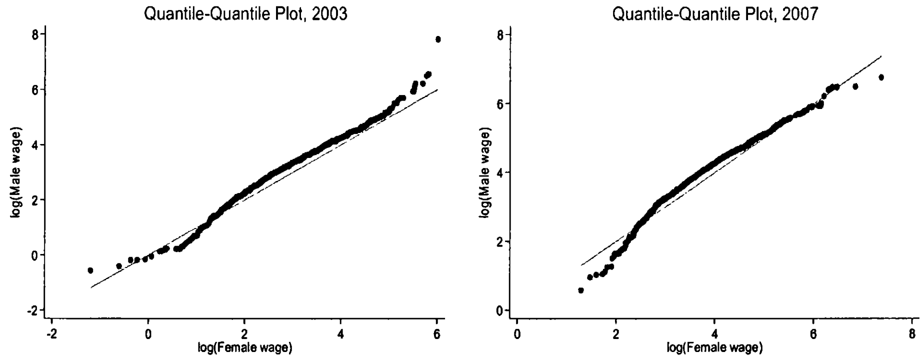


Source: Bulgaria LSMS, 2001 and 2003 data. The plots are created by using qq diagnostic plot in STATA.

Most of the observations in 2001 and 2003 are above the diagonal line, implying that wages for male workers in Bulgaria are higher than wages for female workers at comparable quantiles of the wage distribution. However, the Q-Q plot in 2001 reveals that some low-earnings women have higher wages than men. The plots also indicate that wage inequality is somewhat stronger at low and high income levels in 2003.

The Russian data used in this chapter are taken from the Russian Longitudinal Monitoring Survey (RLMS), which is designed to measure the effect of economic and political reforms on the economic welfare of the Russian population. For comparison purposes we use round 12 and round 16 of the RLMS. In Appendix Table A7.2, we report the descriptive statistics for the samples and the variables used in the empirical analysis. The Russian sample consists of 1,907 working men and 2,261 working women in 2003, and 1,917 working men and 2,315 working women in 2007. There are differences in the characteristics of men and women with respect to both educational qualifications and job status. The data indicates that women earn less than men, with an unconditional mean wage gap of 29% in 2003 and 22% in 2007. Tenure decreases between 2003 and 2007 for both male and female workers. This may point to a relatively mobile labour market. We can see that in both 2003 and 2007 a higher proportion of women than men complete university degree. A significantly higher proportion of working men are married (82%). The wage arrears effect is higher for males than females and decreased in size between 2003 and 2007. Not surprisingly, women have higher participation in the public sector. The Q-Q plots of male/female wage distributions in Russia show that not all of the data points are to the left of the main diagonal (see Figure 7.2). We also see that income inequality is stronger for low and high earners. Moreover, in 2007 there are some low-earners women with higher wages than men.

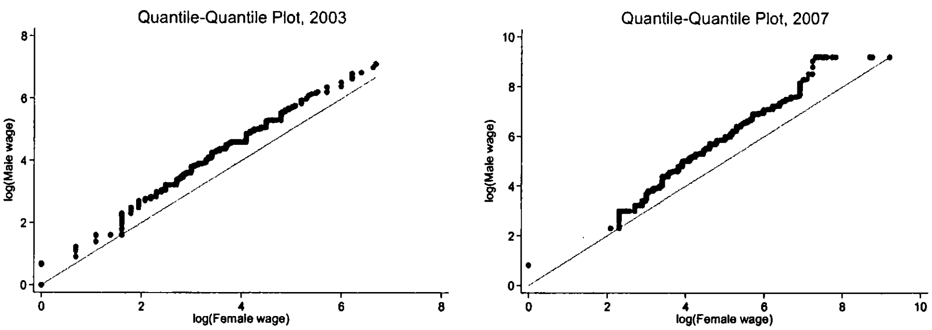
FIGURE 7.2: Comparing empirical wage distributions, Q-Q plot, Russia



Source: Russia RLMS, 2003 and 2007 data.

The Tajikistan sample used in the analysis consists of 3,006 working men and 1,832 working women in 2001 and 4,133 working men and 2,240 working women in 2007 (see Table A7.3). Similarly to the other countries examined, the descriptive statistics for Tajikistan indicate that in both years men are paid much more than women. We can also see that a higher proportion of men have completed a university degree (21%-23% for men compared to 13-15% for women). Around 66% (68%) of males (females) live in rural areas, which could be a barrier to equal opportunities in Tajikistan because the traditional stereotyping of female and male roles are different in the rural compared with urban areas. In addition, there are more opportunities for women to get better education, better paid jobs, and to participate in the decision-making process in urban areas. The Q-Q plot for Tajikistan reveals that all of the observations in 2003 are above the diagonal line, confirming that wages for male workers are higher than wages for females across the distribution. In 2007, at the higher quantiles the male/female wage differential is even larger. As a result we can expect larger differential between male and female wages among the higher-paid than among the lower paid workers.

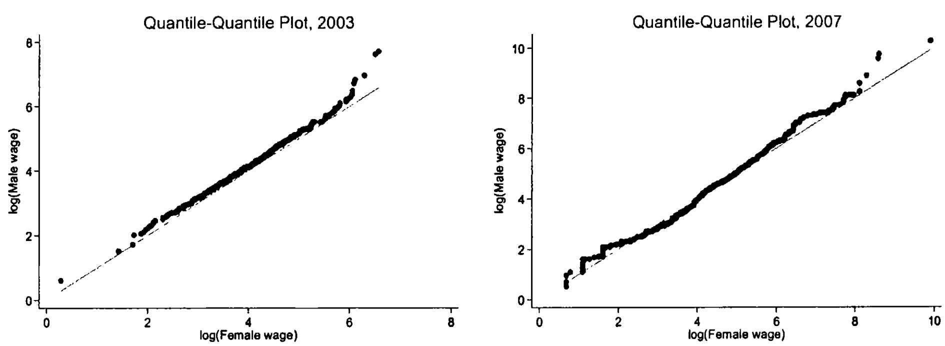
FIGURE 7.3: Comparing empirical wage distributions, Q-Q plot, Tajikistan



Source: Tajikistan TLSS, 2003 and 2007 data.

Finally, the Serbian sample consists of 1,466 working men and 984 working women in 2003, and 2,983 working men and 2,014 working women in 2007. Appendix Table A7.4 reports the means and standard deviations of the main variables used in the study. We observe an increase in the average wage from 2003 to 2007 which is higher for women than for men. In both years, females are better educated than males. The majority of public sector employees are women. Also in both years, at the higher quantiles of the wage distributions, male wages are higher than female wages (see Figure 7.4).

FIGURE 7.4: Comparing empirical wage distributions, Q-Q plot Serbia



Source: Serbia, LSMS, 2003 and 2007 data.

An important feature of these data is that in all four countries (the 2003 Tajikistan sample is an exception) women are more likely than men to be employed in the public sector. Since public-sector wages tend to be higher, this implies that gender differences in wages are probably more muted than they might be otherwise (Appleton *et al.*, 1999). Overall, apart from Tajikistan, women employees are systematically more educated than their male counterparts.

7.4. Methodology

This chapter utilises the Machado and Mata's (2005) technique to decompose the difference in male and female log wage distributions into a component due to differences in labour market characteristics between men and women and a component due to difference in the rewards men and women receive for their labour market characteristics. Based on Koenker-Basset (1978), Machado and Mata (2005) propose a method to extend the traditional Oaxaca-Blinder decomposition. Their main methodological procedure is to simulate the conditional marginal wage distribution estimated through quantile regressions run separately for men and women[110]. The general idea of their methodology is to generate the female wage distribution that would emerge if women retained their own labour market characteristics but were "paid like men". We extend this procedure by generalising the simulation and adopting the Oaxaca-Ransom methodology, therefore taking into account the 'index' number problem suggested in Neumark (1988) and Oaxaca and Ransom (1988 and 1994). Neumark (1988) shows that the non-discriminatory wage structure can be estimated from an earnings equation estimated over a pooled sample of men and women, rather than using the male or female wage structure as a base. This is the non-discrimination equation, which is the benchmark in the estimation of the Oaxaca and Ransom decomposition. The question we address by the proposed Machado and Mata-Oaxaca-Ransom (MM-OR) decomposition is how much would the male-female wage gap change across the distribution if men and women were paid according to a common wage structure, but their work related attributes remain as they are. A positive wage gap implies that the returns to women's characteristics are lower than those of men, and a negative gap implies the reverse.

Formally, our estimation decomposition procedure involves the following steps:

Step 1: Generate a randomly drawn sample of size N from a uniform distribution $U\,[0,1]\; u_1, ..., u_N$. This will give a series of numbers telling us which percentiles are to be estimated.

Step 2: Using the male and female datasets, we estimate, for each percentile derived in the first step, regression coefficients for men, women and the pooled sample:

[110] The wage distribution of the second group is simulated basing on the wage distribution and the characteristics of the first group, and these steps are repeated n times.

$\hat{\beta}^m_{u_i}(\theta), \hat{\beta}^f_{u_i}(\theta), \hat{\beta}^*_{u_i}(\theta), i = 1......N$, where $\hat{\beta}^m_{u_i}(\theta)$ are u_i^{th} quantile regression estimates taken from the log hourly earnings equation for males; $\hat{\beta}^f_{u_i}(\theta)$ are u_i^{th} quantile regression estimates taken from the log hourly earnings equation for females, and $\hat{\beta}^*_{u_i}(\theta)$ are u_i^{th} quantile regression estimates taken from the log hourly earnings equation for all workers.

Step 3: For each percentile, characteristics of a randomly drawn sample of 10,000 men and women (with replacement) are used to predict wages by using the estimated coefficient vectors $\hat{\beta}^m_{u_i}(\theta)$, $\hat{\beta}^f_{u_i}(\theta)$ and $\hat{\beta}^*_{u_i}(\theta)$. This process generates sets of predicted wages covering the whole distribution and enables us to calculate the wage distribution for males, females and for both males and females together. The vectors of characteristics for males $\{\widetilde{X}_i^m\}_{i=1}^N$ and females $\{\widetilde{X}_i^f\}_{i=1}^N$ are then used to predict (log) hourly earnings for males $\{\widetilde{w}^m_{u_i} = \widetilde{X}_i^m \hat{\beta}^m_{u_i}\}_{i=1}^N$ and females $\{\widetilde{w}^f_{u_i} = \widetilde{X}_i^f \hat{\beta}^f_{u_i}\}_{i=1}^N$. These predicted wages are equivalent to a random sample of size N drawn from the marginal wage distributions for males (w^m) and females (w^f). Counterfactual densities for male and females workers being rewarded equally are found as $\{\widetilde{w}_i^{cfm} = \widetilde{X}_i^m \hat{\beta}^*_{u_i}\}_{i=1}^N$ and $\{\widetilde{w}_i^{cff} = \widetilde{X}_i^f \hat{\beta}^*_{u_i}\}_{i=1}^N$ which are the densities that would arise if women and men retained their own labour market characteristics but were paid like a randomly chosen individual from the entire sample[111].

Step 4: The difference between the quantiles of the simulated distributions can be used to quantify gender difference across the distribution. Using the calculated distributions, we decompose the gender wage gap into the contribution of the coefficients and the contribution of the covariates as:

[111] Note that the original counterfactual density is generated as $\{\widetilde{y}_i^{cf} = \widetilde{X}_i^f \hat{\beta}^m\}_{i=1}^N$ which is the density that would arise if women retained their own labour market characteristics, but were paid like men. The decomposition could also be made with the counterfactual $\{\widetilde{y}_i^{cf} = \widetilde{X}_i^m \hat{\beta}^f\}_{i=1}^N$ which is a counterfactual earning density that would have prevailed if women were given men's labour market characteristics, but remained to receive returns to those characteristics like women.

$$Q_\theta(\tilde{w}^m \mid \hat{\beta}^m - \tilde{w}^f \mid \hat{\beta}^f) =$$

$$= \underbrace{Q_\theta[(\underbrace{\tilde{w}^m \mid \hat{\beta}^m - \tilde{w}^{cfm}}_{\textit{male advantage}} \mid \hat{\beta}^*) + [(\underbrace{\tilde{w}^{cff} \mid \hat{\beta}^* - \tilde{w}^f}_{\textit{female disadvantge}} \mid \hat{\beta}^f)]}_{\textit{coeeficien effect}} + Q_\theta[\underbrace{(\tilde{w}^{cfm} \mid \hat{\beta}^* - \tilde{w}^{cff} \mid \hat{\beta}^*]}_{\textit{characteristics component}} + resid \tag{7.1}$$

where Q_θ is the θ^{th} percentile of the earnings distribution and $(\tilde{w}^i \mid \hat{\beta}^j)$ is the estimated marginal earnings distribution for the i^{th} individual based on the j^{th} quantile regression estimates.

The first two terms of the right hand side of equation (7.1) identifies the part of any wage difference explained by differences in the returns that male and female receive for a given set of characteristics (the coefficient effect)[112]. The third term of equation 7.1 identifies the part of the hourly wage difference at the θ^{th} centile, explained by differences in the characteristics of male and female workers (the characteristics effect). This is the portion of the wage differential that may be attributed to differences in the characteristics possessed by men and women. The difference between θ^{th} quantile of the marginal wage densities between male and female distributions weighted by the characteristics of workers randomly chosen in the economy does contain an additional component, which we treat as a residual. The residual term is typically of second order of importance and tends to became smaller with a larger number of simulations. Standard errors for the reported components of the decomposition were obtained using a standard bootstrapping method[113].

7.5. Empirical results

7.5.1. Explaining changes in relative wages

To shed a light on the role the overall wage structure has played in the changing gender gap, we summarise the changes in male and female wage inequality in Bulgaria, Russia, Tajikistan and Serbia in Table 7.2. The 90/10 log wage ratio is given as a measure of overall wage dispersion and the 90/50 and 50/10 log wage ratio provide an indication of whether changes in wage dispersion occurred predominantly in the upper or lower half

[112] Specifically, the technique allows the discrimination component to be disaggregated into overpayment of male workers (male advantage) and underpayment of female workers (female disadvantage). We do not attempt in our study to decompose into these two components.

[113] Bootstrap estimates are based on 800 replications.

of the wage distribution. Analysing these measures allows one to identify which part of the distribution contributes more to the overall earnings inequality. A positive sign for the change in the wage distribution is associated with an increase in earnings dispersion whereas a negative sign indicate that wage dispersion has decreased over time.

TABLE 7.2: Summary measures of the log wage distribution

	Men			Women		
Year	90/10	90/50	50/10	90/10	90/50	50/10
			Bulgaria			
2001	1.487	0.714	0.772	1.288	0.659	0.629
2003	1.384	0.760	0.624	1.193	0.662	0.531
% change	-7%	6%	-19%	-7%	0%	-16%
			Russia			
2003	2.154	0.972	1.182	2.115	1.039	1.076
2007	1.785	0.858	0.927	1.853	1.003	0.850
% change	-17%	-12%	-22%	-12%	-3%	-21%
			Tajikistan			
2003	2.708	1.609	1.099	2.457	1.476	0.981
2007	2.863	1.477	1.386	2.446	1.322	1.124
% change	6%	-8%	26%	0%	-10%	15%
			Serbia			
2003	1.588	0.812	0.777	1.560	0.770	0.790
2007	2.189	0.944	1.245	1.946	0.916	1.030
% change	38%	16%	60%	25%	19%	30%

Note: The log wage at the 90th percentile of the wage distribution minus the log wage at the 10th percentile of the distribution. In a similar way were calculated the 90/50 and 50/10 measures.

Changes in relative wages differ between countries. If we take the ratio of male earnings at the top decile to earnings at the bottom decile (90/10) for the 2003, then this decile ratio has a value of 1.39 in Bulgaria, rising to 2.15 in Russia, 2.71 in Tajikistan and 1.59 in Serbia, which shows that earnings dispersion is distinctly greater in Tajikistan and it is lowest in Bulgaria.

Over the period examined, the wage structure has widened substantially for both males and females in Serbia. In particular, the increase in wage inequality is stronger for male workers at the lower tail of the earnings distribution, where we estimate that the log 50/10 percentile ratio for male labour earnings increased from 0.78 in 2003 to 1.25 in 2007. This is a 60% increase in the 50/10 ratio, which is sizeable for a 4-years period. In Tajikistan, the increase in the wage inequality for both genders has been stronger in the lower earnings tail (50/10), where we find a 26% increase for male workers and 15% increase for female workers. In Bulgaria and Russia for both males and females, the 50/10 log hourly wage ratio actually shrinks. Specifically, the 50/10 log hourly wage ratio in Bulgaria has decreased by 19% for males and 16% for females and similarly the Russian 50/10 log hourly wage ratio shows a decrease of around 21%.

To put the analysis into a perspective, one can see that wages at the 90th percentile of the Bulgarian male distribution in 2001 are about 93 log percentage points higher than wages at the 10th percentile. The corresponding figures in the year 2003 are 122 log percentage points. For Serbia, the 90/10 ratio seems to have flattened out in 2007. The wages at the 90th percentile of the male wage distribution are about 76 percentage points higher than the wages at the 10th percentile, whereas in 2003 the difference between the 90th and 10th percentile was 104 percentage points. We may conclude that between 2003 and 2007 there has been a marked reduction in the earnings inequality in Russia. Overall, taking the period as whole, in all four countries changes in the lower part of the wage distribution (50/10 log hourly ratio) are higher than changes in the upper part.

7.5.2. Gender wage gap decomposition results

The results of the wage decomposition are presented in Table 7.3. In the first three columns, we present the raw wage gap estimates for each year and the changes over time. The raw wage gap is calculated as the difference in log hourly wages between males and females at certain points of the wage distribution. In the next columns, we give the estimates of the proportion of the gap due to difference in the returns that workers receive for a given set of characteristics (the coefficients) and to the effect to different observable characteristics or 'endowments' across groups (the covariates), and residual terms that are due to the differences between the actual and simulated densities. The bootstrapped standard errors for these contributions are given in parenthesis. The estimated coefficient effects are graphed in Appendix Figure A7.5. A value greater than 100% for the proportion of the observed raw gap, explained by the differences in returns to characteristics, implies that women have characteristics that compensate them for 'discrimination' (Arulampalam *et al.*, 2007). Estimates at the 10th, 25th, 50th, 75th and 90th percentile are reported.

There are substantial variations in the decompositions across the countries. Estimates of the changes in the raw wage gap between both years indicate what has been happening to the gender wage premium[114]. The observed negative changes in the raw wage gap across the distribution for Bulgaria, Russia and Serbia imply that the gender wage differential has fallen over time. This is most pronounced in Russia at the higher percentiles where the raw gap decreases by 41%. Similarly, the raw wage gap in Serbia is systematically higher for the 2003 sample compared to the 2007 sample. The previous literature (Brainerd, 2000; Adamchik and Bedi, 2003, Newell and Socha, 2007) attribute this change to improvements in gender specific skills and increase in returns to human capital. In contrast, estimates of the changes in the raw wage gap between 2003 and 2007 in Tajikistan indicate that the observed differential has risen. In 2007, at the bottom of the distribution the observed raw gap increased by 56% (at the 25th percentile level the raw wage gap increases from 0.588 log points in 2003 to 0.916 in 2007). This confirms our initial findings of increased wage inequality in the lower tail (50/10) of the distribution (see discussions in Section 7.5.1). The findings for Tajikistan are in line with Johnes (2002) who shows that gender wage gap in the country is quite large in comparison with many other economies.

Several features are worth mentioning. In all countries, both coefficients and covariates contribute to the actual evolution of the gender wage gap and their effect is significantly different from zero. The largest fraction of the gender wage gap is attributable to the differences in coefficients across the whole wage distribution. In some cases, residuals also have relative high portion of the total gap. In Serbia, for example, at the bottom part of the distribution, the model does not seems to work very well as the residuals account for a significantly high portion of the total gap.

For both years and in all four countries, the estimated coefficient effect is significant and positive across the entire distribution, suggesting that men are paid more than women. The coefficient effect is systematically higher for the 2001 Bulgarian sample compared to the 2003 sample, ranging from 0.187 log points in 2001 to 0.141 in 2003 at the 10th percentile. In addition, it appears that in both years, women at the bottom part of the wage distribution in Bulgaria are subject to less discrimination as the coefficient component is much larger at the top end; hence women in high paying jobs suffer from

[114] The raw wage gap is sometimes called the unadjusted gap to differentiate it from the conditional (adjusted) gender wage gap, which is estimated by taking into account differences between male and female labour market characteristics. For example, a gap of 0.4 at the *i*-th percentiles is interpreted as one group having a log-wage 40% higher than the other at that percentile.

higher wage gaps than those in the lower-middle portions of the conditional wage distribution. The results are in line with Paci and Reilly (2004), who show that discrimination effect in Bulgaria in 2001 appears relatively stable for the 10th and 50th percentile but exhibit a tendency to rise at the 90th percentile.

TABLE 7.3

Wage decomposition of changes in parameters of the distribution

	Raw Gap			Decomposition of changes					
Bulgaria				2001			2003		
Percentiles	**2001**	**2003**	**Changes**	**Coefficients**	**Covariates**	**Residual**	**Coefficients**	**Covariates**	**Residual**
10th	0.041	0.069	0.028	0.187	-0.129	-0.016	0.141	-0.068	-0.004
				(0.0155)	(0.0108)		(0.0116)	(0.0083)	
25th	0.208	0.114	-0.094	0.260	-0.090	0.038	0.154	-0.071	0.031
				(0.0107)	(0.0075)		(0.0092)	(0.0064)	
50th	0.236	0.163	-0.074	0.308	-0.083	0.011	0.208	-0.077	0.032
				(0.0092)	(0.0066)		(0.0101)	(0.0068)	
75th	0.262	0.186	-0.076	0.318	-0.082	0.026	0.269	-0.077	-0.006
				(0.0107)	(0.0072)		(0.0120)	(0.0082)	
90th	0.267	0.266	-0.001	0.329	-0.072	0.010	0.320	-0.070	0.016
				(0.0160)	(0.0108)		(0.0158)	(0.0108)	
Russia	**2003**	**2007**	**Changes**		**2003**			**2007**	
10th	0.241	0.232	-0.009	0.219	-0.036	0.058	0.235	-0.021	0.018
				(0.0363)	(0.0234)		(0.0232)	(0.0172)	
25th	0.318	0.259	-0.059	0.324	-0.007	0.000	0.273	0.013	-0.028
				(0.0223)	(0.0172)		(0.0185)	(0.0121)	
50th	0.346	0.310	-0.036	0.320	0.017	0.009	0.273	0.033	0.005
				(0.0188)	(0.0141)		(0.0174)	(0.0129)	
75th	0.322	0.270	-0.052	0.287	0.034	0.000	0.231	0.034	0.004
				(0.0225)	(0.0161)		(0.0207)	(0.0161)	
90th	0.279	0.164	-0.114	0.265	0.037	-0.023	0.189	0.027	-0.051
				(0.0318)	(0.0215)		(0.0273)	(0.0200)	
Tajikistan	**2003**	**2007**	**Changes**		**2003**			**2007**	
10th	0.511	0.654	0.143	0.453	0.091	-0.034	0.551	0.217	-0.114
				(0.0211)	(0.0146)		(0.0237)	(0.0188)	
25th	0.588	0.916	0.329	0.518	0.109	-0.040	0.658	0.255	0.003
				(0.0162)	(0.0117)		(0.0194)	(0.0144)	
50th	0.629	0.916	0.288	0.571	0.110	-0.052	0.710	0.265	-0.059
				(0.0160)	(0.0112)		(0.0174)	(0.0132)	
75th	0.847	0.981	0.134	0.649	0.103	0.095	0.733	0.248	-0.001
				(0.0216)	(0.0146)		(0.0197)	(0.0138)	
90th	0.762	1.055	0.293	0.626	0.095	0.041	0.680	0.329	0.046
				(0.0281)	(0.0191)		(0.0299)	(0.0225)	
Serbia	**2003**	**2007**	**Changes**		**2003**			**2007**	
10th	0.163	-0.174	-0.338	0.141	-0.063	0.086	0.001	-0.147	-0.028
				(0.0161)	(0.0115)		(0.0326)	(0.0229)	
25th	0.130	0.039	-0.091	0.147	-0.056	0.039	0.129	-0.121	0.031
				(0.0097)	(0.0071)		(0.0153)	(0.0229)	
50th	0.153	0.041	-0.112	0.167	-0.044	0.030	0.148	-0.107	0.000
				(0.0087)	(0.0065)		(0.0113)	(0.0078)	
75th	0.131	0.065	-0.066	0.143	-0.035	0.023	0.164	-0.103	0.004
				(0.0110)	(0.0073)		(0.0135)	(0.0094)	
90th	0.204	0.069	-0.135	0.199	-0.027	0.031	0.153	-0.065	-0.019
				(0.0187)	(0.0129)		(0.0209)	(0.0149)	

Notes: Bootstrapped standard errors in brakets. The differential is calculating by every 5th percentile. We present the main percentile results. The estimation using the proposed decomposition technique takes between 12 to 48 hours for each sample.

In both years, the counterfactual gap in Bulgaria is wider than the raw gap and this reflects the situation where women possess some advantages in characteristics, mainly education, as it is often the case in Central European economies[115]. Specifically, when females are more productive, more educated but are paid less than males the coefficient effect is greater than the raw gender wage gap. As the gap is getting larger once we move towards the top of the distribution, it indicates possible presence of a 'glass ceiling' in Bulgaria suggesting a barrier to further advancement of women once they have attained a certain level of payment. Similarly, the larger 'unexplained' component than the raw differential in Serbia, especially pronounced in 2007, is in line with Kecmanovic (2009), who finds that Serbian women are significantly more endowed compared to men. In Tajikistan, the estimated coefficient effects increase in both years as one moves from the 10^{th} to the 70^{th} quantile and slightly drops after that, which means that in jobs with options for generally higher levels of wages, it is more and more difficult for women to get the same wage as that of men. These results are not surprising having in mind Tajikistan institutional background and labour market characteristics. Nowadays, traditional stereotyping indicating the role of women and men in the Tajik society remains widespread, and often causes difficulties for women trying to balance professional and family duties[116]. The influence of Islam has strengthened since the end of the Soviet rule, and the general view is that women's place is in the home has remained.

Do endowments explain the gender wage gap? The decomposition results indicate that the gender wage differential in Bulgaria, Russia and Serbia is mainly attributed to differences in return, that is the 'male advantage' and the 'female disadvantage' and to a lesser extent to the employee's endowment. The relatively higher unexplained wage differential in Russia is in line with Newell and Reilly (2001) results, where they conclude that most of the differential in transition countries is not explained by the observable characteristics component. In Bulgaria and Serbia, female employees

[115] The counterfactual gap, which is lower than the raw gap is found in Albrecht *et al.* (2003) for Sweden, De la Rica *et al.* (2008) for Spain.
[116] In the Soviet period, girls were required to attend school and women were encouraged to go on to higher education and work outside the home. However, since 1991, poverty and high unemployment have increased the pressure on women to marry as early as possible, especially in rural communities. Also, according to UNICEF Regional Report (1999), violence against women and threats against women is a major barrier to the empowerment of women and their equal participation in society.

actually have more favourable characteristics than do their male counterparts[117]. For instance, Bulgarian women in 2003 have an advantage of nearly 100% at the bottom of the distribution with their own endowments compared to with men's endowments. Similar negative covariate effect across the whole earnings distribution was observed for Serbia. In Russia, women in 2003 are more 'endowed' only at the bottom end of the distribution, where the coefficient component of the gap is much lower as compared to the upper tails. However, the estimated covariate effect in Russia changes to positive at the top end of the distribution, indicating favourable male characteristics of about 16%. Significant positive covariate effect across the whole wage distribution is found in Tajikistan, where on average males dominate with their endowment component, and this was confirmed by our descriptive statistics[118]. One interesting findings for Tajikistan is that the increased gender raw wage gap in 2007 is due mainly to the increased contribution of the covariate effects, which is indication that changes in individual observed characteristics tend to increase the wage inequality. For example, in 2007 at the top end of the distribution, 31% of the raw gap in Tajikistan was due to characteristic component whereas in 2003 covariates contributed with only 12%.

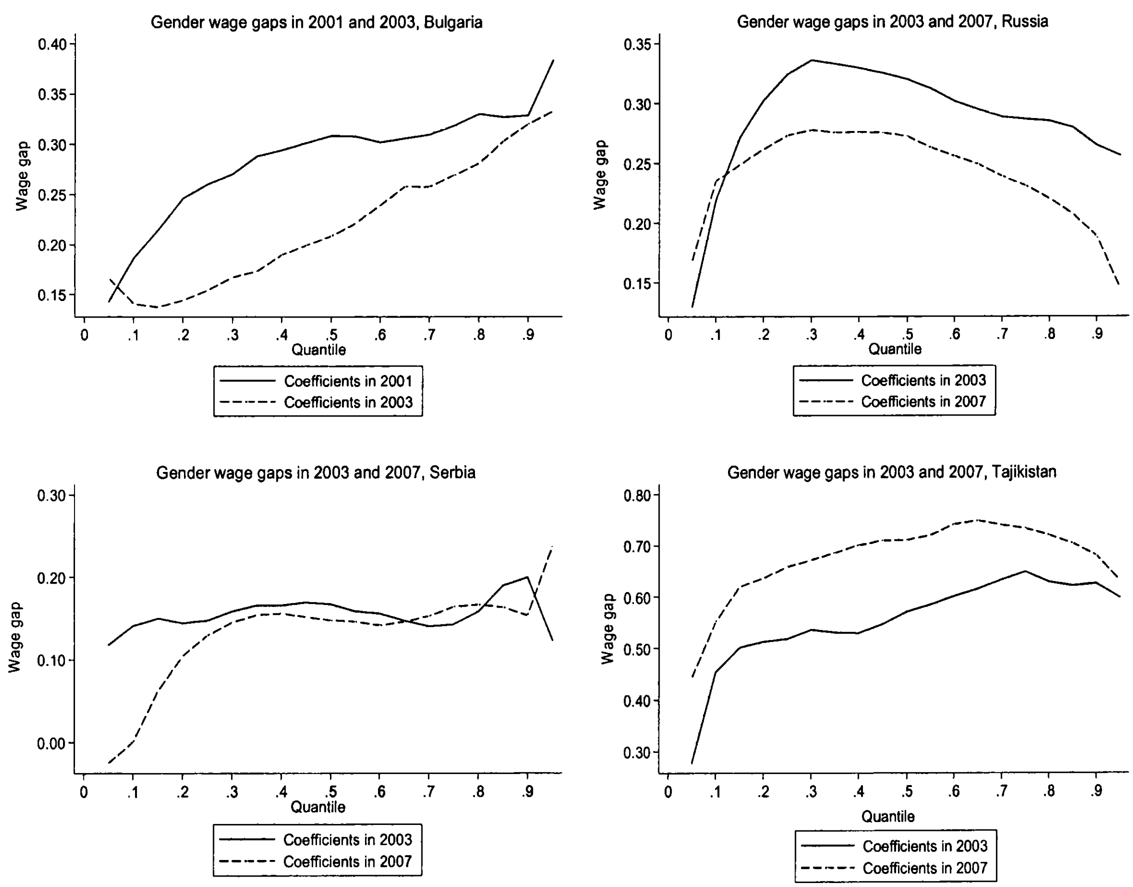
To summarise our findings over time, in three of the countries, the estimated coefficient effects of the gap are lower in 2007 than those in 2003. In Figure 7.5, we compare by country, the estimated differentials for the two time periods. No confidence intervals are plotted to avoid surcharging the figure. The gender wage gap in Bulgaria declined by more than 25% at the lowest decile, and by less than 3% at the highest decile. The gap narrowed considerably at the bottom of the distribution in Serbia. In contrast, the gender wage gap in Tajikistan increased between 2003 and 2007 – by 22% at the bottom of the conditional distribution and by 9% at the top end. This is partly due to the increase in the observable covariate effect. On average, the gender wage gap narrowed by 23% in Bulgaria, 26% in Serbia and 14% in Russia, while it increased on average by 19% in

[117] The covariate effect is negative in all percentiles and for both years, which indicates that women are with higher endowment component. A negative observed characteristics effect also shows that changes in the differences between men and women in observed characteristics, such as education served to diminish the gap. In other words, additional years of education, for example, benefit women to improve their relative earnings.

[118] We can see that higher proportion of men than women have completed secondary technical and university degree (See Table A7.3). In this regard it worth mention that according to Gender Statistics in the Republic of Tajikistan, fewer girls than boys attended school, particularly after primary school. This is related with the parental fears for their daughter's safety as they mature, and this fear is greater in urban than in rural arrears. Enrolment has also fallen due to the rising cost of education. Giving the choice of education a boy or a girl child, it seems that parents may be more willing to invest in a son's education than a daughter's (World Bank, 2000).

Tajikistan. These results point to substantial heterogeneity with regard to the changes in the gender wage gap along the distribution over the observed period and the relevance of the factors that are responsible for these changes. The increase in educational attainment in favour of female workers might contribute to the decline in the gender wage gap in Bulgaria, Russia and Serbia. However, very little is known about the factors that are responsible for distributional changes as our decomposition technique does not identify a separate contribution of the variables. A large part of the changes in the gender wage gap remains unexplained.

FIGURE 7.5: Comparing estimated coefficient effects of the gap by years



Notes: The differential is calculating by every 5th percentile.

Finally, one immediate question that might arise is why there is different degree of gender wage differentials among countries. Although we consider human capital and work experience as a major determinants of gender pay gap, one possible answer to the question is that the degree of gender specific policies and antidiscrimination laws differ from country to country. However, the effect of these policies on the earnings depends on the effectiveness of the legislations and this is difficult to be captured by our model.

7.6. Conclusions

The main novelty in respect to this Chapter lays in the extension of the Machado and Mata (2005) framework for decomposing the gender wage gap and accounting for the 'index' number problem. Applying the method to data from Bulgaria, Serbia, Russia and Tajikistan, positive discrimination effect was found across the whole earnings distribution, indicating that in all these countries otherwise identical men and women receive different returns to their characteristics. Together, changes in personal characteristics, coefficient effects as well as residual wage changes, work toward a reduction of the gender wage gap in Bulgaria, Russia and Serbia. In contrast, observable characteristics change in a way to increase gender wage gap and we conclude that the rise in wage inequality in Tajikistan is driven by the increased characteristics effect in favour of men.

Despite the existence of legal provisions aimed at gender equality, in all four countries men are paid more than women. The estimated differential is not constant along the conditional wage distribution and the decomposition method applied, provided more informative insight than the decomposition at the mean. The coefficient effect was found to be larger at the top of the distribution than at the bottom in both Bulgarian samples, which indicates possible presence of a 'glass ceiling' effect in the country. The magnitude of the coefficient effect in Bulgaria decreased slightly between 2001 and 2003. In Russia, the changes in the gender wage gap, between 2003 and 2007 have been negligible and the gap in 2007 was especially pronounced for the lower paid females. The observed positive and significant differential found in Russia was larger in the middle of the distribution. The estimated differential for Tajikistan confirmed Johnes (2002) findings that the gender wage gap in the country is one of the highest among the countries in transition. In the case of Serbia, the gender wage gap has decreased in 2007, but the discrimination component of the gap was still pronounced especially at the higher part of the wage distribution. However, compared to the rest of the countries examined in this Chapter, women in Serbia seemed to fare quite well in terms of their relative earnings.

Very little of the measured wag gap in Bulgaria, Russia and Serbia can be explained by the variables used in the model and this is in line with previous findings in the literature. The large proportion of the gap in Tajikistan – some about 80% – appears to be the result of the 'unexplained' component, 20% of which may be labelled as female disadvantage.

CHAPTER SEVEN APPENDIX

TABLE A7.1

Descriptive statistics - *Bulgarian* working samples, 2001 and 2003

		2001					2003				
		MALE		FEMALE			MALE		FEMALE		
Variable	Description	Mean	Std. Dev.	Mean	Std. Dev.	t-stat	Mean	Std. Dev.	Mean	Std. Dev.	t-stat
lhwage	log of hourly wage	1.519	0.668	1.342	0.539	5.51	1.622	0.558	1.460	0.519	7.45
exp	potential experience	21.790	11.739	20.250	10.799	2.58	19.972	12.264	18.653	11.179	2.79
expsq	exp squared	612.427	544.666	526.506	440.726	3.28	549.182	558.242	472.788	457.067	3.71
tenure1	1 if <1 year	0.136	0.343	0.159	0.366	-1.26	0.270	0.444	0.226	0.418	2.53
tenure2	1 if 1-2 years of experience	0.071	0.258	0.075	0.263	-0.24	0.177	0.382	0.163	0.369	0.97
tenure3	1 if 3-5 years of experience	0.148	0.355	0.133	0.339	0.84	0.191	0.393	0.206	0.404	-0.94
tenure4	1 if 6-10 years of experience	0.255	0.436	0.240	0.427	0.67	0.124	0.330	0.142	0.349	-1.33
tenure5	1 if >10 years of experience	0.390	0.488	0.394	0.489	-0.15	0.238	0.426	0.263	0.440	-1.46
married	1 if married	0.760	0.427	0.726	0.446	1.45	0.680	0.467	0.728	0.445	-2.65
primary	1 if primary	0.143	0.350	0.102	0.302	2.38	0.154	0.361	0.105	0.307	3.56
secondary	1 if secondary technical	0.635	0.482	0.557	0.497	3.02	0.620	0.486	0.532	0.499	4.42
university	1 if university	0.137	0.344	0.207	0.406	-3.53	0.167	0.374	0.256	0.437	-5.46
public	1 if in public sector	0.432	0.496	0.444	0.497	-0.46	0.306	0.461	0.390	0.488	-4.36
north_west	1 if North West region	0.134	0.341	0.106	0.308	1.67	0.100	0.299	0.116	0.321	-1.35
north_east	1 if North East region	0.133	0.340	0.135	0.342	-0.13	0.148	0.355	0.130	0.336	1.31
central	1 if Central	0.110	0.313	0.126	0.332	-0.92	0.122	0.327	0.118	0.323	0.29
south_west	1 if South West	0.222	0.416	0.267	0.442	-1.95	0.194	0.395	0.226	0.418	-1.97
south_east	1 if South East	0.217	0.412	0.175	0.380	1.99	0.212	0.409	0.185	0.389	1.66
sofia_region	1 if Sofia region	0.184	0.388	0.192	0.394	-0.38	0.225	0.417	0.224	0.417	0.01
N		729		709			1296		1186		

Source: Bulgarian LSMS, 2001 and 2003.
Notes: Descriptive statistics relate to decomposition samples; t-stat-tests for the hypothesis that the means of the two groups are equal.

TABLE A7.2

Descriptive statistics - *Russian* working samples, 2003 and 2007

		2003					2007				
		MALE		*FEMALE*			*MALE*		*FEMALE*		
Variable	Description	Mean	Std. Dev.	Mean	Std. Dev.	t-Stat	Mean	Std. Dev.	Mean	Std. Dev.	t-Stat
lhwage	log of hourly wage	3.149	0.915	2.864	0.857	10.32	3.997	0.747	3.774	0.729	9.81
exp	potential experience	16.158	12.624	16.591	12.521	-1.11	17.935	12.001	18.125	11.876	-0.52
expsq	expsquared	420.386	499.108	431.956	481.583	-0.76	465.597	499.335	469.492	487.815	-0.26
tenure1	=1 if <1 year	0.026	0.158	0.037	0.189	-2.10	0.201	0.401	0.171	0.377	2.48
tenure2	=1 if 1-2 years of experience	0.055	0.227	0.060	0.238	-0.77	0.311	0.463	0.245	0.430	4.27
tenure3	=1 if 3-5 years of experience	0.093	0.290	0.092	0.288	0.14	0.180	0.384	0.156	0.363	1.81
tenure4	=1 if 6-10 years of experience	0.153	0.360	0.116	0.321	3.44	0.170	0.376	0.155	0.362	1.14
tenure5	=1 if >10 years of experience	0.662	0.473	0.682	0.466	-1.41	0.270	0.444	0.389	0.488	-7.45
married	=1 if married	0.823	0.382	0.663	0.473	11.88	0.813	0.390	0.639	0.481	12.96
educ1	=if PTU diploma	0.342	0.475	0.206	0.404	10.04	0.327	0.469	0.198	0.398	9.65
educ2	=1 if vocational with no secondary	0.113	0.316	0.072	0.258	4.61	0.119	0.324	0.066	0.248	6.05
educ3	=1 if vocation secondary	0.243	0.429	0.153	0.361	7.34	0.290	0.454	0.188	0.391	7.86
educ4	=1 if technical college	0.222	0.416	0.376	0.485	-10.90	0.264	0.441	0.460	0.499	-13.41
educ5	=1 if university completed	0.207	0.405	0.271	0.445	-4.85	0.268	0.443	0.358	0.480	-6.32
public	=1 if in public sector	0.585	0.493	0.676	0.468	-5.67	0.485	0.500	0.610	0.488	-8.08
central	=1 if Central region	0.210	0.407	0.207	0.405	0.22	0.201	0.401	0.194	0.395	0.59
moscow	=1 if Moscow	0.200	0.400	0.208	0.406	-0.60	0.173	0.378	0.186	0.389	-1.14
urals	=1 if Ural	0.158	0.365	0.150	0.357	0.75	0.147	0.354	0.158	0.365	-1.04
northern	=1 if Northern region	0.066	0.248	0.067	0.250	-0.16	0.070	0.255	0.079	0.271	-1.18
northcav	=1 if North Cavcas	0.058	0.233	0.060	0.237	-0.28	0.089	0.284	0.086	0.280	0.36
wsiberian	=1 if West Siberian	0.063	0.244	0.062	0.242	0.14	0.060	0.238	0.067	0.251	-0.98
east	=1 if East region	0.072	0.258	0.069	0.254	0.36	0.072	0.259	0.060	0.238	1.50
volga	=1 if Volga	0.100	0.300	0.110	0.313	-1.00	0.093	0.291	0.102	0.302	-0.89
arrears	=1 if arrears effect	0.152	0.359	0.114	0.318	3.42	0.055	0.229	0.043	0.202	1.89
N		1907		2261			1917		2315		

Source: Russia, RLMS, 2003 and 2007.

Notes: Descriptive statistics relate to decomposition samples; *t*-stat-tests for the hypothesis that the means of the two groups are equal.

TABLE A7.3

Descriptive statistics - *Tajikistan* working samples, 2003 and 2007

		2003					2007				
Tajikistan		**MALE**		**FEMALE**			**MALE**		**FEMALE**		
Variables	Description	Mean	Std. Dev.	Mean	Std. Dev.	t-stat	Mean	Std. Dev.	Mean	Std. Dev.	t-stat
lhwage	log of hourly wage	3.555	1.022	2.911	0.947	21.87	5.364	1.090	4.447	1.001	32.99
exp	potential experience	19.444	12.156	18.456	11.576	2.79	21.594	12.380	19.722	11.671	5.87
expsq	exp squared	525.810	584.678	474.539	510.768	3.10	619.539	627.951	525.097	530.108	6.04
tenure1	=1 if < 1 year	0.197	0.398	0.178	0.383	1.58	0.247	0.431	0.126	0.332	11.52
tenure2	=1 if 1-2 years of experience	0.155	0.362	0.160	0.367	2.25	0.176	0.381	0.183	0.387	5.44
tenure3	=1 if 3-5 years of experience	0.205	0.403	0.199	0.400	-0.47	0.211	0.408	0.259	0.438	-0.70
tenure4	=1 if 6-10 years of experience	0.145	0.352	0.138	0.345	0.44	0.120	0.324	0.136	0.343	-4.38
tenure5	=1 if >10 years of experience	0.298	0.457	0.324	0.468	0.72	0.247	0.431	0.296	0.457	-1.86
married	=1 if married	0.814	0.389	0.633	0.482	14.33	0.827	0.379	0.612	0.487	19.48
educ1	=1 if basic	0.020	0.140	0.041	0.198	-4.30	0.018	0.134	0.038	0.191	-4.75
educ2	=1 if primary	0.090	0.286	0.129	0.335	-4.30	0.094	0.291	0.179	0.383	-9.96
educ3	=1 if secondary general	0.490	0.500	0.561	0.496	-4.83	0.440	0.496	0.482	0.500	-3.22
educ4	=1 if secondary special	0.096	0.295	0.101	0.301	-0.51	0.117	0.322	0.116	0.320	0.15
educ5	=1 if secondary technical	0.083	0.276	0.030	0.171	7.38	0.102	0.303	0.029	0.169	10.49
educ6	=1 if university	0.214	0.410	0.131	0.337	7.27	0.226	0.418	0.152	0.359	7.05
public	=1 if public sector	0.465	0.499	0.443	0.497	1.46	0.346	0.476	0.449	0.498	-8.12
oblast1	=1 if Dushanbe	0.085	0.280	0.108	0.311	-2.55	0.166	0.372	0.167	0.373	-0.12
oblast2	=1 if Sogd	0.297	0.457	0.266	0.442	2.29	0.184	0.388	0.166	0.372	1.87
oblast3	=1 if Khatlon	0.368	0.482	0.395	0.489	-1.87	0.334	0.472	0.436	0.496	-8.03
oblast4	=1 if Rrp	0.130	0.336	0.127	0.333	0.26	0.237	0.425	0.134	0.341	9.80
oblast5	=1 if Gbao	0.120	0.325	0.103	0.304	1.79	0.079	0.269	0.097	0.296	-2.49
N		3006		1832			4133		2240		

Source: Tajikistan TLSS, 2003 and 2007.

Notes: Descriptive statistics relate to decomposition samples; *t*-stat-tests for the hypothesis that the means of the two groups are equal.

TABLE A7.4

Descriptive statistics - *Serbian* working samples, 2003 and 2007

		2003 Working Sample					2007 Working Sample				
		MALE		FEMALE			MALE		FEMALE		
Variable	Description	Mean	Std. Dev.	Mean	Std. Dev.	t-stat	Mean	Std. Dev.	Mean	Std. Dev.	t-stat
lhwage	log of hourly wage	4.135	0.713	3.974	0.697	5.52	4.606	1.009	4.570	0.947	1.30
exp	potential experience	25.164	12.970	24.146	12.485	1.93	28.413	12.808	26.904	11.799	4.46
expsq	exp squared	801.330	755.500	738.768	759.920	2.00	971.299	817.072	862.961	709.864	5.07
tenure1	=1 if < 1 year	0.218	0.413	0.263	0.441	-2.60	0.086	0.280	0.089	0.285	-0.37
tenure2	=1 if 1-2 years	0.029	0.169	0.032	0.175	-0.30	0.122	0.327	0.149	0.356	-2.74
tenure3	=1 if 3-5 years	0.080	0.272	0.072	0.259	0.75	0.177	0.381	0.161	0.368	1.43
tenure4	=1 if 6-10 years	0.098	0.297	0.118	0.323	-1.60	0.172	0.378	0.183	0.387	-0.98
tenure5	=1 if >10 years	0.575	0.495	0.515	0.500	2.92	0.443	0.497	0.418	0.493	1.77
married	=1 if married	0.749	0.434	0.709	0.454	2.17	0.724	0.447	0.706	0.456	1.36
educ1	=1 if primary	0.145	0.353	0.150	0.358	-0.35	0.170	0.376	0.134	0.340	3.56
educ2	=1 if vocational (1-2 years)	0.027	0.161	0.025	0.157	0.18	0.006	0.077	0.009	0.094	-1.15
educ3	=1 if secondary (3 years)	0.287	0.453	0.169	0.375	6.80	0.220	0.414	0.116	0.321	9.96
educ4	=1 if secondary (4 years)	0.310	0.463	0.351	0.477	-2.12	0.379	0.485	0.416	0.493	-2.59
educ5	=1 if gymnasium	0.025	0.155	0.041	0.198	-2.25	0.028	0.164	0.053	0.223	-4.26
educ6	=1 if post-secondary	0.074	0.261	0.083	0.277	-0.88	0.059	0.236	0.095	0.294	-4.59
educ7	=1 if university	0.084	0.277	0.121	0.326	-2.92	0.092	0.289	0.137	0.343	-4.80
public	=1 if in public sector	0.457	0.498	0.506	0.500	-2.39	0.320	0.467	0.414	0.493	-6.72
region1	=1 if in Belgrade	0.143	0.350	0.191	0.393	-3.15	0.157	0.363	0.207	0.405	-4.46
region2	=1 if in Vojvodina	0.261	0.439	0.273	0.446	-0.67	0.246	0.431	0.242	0.428	0.32
region3	=1 if in West Serbia	0.119	0.324	0.092	0.290	2.10	0.142	0.350	0.119	0.323	2.47
region4	=1 if in Šumadija i Pomoravlje	0.205	0.404	0.196	0.397	0.56	0.192	0.394	0.176	0.381	1.39
region5	=1 if East Serbia	0.104	0.306	0.098	0.297	0.55	0.113	0.316	0.125	0.331	-1.33
region6	=1 if South-East Serbia	0.166	0.373	0.149	0.357	1.13	0.151	0.358	0.132	0.338	1.93
N		1466		984			2983		2014		

Source: Serbia LSMS, 2003 and 2007.

Notes: Descriptive statistics relate to decomposition samples, t-stat-tests for the hypothesis that the means of the two groups are equal.

FIGURE A7.1: Kernel density of the wage distribution, Bulgaria, 2001 and 2003

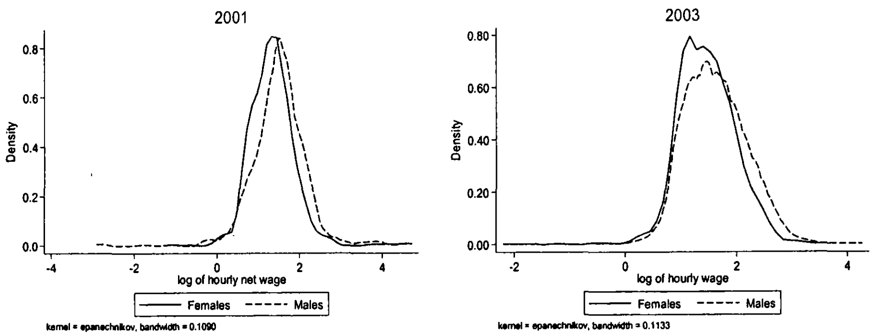


FIGURE A7.2: Kernel density of the wage distribution, Russia, 2003 and 2007

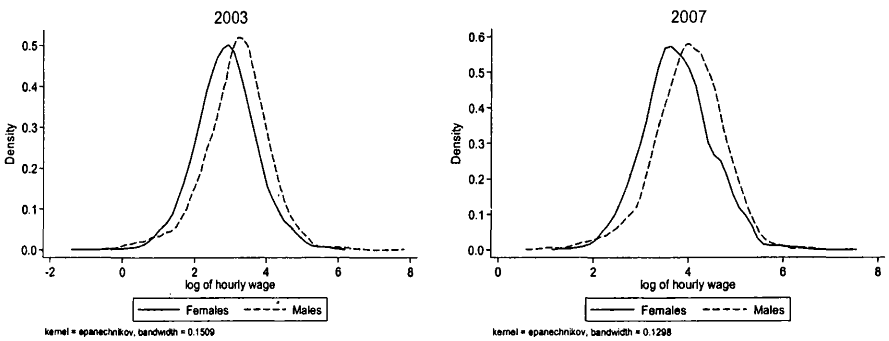


FIGURE A7.3: Kernel density of the wage distribution, Tajikistan, 2003 and 2007

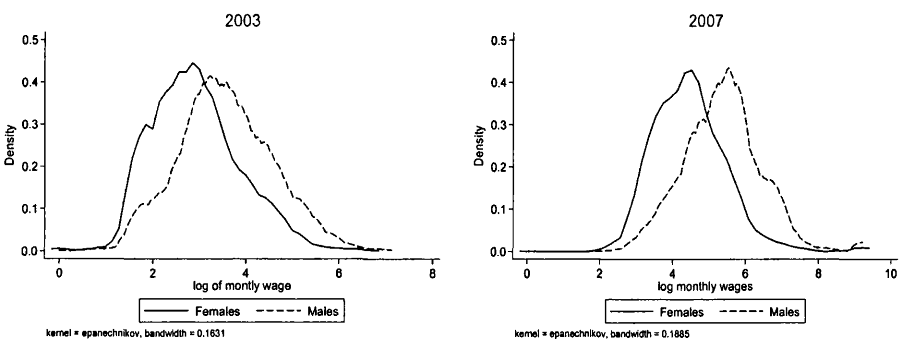
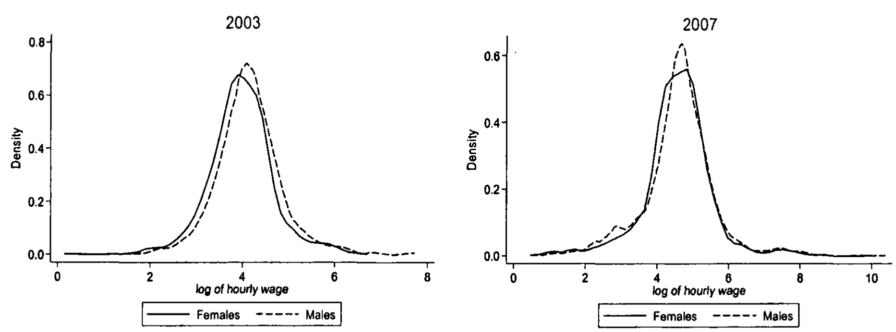
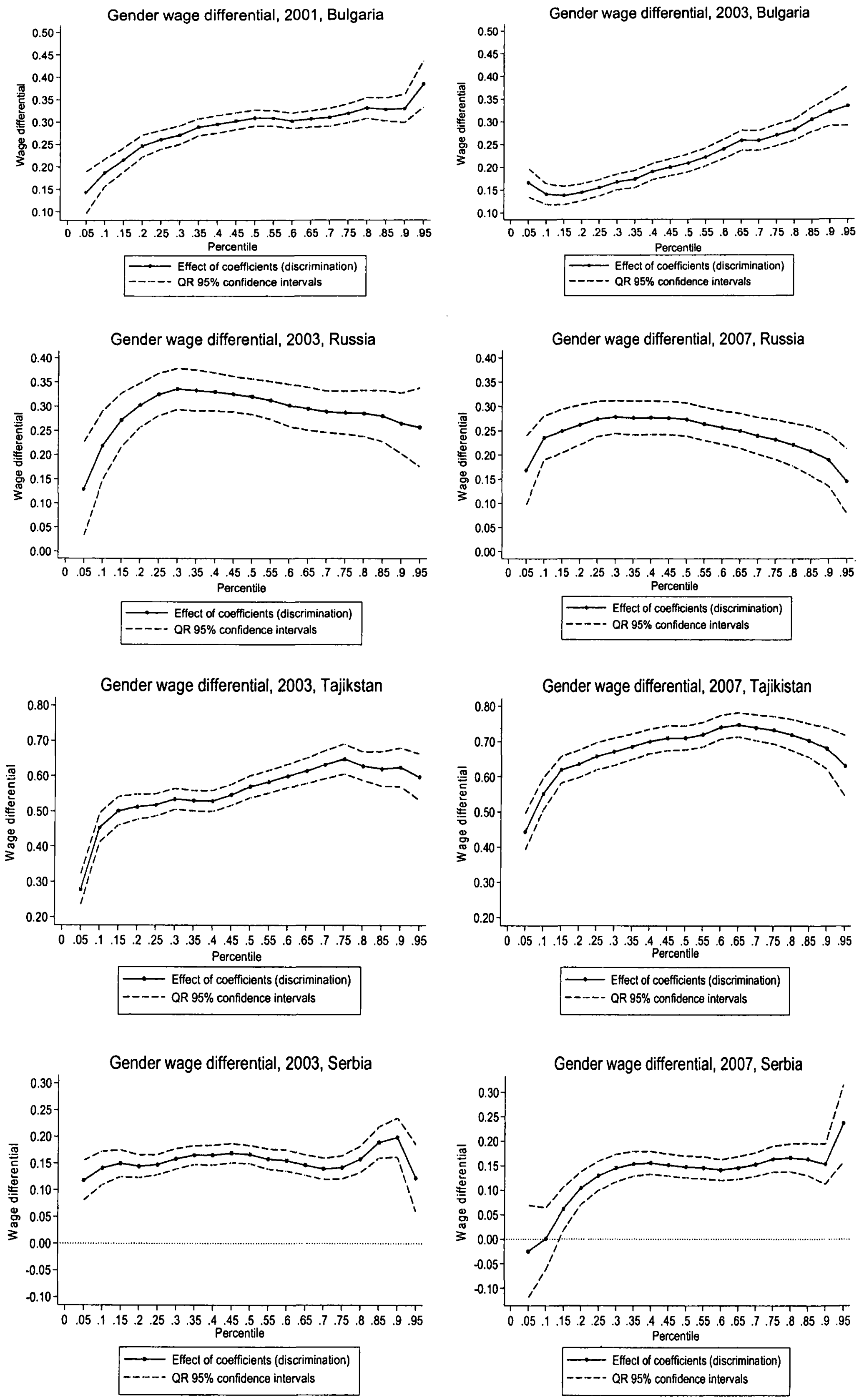


FIGURE A7.4: Kernel density estimates of wage distribution, Serbia, 2003 and 2007



Note: Density functions estimated using an Epanechnikov kernel estimator; we apply Kolmogorov-Smirnov test of the equality of distributions. From the tests it was apparent that the wages across gender do not have the same distribution functions.

FIGURE A7.5: The estimated coefficient effect by countries



ESSAY THREE

PUBLIC – PRIVATE WAGE DIFFERENTIAL: AN ANALYSIS BASED ON CONDITIONAL WAGE DISTRIBUTION

CHAPTER EIGHT

PUBLIC-PRIVATE WAGE DIFFERENTIAL – THEORETICAL REASONS AND DIFFERENT ESTIMATION APPROACHES

8.1. Introduction

Public sector employment accounts for a significant share of total employment and it plays an important role in economic performance. It is often considered that public sector workers are better paid than private sector workers. The idea that public sector workers are overpaid can possibly be attributed to the inelasticity of demand for public services, the inflexibility of public sector wages to market conditions, and the bargaining power of public sector union. However, it is argued in this Essay that looking at the average wages between the public and private sector provides a misleading picture. The recent empirical evidence shows that the highest premium for public sector employees is found at the lower tail of the earnings distribution than at the top where the differential is often negative. Similarly, the findings typically show a higher wage premium for women compared to men[119].

The empirical findings in this study contribute to the literature in several ways. First, as the evidence from transition economies are not sufficiently comprehensive, empirical examination of the public-private wage differential in Bulgaria, Serbia, Russia and Tajikistan is required. As we have discussed, these countries moved from totally centralized economy to a free market economy and this led to dramatic increase in the private sector employment. This might indicate a reallocation of labour resulting from the new process of wage determination. Second, while most of the studies on the public sector wage differential focus only at the mean of the wage distribution, we use QR decomposition procedure to analyse the entire conditional wage distribution. The differential identified by the proposed analysis accounts for important compositional differences between two sectors and allows us to quantify the size of the public sector differential at different points in the earnings distribution. Third, given that there is a choice being made by workers whether to

[119] This literature includes, for instance, Disney and Gosling (1998), Blackaby *et al.* (1999), Elliott and Bender (1997) for UK, Melly (2005) for Germany; Poterba and Rueben (1994) for the U.S.; Mueller (1998) for Canada; Lucifora and Meurs (2006) for Italy, France and UK; Bonjour and Gerfin (2001) for Switzerland.

work in the public or private sector, we allow at the same time for endogeneity of the sector choice and for heterogeneous public sector gap at different points of the distribution. Finally, we adopt a time-wise decomposition framework in order to identify the main factors contributing to the change in the ratio of the Bulgarian and Russian public to private sector earnings over time. It was not possible, however, to obtain a longer time span of data for each country, as required by the analysis over time. Due to a lack of pre-1990 data information for Serbia and Tajikistan, it was not possible to perform the time-wise decomposition analysis for these two countries. By focusing on two periods, we compare the sector wage differentials before transitional reforms (1986 in Bulgaria) or early transition (1994 for Russia) with those in the late transition. As far as it is known, the present study is the first that both controls for endogeneous sector choice and analyses the public sector pay gap across the distribution in the selected transition countires. The empirical analysis will be of interest for policy makers, practitioners and in particular those interested in what has been happening to the public and private sector workers payment in these countries.

This Chapter provides an overview of the empirical literature related to the public-sector wage differential. Section 8.2 explains theoretical background of the public-private wage differences, pointing out several factors which might induce differences between wage rates in the public and private sector. Section 8.3 discusses the main research methodologies used to estimate the differential. Section 8.4 reviews previous empirical studies in this field and finally Section 8.5 provides some conclusion remarks.

8.2. Theoretical explanations for public-private differential

In the literature, a number of reasons have been given for the existence of earnings differentials between public and private sector. First, the public sector organisations usually have different aims and are subject to political constraints rather than profit-maximization[120]. In addition, public sector decision makers could attempt to implement an equal pay policy for their own employees – with the objective of

[120] The private sector labour market is one that is generally based on the premise that firms want to be a successful as possible and aim to maximise their profit. Profit will be not maximised in the case where firms are overcompensating employers for their work and so wages in the private sector generally reflect the value of the employees to firms.

reducing the degree of wage discrimination elsewhere in the economy (Gregory and Borland, 1999). Wages in the public sector may also be used to achieve equity and fairness. Second, the institutional environment for wage setting may differ between two sectors (Cai and Liu, 2010). There could be an imperfect labour market in the public sector. Union density is often higher in the public sector than in the private sector (Bender 1998). On the other hand, public sector may be able to pay more since wages are only subjected to a price floor because of private sector competition (Mueller, 1998). Gunderson (1979) notes that the main difference between the two sectors is that the profit constraint on wages is replaced by the ultimate political constraints of competing with budget allocations with other interest groups and with taxpayers over the size of the government's budget. He explains that employers justify higher pay by having the employees participate in vote-producing activities. The author indicates that there is a significant influence of the relatively inelastic labour demand curve in the public sector labour market which unions exploit to gain higher wages for public sector workers (Bender 1998). Moreover, wages of public and private sector employees can differ because they are paid differently and because they have different skills or work in different jobs. Public sector workers do systematically different jobs from those in the private sector. For example, some jobs are often seen as 'public sector job': public administration, nursing, teaching or security (police, armed force etc.).

Many studies (Schager, 1993; Katz and Kreuger, 1993) indicate that male low skilled worker earned more in the public sector than in the private sector. Although unskilled workers in the public sector typically received much higher rates of pay than public sector counterparts, as we move up the conditional hierarchy to more skilled and highly qualified jobs, the wage advantages slowly declines so that at the top of hierarchy public sector workers are much less well paid than their private sector counterparts. The wage structure is also much more compressed in the public sector than in the private sector (Elliott and Bender, 1997). Moore and Raisian (1991) give other reasons for explaining the differential. They explain that the premium is due to skill differentials which, in the long-run, should be the only reason for wages differential. Other reasons include short run market disequilibria, unions taking advantage of the low labour demand elasticity, political action by unions to increase demand for public sector goods and relative discrimination.

Gregory (1990) indicates four reasons for the existence of public sector wages being different than private sector wages in Britain. The first is the adherence of the 'Good Employer Obligation' policy for the past 100 years. The second reason is the 'Bargaining Power of Public Sector Unions'. Public sector unions have a very strong bargaining position. On the other hand, according Gregory (1990), public sector wages may be lower than private sector due to the nature of public service employment. Public sector employees often have greater job satisfaction, job security, longer holidays, and more generous pensions than their private sector counterparts. The reason is that wages could be subject to 'Government and National Economic Policy'. Borjas (2003) states that given the remarkable changes in the wage structure that occurred over the past 20 years, it is unlikely that the wage structure evolved in similar ways in both sectors, and therefore there could be sizable differences in the trend of the public-private sector pay gap for workers in different skill groups.

Earnings gap may also reflect differences in the nature of compensation packages in the sectors, with the public sector in former communist countries typically offering more fringe benefits. In addition, there are differences in the existing formalized agreements between the public and private sector. Civil servants and public employees normally enjoy generous benefits. Private sector workers most often benefit from pension provisions only if they belong to large firms and different formulas, eligibility rules and funding mechanisms determine actual benefits. The study by Estrin *et al.* (1997) show that new private enterprises in Poland provide significantly fewer numbers of benefits to employees than state-owned firms. Thus, non-wage benefits raise the effective wage paid to the public sector workers. As a result, it may force private firms either to provide the adequate social provision or to compensate for lower non-wage benefits by paying higher wages (Adamchik and Bedi, 2000).

As the theoretical discussion on explanations for public-private wage differential, given in this Chapter focuses on what has been observed in western economies, an important question is do we necessarily expect the same arguments to apply equally to centrally planned economies in their pre-transition stage. They did not operate like western labour markets and characteristics of sectors in the case of transition economies are different from those common for stable western economies. If in stable economies, public sector employee's decision to work in the public sector is explained by relatively generous fringe benefits and job security that the sector offers, in the case of transition, most often additional monetary compensations in a form of informal payments motivate individuals to work in this sector (Chukmaitova, 2011).

To summarise, wage differential between the public and private sector may be explained by several factors. On the one hand, the occupational composition of the public and private sector workforce differs. On the other hand, the distinction between both sectors lies in the different use of the incentive-based payments. Generally, public sector provides more non-wage benefits, such as better job protection or lengthier paid vacations, and better pension schemes. Finally, centrally planned market structure and wage settlements are important elements. While wages in the private sector are based on profit considerations, wages in the public sector are decided by the government.

8.3. Different estimation approaches

A variety of approaches to compare earnings of individual workers in the public and private sector have been applied. Several studies, such as Ehrenberg and Goldstein (1975), Ehrenberg and Schwarz (1983), Ehrenberg and Smith (1981) and Gyourko and Tracy (1988), include a measure of taxes or government ability to pay in the wage equation. A similar argument could be made with the inclusion of grants from central to local governments as in Ashenfelter (1977). Other studies use the effect of unions on public sector wages, although there is very little formal modelling of the effects of public sector union on wages. Some studies utilise occupation-level or individual level data to examine differences in average earnings between public and private sector workers at a particular point in time. Other studies use the same occupational-level data, but have focused on changes over time in relative earnings of public and private sector employees. One approach is to include dummy variables for worker's sector of employment in the earnings regression. The problem with this method is that it models the effect of sector as an 'intercept' effect – returns to other productivity related characteristics and job attributes are restricted to be equal across sectors.

An alternative approach involves estimating separate earnings regression for public and private sector employees and decomposing the wage differential into effects of differences in average worker characteristics and job attributes between sectors, and differences in the returns to worker characteristics and job attributes between sectors. However, the allocation of workers to the public or private sector may not necessarily be exogenous to the wage formation process. There is a choice that workers have to make whether to work in the public sector or in the private sector. A worker will choose to be in the public or private sector depending on the greatest advantage earned in each of these sectors. More educated individuals prefer to work in the private sector, where their performance and skills are better rewarded (Adamchik and Bedi (2000) show this for Poland). More experienced individuals prefer public sector employment, as documented by Falaris (2004) for Bulgaria. Women may also prefer to work in the public sector, where flexibility in hours worked is higher. Therefore, the wage equation is not independent of the selection (sectoral choice) process (Choudhury, 1994). This is the well documented self-selection issue[121] that has been treated in the public-private sector wage differential literature by Van Ophem (1993) and Hartog and Oosterbeek (1993), and Dustmann and van Soest (1998). The econometric solution to this problem is to estimate a sector choice equation together with the wage equations relating to each sector.

To choose appropriate identifying variable the literature attempts to ascertain whether certain characteristics of individual employees increase the probability that they will seek employment in the public sector. Bender and Elliott (1997) argue that identification of the selection equation for sector of employment may be difficult to achieve. To identify the selection equation most studies of worker's choice of sector of employment have used a variable such as age or education and such variables are more appropriate as explanatory variables in the earnings regression. The probability to work in the public sector increases with experience and education. Most authors simply use the same human capital variables in the wage equation and in the sector

[121] This problem is very similar to the one raised by Heckman (1976) for women's wage equation, where participation in the labour market is assumed to depend on their reservation wage.

choice equation, sometimes with arbitrary exclusion restrictions. Others use parent's occupational status or whether father was a civil servant or not, motivating that child learns through imitation of adults living in their neighbourhood (Melly, 2006). Falaris (2004) use as an instrument the amount of land received as a part of Bulgaria's property restitution programme. Jovanovic and Lokshin (2004) use the industry of pre-1992 employment, the number of children and marital status as instruments in the selection equation. Heitmueller (2004) uses union perception as identification in the sector choice equation. In most studies, the data is not rich enough to provide appropriate instruments and identification assumptions are sometimes doubtful. One important non-observable sector choice variable is individual risk aversion. A recent study by Pfeiffer (2008) based on German data show more risk adverse individuals seek public employment and that such risk taking is rewarded with higher payment in the private but not in the public sector. Bargain and Melly (2008) show that public sector premium or penalties are much smaller when correcting for selection. Moreover, their results indicate that across quantiles, the usual result that public sector compresses the wage distribution partly disappears when controlling for selection. Dustmann and Van Soest (1998) conclude that correcting for non-random selection is important, but only useful if appropriate instrumental variables are available.

However, there is a lack of consensus in the literature on the role that sample selection effects might have on sector wage gap estimates. For instance, Choudhury (1994) and Hartog and Oosterbeek (1993) show the selection terms for choice of sector to be insignificant in the private sector wage equation and significant in the public sector equation, whereas Borland *et al.* (1998) find selection to be important only for female private sector employment. Some other authors argue that controlling for one type of selection in the earnings equation only and ignoring non-labour force participants may still lead to biased estimates (Co *et al.*, 1999). Additionally, findings from decomposition analysis applying sample selection correction will be sensitive to the interpretation of the role of the selectivity term in the wage equation. Controlling for selection, Disney and Gosling (2003) find that public sector premium becomes insignificant for men and remains positive for women. More educated men seem to select themselves into the public sector on some negative characteristics, as cross-section estimates of the public pay gap is negative and becomes positive once selection is controlled for (Vinay and Turon, 2007). In these studies, identification of the public wage premium in the presence of non-random selection is achieved with either functional form assumptions or with some instrumental variable procedure appealing to variables such as family background or variations in public sector status arising from privatisation (Disney and Gosling, 2003) as instruments for selection into the public sector. Stelcner *et al.* (1989), Adamchik and Bedi (2000) and Heitmuller (2004) use an endogenous switching regression model which allows controlling for the possible sample selection bias. Results of these estimations show that selection bias between working and not working is not an issue, but the selection of workers between public and private employment is not random (Leping, 2006).

A second set of issues related to findings from decomposition is that estimated size of the public sector wage premium appears to be sensitive to the number and type of explanatory variables included in earnings equation. Increasing the number of explanatory variables in the wage regression generally increase the proportion of the raw wage differential between public and private sector that is attributed to differences in worker characteristics, and reduces the size of the public sector wage premium due to differences in returns between sectors (Bender and Elliott, 1997).

The size of the public sector wage premium also is affected by whether particular explanatory variables are included in the earnings regression. Some studies indicate that including establishment size as an explanatory variable in the earning regression reduces the estimated size of the public sector wage premium[122]. Moulton (1990) and Poterba and Rueben (1999) find that the size of the public sector wage premium is sensitive to whether detailed controls for worker's occupational classification are included in the earnings regression.

Another type of problem arises where employees' earnings are affected by the unobserved job attributes. The effect of differences in unobserved job attributes between sectors will be captured by differences in the intercept terms in the earnings regression. Where compensating payments for job attributes differ between sectors, estimates of the extent of over-payment of public sector employees will be biased

[122] See Belman and Heywood, 1990, 1993.

(Gregory and Borland, 1999). Therefore it is necessary to use a method of controlling for differences in the productivity related characteristics and job attributes of public and private sector employees.

8.4. Previous studies

8.4.1. Previous studies in transition countries

The public-private wage differential has been investigated across different countries and time periods. However, much less is known about the public wage setting behaviour in the case of transition countries[123]. The sector wage differential in transition countries have been estimated by Adamchik and Bedi (2000), Reilly (2003), Jurada (2003), and Leping (2006). Existing evidence on sector differential is provided by Rutkowski (1996), Coricelli *et al.* (1995) in Poland, Lokshin and Jovanovic (2003) in Yugoslavia, Falaris (2004) in Bulgaria, Voinea and Mihaescu (2012) in Romania. Adamchik and Bedi (2000) find private sector wage advantage in Poland, which is particularly pronounced at university level. For males, they observe positive selection in both public and private sectors. Voinea and Mihaescu (2012) find that compared to the private sector, public sector workers in Romania earn, on average, about 10% more. A possible explanation according the latter authors is the higher prevalence of the informal economy in the private sector. Workers in the private sector receive informal payments – twice as many employees are paid with the minimum wage in the private sector compared to the public sector. Disney (2007) indicates that in transition economies, where a public sector with regulated pay and unregulated private sector co-exist, the public pay effect may be negative. He also adds that an essential difficulty in many transition studies of this kind is that household or labour force survey data used to investigate pay structure are disproportionately collected from households that are predominantly formal sector or salaried (tax-paying) earners.

In the case of Bulgaria, Falaris (2004) treats the sector of employment as endogenous and corrects for selectivity in the wage equation. He finds evidence of positive selection in the public sector wage equations for men and women in 1995 but negative selectivity in the private sector wage equation for women. The author

[123] For a survey in developing countries see Van der Gaag *et al.* (1988).

concludes that men and women who work in the public sector receive highest financial benefit, but private sector employed women suffer an implied wage penalty. Falaris (2004) also states that in the private sector women perform mostly low skill work requiring neither experience nor schooling; however women receive positive returns to both experience and schooling in the public sector.

Estimating public-private wage differential in Estonia, Leping (2006) finds negative differential during the early transition, which substantially decreased over time. The author shows that employees with low potential wages tend to gain more from working in the public sector compared to workers with high potential wages. Women in Estonia also appear to benefit more from working in the public sector than men. Furthermore, Gorodnichenko and Sabirianova (2005) find that public sector employees in Ukraine are significantly underpaid compared to workers in the private sector. The average wage gap is 24%-32% and is largest (up to 60%) among the most productive and highly paid workers. Moreover, the gap has not diminished in the period 1997-2003 despite similar rates of voluntary separations and labour mobility in sectors.

Reilly (2000) provides evidence for private sector premium in Serbia using Yugoslav Labour Force Survey data from 1995 to 2000 for male employees. The results suggest that the hourly wage premium for a private sector job at the median of the conditional wage distribution is over 20% in 1995, not significantly different from zero in 1996, 1997 and 1999, and nearly 24% in 1998. The QR estimates show 17% premium for those at the median and almost 70% premium for private workers at the top of distribution. Similarly, Lokshin and Jovanović (2003) find on average private sector premium of 9.4% for males and 4% for females in the former Yugoslavia. Their explanation is that part of this gap may be offset by the benefits public sector employees receive. Laušev (2009) studies public-private wage differentials in Serbia over the period of transition from 1995 to 2006. The author finds that earnings are more equally distributed in the public sector than in the private sector. He finds significant private sector premium across the distribution from 1998 to 2002. However, that initially growing private sector markup disparity over the period tends to transfer into public sector wage premium. Ognjenović (2011) shows that with advance of the transition, the public sector in Serbia generates wage premium for those who work in that sector compared to the employed in the private sector. The public sector overpaid both men and women and the estimated wage premium for women is lower compared to that for men (22.3% and 25.4% respectively). The only group of workers who are penalized for working in the public sector are those with higher education.

In the case of Russia, Jovanović and Lokshin (2004) investigate wage differentials between the state and the private sector in the city of Moscow in 1997. They show that workers in private sector earn 16% higher wages than workers in the state sector. Wage structure differs across the sectors for both genders. The private sector wage premium for women is higher than for men. In addition, the probability of employment in private sector decreases with age and tenure. Gimpelson and Lukiyanova (2009) show that for the 2000-2004 period wages in public sector are on average 40% lower than in the private sector. Very few studies investigate the public-private sector over time. There are some evidence of widening wage gap between the public and private sectors during the transition period in Hungary and other CEE countries (Kertesi and Köllő, 2002 and Keane and Prasad, 2001).

Overall, the evidence from transition are mixed, with a wage premium in favour of the private sector (Adamchick and Bedi, 2000; Leping, 2006), or wage premium for the public sector workers (Falaris, 2004, Voinea and Mihaescu, 2012 and Ognjenović, 2011). The literature review also supports the fact that there are very few studies which exclusively investigate the public-private wage gap in Bulgaria, Serbia, Russia and Tajikistan, and there is none for Tajikistan. The need to conduct a comprehensive research about the public sector wage differential taking the cross-country perspective is therefore quite important.

8.4.2. Previous studies in Western countries

It is interesting to contrast how the situation we have described for the transition countries compares with those in some western economies, and more specifically with that in United Kingdom (UK). A common finding in the UK is that public sector workers typically earn more than comparable workers in the private sector. Most studies in the UK use two main data sources: the New Earning Survey (NES) available since 1970 and the British Household Panel Survey (BHPS) available from 1991 onwards[124]. Earlier studies by Elliott and Murphy (1987), Gregory (1990) and Elliott and Duffus (1996) have used NES data to track movements in the relative pay of employees in the public and private sectors. Distinguishing between manual and non-manual workers, public corporations and the private sector, Elliott and Murphy (1987) establish that male manual workers earned pay premiums of up to 20% and female manual workers are paid penalties of up to 9%. Non-manual males were close to parity or had marginal penalties whereas equivalent females sustained penalties of up to 20% (Bender and Elliott, 1997).

Research in the mid-1980s and early 1990s show the importance of breaking down the public sector into subsectors of central and local government and public corporations because of the changing composition of some parts of the public sector, particularly in the UK. Foster *et al.* (1984) show that in the UK since the 1970s, public corporations have had the highest wages, while central and local government wages are found to be very volatile. Gregory (1990) uses NES data, covering a similar period (1970 to 1982) and splits the public sector into the same three groups (central government, local government and public corporations). The author does not control for age and occupational composition and shows quite different results – with all groups earnings pay premium except men in manual occupations in central and local government who sustained pay penalties of 6.8% and 8.3% respectively. By looking at the differential over time, it becomes smaller, while there is little evidence of this with the public corporation-private sector premium. Rees and Shah (1995) using General Household Survey (GHS), find that for both males and females the public sector wage differential is positive. The authors decompose the positive wage differences into pay premiums and worker characteristics (education, age, job tenure,

[124] Note that the New Earnings Survey was replaced by the Annual Survey of Hours and Earnings in 2004.

health, and marital status and pension provisions). They conclude that male differential is due to the greater productive characteristics of public sector male workers. Although the pay premium is positive for all groups, the implication is that men in the public sector could raise their wages by switching employment to the private sector, but women doing the same would lose around 30% of their current pay (Allington and Morgan, 2002). Rees and Shah (1995) also find substantial changes in the wage structure in both sectors over time, partially explaining the wide variation in the estimated decomposition.

The second GHS study by Elliott *et al.* (1996) indicates a substantial female differential of 17.7 % but the male differential is now revealed smaller at 4.1 %. Study by Potebra and Rueben (1999), and Mueller (1998) show that wage distributions are compressed in the public sector and public sector workers enjoyed a wage premium at the lower tail of the distribution, but a wage penalty at the upper tail. Elliott and Duffus (1996) analyse changes in the relative pay of public sector employees over the period 1970 to 1992. The study finds a considerable dispersion in the earnings growth of those occupations, and despite the substantial real earnings growth enjoyed by some of the occupations over the period, the pay of almost all public sector groups declined relative to that of the private sector over the period 1980 to 1990.

Bender and Elliott (1997) use the NES and the BHPS to investigate the pay differential across the public and private sector in UK over the period 1991 to 1994. They conclude that there is a divergence between return to sector-specific occupational characteristics. They also find evidence that educational and regional returns become more similar between two sectors over the period. This narrowing seems likely to be due, in part, to the reforms in wage settings in some parts of the public sector but it is also due to developments in the private sector, such as the sharp narrowing of regional wage differences in that sector. The authors conclude that by the end of the examined period, public sector workers are on average 'overpaid' relative to similarly qualified private sector workers.

Bender (2003) estimates separate wage equations for public and private sector distinguishing males and females, and taking into account the possible sample selection due to the sector choice of employees. A decomposition technique based on Belman and Heywood (1993) is applied that takes differences in wage distribution into account. Bender (2003) finds that males at the lower end of the public sector pay distribution are better off than their private counterparts whilst high-paid private sectors males earn more than high-paid public employees. Those female public-sector workers who are at the low end of the distribution are paid less than similarly placed private sector females. Moreover, much of the difference in average wages is due to differences in returns rather than characteristics. This is in contrast to Rees and Shah (1995) findings.

Further extension in the literature is the use of QR methods to examine whether public sector pay 'premia' or 'penalties' differ across the sector earnings distribution. A series of papers for the UK have used QR approach to explore the earnings gap across the distribution (Disney and Gosling, 1998; Blackaby *et al.*, 1999; Yu *et al.*, 2005; Lucifora and Meurs, 2006). It is often concluded that the public sector pay distribution is more compressed than the private sector pay distribution, and this should lead naturally to QR methods finding different penalties or premia across the distribution (Disney, 2007).

In a study covering the period 1991 to 1995, Disney and Gosling (1998) investigate the benefits that public sector workers with different educational qualifications might expect from switching employment to the private sector. Using GHS and NES data, they find that most male public service groups would earn more. Comparing percentile pay for different groups of male and female workers in the public and private sector, they conclude that the top 25% of male graduates in the public sector earned a pay premium while the reminder experienced penalties and would have been much better off switching to the private sector. Blackaby *et al.* (1999) use Labour Force Survey (LFS) data to consider the distribution of public sector pay for the period 1993 to 1995. They find positive and only slightly decreasing public sector premium along the entire wage distribution for both males and females. Specifically, the percentile pay differential spreads of 1.5% to -1.9% for men and 3.3% to 0.1% for women. Men in the top two pay percentiles and women in the 50^{th}, 70^{th} and 80^{th} pay percentiles all paid penalties by remaining in the public sector.

Study by Yu *et al.* (2005) shows that after controlling for characteristics associated with productivity, wages are higher for well paid employees in the private sector and vice versa for the lowly paid in the public sector. Mean regression results indicates that the mean return to working in the public sector is positive or the average wage rate is higher among public sector workers. Yu *et al.* (2005) results show that the effect of public sector employment is large and positive at lower quantiles (above 10 % for the 10^{th} quantile) and still generally positive at the median, but it is negative at the 90^{th} quantile. The wage distribution is much more compressed among public sector workers: low wages (i.e. at the lowest decile) are higher in the public sector, whereas high wages (i.e. at the highest decile) are lower in the public sector compared to the private sector (Yu *et al.*, 2005). The authors show that on average, public sector workers are better paid than private sector workers, but at the same time, the chances of obtaining a high pay are higher in the private sector: a high-wage employees in the public sector may be able to get a better pay in the private sector.

Comparing the public sector pay gap across Britain, France and Italy, Lucifora and Meurs (2006) conclude that differential is sensitive to the choice of quantile, which reject the hypothesis of constant wage differential and that the pattern of the premium varies with both gender and skill. The wage gap estimates show that female are better off being in the public sector, particularly at the lowest deciles, whilst the opposite is true for men at the highest deciles. Lucifora and Meurs (2006) argue that differences in unobserved characteristics may be more important for these employees and conclude that the 'low floor effects' is what distinguishes private sector pay of skilled women in Britain.

Among more recent studies focusing on the European context, Portugal and Centeno (2001) use the 1995 wave of the European Community Household Panel (ECHP) to compare public-private wage differential in the European Union member countries. The authors find that the wage gap is wider in Portugal, Ireland, Luxemburg, Spain and Italy. By contrast, in Denmark, Belgium, Germany and Austria, the differential turns out slightly negative.

8.5. Summary and conclusions

In this Chapter, we have described the theoretical explanations for the differences between the wages in the public and private sector. We have reviewed the empirical literature related to the earnings differential between public and private sector. The main findings from the reviewed studies can be summarised as follows:

• The studies that compare the average earnings outcomes in the public and private sectors generally find positive wage differential for the public sector employees.

• Most recent research has concentrated upon the nature and shape of the public sector wage distribution and has found a premium associated with public sector employment in the lower tail of the wage distribution.

• There is consensus of findings in the Western literature that there is a public sector wage premium but it differes between men and women i.e. public sector workers receive returns to their productivity and job-related characteristics which are associated with higher earnings premium, or on average, civil servant earn more than comparable workers in the private sector. The premium varies across the earnings distribution, and it is sensitive to both specification and choice of econometris procedure.

• Research on the public-private wage differential in transition countries is rather limited and in contrast to the Western literature the findings are mixed. The existing empirical studies highlighted different patterns albeit from differnet countries and different time periods, with a wage premium in favour of the private sector after economic liberalization (Adamchick and Bedi, 2000; Leping, 2006), or wage premium for the public sector workers (Falaris, 2004 and Voinea and Mihaescu, 2012). The evidence from the QR shows that there is a negative wage differential for the higher quantiles, but not significant wage difference for the lower quantiles. However, with the advance of transition, the public sector generates wage premium compared to the employed in the private sector (Ognjenović, 2011).

• The earlier studies ignore the potential source of bias that could arise from selection process between sectors. However, there is a lack of consensus in the literature on the role that sample selection effects might have on the sector wage gap estimates. Bargain and Melly (2008) show that public sector premium or penalties are much smaller when correcting for selection. Disney and Gosling (2003) confirm that public sector premium becomes insignificant for men and remains positive for women after the selection is controlled for. Hyder and Reilly (2005) conclude that the differential based on correcting for endogenous sector choice provide a few insights on the magnitude of the public sector premium across the distribution. According to Dustman and Van Soest (1998) correcting for non-random selection is important, but it is only useful if appropriate identifying variable is available.

• Relatively few empirical studies have investigated changes in the public-private wage differential over time. In particular, articles for transition countries are contradictory and do not provide sufficient evidence on how transition process has affected public-private wage differentials.

CHAPTER NINE

PUBLIC-PRIVATE WAGE DIFFERENTIAL - AN EMPIRICAL ANALYSIS BASED ON CONDITIONAL WAGE DISTRIBUTION

9.1. Introduction

In the context of transition economies, the dramatic economic changes raised a number of issues on the direction toward which the two sectors were moving. The transition led to a marked decrease in both output and employment in the public sector. One major aspect of the economic reforms was to encourage development of a non-state sector in the economy. Various forms of non-public ownership such as private-owned, joint-venture, foreign-invested, share-holding and self-employed became alternatives to former state-owned companies (Démurger *et al.*, 2010). There was a huge amount of restructuring of employment by industrial sector in medium and large enterprises in the early 1990s (Sorm and Terrell, 2000). In the Czech Republic for example, industry (i.e., manufacturing plus utilities) shed about one-fifth of its workforce by 1993 and another 10% between 1993 and 1998[125] (Ham, 2000). Employment grew rapidly in construction, wholesale and retail trade, hotels and restaurants. The expansion of the private sector ownership created new job opportunities. During such transformation, earnings in the private sector can be expected to be higher. However, pushed into the informal sphere of the economy, private sector in most transition countries, effectively avoided the wage regulating provisions (Jovanović and Lokshin, 2004).

The motivation for this study is to complement previous research on public-private wage differentials in transition countries and to extend the analysis by providing empirical evidence on the evolution of the public-private pay gap across the distribution and over time. We use decomposition framework applied to the QR estimates to answer the question of whether employees in the public sector in

[125] By comparison, the net declines in the stock of jobs in these sectors over a similar period, 1994 to 1997, were much smaller in Bulgaria, Estonia, Poland and Romania. (Faggio and Konings,1999).

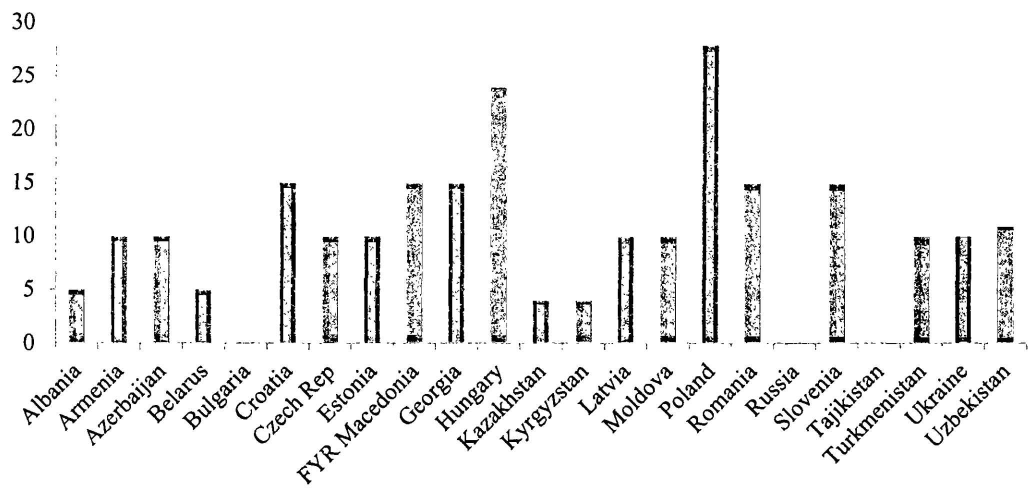
Bulgaria, Serbia, Russia and Tajikistan fare better than those who stayed in the private sector. Given the fact, that there is a choice being made by workers whether to work in the public or private sector, the endogenous sector choice is considered. We also analyse what has been happening to the public sector hourly wage differential at different points in the conditional earnings distribution over time by providing further evidence from Bulgaria and Russia. The main contribution in this Chapter is correcting for endogenous sector choice considering the entire wage distribution.

The structure of the rest of this Chapter is as follows. Section 9.2 looks at the public and private sector employment and earnings in the countries of interest. Section 9.3 describes the data and in Section 9.4 the econometric methodology is presented. Section 9.5 reports the main empirical findings and finally the conclusions and some policy implications are given in Section 9.6.

9.2. Public and private sector employment and wages

We have already discussed that one important conceptual point we should bear in mind in assessing developments in labour relations over the region is that in most transition countries there was an enormous change in the ownership structure of enterprises and organization which led to decline in the share of total employment taken by the public sector. The changing patterns in the labour markets included a great number of people becoming self-employed after losing their jobs and the growing informal sector jobs. The evolution of the private sector was a result of two processes – privatization, by which formerly state owned enterprises were transformed into private ownership and creation of completely new firms. Although the size of the private sector has expanded quite rapidly in most transition countries, this is not true in the case such as Russia, Belarus, Kazakhstan and Kyrgyzstan. The private sector contribution in Bulgaria and Tajikistan for example, were only 10% of 1990 GDP, much lower as compared to other countries such as Poland and Hungary (see Figure 9.1). Salaries and working conditions within public services were allowed to deteriorate in the early 1990s (Standing, 1997). Thus, the pertinent question that needs to be addressed is whether there are wage differential between the public and private sectors in selected countries and what the implications of these differences are.

FIGURE 9.1: Private sector as % of GDP 1990



Source: World Development Indicators (2000) The World Bank.

9.2.1. Bulgaria

To better understand institutional setting in Bulgaria, we briefly discuss the selection of workers into the public sector and how wages are formally settled in both sectors. Bargain and Melly (2008) argue that structural differences between sectors may not necessarily translate into pay level differences. In general, the system of wage formation in Bulgaria is underpinned by several factors like legislation, sector collective bargaining and supplementary company bargaining, the labour market and decisions of company management (Ribarova, 2009). The most important issues subject to negotiation, such as the minimum wage, the average wage in the public sector and the mechanisms of wage setting and determination are discussed at national level and càn consequently be finalized at the branch or at the regional level. Wages are completely centrally set in the activities of the public sector and are entirely subsidized by the state budget (Garibaldi *et al.*, 2001a). Recruitment for public sector occupations in Bulgaria is decentralised and is based on examinations. Wages in the public sector consists of two parts – basic and additional. The basic salary is calculated on the basis of the official's duties. The appointment body defines the wage levels of the basic remuneration of the civil servant taking into account the job level and the evaluation for the individual fulfilment of the obligations after the last attestation. A wage increasing mechanism exists and is based on results of the activity. For the employees under the labour contract some elements of additional pay are negotiated (seniority, overtime work). Some bonuses are awarded because of difficult or dangerous working conditions (Mihaylova, 2008).

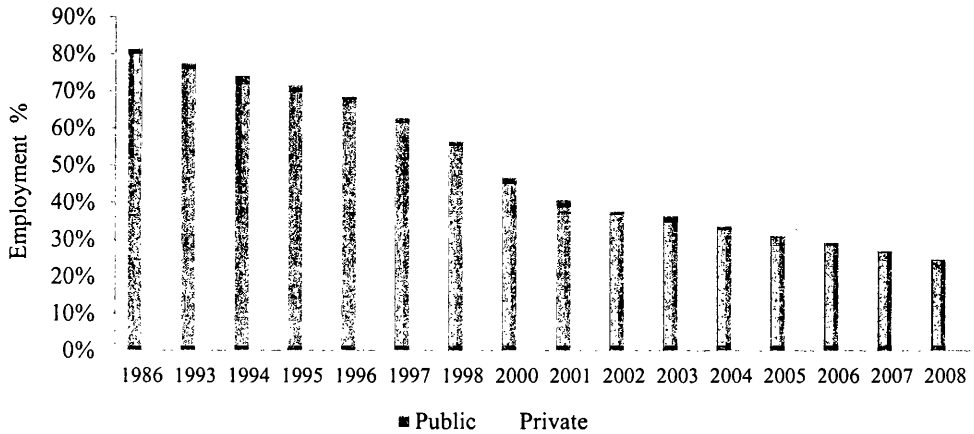
In the private sector, no form of collective bargaining is envisaged and wages tend to be set at the firm level (EFT, 2000). In addition, market forces have determined private sector payment in Bulgaria. Wages in the private sector showed more variation than in the public sector. Although private sector firms tend to pay higher wages than those in public sector, higher wages in the private sector were offset by lower levels of bonuses and non-wage benefits. Employment in the public sector yields at least the minimum wage, whereas reported private-sector wages suggested an average monthly income below the minimum wage (Beleva *et al.*, 1995).

As already discussed, after the transition started wages in the state sector were no longer determined centrally, rather they were influenced by negotiations between the emerging independent labour unions, employers and the government (Falaris, 2004). Jones and Simon (2005) hypothesize that non-competitive forces influenced wage determination process during the planning period. These forces operated through various channels at industry, firm, and regional levels. Based on the Stalinist model of development, with its bias towards heavy industry, there might be industry wage effects. Moreover, coalitions of managers and workers were especially strong in the biggest firms so that their political influence should translate into a firm-size wage premium. Wages in sectors, such as light industry, trade, agriculture and public services, tended to be low and even lower in smaller firms (Jones and Simon, 2005). The last authors also argue that individual differences in human capital in Bulgaria were not expected to play a major role in accounting for wage differences during the planning period and because wages were fixed centrally, employees in more skilled occupations might receive lower wages.

Bulgaria lagged behind most other Eastern European countries in privatisation. Since 1991, the country drafted a series of ambitious privatisation programmes and liberal laws for foreign investment. Privatisation was pursued along three separate lines: restitution of land and urban property to former owners; cash sales of assets and mass (voucher) privatisation. Bulgaria was actually among the first of the Central and Eastern European countries to adopt legislation to attract foreign investors. The Law on Foreign Investment, passed in 1991, was among the most liberal in the region, providing national status to all foreign investors, allowing for 100% foreign ownership, setting very low barriers and offering easy registration. However, as of mid-1996, per capita foreign investment in Bulgaria was less than 6% of that in Hungary and 12% of that in Czech Republic, and it was less than in Albania and Romania (OECD, 1997). The main reason for the low levels of foreign investments was the poor infrastructure and communications. Bulgaria represented a small domestic market with limited endowments of natural resources. Land restitution was problematic and controversial. According to the National Statistics Institute (NSI) data, in 1996 only 18% of the land was officially returned to the owners. An unusually high degree of political instability was another important problem for the privatisation process in Bulgaria.

In the early years of transition and especially in the communist period, the private sector in Bulgaria consisted primarily of small firms. Near the end of the period of central planning, the private sector employed only 6.5% of the workforce which increased to 20% in 1992 (Beleva *et al.*, 1995). After the crisis of 1996 and 1997, stabilization and mass privatization, the private sector in Bulgaria became a larger part of the economy (Falaris, 2004). There was evidence of emergence of significant concentration of ownership in individual Bulgarian firms (Jones and Klinedinst, 2006). At the end of 1996, 313 547 private firms were registered versus 20 177 in the public sector. Approximately 61% of all private firms were in the trade industry (OECD, 1997). There was a massive structural change in the relative proportion of public and private sector employment as shown in Figure 9.2 and Table 9.1. Employment in the public sector decreased from 81% in 1986 to 25% in 2008. Throughout the period 1986 to 2008 public sector employment dropped by 69%. Similarly, employment in private sector rose from 19% in 1986 to 75% in 2008 (see Figure 9.2). The huge restructuring and mass privatization process in public enterprises during the transition might have caused differences in the growth rates of public and private sector wages. Therefore, we could expect that the transition process in Bulgaria may have had an influence on the public-private wage differential.

FIGURE 9.2: Employment as % of total, by sectors in Bulgaria, 1986-2008



Source: Bulgarian National Statistical Institute (NSI) and authors own calculations for 1986.

The decline in the public sector employment was larger for men than for women. Between 1993 and 2006 employment in the public sector dropped from nearly 80% to 33% for women and from 75% to 23% for men. Generally, women in Bulgaria have a higher share of the public sector employment than that of men (see Table 9.1).

TABLE 9.1: Public and private sector employment by gender in Bulgaria

	Males		*Females*	
	Public	Private	Public	Private
1993	75.1	24.7	80.1	19.8
1995	67.2	32.5	75.1	24.6
2000	41.8	57.7	49.9	49.6
2001	37.2	62.3	43.1	56.3
2002	33.4	66.0	39.8	59.9
2003	31.6	68.0	38.0	61.7
2004	27.8	71.9	34.8	65.0
2005	25.3	74.4	33.3	66.6
2006	23.1	76.9	32.8	67.2

Source: UNECE Statistical Division Database.
Notes: Percent of corresponding total for both sectors. The public sector covers all sub-sectors of general government (mainly central, state and local government units, together with social security funds imposed and controlled by those units) and public corporations, i.e. corporations which are subject to control by government units (usually defined by the government owning the majority of shares). The private sector covers private corporations (including those in foreign control), households, and non-profit institutions serving households.

The massive labour shedding, outlined above, was partly associated with reallocation of labour across industries and with the privatization process. As shown in Table 9.2, with the exception of financial services, until 1992 reduction in employment was sizeable across sectors. Employment dynamics by sector started to be more diversified in 1993. While employment decline in industry and construction continued throughout the observed period, job gains were registered in most service sectors and agriculture which was due to the land reform in the country (Garibaldi *et al.*, 2001a).

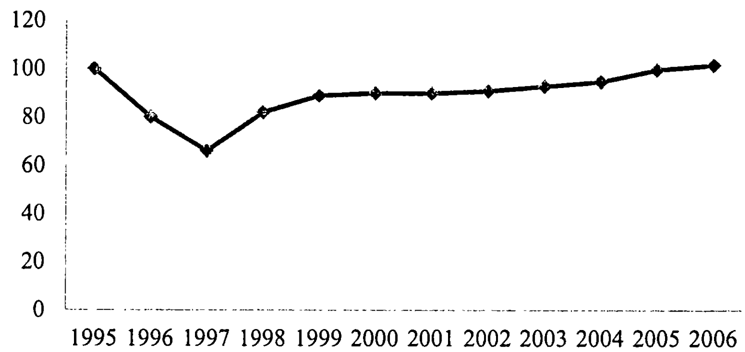
TABLE 9.2: Distribution of employment growth by industry: 1986-1996[1]

Percentage change	1990	1991	1992	1993	1994	1995	1996
Industry	-9.0	-17.9	-13.2	-8.3	-3.7	-2.2	-1.2
Construction	-6.8	-25.0	-19.1	2.3	-8.1	-2.4	-5.6
Agriculture	-6.8	-7.6	-0.3	3.2	5.7	4.3	0.2
Forestry	-11.1	-22.0	-0.9	-17.1	-7.0	-0.2	8.9
Transport	-2.1	-7.8	-13.3	1.9	-4.3	9.3	-2.0
Communications	2.6	-0.7	-1.8	1.1	0.2	1.8	2.8
Trade	-5.4	-8.0	-4.0	0.9	11.1	-3.3	-0.1
Others[2]	14.8	-4.9	-1.6	13.4	5.5	6.5	35.3
Public utilities	-5.1	-12.6	-15.0	-3.2	14.4	8.0	9.2
Science	-6.6	-25.8	-22.0	-30.7	-16.9	-9.7	-8.4
Education	-1.5	-1.7	-1.9	-0.1	-3.0	-0.6	1.1
Arts	3.6	-18.6	-10.6	20.8	-0.7	34.0	1.5
Health care	3.0	-6.3	-1.5	-1.8	-2.6	0.5	2.2
Finance	-3.7	10.0	29.8	5.1	19.4	16.0	1.1
Government	-10.1	-7.4	2.6	29.7	12.1	1.1	1.5
Others[3]	-11.0	-37.1	-31.4	7.4	-28.7	37.9	-26.0
Total	-6.1	-13.0	-8.1	-1.6	0.6	1.3	0.1
Coefficient of variation	1.1	0.8	1.5	9.0	15.9	8.2	-

Source: Bulgarian Authorities.
Notes: [1] The classification of employment by sectors was changed in 1997 to the NACE.
[2] Refers to restaurants and hotels. [3] Refers to local community services.

The distribution of real wages by sectors provides some further lights on the dynamic behaviour of the Bulgarian economy. As we already discussed in Chapter 3, Section 3.2, real wages were found to have significantly declined. The real wages in 1998 reached 40% of their level in 1989, with a cumulative fall of 60% (Garibaldi *et al*, 2001b). Changes in the real wages were mainly due to inflation shocks in particular in 1996-1997 (see Figure 9.3). After the strong erosion in 1996-1997 there was an increase in the real wages with about 52%. The positive trend was a result of changes in the wage policy from restrictive to simulating (significant increase of the minimum wage, gradual increase of salaries in the budget sphere, introduction of the Currency Board in 1997, which decreased the inflation below 10%, and the economic growth (annually 4-6%)). Whereas workers employed in the communication sector continued to enjoy the highest wages across sectors in 1999, the real wages in the service sector experienced dramatic downward shift between 1996 and 1998 (see Table 9.3).

FIGURE 9.3: Real wage dynamics in Bulgaria, 1995-2006, Index 1995=100



Source: Statistical Yearbook, NSI

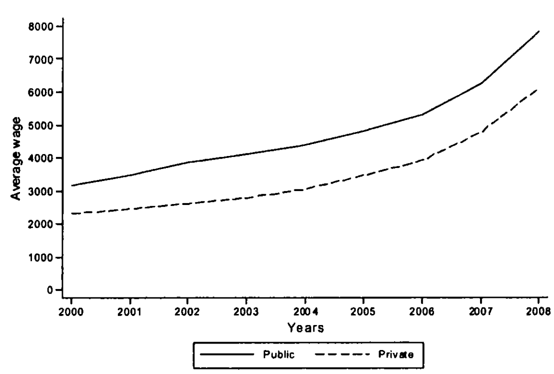
TABLE 9.3: Real wages by sector in Bulgaria, 1991-1999

In Bulgarian levs	1991	1992	1993	1994	1995	1996	1997	1998	1999
Agriculture and forestry	177	176	158	122	116	84	80	103	129
Industry	220	263	240	189	184	161	134	142	152
Construction	234	264	242	195	170	123	98	108	120
Transport	235	279	418	312	198	161	137	146	153
Communications	223	256	172	194	174	126	131	157	193
Trade	189	229	219	172	170	131	99	92	101
Other services	208	219	209	158	144	99	79	98	116
Average monthly wage[1]	212	241	237	192	165	126	108	121	138

Source: Bulgarian Authorities.
Notes: [1] Average monthly wage in constant prices. Units of measure-Bulgarian levs.

According to the National Statistics Institute (NSI) data, earnings growth in the two sectors generally has followed quite similar patterns after 2000 and the average wages in the public sector were higher than in the private sector (see Figure 9.4). That difference increased progressively and was largest in 2008, when the average wages in the public sector were 28% higher than that in the private sector.

FIGURE 9.4: Wages in the public and private sector in Bulgaria, 2000-2008



Source: Bulgaria NSI data.

Since the onset of the economic crisis in mid-2008, wages in the public sector workers have been frozen despite rather low ratios of government debt to GDP (around 15% in 2009, Eurostat 2010). Between 2009 and 2011, the Bulgarian Government demanded a 10% reduction of costs in ministries and public funded bodies, resulting in job cuts and wage freeze for employees working in these institutions (Glassner, 2010). Thus, both public and the private sector policies will need to adjust to the new environment.

9.2.2. Serbia

In 1990, a law governing transformation of social ownership was introduced in Serbia, followed by legislative measures designed to safeguard the inflow of foreign capital. By the end of the year, 23% of the 'social owned' capital entered privatization process[126] (Lokshin and Jovanović, 2003). In 1994 amendments to the 1991 privatisation law introduced the 'obligatory revaluation of privatised property' which took inflation into account. Following implementation of the 1994 legislation, only about 4% of socially-owned capital was actually privatised (Djuricin, 1997). The privatization process was slowed and in some sense it was reversed. In the first quarter of 1992, state control of capital was reintroduced. This move converted about 40% of social capital into a state-controlled capital (Lokshin and Jovanović, 2003).

In July 1997, the Serbian Government adopted the Act on Ownership Transformation, which remained in effect until early 2001 (Bayliss, 2005). Under this legislation, most firms in Serbia were privatised 'autonomously'. However, around 600 large and another 1,500 medium-sized social sector firms, which were responsible for most of the losses in the Serbian economy, were not privatised. This was largely because the privatisation legislations did set financial limits on the amount that could be transferred (Uvalic, 2001). The laws adopted by the Serbian Government in 1994 and 1997 weakened the privatization process and the 1994 law almost completely 're-socialized' privatized equity (Cerovic, 1999).

All these contradictory efforts left Serbia in the late 2000 with two main economic sectors: a dominant state and socially-owned enterprise sector and a private sector.

[126] The privatization model in Serbia was employee share-ownership that allowed firm employees to acquire the firm's shares at huge discount (30-70%) based on the firm's book value and with an 10 years payment term.

While the state sector was thoroughly accustomed to soft budget constraints, making it inefficient, the private sector (mainly composed of small and medium size enterprises) was more profitable (Hansson *et al.*, 2001). Because the private sector has been constrained by overregulation, much of its economic activity has taken place in the informal sector (Lokshin and Jovanović, 2003).

As in other transition countries, the proportion of private sector employees in Serbia increased steadily since 1995 (see Table 9.4). In 1995, 6.3% of men and 9.6% of women were working in the private sector and by 2000 the proportion reached nearly 12% for men and 17% for women (see Figure 9.5). The sector shares were almost equal in 2004-2006 (Laušev 2009).

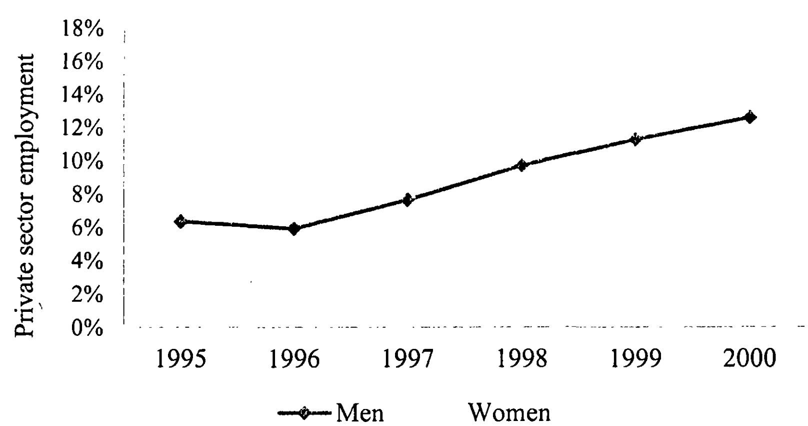
TABLE 9.4: Sectoral distribution of employment in Serbia, 1995-2000

	1995	1996	1997	1998	1999	2000
State	88.9%	88.9%	86.1%	84.1%	82.4%	80.4%
Private	7.7%	7.1%	8.9%	11.2%	12.8%	14.5%
Mixed	3.1%	3.6%	4.6%	4.3%	4.3%	4.7%
Cooperative	0.3%	0.5%	0.5%	0.4%	0.5%	0.5%

Source: Lokshin and Jovanović (2003).

FIGURE 9.5

Changes in proportion of private sector employees in Serbia by gender, 1995-2000



According to the recent official data, employment in the public sector in Serbia has had a rising trend over the second phase of the transition. In 2008 spending on wages of employees in the public sector was more than 25% of total public expenditure and more than 40% of females and 30% of males in 2007 were employed in the public sector (Ognjenović, 2011). Public and socially-owned enterprises traditionally provide stable jobs, and they exhibit very low labour turnover rates. By 2006, fewer than 5% of all workers in non-private ownership firms had been recruited in the past 12 months, while over two-third had been with the same employer for over 10 years (OECD, 2008).

Wage setting in the Serbian public sector is the result of tripartite bargaining of social partners with a dominant role of the government, with a fixed base wage and wage scales based on educational attainment, working conditions and level of responsibilities. Specifically, wages in the budget sector and public enterprises are defined by coefficient multiplies of sector specific base wages with precise coefficients defined for each job title. While industry-based collective agreements determined the base wages, the scales were negotiated at the firm level and determined by so-called collective agreements. About 86% of total wages received by state sector workers come from regular payments, and approximately 15.5% come from subsidies on transportation and meals, whereas in the private sector, only about 4% of total wages come from such subsidies, and 96% come from regular payments. Payments in kind, credits from employers and other kind of payments constitute less than 1% of total wages in Serbia (Lokshin and Jovanović, 2003).

Private sector wages in principal followed the movement of wages in the public sector, while wages at the bottom end of the earnings distribution moved upward very slowly (Ognjenović 2011). Moreover, as we mentioned above, the private sector employees were offered no benefits. In rare cases, private sector employees were given a choice between a higher wage without benefits and a lower wage with benefits (social security and health benefits, in particular).

9.2.3. Russia

A particular reason that makes Russian case interesting is the large public sector in the country that can put strong pressure on private sector wages. The number of public-owned companies in Russia is continuously growing (see Figure 9.6) and the federal government share in the public sector of the economy remains the most important in terms of its impact on the national economy. According to the official data, 14.4 million public sector employees in Russia (22% of total employment) receive their wages from the budgets of various state organisations (Gimpelson and Lukiyanova, 2009).

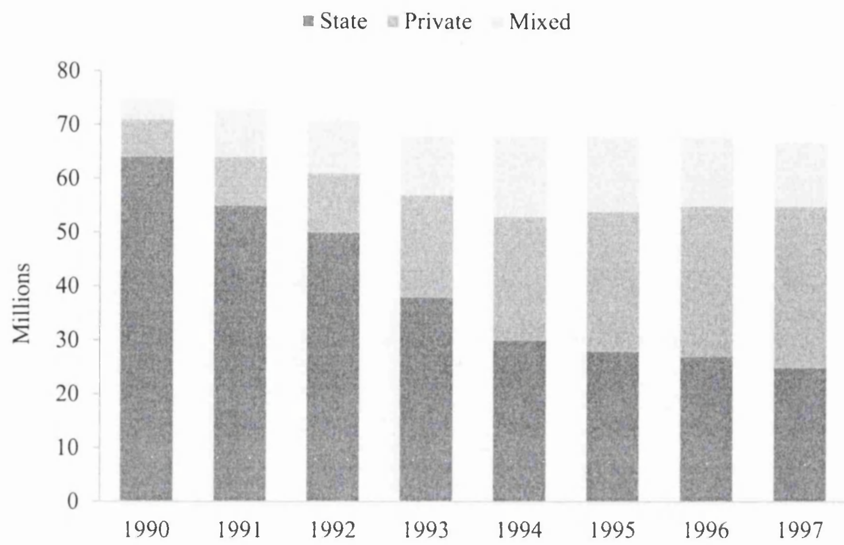
FIGURE 9.6: Total number of public-owned companies in Russia, 1998-2003

Source: Russian Statistical Yearbook, 2004
Notes: Total number in thousands.

Classification of ownership in Russia can be grouped into three major categories: state sector, which includes mainly state-owned firms, municipal firms, collective farms and other collective agricultural enterprises; private sector, which is dominated by private businesses including production co-operatives, partnerships, family-run private businesses, private farms, one-owner enterprises, employees working for private individuals and other forms of self-employment; semi-private sector, which is an intermediate group containing mostly joint-stock companies, consumer associations and other firms with mixed ownership (Gimpelson and Lippoldt, 1999).

Russia's mass privatization officially began on October 1, 1992 and its first stage concluded on July 1, 1994. The immediate goal of the program was to remove politicians' control over firms so that firm's output and management practices would respond to the preferences of consumers and shareholders. Under this program by late 1994, 68% of retail trade firms, 70% of restaurants and 78% of service establishments were privatized (Brainerd, 2002). In 1991, 84% of men and 90% of women were employed in the state sector (Clarke and Kabalina, 1999). The number of employees in the public sector fell sharply in the years that followed (see Figure 9.7). As a consequence of privatization, in 2008, 83% of firms were completely private. Those firms hired 57% of total employed workers (Góra *et al.,* 2010).

FIGURE 9.7: Employment by form of ownership in Russia, 1990-1997



Source: Goskomstat, Russian Statistical Yearbook, 1998.
Notes: The state category includes municipal enterprises; mixed includes privatized enterprises in which the state retains a stake in the charter capital and also includes social organizations and funds, joint ventures and enterprises owned by foreign individuals.

The wage-setting framework in the public sector in Russia is completely different from that in OECD countries. Wages in both sectors usually consist of two parts – guaranteed (basic) and variable component. The basic component of the budgetary sector wage is defined by the Federal authorities via politic-bureaucratic bargaining over the statutory minimum wage and Unified Tariff Scale (UTS). The minimum UTS grade is linked to the minimum wage level which is fixed uniformly by the federal legislation and does not vary across the regions. The UTS and the minimum wage upward adjustments emerge irregularly as a consequence of politico-economic bargaining and expected budget revenues. Further adjustments emerge spontaneously through the variable part as a response to market forces (Gimpelson and Lukiyanova, 2009). The variable part includes various bonuses or wage premiums which indirectly reflect the financial performance of the firm. The public sector (containing public universities and schools, hospitals, museums, research institutions, etc, which are funded from the federal or local budgets), though heavily regulated by the government, follows the same two-tier wage-setting logic as the private sector. In the public sector, the variable part is linked to revenues of regional budgets and if the latter get richer, public sector workers (funded from these budgets) enjoy higher earnings (Gimpelson and Kapeliushnikov, 2011). Various types of bonuses, allowances and benefits continue to occupy a substantial part of the overall remuneration package for most civil servants. These benefits include housing, medical services, vacations, day care, subsidized food and other goods. While as a whole there was a reduction in these benefits over the course of transition (Tratch *et at.*, 1996), their provision differs between sectors.

Private sector employers in Russia are governed by the same labour legislation that regulates the state employers. The Labour Code and Employment Law have no special exceptions for private or small businesses.

Serious financial constraints[127] throughout the 1990s forced the Russian government to cut even the most essential social programmes and to hold down the growth of salaries in the public sector. As a result remuneration of civil servants was downgraded to the level where in many cases it ceased to serve as a motivating factor. Average salary of an average Russian civil servant amounted to an equivalent of \$200 in the mid-1990s which increased to \$400 in 1998 and then declined to \$180-200 following financial crisis and devaluation of the ruble (Kotchegura 2010). However, the real wage growth in the public sector accelerated since that.

The private sector strengthened during the period and substituted for the public sector. The share of both males and females working in the private sector reached about 60% in recent years (Denisova *et al.,* 2007). The flow from unemployment to work in the private sector increased from 15% to 25% for males and from 10% to 17% for females after 2000, which indicates that more unemployed found jobs in the private sector compared to the public sector (Gimpelson and Kapeliushnikov, 2011).

For the present analysis, an issue of central importance is the wage-setting behaviour of newly privatized firms. Table 9.5 shows that the average monthly wages among workers in the state enterprises is less than the average monthly wage for workers in state joint stock companies, which is substantially lower than the average monthly wage of workers in the private sector. Over the period 1993 to 1998, workers in the state enterprises earned 78% to 83% of the wages of their counterparts in the state

[127] The main feature of the economic situation was a huge external debt (\$145 billion on average) and continuing slump in production and investment.

joint stock companies (Brainerd, 2002). There were significant difference in earnings between the state and private sector (see Table 9.5).

TABLE 9.5: Wages by type of ownership in Russia, selected years

In Russian rubles	1993	1994	1997	1998
Average monthly wages				
State enterprises	15,958	102,200	660,336	723,631
State joint stock companies	19,302	106,318	842,101	914,154
Private	33,308	159,664	1,302,977	1,240,059

Source: Survey conducted by Russian Centre for Public Opinion and Research (VTsIOM).

Brainerd (2002) indicates as a potential source of the difference in wages between sectors the differences in hours worked. According to the RLMS data, average hours in the state sector in 1994 were 158.2 per month compared to 166.5 in the private sector. Another potential source of compensation differences between sectors is in the provision of nonwage benefits to employees by firms. As we have discussed, Russian enterprises had traditionally provided a wide range of services and benefits to their workers, either free of charge or highly subsidized prices.

Finally, it worth mentioning that Russia entered the 1990s with a huge old-age pension problem. The state was the sole source of pensions, paid on a pay-as-you-go (PAYG) basis. The pension provision established by 1956 and 1965 laws[128], covered only people employed in the public enterprises or collective farms. There were no additional private or occupational pensions. The system was characterised by low cost for employers (4-12% of payroll depending on the sector), low retirement age and special preferences for certain group of people. Retirement age were low by developed country standards at age 55 for women and age 60 for men. Many small firms used a so called "simplified" taxation regime and did not pay a unified social tax. As a result people hired by small firms were limited in their future pension rights because of the lower contributions paid by the employer. People working at different state firms tended to receive unequal social packages and even at the same organization, those with greater seniority had access to better quality services (Góra *et al.*, 2010).

[128] 1965 Law on state pensions for wage and salary earners and 1965 Law on state pensions for collective farmers.

9.2.4. Tajikistan

The public sector of the Central Asian states still exhibits many of the characteristics of the former Soviet Union. Although the size of the public sector relative to the GDP has been reduced in all countries, governments continue to exert a strong influence on most aspects of economic activity through budgetary operations, quasi-fiscal operations performed by the state-owned financial and non-financial enterprises and extensive regulation on several aspects of economic and social activity (Gürgen 1999).

Tajikistan is a highly agrarian country, with agriculture accounting for 60% of the total employment and around 30% of the GDP (Lerman and Sedik, 2008). The development of Tajikistan's private sector effectively began in 1985, when the Soviet Union allowed the creation of corporatives. The country has enacted laws aimed at developing the market economy. Several laws dealt with privatization and private sector development – Registration of Enterprises (1991), Privatization of State-Owned Enterprises (1992), Joint Stock Companies (1992), Foreign Investment (1992), Entrepreneurship (1992), Bankruptcy (1992) and Enterprises (1993). The Property Law, enacted in 1990 divided properties into three categories: state property; collective property, such as labour collectives, cooperatives and joint stock companies; and 'property of the citizen' – private property, such as individual economic activity and small enterprises. Privatisation in Tajikistan has mainly taken the form of ownership transfers to labour collectives or leasing arrangements. Despite Government intention to promote private sector development, there were several constraints that inhibited its development, such as entry and exit rules, labour regulations, limited access to business information, lack of competition and access to credits. The share of private sector in the GDP was estimated at about 20-30% in 1998 and 40% in 2001 (Ghasimi, 1994).

Similarly to the other transition countries, there was clear evidence of an emerging private sector activity in Tajikistan. In 1985, 16% of workers were employed in the private sector, while in 1998 the private sector employment increased to 33% (see Table 9.6). According to the 2004 Labour Survey data, 27.5% of all employed were occupied in the state enterprises, whereas 60.8% were in the private sector, including family business and personal plots of land. Employment in the private sector was higher in rural areas and marked 63.1%, in comparison with 51.6% employed in urban areas[129].

TABLE 9.6: Sectoral distribution of employment in Tajikistan, 1985-1998

	1985	1990	1991	1992	1993	1994	1995	1996	1997	1998
Government	69.8%	64.7%	54.5%	52.2%	52.5%	50.1%	47.2%	45.3%	42.0%	42.7%
Collective farms	14.1%	13.6%	14.9%	14.9%	16.1%	18.1%	15.7%	15.4%	21.0%	21.0%
Cooperatives	-	2.7%	2.6%	2.6%	2.4%	2.4%	1.3%	0.8%	0.9%	0.7%
Private farming	16.1%	18.9%	18.9%	20.9%	22.6%	22.6%	29.7%	32.9%	33.3%	33.1%

Source: State Statistical Committee.
Notes: Collective farms and cooperatives referred to the state sector.

Public sector wages in Tajikistan were low and not competitive, both in comparison with other countries and with the private sector wages. While there were increases in the government wage levels since the late 1990s, they have not kept pace with the increase in the nominal GDP and total government expenditures and were 2.7% of the GDP in 2003-2004 (IMF, 2001)[130]. The average government wage (including other employment related supplements) in 2004 was about 74 Somoni ($25) per month, just above the poverty line. Average wage for teachers and health care sectors was even lower. The highest average monthly wages belonged to people working in the finance and banking sector (see Table 9.7). Government workers however, had nonmonetary fringe benefits. These benefits include generous travel allowance, cars and mobile phones. Other benefits are off-budgets such as land plots for rural teachers. A key benefit for most government workers was discounted gas and electricity, but this benefit was removed in 2003 following significant energy price increase (IMF, 2001).

129 See Analytical Report on survey of labour force in the Republic of Tajikistan (2005), Dushanbe, State Statistics Committee of the Republic of Tajikistan.

130 The figure indicates implausible low wage bill.

TABLE 9.7: Wages by sector in Tajikistan, 2000-2004

	2000	2001	2002	2003	2004
	In Tajik somoni (SM)				
Total economy	15.6	23.5	32.6	44.6	60.8
Of which:					
Agriculture	7.8	13.7	18.9	27.0	35.0
Industry	47.1	71.2	92.5	114.2	144.1
Transport	31.0	48.9	69.6	101.1	147.6
Construction	38.9	55.4	74.8	100.0	150.8
Finance and banking	76.7	90.5	118.5	174.9	230.9
Private enterprises	-	35.0	62.3	88.6	136.4
General government	23.9	29.9	49.2	63.3	74.2
Health care	6.7	8.9	12.7	17.1	22.8
Education	11.6	17.3	25.6	34.4	43.4
	In percentage				
Total economy	34.1	50.9	38.5	37.1	36.3
Of which:					
Agriculture	44.6	76.0	38.3	42.6	29.5
Industry	31.5	51.2	29.8	23.5	26.2
Transport	33.3	57.8	42.5	45.2	46.0
Construction	5.0	42.4	35.0	33.8	50.8
Finance and banking	56.8	18.0	30.9	47.6	32.0
Private enterprises	-	-	77.7	42.3	54.0
General government	4.7	25.5	64.2	28.7	17.3
Health care	50.5	31.5	43.4	34.8	33.2
Education	51.7	49.2	48.4	34.5	26.1

Source: State Statistical Committee.
Notes: In 2004, SM3 exchanged for US1.

According to the Law on the Civil Service, the monthly salary of a state worker consist of a basic rate and a qualification increase based on a class grade, working years, honours, awards, and academic achievements (generally 30% of the monthly salary is paid for academic achievements). The qualification increase is distributed as follows: between 1 and 3 years in state service would result in 5% salary increase; more than 3 years in service – 5-10% increase; more than 5 years in service – 10-20% increase; more than 10 years in service – 20-30% increase and more than 20 years in service relates to 45% increase.

The short overview of the changing employment and sector wage formation brings us to formulate a few hypotheses related to the public-private wage gap in all these countries. We may draw a conclusion that transition process may have caused differences in the growth rates of public and private sector wages, provided that the two sectors have responded to the economic changes differently in terms how wages were set to attract skilled workers. In addition, the non-wage advantages, such as pension plans and other benefits may have compensated for lower wages in the public sector.

9.3. Descriptive analysis

We now look at the public and private sector employment conditions and wages in more details. Our analysis draws on data from the 1986 and 2003 Bulgarian household surveys, 2003 Serbian LSMS, 1994 Russian RLMS and 2003 Russian NOBUS Survey, and 2003 Tajikistan LSMS. The data are restricted to the sample of employees and omits all observations for which there was any missing information. We exclude from the analysis own account workers, and never worked categories. Since construction of the data and definitions of the variables for the four transition countries, have already been explained in Section 3.3.1, such discussion is not repeated here.

As in previous chapters, the analysis in the current Chapter focuses on the natural logarithm of hourly wage rate. We distinguish between two main sectors: public (state) and private. We define public sector workers in Bulgaria as equal to 1 if the individual is employed in a government public sector or army and 0 if employed in a private company. The type of ownership of the company in which individuals perform their main job is used to define public sector variable in the case of Russia and Serbia. The binary variable takes value 1 if respondents are employed in the state owned enterprises or public associations and 0 if they are employed in the private and foreign ownerships firms. The public sector variable in Tajikistan takes value 1 if individuals are paid public workers or employed in the state administration and state-owned enterprises and 0 if individuals are employed in private firms including agriculture, collective farms, joint ventures and foreign firms. In order to identify the various factors involved in the wage determination in the public and private sectors, a range of other variables were also included in the analysis. The vector of regressors in Bulgarian, Serbia, Russian and Tajikistan earning equations includes experience and its square, job tenure with the current employer, education, marital status, managerial responsibility and region of residence. In contrast with some previous studies of wage determination in transition countries and following Falaris (2004), our main results do not include controls for industry or occupation in the wage equations. Especially in the private sector and to a lesser extent in the public sector, workers have some choice over the type of heir work. When moving from the public sector to the private sector, workers often change their occupations. Therefore, these variables are likely to be endogenous and testing and correcting for simultaneous equation bias arising from the endogeity of occupations is not feasible because we do not have a good exogenous instrument, consistent for all four countries. Only in the case of Bulgaria (1986), a sensitivity analysis is undertaken by including information on occupation in the decomposition estimates.

We should mention some other factors that may affect sector wage differences and which have not been incorporated into our analysis. In particular, wages are incomplete measure of the total return to labour. Workers may be willing to accept lower earnings in exchange for some other nonwage benefits. Superannuation and paid maternity leave entitlement may be particularly important considerations. Other sector differences that have not been accounted may include job security and flexibility, pensions, fringe benefits, working conditions and other non-wage forms of pecuniary compensations. Despite the availability of detailed information in the data files, we can never fully capture all workers specific differences. One potentially important variable not included in our analysis is union status. The variable is not presented in three of the samples. As in Venti (1987) we consider this exclusion on the ground that it is preferable to let union effects implicitly to enter the model in reduced form rather than deal directly with the endogeneity of union status. In addition, we consider the fact that unions' negotiation power is relatively low in most of the transition countries (Cazes, 2002)[131].

The descriptive statistics for the variables used are listed in Appendix Tables A9.1 to A9.4. Most important for the purposes of our analysis is a comparison of wages across the two sectors. While males in Serbia earn higher wages in the private sector, the public wage premium is higher for females, who earn on average 7% more in the public sector than their private sector counterparts (see Table A9.1). In Bulgaria, both men and women employed in the public sector earn more than the private sector

[131] According to available data, the percentage of trade union membership ranges from about 34% in Poland to 74% in Russian Federation. The level of coordination was high until the mid-1990s due to tax based income policy imposed by the government. When the tax-based income policy was abolished in 1995, the coordination ability of the trade unions has decreased (Cazes, 2002).

workers. The average difference is particularly larger in 2003, when the public sector workers earn about 34% higher wages than workers in the private sector (see Table A9.2). In contrast, the average hourly earnings are higher in the private sector than in the public sector for both males and females in the 1994 Russian and 2003 Tajik samples.

Further, the 2003 raw wage distributions in the public and private sectors by countries are described in Table 9.8. In Bulgaria, for both men and women, all major percentiles record higher wages in the public sector. By contrast, for men in Serbia, apart from the 10^{th} and 25^{th} percentiles, wages are higher in the private sector. Wages are higher in the private sector at the top percentile levels and for both males and females in Tajikistan. Based on the ratio of the 90^{th} to the 10^{th} deciles, we find that apart from males in Bulgaria, the private sector wages for both genders in Serbia, Russia and Tajikistan seem to be more dispersed (see Table 9.8).

TABLE 9.8: Summary measures of log hourly wage distributions, 2003

	Bulgaria				Serbia				Russia				Tajikistan			
	Males		Females		Males		Females		Males		Females		Males		Females	
	Public	*Private*	*Public*	*Private*	*Public*	*Private*	*Public*	*Private*	*Public*	*Private*	*Public*	*Private*	*Public*	*Private*	*Public*	*Private*
Major percentiles																
10^{th}	1.179	0.916	1.099	0.829	3.342	3.324	3.397	3.031	1.932	1.701	1.750	1.568	2.485	2.079	1.946	1.609
25^{th}	1.504	1.099	1.322	1.012	3.770	3.704	3.767	3.401	2.443	2.280	2.155	2.086	2.996	2.708	2.303	2.197
50^{th}	1.910	1.447	1.609	1.322	4.107	4.135	4.085	3.816	2.954	2.848	2.666	2.548	3.401	3.555	2.890	2.708
75^{th}	2.277	1.833	1.946	1.650	4.463	4.541	4.415	4.271	3.359	3.253	3.136	3.003	3.912	4.500	3.401	3.689
90^{th}	2.526	2.169	2.247	2.015	4.802	5.063	4.789	4.674	3.829	3.696	3.513	3.424	4.605	5.011	3.912	4.488
Decile ratios																
90/10	1.347	1.253	1.148	1.186	1.460	1.739	1.392	1.643	1.897	1.995	1.763	1.856	2.120	2.932	1.966	2.879
50/10	0.731	0.531	0.510	0.493	0.765	0.811	0.688	0.785	1.022	1.147	0.916	0.980	0.916	1.476	0.944	1.099
75/50	0.367	0.386	0.337	0.328	0.356	0.406	0.330	0.455	0.405	0.405	0.470	0.455	0.511	0.945	0.511	0.981
90/50	0.616	0.722	0.638	0.693	0.695	0.928	0.704	0.858	0.875	0.848	0.847	0.876	1.204	1.456	1.022	1.780

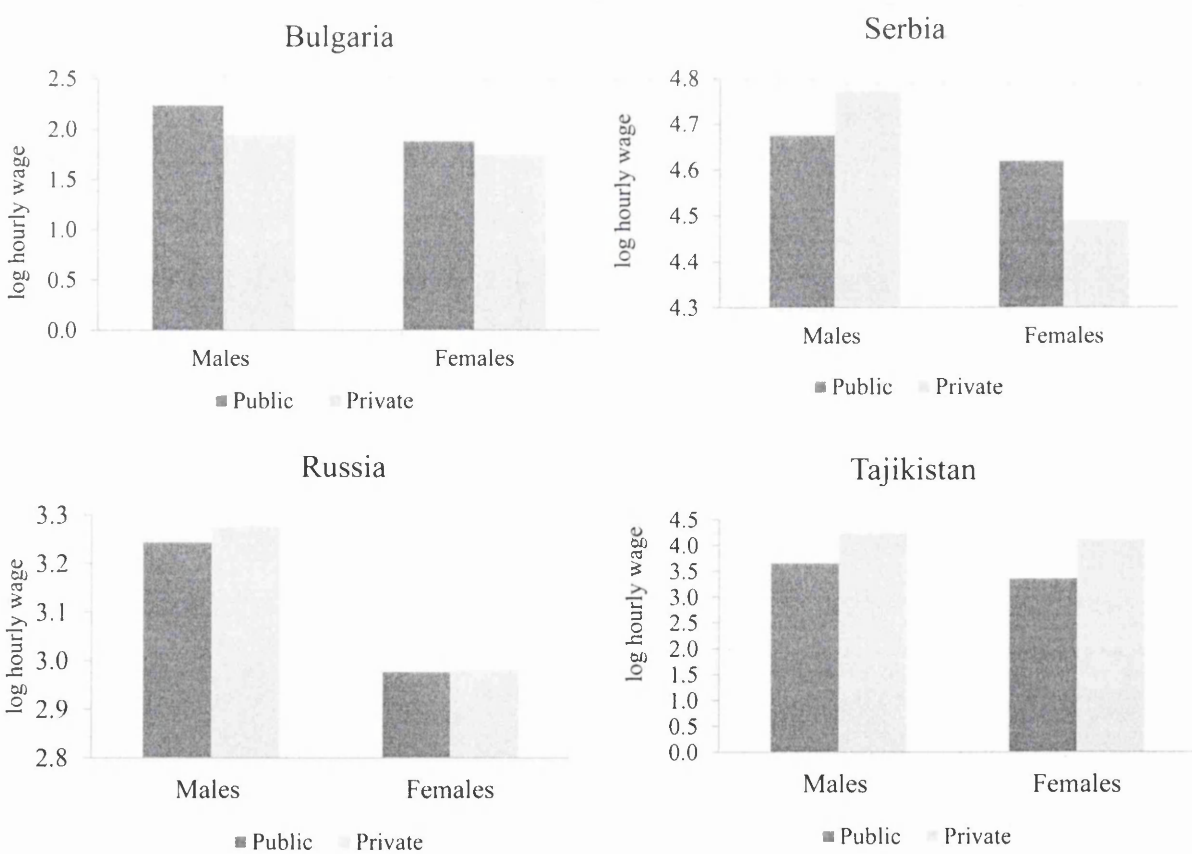
Source: Authors' calculations.

Overall, in all four countries, public sector employees are on average better educated and have more labour market experience than private sector employees. These differences in education and work experience might be expected to result in higher average wages for the public sector employees. For instance, 28% of public sector male employees in Bulgaria in 2003 have completed university degree, compared to 11% for private sector workers (see Table A9.2). In both sectors, male employees in Tajikistan achieve a higher university level than female employees.

In Figure 9.8, we show the distribution of hourly earnings for university graduates males and females by countries. Public sector pay is higher for university male and female graduates in Bulgaria and for females in Serbia. However, men with university degree in Russia and both men and women inTajikistan earn more in the private sector.

FIGURE 9.8

Log hourly earnings for university graduates by sector and by country in 2003



Source: Bulgarian Household survey 2003, Serbian SLMS 2003, Russian NOBUS 2003, Tajikistan LSMS 2003.

In all four countries, we have documented existence of differences in the public and private sector wages, but the interesting question is whether it persists even after controlling for relevant characteristics determining the wage process. We do this in the rest of the present Chapter.

9.4. Empirical methodology

9.4.1. Dummy variable method

In order to estimate the public-private sector wage differential, we utilise the QR approach. We start with the following quantile regression equation:

$$Q_{w_i}(\tau \mid x) = x_i\beta_\tau + Pub_i\chi_\tau \qquad (9.1)$$

where w_i is the log-hourly wage for worker *i*, Pub_i is a dummy variable, which takes value 1 if the individual works in the public sector and 0 if employed in the private sector, x_i is a vector of individual characteristics for worker *i*, β_τ and χ_τ are the parameters of the model for estimation of the τ^{th} quantile. We focus on the coefficient attached to Pub_i, which is the return on public sector employment. Earnings functions are estimated separately for men and women. All other things equal, if the public sector coefficient is positive there is a wage premium for working in the sector. It has to be kept in mind that using this approach imposes a restriction that the returns to observed characteristics are the same for the two sectors and that public-private differences depend only on a shift factor. Therefore, a further decomposition analysis will follow which decompose the sector wages into a component due to differences in labour market characteristics between the sectors and a component due to differences in the rewards public and private sector workers receive for those labour market characteristics.

9.4.2. Machado and Mata decomposition based on Oaxaca-Ransom (1994) approach

As in the previous chapter on gender wage differential, we are able to compare the sector wage differential at various points in the distribution. This comparison provides a much richer description of the wage differential and allows us to see whether there are differences in the differential at different points of earnings distribution. We also investigate how the returns to various characteristics have changed over time at various points in the conditional earnings distribution. Specifically, we examine development of the public-private wage gap in selected countries at conditional quantiles of the wage distribution based on the extended version of Machado-Mata procedure (2005), described in Section 7.3, where for the non-discriminatory wage structure we have followed the Neumark (1988) and Oaxaca and Ransom (1994) procedure. This technique is a natural choice since it is well suited to capture heterogeneity in both the coefficients and covariates across the entire wage distribution. Moreover, because there is a substantial wage differential between men and women, in both the public and private sectors, we decompose the sector pay differential separately for each gender.

The marginal earnings distributions produced by the MM method and how these are used in the Oaxaca-Ransom decomposition can be briefly described as follows:

Step 1: Generate a random drawn sample of size m from a uniform distribution $U\,[0,1]\, u_1 \dots, u_m$.

Step 2: Estimate for the public and private sector separately and for all workers combined the quantile regression coefficients: $\hat{\beta}^{pub}_{u_i}(\theta), \hat{\beta}^{priv}_{u_i}(\theta), \hat{\beta}^{*}_{u_i}(\theta), i = 1 \dots. m$, where $\hat{\beta}^{pub}_{u_i}(\theta)$ are u_i^{th} quantile regression estimates taken from a (log) hourly earnings equation for public sector workers; $\hat{\beta}^{priv}_{u_i}(\theta)$ are u_i^{th} quantile regression estimates taken from the hourly earnings equation for private sector workers, and $\hat{\beta}^{*}_{u_i}(\theta)$ are u_i^{th} quantile regression estimates taken from the hourly earnings equation for all workers. The coefficients in the vector $\hat{\beta}^{*}_{u_i}$ are taken as a baseline wage structure in which characteristics in the public and private sector are rewarded equally.

Step 3: Sampling with replacement, a random sample of size m is taken of public and private sector workers characteristics that were used to estimate quantile regression coefficients. The vectors of characteristics for public $\{\widetilde{X}_i^{pub}\}_{i=1}^{m}$ and private $\{\widetilde{X}_i^{priv}\}_{i=1}^{m}$

sector workers are then used to predict (log) hourly earnings in the private $\{\widetilde{w}_{u_i}^{priv} = \widetilde{X}_i^{priv}\hat{\beta}_{u_i}^{priv}\}_{i=1}^{m}$ and public $\{\widetilde{w}_{u_i}^{pub} = \widetilde{X}_i^{pub}\hat{\beta}_{u_i}^{pub}\}_{i=1}^{m}$ sectors. These predicted wages are equivalent to a random sample of size m drawn from the marginal wage distributions of private (w^{priv}) and public (w^{pub}) sector workers. Counterfactual densities for public and private sector workers being rewarded equally in both sectors are found as $\{\widetilde{w}_i^{cfpub} = \widetilde{X}_i^{pub}\hat{\beta}_{u_i}^{*}\}_{i=1}^{m}$ and $\{\widetilde{w}_i^{cfpriv} = \widetilde{X}_i^{priv}\hat{\beta}_{u_i}^{*}\}_{i=1}^{m}$.

Step 4: Differences in θ^{th} percentiles of the estimated marginal wage distribution are then used to decompose the public-private sector wage gap into an effect due to characteristics in the public and private sector being rewarded differently (coefficient effect) and an effect due to differences in the distribution of worker characteristics in the two sectors (characteristic effect). The difference in the logarithm of hourly earnings between public and private sector workers at the θ^{th} percentile is given by:

$$Q_\theta(\widetilde{w}^{pub} \mid \hat{\beta}^{pub} - \widetilde{w}^{priv} \mid \hat{\beta}^{priv}) =$$
$$= \underbrace{Q_\theta[\underbrace{(\widetilde{w}^{pub} \mid \hat{\beta}^{pub} - \widetilde{w}^{cfpub} \mid \hat{\beta}^{*})}_{public\ advantage} + [\underbrace{(\widetilde{w}^{cfpriv} \mid \hat{\beta}^{*} - \widetilde{w}^{priv} \mid \hat{\beta}^{priv})}_{private\ disadvantge}]}_{coefficient} + Q_\theta\underbrace{[(\widetilde{w}^{cfpub} \mid \hat{\beta}^{*} - \widetilde{w}^{cfpriv} \mid \hat{\beta}^{*}]}_{characteristics} + resid \quad (9.1)$$

where Q_θ is the θ^{th} percentile of the earnings distribution and $(\widetilde{w}^{i} \mid \hat{\beta}^{j})$ is the estimated marginal earnings distribution for the i^{th} sector based on the j^{th} quantile regression estimates.

The first term of the right hand side of expression (9.1), the coefficient component, is a measure of the wage advantage enjoyed by public sector workers and is comprised of two components that indicates the degree of which characteristics of employees in the public and private sectors are rewarded differently relative to the baseline wage structure given by the coefficients contained in the vector $\hat{\beta}^{*}$ for each qantile. The second term of the right hand side of equation (9.1), the characteristic component, is a measure of the difference in logarithm of earnings in the public and private sector attributable to differences in the characteristics or composition of employees in the public and private sectors as reflected by being rewarded according to the $\hat{\beta}^{*}$

baseline wage structure. Standard errors for the reported components of the decomposition were obtained using a standard bootstrapping method[132].

9.4.3. *Correcting for endogenous sector choice*

Given that there is a choice being made by workers whether to work in the public or private sector, there is a potential for the sample selection bias. This bias is due to the possibility that sorting of employees between sectors is not random and occurs on the basis of unobserved productivity-related characteristics. Either through the process of self-selection by individuals or by employers, the location of individuals in either sector may not be interpretable as a random process (Hyder and Reilly, 2005). The probability of being selected within a particular sector depends on the individual's characteristics (observed and unobserved) as well as on characteristics of the employer.

The problem has been typically addressed by jointly estimating equations for a worker's sector of employment and earnings, relying a set of exclusion restrictions (Depalo and Giordano, 2010). For instance, the issue of sample selection bias for sector choice has been addressed by Hyder and Reilly (2005), where the sample selection correction term for sector choice is included in the wage equations estimated for the separate samples of public and private sector workers. The authors however, conclude that the differentials based on correcting for selection bias provide few new insights on the magnitude of the public sector premium across the conditional wage distribution.

The present analysis formulates and estimates a model of sectoral attachment of individuals (the binary choice model between the public *(D=1)* and private sector work *(D=0)* is considered). A worker will be employed in a particular sector, if the offered wage exceeds the reservation wage. We follow Buchinsky's (1998) approach for the selection correction in a quantile regression model by using the power series expansion of the inverse Mill's ratio of the normalized estimated index. Since the probit depends heavily on the distributional assumption, we also estimate the sector choice equation by single index model (Ichimura, 1993). Probit should be used in the

[132] Bootstrap estimates are based on 800 replications.

first step on the selection when errors are normally distributed; the single index estimator should be used otherwise (Badel and Pena, 2010).

There are two issues that need to be addressed. First, there is the identification of the selection effect and the crucial assumption is the presence of instrument and the exclusion of this instrument from the outcome equation (Melly, 2006). Various family structure and social background variables have been used and their use for identification purposes is criticised as being arbitrary. We construct our identifying variable as a share of public sector workers within a certain industry, calculated by different regions[133]. The resulting variable reflects the share of individuals who answered public/state enterprise in a specific industry sector and in a specific macro-region (for example, we generate 133 industry sector-macro region cells in Russia – 19 industries by 7 regions). This allows for a possible regional variation in the public sector employment. Importantly the constructed instrument must be correlated with choice of sector, but should not affect the worker's wage. Second, we have discussed in Section 4.2.4 that Buchinsky (1998) correction of sample selection in the context of quantile regression requires higher order selection terms be included in the model. There is however, the added complication that the constant term in the quantile wage regression is not identified in this instance, as it is conflated with the constant term of the higher order power series used to control for sample selection. The problem is one of distinguishing between the intercept and the first term in the power series approximation to the selection correction term. Buchinsky (2001) provides a detailed discussion of ways to identify the constant along the lines suggested by Andrews and Schafgans (1998). In our study an assessment of the importance of the selectivity of the sector choice is made following Buchinsky (1998) based on Andrews and Schafgans (1998) idea of 'identification at infinity'. The intuition is as follows: if we choose a subsample of workers with labour market characteristics such that the probability of being in a given sector is arbitrary close to 1, we can use this subsample to estimate the intercept in the earnings equation, without adjusting for selection.

[133] We also tried to use as identifying variable a number of jobholders and the number of public sector employees in the household, following the assumption that personal example from close family member may convince an employee to choose a career in the public sector. However, the variables were insignificantly different from zero in the selection equation for all countries.

Specifically, Andrews and Schafgans (1998) estimator replaces Heckman's (1990) indicator function by a smooth function. The authors introduce a weighting scheme for those observations, where observations exceeding this threshold are weighted by a smooth monotone [0,1] valued function $s(\cdot)$, such as distribution function:

$$\hat{\mu}_1 = \frac{\sum_{i=1}^{n}(Y_i - Z_i'\hat{\theta})D_i s(X_i'\hat{\beta} > \gamma_n)}{\sum_{i=1}^{n} D_i s(X_i'\hat{\beta} > \gamma_n)} \tag{9.2}$$

where $s(\cdot)$ is a non-decreasing [0,1]-valued function that has three derivatives bounded over R and for which $s(x) = 0$ for $x \leq 0$ and $s(x) = 1$ for $x \geq b$ for some $0 < b < \infty$; $(\hat{\theta}, \hat{\beta})$ are root-n consistent preliminary estimators of (θ_0, β_0) and the parameter γ_n is the bandwidth or smoothing parameter, $\gamma_n \to \infty$ as $n \to \infty$ (Schafgans and Zinde-Walsh, 2002).

As suggested by Andrewes and Schafgans (1998), we adopt the following smoothing function $s(\cdot)$:

$$s(x) = \begin{cases} 1 - \exp(-\frac{x}{b - x}) & for\ x \in (0, b) \\ 0 & for\ x \leq 0 \\ 1 & for\ x \geq b \end{cases} \tag{9.3}$$

where b is set equal to 1. The bandwidth or smoothing parameter γ_n is chosen in such a way that $\gamma_n \to \infty$ as $n \to \infty$, which should guarantee that only observations for which the probability of selection is close to one are included (Schafgans 2000). The obvious disadvantage of the intercept estimator is that only a few observations are used for identification. A further handicap is that there exists no formal rule to determine the threshold values for b.

9.4.4. Decomposing public sector wage gap accounting for endogenous sector choice

In order to analyse the public-sector wage differential accounting for endogenous sector choice, the MM-OR decomposition algorithm is modified. Following Badel and Pena (2010), we calculate the selection effect as the difference between the distributions of observed and the potential wages of public and private sector workers. We generate random samples of public and private sector workers using the public and private sample selection adjusted coefficients combined with the labour market characteristics of all workers – not just those who work in the public sector. The two selection effects are given as:

$Q_\theta(\widetilde{w}^{pub}) - Q_\theta(\widetilde{w}^{ctfpub})$ – selection effect for public sector workers and

$Q_\theta(\widetilde{w}^{priv}) - Q_\theta(\widetilde{w}^{ctfpriv})$ – selection effect for private sector workers;

$\widetilde{w}_i^{ctfpub} = \hat{\beta}_{u_i}^{pub} \widetilde{X}_i^* + e^{ctf}$ and $\widetilde{w}_i^{ctfpriv} = \hat{\beta}_{u_i}^{priv} \widetilde{X}_i^* + e^{ctf}$ with $Q_\theta(e^{ctf}) = 0$

where $\widetilde{w}^{pub}$ and $\widetilde{w}^{priv}$ are the observed distributions of public and private sector workers; $\widetilde{w}_i^{ctfpub}$ and $\widetilde{w}_i^{ctfpriv}$ are simulated counterfactual distributions for the public and private sector workers using public and private sector selection adjusted coefficients and being sampled from the whole population of workers.

The endogeneity corrected differential or the differential that we would observe if the employees sorted randomly between sectors conditional on their characteristics is decomposed as follows:

$$[Q_\theta(\widetilde{w}^{pub}) - Q_\theta(\widetilde{w}^{ctfpub})] - [Q_\theta(\widetilde{w}^{priv}) - Q_\theta(\widetilde{w}^{ctfpriv})] \qquad (9.4)$$

Further, the selection effect can be decomposed into a portion due to observable labour market characteristics and a portion due to unobservable characteristics by modifying the MM-OR algorithm. This produces another two distributions: the distribution of public sector's wages (respectively the distribution of private sector's wages) that would have prevailed if prices accounted for selection, and workers had the distribution of labour market characteristics of public (private) sector workers – not of all workers.

$$Q_\theta(\widetilde{w}^{pub}) - Q_\theta(\widetilde{w}^{ctfpub}) = [Q_\theta(\widetilde{w}^{pub}) - Q_\theta(\widetilde{w}^{s-pub})] - [Q_\theta(\widetilde{w}^{ctfpub}) - Q_\theta(\widetilde{w}^{s-pub})] \qquad (9.5)$$

where $\widetilde{w}_i^{s-pub} = \hat{\beta}_{u_i}^{pub} \widetilde{X}_i^{pub} + e^{s-pub}$ with $Q_\theta(e^{s-pub}) = 0$

Note that in this exercise $\hat{\beta}_{u_i}^{pub}$ needs to be estimated controlling for sample selection bias. The first term in the brackets is interpreted as the effect of unobservable, while the second term is interpreted as the effect of observables. Specifically, the difference between the public sector workers adjusting for selection counterfactual and the potential distribution tells us how much of the selection effect can be explained by differences in the distribution of characteristics between those who are in the public sector and those who are not. The selection effect for the private sector is given in a similar way:

$$Q_\theta(\widetilde{w}^{priv}) - Q_\theta(\widetilde{w}^{ctfpriv}) = [Q_\theta(\widetilde{w}^{priv}) - Q_\theta(\widetilde{w}^{s-priv})] - [Q_\theta(\widetilde{w}^{ctfpriv}) - Q_\theta(\widetilde{w}^{s-priv})] \quad (9.6)$$

where $\widetilde{w}_i^{s-priv} = \hat{\beta}_{u_i}^{priv} \widetilde{X}_i^{priv} + e^{s-priv}$ with $Q_\theta(e^{s-priv}) = 0$

The endogeneity corrected differential due to the unobservable characteristics is given as the difference between the effect of unobservables for public and private sector workers:

$$[Q_\theta(\widetilde{w}^{pub}) - Q_\theta(\widetilde{w}^{s-pub})] - [Q_\theta(\widetilde{w}^{priv}) - Q_\theta(\widetilde{w}^{s-priv})] \quad (9.7)$$

Finally, the endogeneity corrected differential due to the observable characteristics is given as the difference between the observable components for public and private sector workers:

$$[Q_\theta(\widetilde{w}^{ctfpub}) - Q_\theta(\widetilde{w}^{s-pub})] - [Q_\theta(\widetilde{w}^{ctfpriv}) - Q_\theta(\widetilde{w}^{s-priv})] \quad (9.8)$$

9.4.5. Accounting for the evolution over time

Finally, to analyse what had happened to the sector earnings differential at different points in the conditional earnings distribution over time, we modify the decomposition procedure following the framework outlined in Donohue and Heckman (1991) and first suggested by Smith and Welch (1986). Specifically, the proposed framework identifies the main factors contributing to the changes in the ratio of the public to private sector earnings in Bulgaria and Russia for two periods.

The changes in the conditional earnings distribution from time period *t* to time period *t+1* in the wage gap between public and private sector can be examined using a decomposition framework outlined in Donohue and Heckman (1991):

$$\begin{aligned}
\Delta Q_\theta(\widetilde{w}^{pub} - \widetilde{w}^{priv}) = \\
& Q_\theta[(\widetilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_t^* - \widetilde{w}_{t+1}^{cfprv} \mid \hat{\beta}_t^*) - (\widetilde{w}_t^{cfpub} \mid \hat{\beta}_t^* - \widetilde{w}_t^{cfprv} \mid \hat{\beta}_t^*)] + \\
& Q_\theta[(\widetilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_{t+1}^* - \widetilde{w}_{t+1}^{cfprv} \mid \hat{\beta}_{t+1}^*) - (\widetilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_t^* - \widetilde{w}_{t+1}^{cfprv} \mid \hat{\beta}_t^*)] + \\
& Q_\theta[\{(\widetilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_t^{pub} - \widetilde{w}_t^{pub} \mid \hat{\beta}_t^{pub}) - (\widetilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_t^* - \widetilde{w}_t^{cfpub} \mid \hat{\beta}_t^*)\} + \\
& \{(\widetilde{w}_{t+1}^{cfpriv} \mid \hat{\beta}_t^* - \widetilde{w}_t^{cfprv} \mid \hat{\beta}_t^*) - (\widetilde{w}_{t+1}^{cfpriv} \mid \hat{\beta}_t^{priv} - \widetilde{w}_t^{prv} \mid \hat{\beta}_t^{priv})\}] + \qquad (9.9) \\
& Q_\theta[\{(\widetilde{w}_{t+1}^{cfpriv} \mid \hat{\beta}_{t+1}^* - \widetilde{w}_{t+1}^{priv} \mid \hat{\beta}_{t+1}^{priv}) - (\widetilde{w}_{t+1}^{cfpriv} \mid \hat{\beta}_t^* - \widetilde{w}_{t+1}^{cfpriv} \mid \hat{\beta}_t^{priv})\} + \\
& \{(\widetilde{w}_{t+1}^{pub} \mid \hat{\beta}_{t+1}^{pub} - \widetilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_{t+1}^*) - (\widetilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_t^{pub} - \widetilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_t^*)\}]
\end{aligned}$$

where $(\widetilde{w}_t^i \mid \hat{\beta}_s^j)$ is the estimated marginal earnings distribution for the i^{th} sector based on the distribution of characteristics in that sector at time t and the j^{th} quantile regression estimates at time s.

Following Donohue and Heckam (1991) each component of the right hand side of (9.9) can be given the following interpretation:

1) $[(\widetilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_t^* - \widetilde{w}_{t+1}^{cfpriv} \mid \hat{\beta}_t^*) - (\widetilde{w}_t^{cfpub} \mid \hat{\beta}_t^* - \widetilde{w}_t^{cfpriv} \mid \hat{\beta}_t^*)]$ shows the change in relative hourly earnings attributable to between sector changes in the characteristics of public and private sector workers, evaluated at the baseline wage structure;

2) $[(\widetilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_{t+1}^* - \widetilde{w}_{t+1}^{cfprv} \mid \hat{\beta}_{t+1}^*) - (\widetilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_t^* - \widetilde{w}_{t+1}^{cfprv} \mid \hat{\beta}_t^*)]$ shows the change in relative hourly earnings explained by changes in the baseline wage structure, holding characteristic differences constant at $t+1$ levels;

3) $\{(\widetilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_t^{pub} - \widetilde{w}_t^{pub} \mid \hat{\beta}_t^{pub}) - (\widetilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_t^* - \widetilde{w}_t^{cfpub} \mid \hat{\beta}_t^*)\} +$
$[(\widetilde{w}_{t+1}^{cfpriv} \mid \hat{\beta}_t^* - \widetilde{w}_t^{cfprv} \mid \hat{\beta}_t^*) - (\widetilde{w}_{t+1}^{cfpriv} \mid \hat{\beta}_t^{priv} - \widetilde{w}_t^{prv} \mid \hat{\beta}_t^{priv})]$
shows the change in relative hourly earnings produced by within sector changes in characteristics when each group's wage structure is measured relatively to the baseline wage structure;

4) $[\{(\widetilde{w}_{t+1}^{cfpriv} \mid \hat{\beta}_{t+1}^* - \widetilde{w}_{t+1}^{priv} \mid \hat{\beta}_{t+1}^{priv}) - (\widetilde{w}_{t+1}^{cfpriv} \mid \hat{\beta}_t^* - \widetilde{w}_{t+1}^{cfpriv} \mid \hat{\beta}_t^{priv})\} +$
$\{\widetilde{w}_{t+1}^{pub} \mid \hat{\beta}_{t+1}^{pub} - \widetilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_{t+1}^*) - (\widetilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_t^{pub} - \widetilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_t^*)\}]$

shows the effect on the relative hourly earnings differential produced by changes in the wage advantage of public sector workers and the wage disadvantage of private sector workers, holding sector characteristics constant at t+1 levels;

While these individual components are of interest in identifying the individual source of any change in the relative hourly earnings of public and private sector workers, it is relatively easy to demonstrate that the addition of the first two components in the decomposition represents the change in relative hourly earnings due to changes in the characteristic component of the Oaxaca-Ransom decomposition, while the addition of the last two components gives the changes attributable to changes in the coefficient component of the Oaxaca-Ransom decomposition (O'Leary *et al.*, 2010).

9.5. Empirical results

9.5.1. Dummy variable estimates

The first method used is a simple dummy variable approach. We have regressed the logged hourly earnings on a set of explanatory variables and on the public sector dummy with quantile regression. Table 9.9 reports only the coefficients attached to the public sector. The whole set of results is given in Appendix Tables A9.5 to A9.11. A separate analysis on male and female workers have been undertaken and estimates at the 10^{th}, 25^{th}, 50^{th},75^{th} and 90^{th} percentiles are reported.

TABLE 9.9 Quantile regression estimates for males and females – public sector as a dummy variable

	10th	25th	50th	75th	90th	10th	25th	50th	75th	90th
						Bulgaria, 2003				
			Males					*Females*		
Public	0.1367***	0.1720***	0.1762***	0.1915***	0.2218***	0.1801***	0.0778***	0.025	-0.0171	-0.0553
	(0.0363)	(0.0335)	(0.0472)	(0.0530)	(0.0500)	(0.0462)	(0.0285)	(0.0385)	(0.0427)	(0.0483)
						Serbia, 2003				
			Males					*Females*		
Public	0.0121	-0.0213	-0.0845**	-0.2254***	-0.3478***	0.1785*	0.1850***	0.0874***	-0.0486	-0.1462
	(0.0622	(0.0326	(0.0344)	(0.0515)	(0.0853)	(0.0923)	(0.0494)	(0.0328)	(0.0515)	(0.0930)
						Russia, 2003				
			Males					*Females*		
Public	0.0772***	0.0310**	-0.0119	-0.0451***	-0.0791***	-0.0113	-0.0932***	-0.1050***	-0.0785***	-0.0953***
	(0.0205)	(0.0123)	(0.0106)	(0.0119)	(0.0147)	(0.0161)	(0.0117)	(0.0098)	(0.0129)	(0.0143)
						Tajikistan, 2003				
			Males					*Females*		
Public	-0.0132	-0.0869*	-0.1991***	-0.3818***	-0.4338***	-0.0625*	-0.1562***	-0.2424***	-0.2385***	-0.3144***
	(0.0683)	(0.0498)	(0.0426)	(0.0530)	(0.0695)	(0.0372)	(0.0536)	(0.0578)	(0.0616)	(0.0970)

Notes: Standard errors in parentheses * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

The estimates indicate that the wage premiums vary by quantiles. The public sector males in Bulgaria enjoy significant wage premium at all decile points of the earnings distribution. In 2003 the premium for Bulgarian men increases as going from lower to the higher quantiles – the premium is about 14% at the 10th percentile and above 22% at the 90th percentile. For females in Bulgaria and Serbia, the premium at the top end of the distribution appears to be insignificant. By contrast, the public sector males in Serbia and both males and females in Russia and Tajikistan, suffer a significant wage penalty at the higher part of the distribution. For instance, public sector employment for male employees in Russia shows a wage premium of 7.7% at the 10th percentile and significant wage penalty is evident at the 75th and 90th percentiles. A common result (apart from estimates for males in Bulgaria) is that at the very high quantiles, there is no premium from the public sector employment.

Reported results for other characteristics are also interesting and can briefly be described as follows. A significant and positive return to potential experience across the whole earnings distribution is found for females in Russia and for both males and females in Tajikistan. The returns to experience also vary over the distribution, tending to be higher at the top of the distribution than at the bottom for females in Russia and Tajikistan and are insignificant in Bulgaria and Serbia. The estimates of returns to qualifications at each quantile reveal that a higher rate of return is generally found for workers of both sexes with university qualification. Public sector employees with university qualification, for example, are paid more than those with secondary qualification. There is strong negative wage arrears effect in Russia. Significant and positive wage premium is found for workers with more than 10 years tenure within the firm in Bulgaria and Russia.

These results however, are correct only if the returns to human capital characteristics are the same in both sectors. To see whether the model should be estimated separately for each group of workers, we experiment by estimating a fully interacted model where all characteristics are interacted with the public sector dummy and we test if the interaction terms are significantly different from zero. The F-statistics on the joint significance of the interaction variables reject the hypothesis that workers in both sectors are subject to the same wage determination process[134]. Therefore, the separate quantile regressions in each sector are utilising in decomposing the public-private sector wage difference into a part explained by different distribution of characteristics and a part explained by different coefficients.

9.5.2. Decomposition results

As a benchmark, we first estimate the public-private sector wage differential assuming that the sector choice is exogenous. Estimates of the public-private sector hourly earnings differential for each sample are reported in Table 9.10. The table shows the raw public-private wage gap, the contribution of the coefficients and the covariates to the difference between the θ^{th} quantile of the public sector wage distribution and θ^{th} quantile of the private sector wage distribution and residual terms due to the differences between the empirical and simulated densities. The bootstrapped standard errors for these contributions are given in parenthesis. Separate analysis on male and female workers have been undertaken and estimates at the 10th, 25th, 50th, 75th and 90th percentile are reported.

There are substantial variations in the public sector wage differential across countries. First, the raw public-private wage gap generally decreases as one move up the earnings distribution in Serbia, Russia and Tajikistan. This indicates that employees in these countries with low potential wages tend to gain more from working in the public sector than workers with high potential wages. The raw gap is positive across the entire wage distribution in Bulgaria, Russia and for women in Serbia. In Serbia and Tajikistan at the top of the distribution, the raw gap is negative. For men in Tajikistan it starts at 41% for the first decile of earnings and drops to minus 41% at the 90th percentile.

[134] The F-test on the joint significance of the interaction variables, conducted on a standard OLS, yield an F-statistics of 2.54 for males and 4.98 for females in 2003 Bulgarian sample.

TABLE 9.10: Public-private wage decomposition results, 2003

BULGARIA	MALES				FEMALES			
Percentile	**Raw Gap**	**Coefficients**	**Covariates**	**Residual**	**Raw Gap**	**Coefficients**	**Covariates**	**Residual**
10th	0.263	0.206	0.119	-0.062	0.269	0.075	0.226	-0.032
		(0.000)	(0.000)			(0.011)	(0.008)	
25th	0.405	0.214	0.156	0.035	0.310	0.085	0.213	0.012
		(0.011)	(0.008)			(0.008)	(0.006)	
50th	0.463	0.224	0.202	0.037	0.288	0.081	0.219	-0.012
		(0.009)	(0.006)			(0.009)	(0.007)	
75th	0.445	0.199	0.231	0.015	0.295	0.043	0.210	0.042
		(0.010)	(0.007)			(0.010)	(0.007)	
90th	0.357	0.135	0.242	-0.020	0.223	0.023	0.209	-0.009
		(0.012)	(0.009)			(0.016)	(0.011)	
SERBIA	MALES				FEMALES			
10th	0.068	-0.027	0.108	-0.012	0.412	0.154	0.239	0.019
		(0.0245)	(0.0160)			(0.0206)	(0.0156)	
25th	0.079	-0.001	0.045	0.035	0.361	0.118	0.215	0.028
		(0.0111)	(0.0079)			(0.0115)	(0.0082)	
50th	0.027	-0.015	0.020	0.022	0.283	0.101	0.184	-0.002
		(0.0090)	(0.0059)			(0.0102)	(0.0073)	
75th	-0.119	-0.102	-0.020	0.003	0.189	0.045	0.158	-0.014
		(0.0130)	(0.0089)			(0.0111)	(0.0079)	
90th	-0.223	-0.164	-0.039	-0.020	0.109	-0.060	0.171	-0.002
		(0.0187)	(0.0124)			(0.0205)	(0.0141)	
RUSSIA	MALES				FEMALES			
10th	0.231	0.130	0.135	-0.034	0.183	-0.014	0.183	0.014
		(0.0238)	(0.0173)			(0.0182)	(0.0133)	
25th	0.162	0.045	0.128	-0.011	0.134	-0.072	0.184	0.022
		(0.0157)	(0.0109)			(0.0159)	(0.0106)	
50th	0.148	0.009	0.121	0.018	0.118	-0.077	0.165	0.030
		(0.0134)	(0.0092)			(0.0112)	(0.0081)	
75th	0.087	-0.015	0.116	-0.014	0.133	-0.067	0.157	0.043
		(0.0128)	(0.0090)			(0.0126)	(0.0092)	
90th	0.168	-0.036	0.112	0.092	0.090	-0.081	0.151	0.020
		(0.0131)	(0.0095)			(0.0169)	(0.0121)	
TAJIKISTAN	MALES				FEMALES			
10th	0.405	0.179	0.188	0.038	0.336	0.118	0.126	0.092
		(0.0247)	(0.0174)			(0.0173)	(0.0127)	
25th	0.288	0.033	0.159	0.096	0.251	0.035	0.177	0.039
		(0.0153)	(0.0107)			(0.0164)	(0.0112)	
50th	-0.154	-0.228	0.134	-0.060	0.183	-0.139	0.228	0.094
		(0.0182)	(0.0121)			(0.0145)	(0.0101)	
75th	-0.588	-0.476	0.054	-0.167	-0.446	-0.546	0.308	-0.208
		(0.0200)	(0.0136)			(0.0289)	(0.0187)	
90th	-0.405	-0.451	-0.009	0.055	-0.471	-0.742	0.255	0.016
		(0.0267)	(0.0201)			(0.0319)	(0.0227)	

Notes: Standard errors in brakets. The differential is calculating by every 5th percentile. The main percentile results are presented.

In Bulgaria, the estimated public sector wage differential in 2003 tends to be positive across the entire earnings distribution for both male and female employees, indicating that a positive wage premium is given to the public sector employees. It also decreases over the distribution. For instance, the sector wage differential for males decreases from approximately 21% at the 10th percentile to 14% at the 90th percentile level and respectively from 7.5% at the 10th percentile to 2.3% at the 90th percentile for females. These findings of positive public-private pay differential are in line with Falaris (2004), who shows that men and women who work in the public sector in Bulgaria are the ones who derive the greatest financial benefit from such employment. A possible reason for the public sector wage premium is that public sector workers in Bulgaria are better educated and take managerial positions more frequently than in the private sector workers. Moreover, competition for employment in the public sector in Bulgaria remains high because of the attached benefits such as pension rights. In adition, the advantage enjoyed by public sector workers tends to be higher for men than for women. For instance, the public sector wage differential for men at the 10^{th} is 20.6% compared to 7.5% differential for women at the same percentile level. In addition, for both men and women in Bulgaria, the wage differential caused by differences in observed characteristics is positive along the distribution and we may draw a conclusion that the advantage from greater productive characteristics of public sector workers in Bulgaria justifies the positive wage differential we found. These results are broadly in line with Voinea and Mihaescu (2012), who show that in Romania public sector offers higher wages. Decomposing the wage premium into the effect of personal characteristics, coefficients and residuals, the authors find that half of the premium can be attributed to personal characteristics.

An examination of the estimated coefficient effect of the gap for Serbia reveals a number of interesting patterns. The differential is negative across the entire earnings distribution for men indicating a comparative advantage enjoyed by private sector males. This is in line with Ognjenović (2011) who shows that the average worker in Serbia in 2003 employed in the public sector was underpaid compared to his or her counterparts in the private sector. This evidence suggests that public sector provides lower returns to skills for workers at the bottom of the distribution. Moreover, our estimates indicate that employment in the private sector is more beneficial for 'high-paid' men. At the top of the male distribution, a large proportion of the gap (74%) is due to differences in returns and this is in line with Lokshin and Jovanovic (2003) who show that male workers in the private sector in Serbia earn on average 9.4% higher wages than those in the public sector. Serbian women seems do better in the public sector as positive public sector wage differential is found at much higher percentile levels. For instance, at the 10^{th} percentile of the earnings distribution, there is a wage premium for public sector employment of 15.4% for females compared to a wage penalty of 2.7% for males. At the very top of the estimated conditional earnings distribution, however, women like men, tend to face a wage penalty from being employed in the public sector. For instance, the wage penalty at the 90th percentile is about 6% for females and 16% for males. These findings are broadly in line with Leping (2006), who shows positive wage differential at the lower part of the distribution in Estonia and negative for the median and higher parts of the distribution. Similarly, Blackaby *et al.* (1999) find positive public sector premium at the bottom of the distribution and negative at the top end for the UK. In addition, the estimated characteristics effect for males tends to be positive at the bottom and negative at the top, indicating favourable private sector characteristics for top earners. The characteristics effect for females actually outweighs the coefficient effects. It is difficult to infer about relative importance of the coefficient and covariate effects in male estimates for Serbia as the presence of a large residuals term.

Interestingly, females working in the public sector in Russia face a negative wage premium across the entire wage distribution. For example, women in the public sector at the top of the distribution earn around 8% less than those at the same percentile in the private sector. These findings are comparable with the results of an earlier analysis of the public sector differential in Russia. According to Brainerd (1998) wages of women employed in the state sector were 7.5 % lower than wages of women employed in the private sector in 1994 and 20.2% lower in 1997. Nesterova and Sabirianova (1998) find a wage premium of 20.8% for private sector workers in 1994. Similarly, Gimpelson and Lukiyanova (2009) report a wage penalty for public sector workers in Russia over the period 2000-2004. The authors show that in 46 regions in Russia, wages are markedly higher in the private sector and consequently conclude that to eliminate the negative wage gap, wages in the public sector should be linked to the private sector wages at regional level.

In contrast, males' wage premium declines more or less monotonically from high positive values at the lower end of the distribution to negative values at the higher end of the distribution. Male workers in Russia who are at the lower tail of the distribution earn about 13% more in the public sector than the corresponding private sector counterparts. However, those at the upper tail earn up to 4% less. These results are in line with some other findings (Blackaby *et al.*, 1999) and show that low earning males tend to gain more in the public sectors but higher earners tend to be better in the private sector. For both males and females in Russia, the effect of observable individual and job characteristics is positive and stable across the distribution. This suggests that public sector employees in Russia have individual characteristics that are more conducive to higher remuneration. This was confirmed by the descriptive statistics, where we found Russian public sector workers are more likely to have university degree (see Table A9.3). Public sector workers have on average 22 years of potential experience compared to 20 years for private sector workers. Longer tenure, associated for public sector workers, might suggests lower inter and intra-sector mobility. Finally, the duration of working hours in the public sector is shorter by 2 hours. Higher education and longer tenure and experience, other things being equal, tend to increase relative wages for public sector workers. Meanwhile, shorter working hours may have the opposite effect.

In Tajikistan for both males and females, we find a wage premium for the public sector at the lower part of the conditional wage distribution and significant wage penalty at the upper part of the distribution. A woman at the 10^{th} percentile level earns approximately 12% more than a woman at the same percentile in the private sector wage distribution. The difference declines rapidly toward higher percentiles and we can see that at the median of the public sector wage distribution women earn 13.9% less than those at the same percentile of the private wage distribution and the magnitude of the negative effect is larger at the top of the conditional wage distribution. At the top of the distribution, the advantage enjoyed by private sector workers is higher for women than for men. Moreover, the part explained by characteristics is not constant across the distribution and for males it decreases monotonically as we move across the conditional earnings distribution. A potential explanation to this negative and significant differential found at the top end of the distribution, is that fringe benefits and job security in Tajikistan are all known to be better for workers in the public sector and these non-wage benefits potentially could raise the effective wage paid to the public sector workers.

It is tempting to conclude that public sector in Serbia, Russia and Tajikistan runs the risk of being unattractive to highly qualified workers and that 'low-wage workers' might select themselves in the public sector. A commonly cited reason for public wage penalty at the upper end of the conditional wage distribution is public opposition to high pay for public servants (Katz and Kreuger, 1991; Lucifora and Meurs, 2006), while private sector is not subject to such opposition. This allows the private sector to use high pay to attract high-skilled workers. Some other studies find that overall satisfaction in the public sector is higher than in the private sector (Gardner and Oswald, 1999; Jürges, 2002). Jürges (2002) argues that the positive wage premium to private sector employment might in part be explained as a compensating differential. Since job satisfaction appears to be closely related to job quits, shirking or absenteeism, private sector employers might pay higher wages in order to reduce turnover (Jürges 2002). Similarly, Adamchik and Bedi (2000) argue that higher wages in the private sector may be compensating for the lower level of social benefits. Krishnan (2000) adds that in the absence of regulation or union pressure on wages, there are a number of different models consistent with the existence of a private sector premium. First, in a standard competitive model, wage premiums are likely to be observed if workers differ in their reservation wages or in ability (observable only to the employer), thus creating dispersion in the offered wage. In any empirical analysis, introducing controls for such differences should therefore eliminate evidence of a premium. The relatively higher risk aversion of workers might also induce premiums, if firms keep wages stable despite changing labour market conditions in order to protect workers against fluctuations in income. However, we are not able to account for all potential factors that might cause the private sector wage premium at the top of the distribution found in these countries. As our data have no information on work efforts (which are often lower in the public sector) or non-wage benefits (often higher in the public sector) the estimates are likely to underestimate the true public sector premium. Finally, as Leping (2006) notes, we should keep in mind that illegal employment and tax evasion have been problems during the early transition period, and as these problems occurred in the private sector, actual labour income in the private sector might be underestimated both by official wage statistics and probably by reported wages in the surveys.

9.5.3. The sector wage gap correcting for endogeneity

Endogenous sector choice

To describe the selection process between both sectors, we estimate probability to work in the public sector conditionally on a set of personal characteristics. The first results presented in this section pertain to the probit and single index analysis for the public sector choice process. Tables 9.11 to 9.14 provide results from the standard probit and single index (Ichimura, 1993) semi-parametric estimates (SLS) of the sector choice decision. The constant and the coefficient on the first reported continuous variable (years of potential work experience) are not identified in the single index model so they are normalized by setting them equal to the corresponding values in the probit model, thereby making the results of the two models comparable. At the bottom end of these tables the pseudo-likelihood ratio statistic is reported. The analysis is carried out separately by gender.

Bulgarian workers with higher education such as university and college degree are more likely to desire employment in the public sector. Secondary level of education for both males and females does not have significantly different effect on probability of being employed in the public sector. Potential experience plays a positive and significant role for females but it is not statistically significant for males. In line with Falaris (2004), in both models males and females in Bulgaria have a higher probability of employment in the public sector if they have more work experience within the firm. Secondary levels of education do not have significantly different effects on the probability of being employed in the public sector. Being married is positively associated with being in the public sector for males but it is insignificant in the female probit specification. The constructed identifying variable is insignificant in the probit estimates. However, it is statistically significant in the single index model. Finally, whether or not individuals live in the capital seems to have no impact on the incidence of public sector employment for females in Bulgaria.

Following Badel and Pena (2010), we test the null hypothesis of normal errors, given the existence of the single index which is consistent under both null and alternative hypothesis. The null hypothesis of normal errors (Bera-Jarque-Lee test (1984)) is clearly rejected at the 5% significance level for males; hence the residual terms are not normally distributed and the single index model should be used in the first step of selection for males. However, the null hypothesis is not rejected in the female's probit model and the power series expansion using the probit estimates is used.

TABLE 9.11: Estimates of the incidence of public sector employment, Bulgaria 2003

Pr(public=1)	*Males*		*Females*	
	Probit	SLS	Probit	SLS
exp	0.0165		0.0273*	
	(0.0126)		(0.0144)	
expsq	-0.0002	0.000	-0.0002	0.000***
	(0.0003)	(0.000)	(0.0003)	(0.000)
university	0.7270***	1.290***	0.7796***	0.959***
	(0.1471)	(0.237)	(0.1695)	(0.240)
college	0.7139***	1.889***	0.9998***	1.427***
	(0.2419)	(0.328)	(0.1959)	(0.332)
secondary	0.0862	-0.059	-0.0105	0.012
	(0.1203)	(0.121)	(0.1434)	(0.118)
1-2 years	0.1577	0.493***	0.2291	0.339***
	(0.1400)	(0.137)	(0.1514)	(0.138)
3-5 years	0.5579***	1.599***	0.5420***	1.168***
	(0.1293)	(0.273)	(0.1397)	(0.271)
6-10 years	0.9454***	2.697***	0.9767***	1.902***
	(0.1420)	(0.453)	(0.1511)	(0.439)
>10 years	1.3799***	3.452***	1.6013***	2.416***
	(0.1257)	(0.574)	(0.1441)	(0.547)
married	0.1963**	0.215***	-0.0049	0.160**
	(0.0970)	(0.074)	(0.1018)	(0.073)
urban	0.0905	-0.019	0.0741	0.012
	(0.1038)	(0.086)	(0.1173)	(0.088)
sofia	-0.1586	0.529***	0.2669	0.286
	(0.1925)	(0.171)	(0.2143)	(0.180)
instrument	-0.0042	-0.009***	-0.0019	-0.006***
	(0.0026)	(0.002)	(0.0027)	(0.002)
constant	-1.7201***	-1.7201	-1.7771***	-1.7771
	(0.1776)		(0.1955)	
N	1296	1296	1186	1186
Pseudo-likelihood	-645.6011		-590.6521	
Pseudo R^2	0.1915		0.2551	
Bera-Jarque-Lee test				
LM chi2(2)	9.196		0.737	
Prob > chi2	0.0100		0.691	

Notes: Standard errors in parentheses * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. The SLS estimator is Ichimura's (1993) estimator and it is performed by using *single.ado* programme for STATA. The constant and the potential experience coefficients in the single index model are normalized. The SLS procedure is highly computer intensive. For comparison, the estimation using the SLS estimator takes between 10 to 16 hours, while the estimation of the probit model, with the same specifications, takes a few seconds. The Bera-Jarque-Lee test is performed by using *bjltest.ado* programme for STATA.

In Serbia, university education increases the probability of being employed in the public sector for both males and females. However, there is an insignificant effect for the secondary and primary education levels and this is in line with Lokshin and Jovanovic (2003) findings. For women, (potential) experience significantly increases the probability of employment in the public sector. Similar to the findings in Bulgaria, the probability of being employed in the public sector increases with worker's tenure. The null hypothesis of normal probit errors is rejected for both genders (see Bera-Jarque-Lee test in Table 9.12). The SLS estimates are not reported as the estimation of the single index model with the data at hand was highly computer intensive and took much more time as compared to the other samples[135]. Note that the probit estimates are consistent under normally distributed errors.

TABLE 9.12: Estimates of the incidence of public sector employment, Serbia 2003

Pr(public=1)	*Males* Probit	*Females* Probit
exp	0.0024	0.0398**
	(0.0140)	(0.0162)
expsq	-0.0002	-0.0007***
	(0.0002)	(0.0003)
university	0.6479***	0.6383**
	(0.2307)	(0.3119)
gymnasium	0.0577	-0.0242
	(0.2913)	(0.3585)
secondary (4 years)	0.2979	0.0601
	(0.2027)	(0.2835)
secondary (1-2 years)	0.4552*	0.1368
	(0.2722)	(0.3644)
primary	0.1175	0.0533
	(0.2105)	(0.2653)
1-2 years	0.4331**	0.8009***
	(0.2184)	(0.2688)
3-5 years	0.5610***	0.8844***
	(0.1434)	(0.1907)
6-10 years	0.7232***	0.9279***
	(0.1417)	(0.1627)
>10 years	1.1191***	1.2633***
	(0.1186)	(0.1390)
married	-0.0096	0.0831
	(0.0964)	(0.1083)
urban	-0.0807	-0.0176
	(0.0752)	(0.1028)
Belgrade	-0.1997	-0.1654
	(0.1283)	(0.1578)
Vojvodina	-0.2365**	-0.2491*
	(0.1105)	(0.1438)
West Serbia	-0.2857**	-0.0051
	(0.1308)	(0.1826)
Šumadija and Pomoravlje	-0.0633	0.0371
	(0.1115)	(0.1546)
East Serbia	-0.0938	-0.1585
	(0.1348)	(0.1799)
instrument	0.0016*	0.0067***
	(0.0009)	(0.0012)
constant	-0.8450***	-1.7733***
	(0.2462)	(0.3329)
N	1466	984
Pseudo-likelihood	-910.4473	-538.1005
Pseudo R^2	0.1020	0.2090
Bera-Jarque-Lee test		
LM chi2(2)	13.029	4.939
Prob > chi2	0.019	0.0845

[135] We also tried a reduced specification of the employment selection equation for Serbia to see if the convergence problem was associated with particular variables and trying to reduce the computational time. However, the convergence has not been resolved.

TABLE 9.13: Estimates of the incidence of public sector employment, Russia 2003

Pr(public=1)	*Males*		*Females*	
	Probit	SLS	Probit	SLS
exp	-0.0046		-0.0055	
	(0.0034)		(0.0034)	
expsq	0.0001*	0.000***	0.0002**	0.000***
	(0.0001)	(0.000)	(0.0001)	(0.000)
university	0.0413	0.021	-0.0430	-0.019
	(0.0304)	(0.023)	(0.0313)	(0.031)
college	-0.1350**	0.001	-0.0772	-0.090**
	(0.0529)	(0.043)	(0.0517)	(0.054)
secondary	-0.0055	-0.013	-0.0776***	-0.094***
	(0.0229)	(0.018)	(0.0247)	(0.022)
primary	0.0036	-0.006	-0.0400	-0.110***
	(0.0316)	(0.026)	(0.0395)	(0.038)
married	-0.0007	-0.010	0.0716***	0.020
	(0.0246)	(0.015)	(0.0201)	(0.019)
1-2 years	0.2021***	0.019	0.1664***	0.626***
	(0.0308)	(0.024)	(0.0323)	(0.037)
3-5 years	0.4280***	0.045**	0.3931***	0.631***
	(0.0343)	(0.024)	(0.0369)	(0.043)
6-10 years	0.6392***	1.475***	0.6725***	2.527***
	(0.0333)	(0.066)	(0.0362)	(0.115)
>10 years	0.7412***	1.543***	0.8233***	3.210***
	(0.0311)	(0.075)	(0.0337)	(0.145)
arrears	0.1174***	0.031	0.1876***	1.336***
	(0.0240)	(0.018)	(0.0288)	(0.061)
Central region	-0.0590*	0.133***	0.0620*	0.283***
	(0.0319)	(0.023)	(0.0336)	(0.037)
North West	0.0200	0.062	0.0766**	0.215***
	(0.0352)	(0.027)	(0.0369)	(0.039)
Siberia	-0.1234***	0.014	-0.0254	0.122***
	(0.0355)	(0.028)	(0.0370)	(0.040)
Far-East	0.1879***	-0.107***	0.1944***	0.673***
	(0.0362)	(0.028)	(0.0386)	(0.049)
Urals	-0.2052***	-0.146***	-0.1430***	-0.595***
	(0.0402)	(0.034)	(0.0427)	(0.034)
Volga	-0.0737**	-0.210***	-0.0632*	-0.224***
	(0.0338)	(0.027)	(0.0355)	(0.036)
instrument	0.0162***	0.050***	0.0204***	0.088***
	(0.0003)	(0.002)	(0.0004)	(0.004)
constant	-0.9533***	-0.9533	-1.3305***	-1.3305
	(0.0476)		(0.0516)	
N	21874	21874	24318	24318
Pseudo-likelihood	-12848.138		-11436.95	
Pseudo R^2	0.1297		0.2362	
Bera-Jarque-Lee test				
LM chi2(2)	384.16		626.19	
Prob > chi2	0.0000		0.0000	

Notes: Standard errors in parentheses * $p < 0.10$, **$p < 0.05$, *** $p < 0.01$. The SLS estimator is Ichimura's (1993) estimator and it is performed by using *single.ado* programme for STATA. The constant and the potential experience coefficients in the single index model are normalized. The estimation using the SLS procedure took approximately 2 days for each gender. The Bera-Jarque-Lee test is performed by using *bjltest.ado* programme for STATA.

In Russia, university education does not seem to have significant effect on the probability of being employed in the public sector. Simply acquiring higher education does not increase chances of public sector employment in Russia. Male employees with college degree and females with secondary education are less likely to be employed in the public sector. This result might be reflected by the type of work available in both sectors. Other estimates reveal that the probability of working in the public sector is lower for employees in Urals and Volga regions. Females in the Central region are more likely to desire public sector employment. The constructed identifying variable is significant and positive in both male and female specifications. Hence, it appears that this instrument is successful at enabling identification. The null hypothesis of normal probit errors is clearly rejected at the 1% significance level in both male and female probit specifications and for this reason the set of quantile regression estimates are based on series expansions using the SLS estimates (see Table 9.13).

In Tajikistan's probit specification, not all levels of educational attainment are statistically significant. University educated workers are more likely to desire employment in the public sector. This variable shows significant and positive sign in both probit and single index models. Females with secondary general education are less likely to be in the public sector. The probability of being employed in the public sector also increases with worker's tenure. In addition, workers in Dushanbe, Khalton and Gbao regions are less likely to be offered public sector jobs. The coefficients on the identifying variable show significant and positive sign. Finally, the null hypothesis of normal errors is rejected for both males and females; hence the residual terms are not normally distributed. The estimated single index is used to form a series approximation for the unknown selection bias term.

TABLE 9.14

Estimates of the incidence of public sector employment, Tajikistan 2003

Pr(public=1)	*Males*		*Females*	
	Probit	SLS	Probit	SLS
exp	-0.003		0.019*	
	(0.008)		(0.011)	
expsq	0.000	0.000**	0.000**	0.000
	(0.000)	0.000)	(0.000)	(0.000)
university	0.569***	0.613***	0.549***	0.865***
	(0.104)	(0.078)	(0.153)	(0.090)
secondary technical	0.067	0.173*	0.367	-0.111
	(0.121)	(0.095)	(0.237)	(0.092)
secondary special	0.017	0.208**	0.460***	-0.129
	(0.116)	(0.085)	(0.154)	(0.082)
secondary general	-0.100	-0.032	0.025	-0.127**
	(0.091)	(0.059)	(0.104)	(0.064)
primary	0.130	0.183	0.235	0.269
	(0.202)	(0.127)	(0.186)	(0.179)
1-2 years of tenrue	0.193**	0.189**	0.109	-0.149
	(0.113)	(0.079)	(0.165)	(0.080)
3-5 years of tenure	0.389***	0.345***	0.203	0.016
	(0.108)	(0.076)	(0.160)	(0.077)
6-10 years of tenure	0.679***	0.587***	0.352**	0.854***
	(0.115)	(0.084)	(0.171)	(0.096)
>10 years of tenure	0.878***	0.796***	0.582***	0.960***
	(0.110)	(0.078)	(0.161)	(0.096)
married	0.040	-0.072	-0.152*	-0.034
	(0.088)	(0.050)	(0.082)	(0.048)
urban	-0.041	-0.029	-0.029	-0.097*
	(0.066)	(0.053)	(0.093)	(0.051)
Dushanbe	-0.549***	-0.449***	-0.270*	-0.999***
	(0.103)	(0.068)	(0.130)	(0.104)
Khalton region	-0.749***	-0.712***	-0.587***	-1.319***
	(0.102)	(0.066)	(0.135)	(0.122)
Rrg region	-0.059	-0.063	-0.004	-0.103
	(0.123)	(0.088)	(0.158)	(0.075)
Gbao region	-0.532***	-0.448***	-0.259*	-0.980***
	(0.118)	(0.081)	(0.155)	(0.121)
instrument	0.013***	0.014***	0.017***	0.025***
	(0.001)	(0.001)	(0.001)	(0.002)
constant	-0.424***	-0.424	-0.585***	-0.585
	(0.163)		(0.221)	
N	3006	3006	1832	1832
Pseudo-likelihood	-1630.839		-926.815	
Pseudo R^2	0.2145		0.2633	
Bera-Jarque-Lee test				
LM chi2(2)	13.517		7.990	
Prob > chi2	0.0011		0.0183	

Notes: Standard errors in parentheses * $p < 0.10$, ** $p < 0.05$, ***$p < 0.01$. The SLS estimator is Ichimura's (1993) estimator and it is performed by using *single.ado* programme for STATA. The constant and the potential experience coefficients in the single index model are normalized. The Bera-Jarque-Lee test is performed by using *bjltest.ado* programme for STATA.

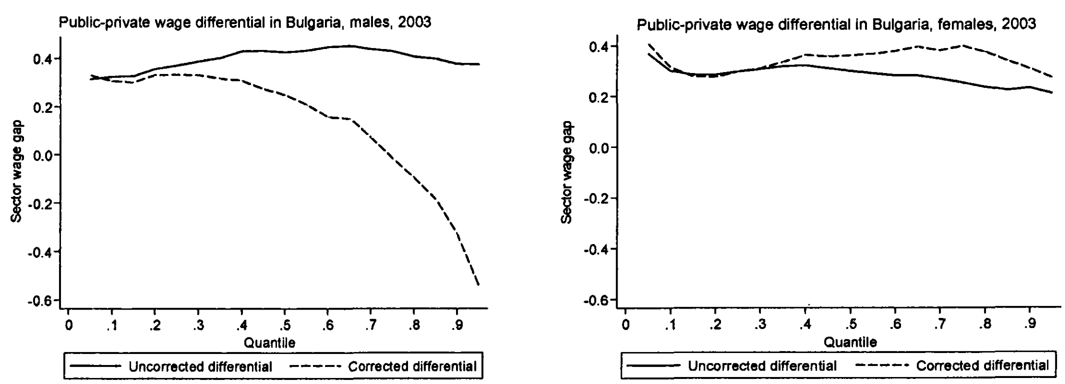
Endogeneity corrected sector wage differential

Once wages are consistently estimated for the endogenous sector choice, differences in the public and private sector are decomposed. The simulated differentials corrected for endogenous sector choice, presented in Tables 9.15 to 9.18, are the differentials that we observe in each country if the employees were sorted randomly between sectors conditionally on their labour makret characteristics. Further, we present the gap that remains due to the unobservable characteristics and gap due to observable characteristics, as described in Section 9.4.4. The bootstrapped standard errors for these contributions are given in parenthesis. Bootstrap estimates are based on 800 replications. The parameter estimates of the intercept of the wage functions follow the Andrews and Schafgans (1998) estimator, with the smoothing function $s(\cdot)$ as given in Section 9.4.3 and b set equal to one. The second stage quantile regression results are reported in Appendix Tables A9.12-A9.18[136].

In Figures 9.9 to 9.12, we compare by country, the uncorrected simulated public sector wage differential with the endogeneity corrected differential. The uncorrected differentials are taken from Table 9.10 and summarises the estimated coefficient and covariate components, and excludes residual terms[137]. No confidence intervals are plotted to avoid to surchange the figures.

FIGURE 9.9

Uncorrected and endogeneity corrected sector wage differential in Bulgaria, 2003



Notes: The uncorrected differential is the differential simulated without correcting for endogeneity, equals the coefficient and covariate components from Table 9.10, and excludes residual terms. The corrected differential is the differential that we would observe if the employees were sorted randomly between sectors conditional on their characteristics.

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136 Analysis of these estimates is beyond the scope of the chapter.

137 This is the predicted wage distribution based on the MM-OR procedure conditional on the variables of interest and does not equals the actual wage distribution as it excludes the residual terms.

Overall, the estimates indicate that non-random sector selection is an issue in the calculation of the sector wage differential. The inclusion of the selectivity corrected power series terms in the computation of the sector wage gap seems to produce different estimates compared to the uncorrected wage gap estimates. We note that the endogeneity corrected differential for males in Bulgaria is much lower than the uncorrected one. The premium for the public sector workers decreases throughout the wage distribution and is negative at the top. It therefore follows that when the random sector choice is allowed for, men at the top of the wage distribution are rewarded better in the private sector. The reverse is true for females, where the corrected differential is higher than the uncorrected one (see Figure 9.9). Hence, a randomly drawn woman employed in the private sector is predicted to earn less in that sector than would be earned in the public sector. Allowing for the endogeneity in sector choice, the advantage enjoyed by public sector workers is higher for females, which is in contrast with the uncorrected estimates, where we found the advantage enjoyed by public sector workers was higher for males. Hence, correcting for endogenous sector choice it is more worthwhile for women to remain in the public sector than for men[138].

TABLE 9.15: Decomposition of public-private wage differential in Bulgaria correcting for endogeneity, 2003

	10th	25th	50th	75th	90th
	MALES				
Gap due to unobservables $[Q_\theta(\tilde{w}^{pub})-Q_\theta(\tilde{w}^{s\text{-}pub})]-[Q_\theta(\tilde{w}^{priv})-Q_\theta(\tilde{w}^{s\text{-}priv})]$	0.315 (0.0094)	0.378 (0.0091)	0.873 (0.0124)	0.563 (0.0416)	-0.300 (0.0309)
Gap due to observables $[Q_\theta(\tilde{w}^{ctfpub})-Q_\theta(\tilde{w}^{s\text{-}pub})]-[Q_\theta(\tilde{w}^{ctfpriv})-Q_\theta(\tilde{w}^{s\text{-}priv})]$	0.008 (0.0092)	0.046 (0.0087)	0.625 (0.0184)	0.578 (0.0451)	0.028 (0.0361)
	FEMALES				
Gap due to unobservables $[Q_\theta(\tilde{w}^{pub})-Q_\theta(\tilde{w}^{s\text{-}pub})]-[Q_\theta(\tilde{w}^{priv})-Q_\theta(\tilde{w}^{s\text{-}priv})]$	0.289 (0.0077)	0.296 (0.0062)	0.562 (0.0080)	0.285 (0.0183)	0.151 (0.0212)
Gap due to observables $[Q_\theta(\tilde{w}^{ctfpub})-Q_\theta(\tilde{w}^{s\text{-}pub})]-[Q_\theta(\tilde{w}^{ctfpriv})-Q_\theta(\tilde{w}^{s\text{-}priv})]$	-0.025 (0.0072)	-0.001 (0.0067)	0.201 (0.0103)	-0.113 (0.0176)	-0.158 (0.0186)

Notes: The bootstrapped standard errors in brakets. The differential is calculating by every 5th percentile. The main percentile results are presented.

The effect of the unobservable is positive across the distribution and turns to negative at the 90th percentile for males. For females, the unobservable gap is lower as compared to the gap for males with a more obvious decline at the 90th percentile. A

[138] See Lucifora and Meurs (2004) for similar results.

large part of the selection effect is due to unobservables. Comparing the present results with those presented in Table 9.10, accounting for differences in observables between public and private sector and considering that workers were sorted randomly between sectors, has more of an impact on the sector wage gap, especially at the 50th and 75th percentiles (see Table 9.15).

The endogeneity corrected sector wage differential for males in Serbia is lower than the uncorrected one. It also is negative across the entire wage distribution (see Figure 9.10). The gap declines monotonically toward higher percentiles and shows that the premium in the public sector decreases throughout the wage distribution. Similar to the uncorrected estimates, men at the top of the distribution are still rewarded better in the private sector. At the top of the distribution, a large proportion of the gap is due to the differences in unobservable characteristics. The effect of observables is positive across the male distribution and insignificantly different from zero at the top (see Table 9.16). In contrast, the estimated endogeneity corrected sector wage differential for females is slightly larger at the bottom of the distribution. The advantage enjoyed by public sector workers is much higher for women than for men.

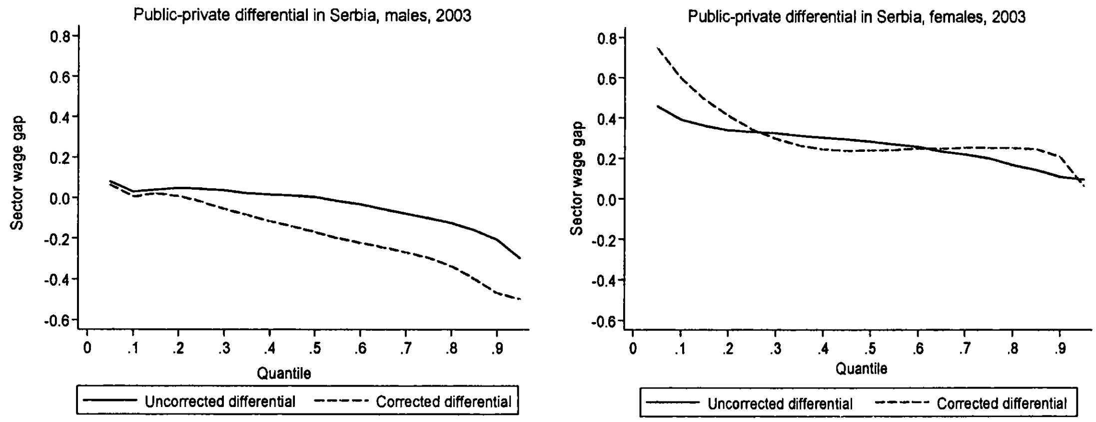
TABLE 9.16: Decomposition of public-private wage differential in Serbia correcting for endogeneity, 2003

	10th	**25th**	**50th**	**75th**	**90th**
	MALES				
Gap due to unobservables	0.038	0.039	-0.028	-0.190	-0.463
$[Q_\theta(\tilde{w}^{pub})-Q_\theta(\tilde{w}^{s-pub})]-[Q_\theta(\tilde{w}^{priv})-Q_\theta(\tilde{w}^{s-priv})]$	(0.0173)	(0.0121)	(0.0088)	(0.0192)	(0.0269)
Gap due to observables	0.032	0.058	0.139	0.103	0.004
$[Q_\theta(\tilde{w}^{ctfpub})-Q_\theta(\tilde{w}^{s\_pub})]-[Q_\theta(\tilde{w}^{ctfpriv})-Q_\theta(\tilde{w}^{s\_priv})]$	(0.0075)	(0.0074)	(0.0075)	(0.0173)	(0.0278)
	FEMALES				
Gap due to unobservables	0.023	-0.423	-0.431	-0.483	-0.463
$[Q_\theta(\tilde{w}^{pub})-Q_\theta(\tilde{w}^{s-pub})]-[Q_\theta(\tilde{w}^{priv})-Q_\theta(\tilde{w}^{s-priv})]$	(0.0572)	(0.0700)	(0.0527)	(0.1248)	(0.1916)
Gap due to observables	-0.577	-0.777	-0.672	-0.737	-0.674
$[Q_\theta(\tilde{w}^{ctfpub})-Q_\theta(\tilde{w}^{s\_pub})]-[Q_\theta(\tilde{w}^{ctfpriv})-Q_\theta(\tilde{w}^{s\_priv})]$	(0.0388)	(0.0108)	(0.0079)	(0.1340)	(0.1557)

Notes: The bootstrapped standard errors in brakets. The differential is calculating by every 5th percentile. The main percentile results are presented.

FIGURE 9.10

Uncorrected and endogeneity corrected sector wage differentials in Serbia, 2003



Notes: The uncorrected differential is the differential simulated without correcting for endogeneity, equals the coefficient and covariate components from Table 9.10, and excludes residual terms. The corrected differential is the differential that we would observe if the employees were sorted randomly between sectors conditional on their characteristics.

Similar results were found for males in Russia, where the corrected differential is much lower than the uncorrected one. Hence, a randomly drawn man employed in the public sector is predicted to earn less in that sector than would be earned by an observationally identical man, assuming that sector choice is exogenous. Both the gap due to the unobservables and observables tends to be negative at the top of the wage distribution. Unfortunately, we have not been able to estimate the model for females due to a convergence problem with the female sample due to inclusion of the power series in the QR model. The convergence rate was very slow in particular at the lower part of the distribution – 0.05 and 0.1 percentiles. We have tried experimenting with different specifications of the power series produced from the single index and probit models and with increasing the number of simulations; however the problem has not been resolved. With the Stata *capture* option, we should be able to overcome the convergence problem at the lower part of the distribution and move on the next percentiles level, however, this could affect the final simulation results[139].

[139] Kwak (2010) propose that the convergence could be achived for a large enough sample size.

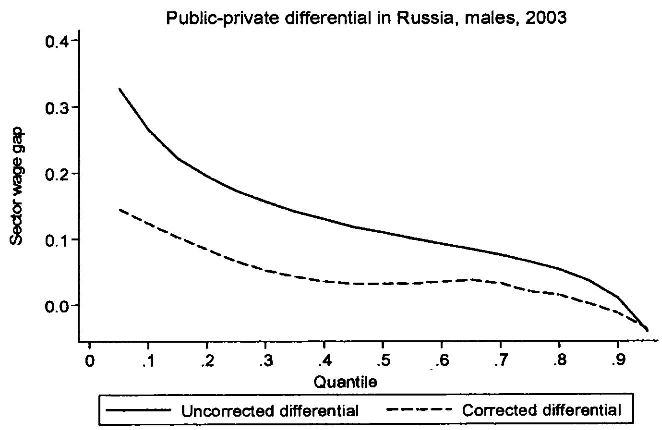
TABLE 9.17: Decomposition of public-private wage differential for males in Russia correcting for endogeneity, 2003

	10th	25th	50th	75th	90th
Gap due to unobservables $[Q_\theta(\tilde{w}^{pub})-Q_\theta(\tilde{w}^{s\text{-}pub})]-[Q_\theta(\tilde{w}^{priv})-Q_\theta(\tilde{w}^{s\text{-}priv})]$	0.265 (0.0244)	0.173 (0.0159)	-0.322 (0.0181)	-0.332 (0.0179)	-0.262 (0.0205)
Gap due to observables $[Q_\theta(\tilde{w}^{ctfpub})-Q_\theta(\tilde{w}^{s\text{-}pub})]-[Q_\theta(\tilde{w}^{ctfpriv})-Q_\theta(\tilde{w}^{s\text{-}priv})]$	0.141 (0.0089)	0.107 (0.0072)	-0.354 (0.0149)	-0.352 (0.0152)	-0.249 (0.0157)

Notes: The bootstrapped standard errors in brakets. The differential is calculating by every 5th percentile. The main percentile results are presented.

FIGURE 9.11

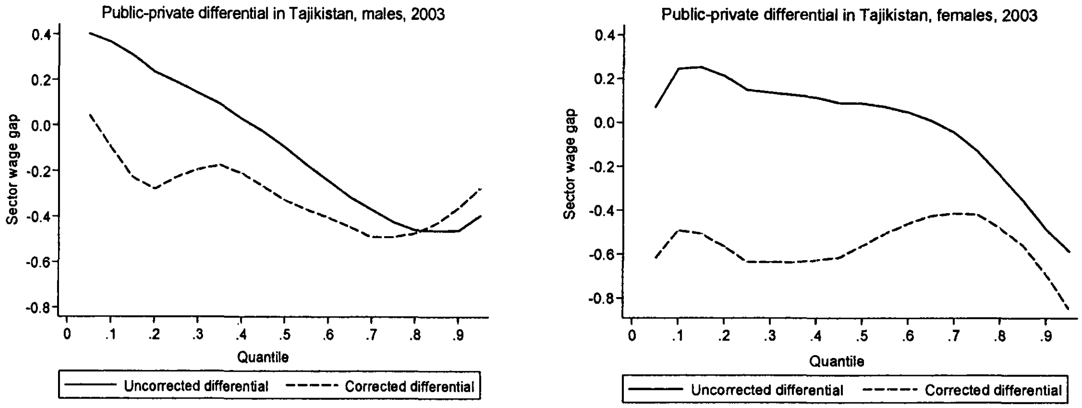
Uncorrected and endogeneity corrected sector wage differentials for males in Russia, 2003



The selection correction seems to be very important for the Tajikistan data. We find that the corrected differential for both males and females is negative across the entire wage distribution and rather low as compared to the uncorrected one. The lowest wage gap for males is around -25% and it is observed in the lower-end of the distribution. The public sector wage penalty for males declines more monotonically toward higher percentiles where corrected differential tends to be higher compared to the uncorrected one. Allowing for the endogeneity in the sector choice the advantage enjoyed by private sector workers is higher for females. In addition, for males in Tajikistan, the portion of selection effect related to the unobservable characteristics is similar in size and shape across the distribution with the uncorrected unexplained component with a more obvious decline between the 75th and 80th percentiles (see Figure A9.1). Similar to the uncorrected estimates, the advantage enjoyed by private sector workers is higher for women than for men.

FIGURE 9.12

Uncorrected and endogeneity corrected sector wage differential in Tajikistan, 2003



Notes: The uncorrected differential is the differential simulated without correcting for endogeneity, equals the coefficient and covariate components from Table 9.10, and excludes residual terms. The corrected differential is the differential that we would observe if the employees were sorted randomly between sectors conditional on their characteristics.

TABLE 9.18: Decomposition of public-private wage differential in Tajikistan correcting for endogeneity, 2003

	10th	25th	50th	75th	90th
	MALES				
Gap due to unobservable	-0.040	0.197	-0.117	-0.530	-0.415
$[Q_\theta(\tilde{w}^{pub})-Q_\theta(\tilde{w}^{s\_pub})]-[Q_\theta(\tilde{w}^{priv})-Q_\theta(\tilde{w}^{s\_priv})]$	(0.0327)	(0.0157)	(0.0196)	(0.0255)	(0.0302)
Gap due to observables	0.056	0.423	0.209	-0.042	-0.055
$[Q_\theta(\tilde{w}^{ctfpub})-Q_\theta(\tilde{w}^{s\_pub})]-[Q_\theta(\tilde{w}^{ctfpriv})-Q_\theta(\tilde{w}^{s\_priv})]$	(0.0238)	(0.0155)	(0.0095)	(0.0190)	(0.0199)
	FEMALES				
Gap due to unobservable	-0.428	-0.223	0.081	-0.716	-1.040
$[Q_\theta(\tilde{w}^{pub})-Q_\theta(\tilde{w}^{s\_pub})]-[Q_\theta(\tilde{w}^{priv})-Q_\theta(\tilde{w}^{s\_priv})]$	(0.0174)	(0.0181)	(0.0142)	(0.0261)	(0.0362)
Gap due to observables	0.063	0.413	0.641	-0.299	-0.349
$[Q_\theta(\tilde{w}^{ctfpub})-Q_\theta(\tilde{w}^{s\_pub})]-[Q_\theta(\tilde{w}^{ctfpriv})-Q_\theta(\tilde{w}^{s\_priv})]$	(0.0117)	(0.0126)	(0.0126)	(0.0213)	(0.0331)

Notes: The bootstrapped standard errors in brakets. The differential is calculating by every 5^{th} percentile. The main percentile results are presented.

9.5.4. Evolution in the public-private wage differential

A further question of interest is the public-private wage differential over time. In Table 9.19 we compare Bulgarian sector wage differential estimates, obtained by running the MM-OR separately for two years (1986 and 2003). In the first three columns, we present the raw public-private gap estimates for selected years and for the change between them. Thus a negative difference indicates that the public sector earnings gap decreases over time, while the reverse is true if we observe the positive wage difference. In the next columns, we present the contribution of the coefficients and the covariates for both genders at the 10^{th}, 25^{th}, 50^{th}, 75^{th} and 90^{th} percentiles. The bootstrapped standard errors for these contributions are given in parenthesis[140]. In Figures 9.13 and 9.14 we compare the estimated coefficient components of the gap for the two time periods.

Overall, the observed positive change in the raw wage gap implies that the sector wage differential in Bulgaria has risen over time, this is true for both men, and women (see Table 9.19). There seems to have been an improvement in the relative pay of the public sector and the differential has improved at the upper part of the distribution. In particular, the change in the raw gap is higher at the top of the earnings distribution, where the gap for males increases from minus 14.7 log percentage points in 1986 to 35.7 log percentage points in 2003, and from minus 16.2 log percentage points to 22.3 log percentage points for females. This suggests that the gains to the public sector employees tend to be concentrated among workers at the top half of the conditional earnings distribution. Only at the lower percentile levels, the coefficient effect (unexplained part) of the sector wage differential decreases over time. For example, the wage premium from working in the public sector for the low-waged male employees (the 10^{th} percentile) declines from 32.6% in 1986 to 20.6% in 2003 and from 12.7% in 1986 to 7.5% in 2003 for females (see Table 9.19).

140 In bootstrapping the 95 % confidence intervals, 800 replications were used.

TABLE 9.19

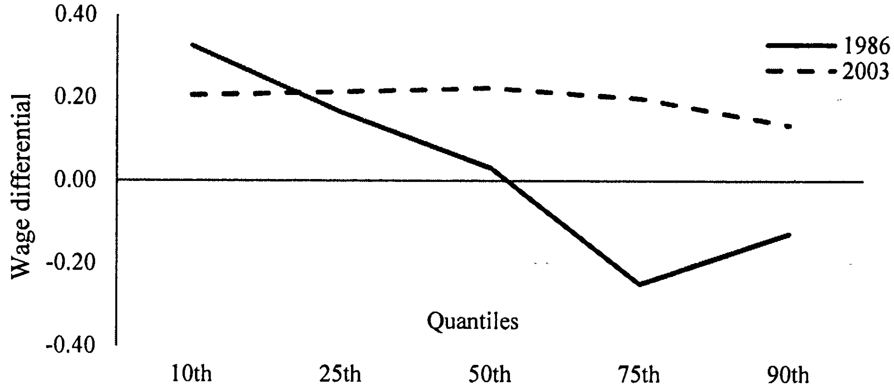
Decomposition of changes in the public-private sector wage differential for male and female employees in Bulgaria, 1986 and 2003

Sample estimates				Decomposition of changes:					
				MALES					
	Raw differential			1986			2003		
Percentiles	1986	2003	Change	Coefficients	Covariates	Residuals	Coefficients	Covariates	Residuals
10th	0.503	0.262	-0.241	0.326	0.119	0.058	0.206	0.119	-0.062
				(0.017)	(0.011)		(0.000)	(0.000)	
25th	0.230	0.405	0.175	0.165	0.084	-0.019	0.214	0.156	0.035
				(0.010)	(0.006)		(0.011)	(0.008)	
50th	0.080	0.463	0.383	0.031	0.039	0.010	0.224	0.202	0.037
				(0.006)	(0.004)		(0.009)	(0.006)	
75th	-0.287	0.445	0.732	-0.250	0.001	-0.038	0.199	0.231	0.015
				(0.010)	(0.004)		(0.010)	(0.007)	
90th	-0.147	0.357	0.504	-0.128	-0.035	0.016	0.135	0.242	-0.020
				(0.007)	(0.004)		(0.012)	(0.009)	
				FEMALES					
				1986			2003		
Percentiles	1986	2003	Change	Coefficients	Covariates	Residuals	Coefficients	Covariates	Residuals
10th	0.223	0.269	0.046	0.127	0.074	0.022	0.075	0.226	-0.032
				(0.018)	(0.012)		(0.011)	(0.008)	
25th	0.179	0.310	0.131	0.096	0.067	0.016	0.085	0.213	0.012
				(0.009)	(0.006)		(0.008)	(0.006)	
50th	0.077	0.288	0.211	0.030	0.045	0.002	0.081	0.219	-0.012
				(0.005)	(0.004)		(0.009)	(0.007)	
75th	-0.157	0.296	0.453	-0.151	0.007	-0.013	0.043	0.210	0.042
				(0.006)	(0.004)		(0.010)	(0.007)	
90th	-0.162	0.223	0.385	-0.121	-0.050	0.009	0.023	0.209	-0.009
				(0.008)	(0.005)		(0.016)	(0.011)	

Notes: Bulgaria Household surveys 1986 and 2003.

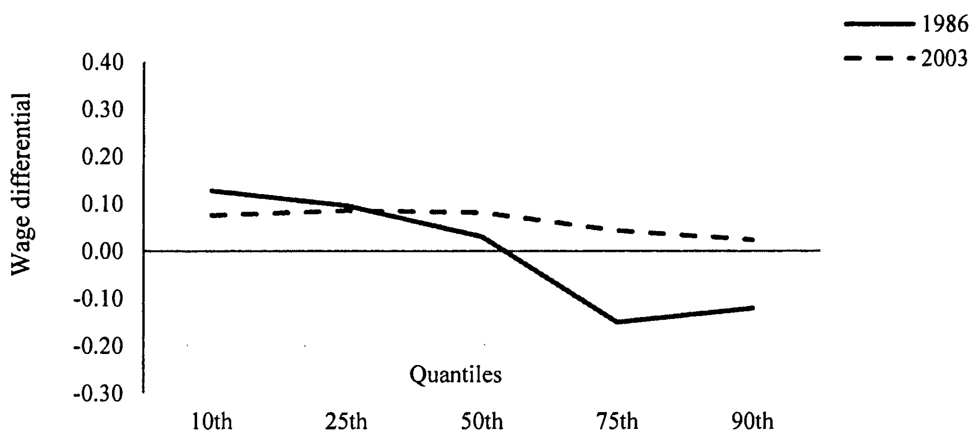
The results of the decomposition clearly show that for both genders, the unexplained part of the differential in Bulgaria declines from the lower to the upper deciles. In 1986, the estimated public sector wage differential tends to be positive at the bottom of the distribution but negative at the top, which indicate that public sector workers at the lower part of the conditional distribution enjoy significant wage premium, but at the top they face a wage penalty. For example, in the top of the wage distribution, the negative effect is about 12%. This is true for both men and women in 1986. Similarly to the 2003 findings, the advantage enjoyed by public sector workers at the bottom of the distribution in 1986 is higher for men than for women. For males, the estimated coefficient effect in 1986 decreases across the distribution from 32.6% at the 10th percentile to minus 12.8% at the 90th percentile. For females, the differential in the same year varies from 12.7% at the 10th percentile to a minus 12.1% at the 90th percentile (see Table 9.19). These results show that lower earners in 1986 tended to gain more in the public sector but higher earners tend to be better in the private sector.

FIGURE 9.13: Estimated public-private wage differentials for males, Bulgaria



Source: Authors' decomposition estimates, 1986 and 2003 Bulgarian data.

FIGURE 9.14: Estimated public-private wage differentials for females, Bulgaria



Source: Authors' decomposition estimates, 1986 and 2003 Bulgarian data.

There have been relative improvements in the quality of public sector workers as differences in endowments seem to drive the increased wage premium across the distribution. While for both genders in 1986 the individual covariates only explain a small share of the observed gap, the effect is larger in 2003 and we conclude that changes in the characteristic component over time contributed to the increase in the relative earnings of public sector workers in Bulgaria. For example, at the top of the male earnings distribution 68% of the observed gap in 2003 was due to the characteristic effect while the same effect was about 24% in 1986. This suggests that public sector employees gained individual and job characterises that are more conductive to higher remuneration.

Another important fact is that the position of both men and women in the Bulgarian labour market has changed during the economic transition to a market economy. As we have discussed, many jobs have been lost due to the economic restructuring and privatization and small number of new jobs has been created. Our results indicate that public sector in Bulgaria has to have taken a role of the employment generator. Finally, the increased public sector premium over time that we found might also be a result of some unobserved factors such as working conditions, job security and other non-wage aspects of employment.

A detailed examination of the public-private wage differential in Bulgaria over time using the modified MM-Donohue-Heckman procedure is presented in Table 9.20. Results on male workers have been undertaken and estimates at the 10^{th}, 20^{th}, 30^{th}, 40^{th}, 50^{th}, 60^{th}, 70^{th}, 80^{th} and 90^{th} percentile are shown. In order to identify the source of any change in the estimated differential, the effects of all four components given in methodology Section 9.4.5 are reported. Bootstrapped standard errors are given in parentheses. Again the convergence problem with the female sample does not allow us to estimate the model for females.

The MM-Donohue-Heckman estimates confirm the findings that for the period of 1986 to 2003, the sector wage differential for males in Bulgaria has increased across the entire earnings distribution. The rise in the relative earnings of public sector workers tends to be higher at the upper part of the distribution (70^{th} and 80^{th} percentiles) than at the bottom, and the contribution made by changes in the relative workers' characteristics tends to be higher at the top of the male distribution than at the bottom. Within the 'characteristic effect' changes in the parameters used to generate the baseline wage structure, accounted for significant part of the overall characteristics effect. Within the 'coefficient effect' changes in the wage advantage of public sector workers and disadvantage of private sector, holding sector characteristics constant at 2003 level, accounts for a major part of the coefficient effect (see Table 9.20).

TABLE 9.20

DHMM decomposition of changes in the public-private wage gap at different points in the earnings distribution for male workers, Bulgaria

	1986 to 2003								
	10th	20th	30th	40th	50th	60th	70th	80th	90th
1. $[(\widetilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_t^* - \widetilde{w}_{t+1}^{cfpriv} \mid \hat{\beta}_t^*) - (\widetilde{w}_t^{cfpub} \mid \hat{\beta}_t^* - \widetilde{w}_t^{cfpriv} \mid \hat{\beta}_t^*)]$	-0.005	-0.004	0.004	-0.008	-0.001	0.004	0.012	0.037	0.048
	(0.020)	(0.015)	(0.012)	(0.010)	(0.009)	(0.010)	(0.011)	(0.011)	(0.008)
2. $[(\widetilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_{t+1}^* - \widetilde{w}_{t+1}^{cfprv} \mid \hat{\beta}_{t+1}^*) - (\widetilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_t^* - \widetilde{w}_{t+1}^{cfprv} \mid \hat{\beta}_t^*)]$	0.005	0.004	0.043	0.106	0.161	0.168	0.189	0.209	0.222
	(0.014)	(0.011)	(0.010)	(0.009)	(0.009)	(0.010)	(0.010)	(0.009)	(0.012)
3. $\{(\widetilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_t^{pub} - \widetilde{w}_t^{pub} \mid \hat{\beta}_t^{pub}) - (\widetilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_t^* - \widetilde{w}_t^{cfpub} \mid \hat{\beta}_t^*)\} + [(\widetilde{w}_{t+1}^{cfpriv} \mid \hat{\beta}_t^* - \widetilde{w}_t^{cfprv} \mid \hat{\beta}_t^*) - (\widetilde{w}_{t+1}^{cfpriv} \mid \hat{\beta}_t^{priv} - \widetilde{w}_t^{prv} \mid \hat{\beta}_t^{priv})\}]$	0.094	0.099	0.076	0.042	0.039	0.087	0.208	0.299	0.109
	(0.027)	(0.021)	(0.015	(0.013)	(0.014)	(0.016)	(0.018)	(0.016)	(0.008)
4. $[\{(\widetilde{w}_{t+1}^{cfpriv} \mid \hat{\beta}_{t+1}^* - \widetilde{w}_{t+1}^{priv} \mid \hat{\beta}_{t+1}^{priv}) - (\widetilde{w}_{t+1}^{cfpriv} \mid \hat{\beta}_t^* - \widetilde{w}_{t+1}^{cfpriv} \mid \hat{\beta}_t^{priv})\} + \{\widetilde{w}_{t+1}^{pub} \mid \hat{\beta}_{t+1}^{pub} - \widetilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_{t+1}^*) - (\widetilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_t^{pub} - \widetilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_t^*)\}]$	-0.210	-0.072	0.071	0.159	0.156	0.219	0.202	0.165	0.156
	(0.020)	(0.017)	(0.015	(0.012)	(0.013)	(0.014)	(0.013)	(0.012)	(0.016)
Change in covariates component	0.000	0.000	0.047	0.098	0.160	0.172	0.201	0.246	0.270
Change in coefficient component	-0.116	0.027	0.147	0.201	0.195	0.306	0.410	0.464	0.265

Notes: The *component 1* shows the change in relative hourly earnings attributable to between sector changes in characteristics of public and private sectors, evaluated at the baseline wage structure; *component 2* shows the change in the relative hourly earnings explained by changes in the baseline wage structure, holding characteristic differences constant at *t+1* (2003) level; *component 3* shows the change in relative hourly earnings produced by within sector changes in characteristics when each group's wage structure is measured relatively to the baseline wage structure; *component 4* shows the effect on the relative hourly earnings differential produced by changes in the wage advantage of public sector workers and the wage disadvantage of private sector workers, holding sector characteristics constant at level *t+1* (2003).

The Russian decomposition results are given in Table 9.21. The estimates are obtained by running the MM-OR decomposition separately for the two periods (1994 and 2003). The results over the whole distribution are best viewed graphically. For reference, recall that Figure 9.15 and Figure 9.16 present the estimated coefficient effects by gender for the sector wage differential at different points of the conditional earnings distribution for 1994 and 2003. No confidence intervals are plotted to avoid to surchange the figures.

The observed positive change in the raw wage gaps in Russia implies that the raw public-private wage differential has risen over time. For both men and women and similar to the Bulgarian findings, the change in the raw gaps is higher at the top of the earnings distribution, where it turns from negative to positive. However, at the higher percentiles the coefficient effect stays negative in both years, indicating that workers with high potential wages tend to gain more from working in the private sector. Low-paid workers in Russia (generally those with low skills) benefit more from employment in the public sector.

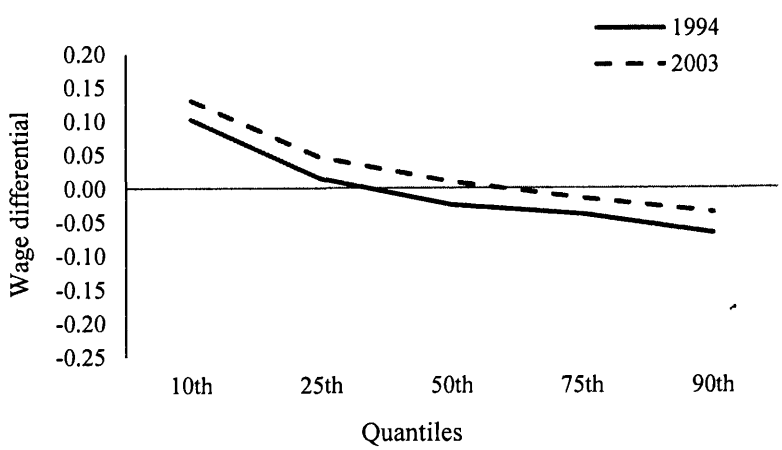
For males, the sector wage differential slightly increases over time but overall it stays fairly stable (see Figure 9.15). In 2003, the differential is more positive for the lower percentiles (the wage premium from working in the public sector for the low-waged male employee's raises from 10% in 1994 to 13% in 2003) and it is less negative at the top of the distribution. This is in sharp contrast to female distribution, where public sector differential changes substantially over the observed period and the changes for the extremes of the wage distribution are more pronounced. For instance, the public sector wage differential in 1994 is 14% at the lower part of the female's distribution and become insignificantly below zero in 2003 (see Figure 9.16).

TABLE 9.21: Decomposition of changes in the public-private sector wage differential for male and female employees in Russia, 1994 and 2003

Sample estimates				Decomposition of changes:					
						MALES			
	Raw Differential			1994			2003		
Percentiles	1994	2003	Change	Coefficients	Covariates	Residuals	Coefficients	Covariates	Residuals
10th	0.088	0.231	0.143	0.102	-0.037	0.023	0.130	0.135	-0.034
				(0.0274)	(0.019)		(0.023)	(0.017)	
25th	0.016	0.163	0.146	0.014	-0.061	0.063	0.045	0.128	-0.010
				(0.019)	(0.013)		(0.015)	(0.010)	
50th	-0.113	0.148	0.261	-0.025	-0.090	0.002	0.009	0.121	0.018
				(0.015)	(0.011)		(0.013)	(0.009)	
75th	-0.135	0.087	0.222	-0.039	-0.092	-0.004	-0.015	0.116	-0.014
				(0.016)	(0.009)		(0.012)	(0.009)	
90th	-0.159	0.168	0.327	-0.067	-0.100	0.008	-0.036	0.112	0.092
				(0.021)	(0.014)		(0.013)	(0.0095)	
						FEMALES			
				1994			2003		
Percentiles	1994	2003	Change	Coefficients	Covariates	Residuals	Coefficients	Covariates	Residuals
10th	0.134	0.183	0.049	0.144	-0.038	0.028	-0.014	0.183	0.014
				(0.025)	(0.017)		(0.0182)	(0.0133)	
25th	-0.069	0.134	0.203	-0.022	-0.043	-0.004	-0.072	0.184	0.022
				(0.018)	(0.011)		(0.0159)	(0.0106)	
50th	-0.117	0.118	0.235	-0.115	-0.042	0.040	-0.077	0.165	0.030
				(0.014)	(0.009)		(0.0112)	(0.0081)	
75th	-0.159	0.133	0.292	-0.104	-0.051	-0.004	-0.067	0.157	0.043
				(0.015)	(0.010)		(0.0126)	(0.0092)	
90th	-0.324	0.090	0.414	-0.214	-0.067	-0.043	-0.081	0.151	0.020
				(0.022)	(0.014)		(0.0169)	(0.0121)	

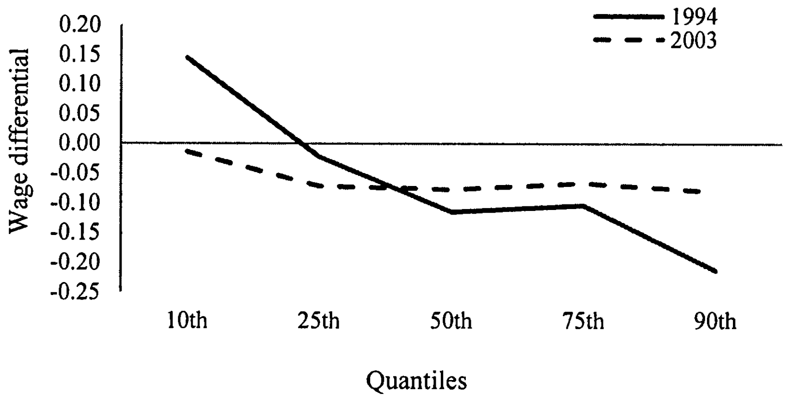
Source: 1994 RLMS and 2003 NOBUS data.

FIGURE 9.15: Estimated public-private wage differentials for males, Russia



Source: Authors' decomposition estimates based on the 1994 and 2003 Russian data.

FIGURE 9.16: Estimated public-private wage differentials for females, Russia



Source: Authors' decomposition estimates based on the 1994 and 2003 Russian data.

The remaining discussion and Table 9.22 identify the source of any change in the estimated public-private wage differential. For the period of 1994 to 2003, changes in the characteristics component actually contributed to an increase in the relative earnings for both male and female workers in the public sector. More specifically, changes in the relative hourly earnings attributable to between sectors changes in characteristics of public and private sector workers accounted for a major part of the overall characteristics effect. The contribution made by changes in the relative worker characteristics was not only larger for men than for women in this period but it also tended to be higher at the top of the distribution. Further investigation of what lay behind changes in the coefficient effect for this period suggest that both changes in the wage advantages/disadvantages of public and private sector and changes in the characteristics of workers at which the advantage/disadvantage was measured, contributed to the overall coefficient effect identified in Figure 9.15 and Figure 9.16.

TABLE 9.22

DHMM decomposition of changes in the public-private wage gap at different points in the conditional earnings distribution, Russia

	10th	20th	30th	40th	50th	60th	70th	80th	90th
1994 to 2003									
Males									
1. $[(\tilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_t^* - \tilde{w}_{t+1}^{cfpriv} \mid \hat{\beta}_t^*) - (\tilde{w}_t^{cfpub} \mid \hat{\beta}_t^* - \tilde{w}_t^{cfpriv} \mid \hat{\beta}_t^*)]$	0.098	0.142	0.172	0.199	0.197	0.205	0.219	0.221	0.271
	(0.034)	(0.025)	(0.023)	(0.021)	(0.020)	(0.018)	(0.018)	(0.020)	(0.030)
2. $[(\tilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_{t+1}^* - \tilde{w}_{t+1}^{cfprv} \mid \hat{\beta}_{t+1}^*) - (\tilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_t^* - \tilde{w}_{t+1}^{cfprv} \mid \hat{\beta}_t^*)]$	0.075	0.039	0.026	0.017	0.012	0.002	-0.009	-0.011	-0.046
	(0.024)	(0.019)	(0.016)	(0.014)	(0.013)	(0.012)	(0.014)	(0.015)	(0.021)
3. $\{(\tilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_t^{pub} - \tilde{w}_t^{pub} \mid \hat{\beta}_t^{pub}) - (\tilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_t^* - \tilde{w}_t^{cfpub} \mid \hat{\beta}_t^*)\} + [(\tilde{w}_{t+1}^{cfprv} \mid \hat{\beta}_t^* - \tilde{w}_t^{cfprv} \mid \hat{\beta}_t^*) - (\tilde{w}_{t+1}^{cfpriv} \mid \hat{\beta}_t^{priv} - \tilde{w}_t^{prv} \mid \hat{\beta}_t^{priv})\}]$	0.007	-0.049	-0.050	-0.060	-0.073	-0.069	-0.041	0.031	0.024
	(0.044)	(0.034)	(0.029)	(0.026)	(0.025)	(0.023)	(0.026)	(0.026)	(0.037)
4. $[\{(\tilde{w}_{t+1}^{cfpriv} \mid \hat{\beta}_{t+1}^* - \tilde{w}_{t+1}^{priv} \mid \hat{\beta}_{t+1}^{priv}) - (\tilde{w}_{t+1}^{cfpriv} \mid \hat{\beta}_t^* - \tilde{w}_{t+1}^{cfpriv} \mid \hat{\beta}_t^{priv})\} + \{\tilde{w}_{t+1}^{pub} \mid \hat{\beta}_{t+1}^{pub} - \tilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_{t+1}^*) - (\tilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_t^{pub} - \tilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_t^*)\}]$	0.020	0.069	0.067	0.082	0.096	0.078	0.031	0.029	0.009
	(0.036)	(0.027)	(0.022)	(0.020)	(0.018)	(0.018)	(0.019)	(0.022)	(0.029)
Change in characteristic component	0.173	0.181	0.198	0.216	0.209	0.207	0.210	0.210	0.225
Change in coefficient component	0.027	0.020	0.017	0.022	0.023	0.009	-0.010	0.060	0.033
Females									
1. $[(\tilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_t^* - \tilde{w}_{t+1}^{cfpriv} \mid \hat{\beta}_t^*) - (\tilde{w}_t^{cfpub} \mid \hat{\beta}_t^* - \tilde{w}_t^{cfpriv} \mid \hat{\beta}_t^*)]$	0.183	0.203	0.202	0.200	0.196	0.198	0.194	0.198	0.247
	(0.028)	(0.022)	(0.020)	(0.018)	(0.019)	(0.018)	(0.019)	(0.021)	(0.027)
2. $[(\tilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_{t+1}^* - \tilde{w}_{t+1}^{cfprv} \mid \hat{\beta}_{t+1}^*) - (\tilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_t^* - \tilde{w}_{t+1}^{cfprv} \mid \hat{\beta}_t^*)]$	0.037	0.031	0.020	0.013	0.011	0.008	0.007	0.008	-0.019
	(0.020)	(0.017)	(0.015)	(0.014)	(0.013)	(0.013)	(0.014)	(0.016)	(0.020)
3. $\{(\tilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_t^{pub} - \tilde{w}_t^{pub} \mid \hat{\beta}_t^{pub}) - (\tilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_t^* - \tilde{w}_t^{cfpub} \mid \hat{\beta}_t^*)\} + [(\tilde{w}_{t+1}^{cfprv} \mid \hat{\beta}_t^* - \tilde{w}_t^{cfprv} \mid \hat{\beta}_t^*) - (\tilde{w}_{t+1}^{cfpriv} \mid \hat{\beta}_t^{priv} - \tilde{w}_t^{prv} \mid \hat{\beta}_t^{priv})\}]$	-0.191	-0.202	-0.168	-0.152	-0.127	-0.091	-0.031	-0.003	-0.026
	(0.039)	(0.030)	(0.026)	(0.023)	(0.024)	(0.024)	(0.026)	(0.026)	(0.037)
4. $[\{(\tilde{w}_{t+1}^{cfpriv} \mid \hat{\beta}_{t+1}^* - \tilde{w}_{t+1}^{priv} \mid \hat{\beta}_{t+1}^{priv}) - (\tilde{w}_{t+1}^{cfpriv} \mid \hat{\beta}_t^* - \tilde{w}_{t+1}^{cfpriv} \mid \hat{\beta}_t^{priv})\} + \{\tilde{w}_{t+1}^{pub} \mid \hat{\beta}_{t+1}^{pub} - \tilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_{t+1}^*) - (\tilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_t^{pub} - \tilde{w}_{t+1}^{cfpub} \mid \hat{\beta}_t^*)\}]$	0.040	0.113	0.161	0.166	0.160	0.129	0.074	0.062	0.136
	(0.030)	(0.025)	(0.021)	(0.019)	(0.018)	(0.019)	(0.020)	(0.023)	(0.030)
Change in characteristic component	0.220	0.234	0.222	0.213	0.207	0.206	0.201	0.206	0.228
Change in coefficient component	-0.151	-0.089	-0.007	0.014	0.033	0.038	0.043	0.059	0.110

Notes: The *component 1* shows the change in relative hourly earnings attributable to between sector changes in characteristics of public and private sectors, evaluated at the baseline wage structure; *component 2* shows the change in the relative hourly earnings explained by changes in the baseline wage structure, holding characteristic differences constant at *t+1* level; *component 3* shows the change in relative hourly earnings produced by within sector changes in characteristics when each group's wage structure is measured relatively to the baseline wage structure; *component 4* shows the effect on the relative hourly earnings differential produced by changes in the wage advantage of public sector workers and the wage disadvantage of private sector workers, holding sector characteristics constant at level *t+1*.

9.5.5. *Sensitivity analysis (occupational effect)*

One cause of disparity may be the different concentration of professionals and technicians between sectors. For instance, both male and female public workers in Bulgaria are more likely to be in a higher and lower managerial positions, whereas the private sector workers tend to be in semi-routine occupations. Six different occupational categories are used in the 1986 Bulgarian decomposition analysis: higher managerial and professional occupations (includes large establishments); lower managerial and professional occupations; intermediate occupations (intermediate clerical and administrative); lower supervisory and technical occupations; semi-routine occupations and routine occupations.

As we expected, including occupational controls reduces significantly the size of the sector wage differential (see Table 9.23). The overall pattern of results, however, when looking at the differential at different points in the conditional distribution remains largely unchanged. Controlling for occupation, public sector male workers at the lower part of the distribution still enjoy a wage premium over their private sector counterparts. At the top of the distribution the differential turns to be negative. Moreover, the differential is less negative at the 90th percentile when we control for occupation, implying that there are differences in individual characteristics and occupations across sectors. It may be argued that transferability of skills between public and private sectors were different between occupations, especially at the top of the distribution.

TABLE 9.23

Decomposition of changes in the public-private sector wage differential for male and female employees in Bulgaria with occupational controls, 1986

1986	MALES				FEMALES			
Percentiles	**Raw gap**	**Coefficients**	**Covariates**	**Residuals**	**Raw gap**	**Coefficients**	**Covariates**	**Residuals**
10th	0.503	0.211	0.231	0.061	0.212	0.038	0.158	0.016
		(0.016)	(0.012)			(0.018)	(0.013)	
25th	0.230	0.106	0.142	-0.018	0.179	0.022	0.128	0.029
		(0.011)	(0.007)			(0.009)	(0.007)	
50th	0.080	0.011	0.050	0.019	0.073	0.005	0.066	0.002
		(0.007)	(0.005)			(0.006)	(0.005)	
75th	-0.282	-0.150	-0.091	-0.041	-0.153	-0.045	-0.082	-0.026
		(0.008)	(0.007)			(0.009)	(0.008)	
90th	-0.151	-0.039	-0.121	0.009	-0.163	-0.017	-0.153	0.007
		(0.006)	(0.005)			(0.006)	(0.005)	

Notes: Bulgaria Household Survey 1986.

9.6. Conclusions

In this Chapter we have examined the differences in the wage distributions between the public and private sector in Bulgaria, Serbia, Russia and Tajikistan. The effect of the endogenous sector selection on the sector wage differential was analysed. The Chapter also examined what has happened to the public-private wage differential over time by providing evidence for Bulgaria and Russia.

Our results support the use of the quantile regression decomposition method, including allowance for possible non-random selection between sectors. Taking the log wage gap at each quantile of the public and private sector distributions, in three of the countries, we found a wage premium for the public sector workers at the lower part of conditional wage distribution and wage penalty at the upper part of the distribution, a result similar to a number of international studies. This means that the public sector successfully attracted better employees in the lower part of the distribution, but was unsuccessful in retaining the most productive employees at the top. The public wage premium at the lower end of the conditional wage distribution might be due to the more effective implementation of equal opportunity and anti-discrimination policies since the government may use the public sector pay to achieve objectives such as equity and to be a 'good' employer (Bender and Elliot, 1997).

Since 2002, the Bulgarian government has tried to force the private sector to increase wages by rising payment in the public sector and increasing the minimum wages. However, we found that in 2003 the public sector workers in Bulgaria still earn more than their private sector counterparts across the conditional earnings distribution. Although the public sector premium in 2003 decreases as we move to the higher quantiles, it remains positive throughout the entire wage distribution. The public sector premium is higher in the lower tail of the wage distribution as compared to the top. Moreover, the sector wage differential for men in 2003 was invariably higher than those found for women. Specifically, the public sector wage advantage faced by men at the 10^{th} percentile was about 20.6% compared to 7.5% for females.

The covariate effect for both males and females in Bulgaria was positive across the whole earnings distribution. The part of the wage gap due to differences in observable individual and job characteristics was substantial. The characteristics effect was much lower in 1986 as compared to 2003, and we concluded that the gain enjoyed by public sector workers over time was due to their increased endowment component. Public sector workers with their individual and job characteristics showed to be more conducive to higher remuneration. These results refute the widely held view by government employees in Bulgaria that they are underpaid compared to the private sector employees.

A negative public sector wage differential throughout the entire conditional earnings distribution was found for males in Serbia and for females in Russia. In Tajikistan, a positive wage premium was found at the lower part of the distribution for both men and women but at the higher quantiles the differential was significant and negative. The advantage enjoyed by private sector workers at the top of the distribution in Tajikistan was higher for women than for men. The findings of negative earnings differential for both males and females, especially in the upper part of the distribution in these three countries has important implications for the quality of public sector workers (since the high skilled workers may switching to the private sector) and the ability of government to retain the services of high quality public sector workers.

A number of important conclusions can be drawn from the Serbian estimates. As highlighted earlier, public sector wage differential for Serbian women was higher than those found for men indicating that the advantage enjoyed by female public sector workers tended to be higher. The estimated differential for females was especially high at the lower deciles, where public sector female employees earned about 15% more compared to their private sector counterparts. At the top of the conditional earnings distribution, however, Serbian women, like men, tended to face a wage penalty from being employed in the public sector. This results indicate that public sector in Serbia is likely to have a problem relating to recruitment.

The study also found that in Bulgaria and Serbia, the probability of being employed in the public sector increased with university and college levels of education. These findings can also be compared to findings of studies of other transition countries. Adamchik and Bedi (2000) find that workers with higher education in Poland are less likely to work in the private sector. By contrast, university education in Russia did not show significant effect on the probability of being employed in the public sector. After allowing for possible selection into sectors, a lower sector wage gap is found for males in Bulgaria and Serbia and for both males and females in Russia and Tajiksitan. This is in line with Barbosa and Filho (2008) who show the lower wage gap (around 16.7%) for Brazil when endogenous sector choice is considered. Similarly Bargian and Melly (2008) find that small wage differential remain between the sectors after taking into account the non-observable heterogeneity. However, the sector wage gap widened for females in Bulgaria, when we control for non-random endogenous sector selection, which suggest that Bulgarian women were positively selected into the public sector. In line with Badel and Pena (2010), and in contrast Albrecht *et al.* (2009), we found that endogenous selection effect in selected transition countries was mostly due to the unobservables, which is the difference between the potential wage distributions of public and private sector workers and the endogeneity adjusted counterfactual distributions. A small part of the selection effect was due to the effect in observables, calculated as the difference between the simulated counterfactual distributions for the sector wage workers using selection adjusted coefficients and being sampled from the whole population of workers.

The present Chapter also answered the question what has been happening to the relative wages of public and private sector workers over time by providing further evidence from Bulgaria and Russia. The Chapter contributed with adaptation of the Donohue and Heckman (1991) time-wise decomposition framework across the entire wage distribution, which identifies the sources of changes in the estimated public-private wage differential. Public sector wage differential has widened over time in Bulgaria and this was observed for both genders. The rise in the public sector earnings tended to be higher at the top of the conditional earnings distribution. The results showed that differences in endowments for Bulgarian workers gained importance over time in accounting for the general increasing trend in the public sector earnings distribution.

In Russia, the public-private wage differential increased over time for males but overall the increase was fairly stable. In contrast, the differential changed substantially over the observed period for females. Further investigation of what lay behind this change suggest that both changes in the wage advantages/disadvantages of public and private sector in Russia and changes in the characteristics of workers at which the advantages/disadvantages was measured, contributed to the overall coefficient effect.

Some of the limitations of the present study are that we have not accounted for all cash and non-pecuniary advantages attached to a particular sector. There may be other factors, such as occupational pension benefits, or payments in kind, that can contribute to explaining the public sector wage differential. Moreover, the study has documented the existence of the public sector wage differential in all these countries, but has explained little about its causes and consequences. While understanding whether it depends on institutional features and what policy implications are is of crucial importance, we take the analysis in this study as a necessary step and leave the other questions for future research.

CHAPTER NINE APPENDIX

TABLE A9.1

Descriptive statistics by public and private sector: males and females, Serbia 2003

Description	*MALES*				*FEMALES*			
	Public		**Private**		**Public**		**Private**	
	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.
log of hourly wage	4.106	0.627	4.154	0.784	4.099	0.639	3.830	0.733
1 if primary	0.126	0.332	0.164	0.371	0.123	0.329	0.182	0.386
1 if vocational (1-2 years)	0.030	0.171	0.023	0.150	0.025	0.155	0.026	0.160
1 if secondary (3 years)	0.272	0.445	0.303	0.460	0.157	0.364	0.182	0.386
1 if secondary (4 years)	0.326	0.469	0.293	0.455	0.360	0.480	0.340	0.474
1 if gymnasium	0.022	0.147	0.027	0.163	0.034	0.182	0.048	0.215
1 if post-secondary	0.084	0.277	0.064	0.245	0.110	0.313	0.053	0.224
1 if university	0.106	0.308	0.049	0.216	0.167	0.373	0.055	0.228
potential exp	25.781	11.220	24.554	14.478	24.354	9.547	23.906	15.202
exp squared	790.348	598.075	812.193	884.358	684.097	446.200	802.072	1004.751
1 if married	0.796	0.404	0.703	0.457	0.744	0.437	0.669	0.471
1 if < 1 year	0.092	0.289	0.342	0.475	0.087	0.282	0.467	0.499
1 if 1-2 years	0.021	0.142	0.038	0.191	0.027	0.161	0.037	0.190
1 if 3-5 years	0.074	0.262	0.087	0.282	0.066	0.249	0.079	0.270
1 if 6-10 years	0.091	0.287	0.104	0.306	0.134	0.341	0.099	0.299
1 if >10 years	0.723	0.448	0.429	0.495	0.686	0.465	0.318	0.466
1 if managerial responsibility	0.021	0.142	0.134	0.341	0.008	0.087	0.083	0.277
1 if in Belgrade	0.140	0.347	0.147	0.354	0.208	0.407	0.171	0.377
1 if in Vojvodina	0.265	0.442	0.258	0.438	0.252	0.435	0.298	0.458
1 if in West Serbia	0.096	0.295	0.142	0.350	0.091	0.288	0.094	0.293
1 if in Šumadija i Pomoravlje	0.209	0.407	0.202	0.402	0.193	0.395	0.200	0.400
1 if East Serbia	0.108	0.311	0.100	0.301	0.093	0.290	0.103	0.304
1 if South-East Serbia	0.182	0.386	0.151	0.358	0.163	0.370	0.134	0.341
N	729		737		528		456	

Source: Serbia LSMS, 2003.

TABLE A9.2

Descriptive statistics by public and private sector: males and females workers, Bulgaria, 1986 and 2003

	1986								2003							
	Males				*Females*				*Males*				*Females*			
	Public		**Private**		**Public**		**Private**		**Public**		**Private**		**Public**		**Private**	
Variable	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.
lhwage	0.971	0.404	0.889	0.647	0.946	0.433	0.911	0.563	1.490	0.535	1.097	0.525	1.235	0.442	0.941	0.532
1 if primary	0.272	0.445	0.441	0.497	0.248	0.432	0.402	0.491	0.106	0.308	0.175	0.380	0.071	0.258	0.127	0.333
1 if secondary	0.540	0.498	0.349	0.477	0.521	0.500	0.356	0.479	0.547	0.498	0.652	0.477	0.407	0.492	0.612	0.488
1 if university	0.110	0.313	0.048	0.213	0.109	0.311	0.029	0.168	0.282	0.451	0.117	0.321	0.357	0.480	0.192	0.394
exp	30.539	13.889	28.118	17.331	31.884	13.061	29.083	22.460	22.584	11.253	18.819	12.519	21.348	9.918	16.932	11.597
expsq	1125.481	749.607	1090.507	820.968	1187.098	694.431	1349.462	959.952	636.363	563.859	510.683	551.686	553.920	423.245	421.015	470.419
1 if<1 years	0.035	0.184	0.023	0.150	0.042	0.201	0.026	0.158	0.098	0.298	0.346	0.476	0.071	0.258	0.325	0.469
1 if 1-2 years	0.031	0.175	0.021	0.145	0.028	0.166	0.020	0.142	0.083	0.276	0.219	0.414	0.078	0.268	0.217	0.412
1 if 3-5 years	0.102	0.303	0.064	0.245	0.095	0.294	0.066	0.249	0.169	0.375	0.200	0.400	0.165	0.371	0.232	0.422
1 if 6-10 years	0.159	0.366	0.153	0.360	0.170	0.376	0.167	0.373	0.174	0.379	0.102	0.303	0.177	0.382	0.120	0.325
1 if>10 years	0.672	0.470	0.737	0.441	0.664	0.473	0.721	0.449	0.476	0.500	0.132	0.339	0.509	0.500	0.106	0.309
1 if managers	0.256	0.436	0.132	0.338	0.173	0.378	0.082	0.274	0.086	0.280	0.022	0.148	0.043	0.204	0.015	0.122
1 if married	0.814	0.389	0.840	0.366	0.850	0.357	0.886	0.318	0.793	0.405	0.630	0.483	0.773	0.420	0.700	0.458
1 if bulgarian	0.875	0.331	0.753	0.431	0.899	0.302	0.719	0.450	0.937	0.243	0.885	0.319	0.939	0.239	0.869	0.338
1 if urban	0.586	0.493	0.266	0.442	0.595	0.491	0.279	0.449	0.824	0.382	0.741	0.438	0.857	0.350	0.776	0.417
1 if Sofia	0.129	0.335	0.030	0.170	0.162	0.369	0.037	0.190	0.035	0.185	0.047	0.211	0.058	0.235	0.032	0.176
N	*2604*		*608*		*2650*		*587*		*397*		*899*		*462*		*724*	

Source: Bulgaria 1986 and 2003.

TABLE A9.3

Descriptive statistics by public and private sector: males and females in Russia, 1994 and 2003

	1994								2003							
	MALES				FEMALES				MALES				FEMALES			
	Public		Private		Public		Private		Public		Private		Public		Private	
Variable	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.
lhwage	7.237	1.029	7.325	1.07	6.998	0.904	7.116	0.991	2.911	0.768	2.789	0.868	2.637	0.711	2.532	0.766
1 if university	0.279	0.449	0.265	0.442	0.325	0.469	0.275	0.447	0.212	0.409	0.139	0.346	0.275	0.446	0.166	0.372
1 if college	0.250	0.433	0.262	0.440	0.464	0.499	0.486	0.501	0.032	0.176	0.035	0.184	0.038	0.191	0.040	0.197
1 if secondary	0.218	0.413	0.286	0.453	0.122	0.328	0.165	0.372	0.452	0.462	0.466	0.460	0.467	0.485	0.500	0.484
1 if primary	0.151	0.358	0.129	0.336	0.082	0.274	0.071	0.257	0.007	0.085	0.012	0.108	0.006	0.077	0.006	0.075
exp	23.928	12.375	20.105	10.951	22.42	10.826	21.286	10.459	22.258	11.396	20.599	11.491	22.164	10.854	19.955	11.179
expsq	725.511	683.648	523.742	537.124	619.736	541.018	562.063	483.321	625.275	524.232	556.359	510.006	609.06	489.398	523.153	468.153
1 if managers	0.032	0.176	0.252	0.435	0.032	0.177	0.180	0.385	0.340	0.474	0.218	0.413	0.527	0.499	0.246	0.431
1 if < 1 year	0.163	0.369	0.255	0.437	0.129	0.336	0.231	0.423	0.109	0.312	0.209	0.407	0.084	0.278	0.201	0.401
1 if 1-2 years	0.200	0.401	0.292	0.456	0.187	0.390	0.247	0.432	0.172	0.377	0.245	0.430	0.158	0.365	0.267	0.442
1 if 3-5 years	0.136	0.343	0.163	0.370	0.097	0.296	0.129	0.336	0.133	0.340	0.137	0.344	0.115	0.318	0.133	0.340
1 if 6-10 years	0.108	0.311	0.086	0.281	0.152	0.360	0.106	0.308	0.195	0.397	0.137	0.344	0.186	0.389	0.130	0.336
1 if >10 years	0.393	0.489	0.203	0.403	0.435	0.496	0.286	0.453	0.390	0.488	0.272	0.445	0.458	0.498	0.270	0.444
1 if arrears	0.343	0.475	0.280	0.450	0.28	0.449	0.224	0.417	0.182	0.386	0.199	0.399	0.137	0.343	0.143	0.350
1 if Central	0.204	0.403	0.262	0.440	0.188	0.391	0.239	0.427	0.204	0.403	0.229	0.420	0.216	0.412	0.234	0.424
1 if North-West	0.075	0.264	0.068	0.252	0.089	0.285	0.075	0.263	0.143	0.350	0.133	0.340	0.141	0.348	0.137	0.344
1 if Siberia	0.045	0.207	0.052	0.223	0.041	0.199	0.067	0.250	0.124	0.330	0.141	0.348	0.131	0.337	0.131	0.338
1 if Far-East	0.068	0.253	0.062	0.241	0.057	0.231	0.067	0.250	0.155	0.362	0.103	0.304	0.144	0.352	0.100	0.300
1 if Urals	0.133	0.340	0.117	0.322	0.128	0.334	0.122	0.327	0.075	0.263	0.094	0.291	0.077	0.266	0.090	0.286
1 if Volga	0.098	0.297	0.086	0.281	0.124	0.330	0.094	0.293	0.165	0.371	0.176	0.380	0.164	0.371	0.181	0.385
N	*848*		*325*		*991*		*255*		13020		8854		16880		7438	

Source: Russia 1994 LSMS and 2003 NOBUS data.
Notes: The omitted categories are South region, tenure less than 1 years and primary school and less.

TABLE A9.4

Descriptive statistics by public and private sector: males and females workers in Tajikistan, 2003

Variables	*MALES*				*FEMALES*			
	Public		**Private**		**Public**		**Private**	
	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.
log of hourly wage	3.512	0.881	3.593	1.130	2.897	0.813	2.922	1.041
1 if university	0.348	0.476	0.098	0.297	0.239	0.427	0.045	0.208
1 if secondary technical	0.091	0.288	0.076	0.265	0.044	0.206	0.019	0.135
1 if secondary special	0.100	0.300	0.093	0.291	0.174	0.379	0.043	0.203
1 if secondary general	0.368	0.482	0.595	0.491	0.431	0.496	0.665	0.472
1 if primary	0.067	0.251	0.109	0.312	0.075	0.264	0.172	0.377
1 if basic	0.015	0.122	0.024	0.154	0.028	0.166	0.051	0.220
potential exp	21.513	11.898	17.649	12.096	19.135	10.914	17.915	12.054
exp squared	604.238	599.372	457.715	563.004	485.145	467.517	466.095	542.821
1 if married	0.880	0.325	0.758	0.429	0.627	0.484	0.638	0.481
1 if < 1 year	0.120	0.325	0.264	0.441	0.123	0.329	0.223	0.416
1 if 1-2 years	0.117	0.322	0.188	0.391	0.138	0.345	0.178	0.383
1 if 3-5 years	0.188	0.390	0.219	0.414	0.174	0.379	0.220	0.414
1 if 6-10 years	0.163	0.370	0.129	0.336	0.147	0.354	0.130	0.337
1 if >10 years	0.412	0.492	0.199	0.399	0.419	0.494	0.249	0.433
1 if urban	0.390	0.488	0.291	0.454	0.440	0.497	0.262	0.440
1 if Dushanbe	0.130	0.337	0.047	0.211	0.180	0.384	0.051	0.220
1 if Sogd	0.267	0.443	0.323	0.468	0.277	0.448	0.258	0.438
1 if Khatlon	0.288	0.453	0.438	0.496	0.254	0.435	0.508	0.500
1 if Rrp	0.186	0.389	0.081	0.273	0.172	0.378	0.091	0.288
1 if Gbao	0.128	0.334	0.112	0.316	0.117	0.322	0.092	0.289
N	1397		1609		812		1020	

Source: Tajikistan TLSS, 2003.

TABLE A9.5: QR estimates for males and females – public sector in Bulgaria, 2003

	Males					Females				
	θ=0.10	θ=0.25	θ=0.50	θ=0.75	θ=0.90	θ=0.10	θ=0.25	θ=0.50	θ=0.75	θ=0.90
public	0.1367***	0.1720***	0.1762***	0.1915***	0.2218***	0.1801***	0.0778***	0.0250	-0.0171	-0.0553
	(0.0363)	(0.0335)	(0.0472)	(0.0530)	(0.0500)	(0.0462)	(0.0285)	(0.0385)	(0.0427)	(0.0483)
university	0.5513***	0.4971***	0.5033***	0.4833***	0.5579***	0.6580***	0.4863***	0.6273***	0.6888***	0.5126***
	(0.0548)	(0.0530)	(0.0758)	(0.0852)	(0.0862)	(0.0733)	(0.0471)	(0.0625)	(0.0719)	(0.0892)
college	0.0957	0.1943**	0.1158	0.0460	0.1376	0.4581***	0.2103***	0.2558***	0.2071**	0.0618
	(0.0751)	(0.0879)	(0.1254)	(0.1389)	(0.1289)	(0.0887)	(0.0578)	(0.0753)	(0.0855)	(0.1052)
secondary	0.1841***	0.1378***	0.1582***	0.1382**	0.1857***	0.3638***	0.1074***	0.1889***	0.2218***	0.0836
	(0.0423)	(0.0397)	(0.0562)	(0.0634)	(0.0639)	(0.0637)	(0.0397)	(0.0521)	(0.0581)	(0.0701)
exp	0.0111**	0.0010	0.0058	0.0014	-0.0009	0.0249***	0.0192***	0.0204***	0.0085	0.0094
	(0.0047)	(0.0044)	(0.0058)	(0.0063)	(0.0060)	(0.0060)	(0.0038)	(0.0051)	(0.0054)	(0.0064)
expsq	-0.0002**	-0.0001	-0.0002*	-0.0002	-0.0001	-0.0006***	-0.0004***	-0.0004***	-0.0002	-0.0002
	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)	(0.0001)
1-2 years	0.0897**	0.1506***	0.1424**	0.0587	-0.0252	-0.0100	0.0347	0.0672	0.0994*	0.0274
	(0.0436)	(0.0419)	(0.0592)	(0.0662)	(0.0626)	(0.0607)	(0.0385)	(0.0513)	(0.0579)	(0.0678)
3-5 years	0.1028**	0.1587***	0.2190***	0.1941***	0.0618	0.1449**	0.1500***	0.1670***	0.2294***	0.1844***
	(0.0437)	(0.0415)	(0.0593)	(0.0666)	(0.0617)	(0.0598)	(0.0375)	(0.0494)	(0.0553)	(0.0671)
6-10 years	0.2043***	0.2448***	0.3641***	0.3589***	0.2120***	0.0716	0.2085***	0.2184***	0.2470***	0.2571***
	(0.0510)	(0.0485)	(0.0695)	(0.0771)	(0.0731)	(0.0690)	(0.0431)	(0.0567)	(0.0639)	(0.0757)
>10 years	0.2209***	0.2722***	0.3938***	0.3791***	0.2404***	0.1513**	0.2468***	0.2384***	0.4082***	0.3784***
	(0.0475)	(0.0446)	(0.0623)	(0.0696)	(0.0647)	(0.0647)	(0.0408)	(0.0545)	(0.0609)	(0.0684)
married	-0.0683*	0.0220	0.0249	0.0469	0.0600	-0.0215	-0.0158	-0.0542	-0.0656	-0.0335
	(0.0354)	(0.0336)	(0.0464)	(0.0499)	(0.0477)	(0.0459)	(0.0281)	(0.0372)	(0.0400)	(0.0483)
menagers	0.3138***	0.2710***	0.2205**	0.3352***	0.2882***	0.2141*	0.2127***	0.1855*	0.2336**	0.2414*
	(0.0738)	(0.0696)	(0.1003)	(0.1106)	(0.1021)	(0.1188)	(0.0728)	(0.0973)	(0.1096)	(0.1408)
urban	0.0760**	0.0815**	0.1855***	0.2544***	0.2421***	0.0252	0.0945***	0.0652	0.0972**	0.1656***
	(0.0357)	(0.0342)	(0.0484)	(0.0541)	(0.0511)	(0.0517)	(0.0322)	(0.0419)	(0.0469)	(0.0547)
sofia	0.0862	0.0990	0.1392	0.2507**	0.4778***	-0.0350	0.1205**	0.1185	0.0942	0.1606*
	(0.0679)	(0.0650)	(0.0948)	(0.1063)	(0.0940)	(0.0888)	(0.0568)	(0.0781)	(0.0839)	(0.0952)
constant	0.5933***	0.8494***	0.9841***	1.3092***	1.6190***	0.2124***	0.6086***	0.7967***	1.0776***	1.4461***
	(0.0568)	(0.0538)	(0.0771)	(0.0876)	(0.0838)	(0.0757)	(0.0503)	(0.0682)	(0.0769)	(0.0866)
N	1296	1296	1296	1296	1296	1186	1186	1186	1186	1186
Pseudo R^2	0.1381	0.1775	0.1959	0.2139	0.2047	0.1584	0.1787	0.2033	0.1946	0.1967

Standard errors in parentheses * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

TABLE A9.6: QR estimates for males– public sector in Serbia, 2003

	Males				
	θ=0.10	θ=0.25	θ=0.50	θ=0.75	θ=0.90
public	0.0121	-0.0213	-0.0845***	-0.2254***	-0.3478***
	(0.0622)	(0.0326)	(0.0344)	(0.0515)	(0.0853)
university	0.7751***	0.6451***	0.4191*	0.8599***	0.8288***
	(0.1592)	(0.1873)	(0.2221)	(0.2984)	(0.2203)
post-secondary	0.6882***	0.7404***	0.6859***	0.6788***	0.7784***
	(0.1274)	(0.0692)	(0.0741)	(0.1119)	(0.1796)
gymnasium	0.4846***	0.5571***	0.4049***	0.4080***	0.3073*
	(0.1298)	(0.0699)	(0.0736)	(0.1122)	(0.1835)
secondary (4 years)	0.0666	0.3050***	0.3648***	0.3123**	0.5865**
	(0.1705)	(0.1075)	(0.1118)	(0.1572)	(0.2623)
secondary (1-2 years)	0.2209**	0.3156***	0.2215***	0.2163***	0.2090
	(0.0944)	(0.0499)	(0.0528)	(0.0778)	(0.1277)
primary	-0.0867	0.0847	0.0328	0.0104	0.0731
	(0.1773)	(0.0983)	(0.1058)	(0.1576)	(0.2464)
exp	0.0081	0.0138***	0.0091	0.0115	0.0081
	(0.0098)	(0.0053)	(0.0056)	(0.0087)	(0.0153)
expsq	-0.0003**	-0.0003***	-0.0002*	-0.0002	-0.0001
	(0.0002)	(0.0001)	(0.0001)	(0.0001)	(0.0002)
1-2 years of tenure	0.1342	0.0583	0.0108	-0.1239	0.0094
	(0.1839)	(0.0947)	(0.1010)	(0.1497)	(0.2499)
3-5 years of tenure	0.1146	0.0769	-0.0146	-0.0993	-0.1684
	(0.1179)	(0.0642)	(0.0681)	(0.1029)	(0.1643)
6-10 years of tenure	0.1235	0.0151	0.0247	-0.0659	-0.0938
	(0.1147)	(0.0620)	(0.0652)	(0.0980)	(0.1498)
>10 years of tenure	0.1232	0.0381	0.0312	-0.0515	0.0201
	(0.0909)	(0.0493)	(0.0524)	(0.0792)	(0.1390)
married	0.0919	0.0890**	0.0432	0.0037	0.0471
	(0.0794)	(0.0422)	(0.0446)	(0.0667)	(0.1079)
urban	0.0689	0.0819**	0.0116	0.0443	-0.1251
	(0.0616)	(0.0335)	(0.0354)	(0.0535)	(0.0881)
Belgrade	0.5476***	0.4425***	0.3216***	0.3130***	0.4179***
	(0.1099)	(0.0561)	(0.0594)	(0.0894)	(0.1463)
Vojvodina	0.2711***	0.1648***	0.0742	0.0317	0.1090
	(0.0916)	(0.0483)	(0.0507)	(0.0765)	(0.1244)
West Serbia	0.2588**	0.0925	0.0236	0.0934	0.1397
	(0.1124)	(0.0579)	(0.0613)	(0.0929)	(0.1479)
Šumadija and Pomoravlje	0.3088***	0.1618***	0.1885***	0.1617**	0.2464*
	(0.0950)	(0.0501)	(0.0529)	(0.0792)	(0.1312)
East Serbia	0.1641	0.0179	0.0443	0.0017	-0.0821
	(0.1124)	(0.0599)	(0.0638)	(0.0955)	(0.1526)
constant	2.8091***	3.1251***	3.6899***	4.1749***	4.6327***
	(0.1517)	(0.0831)	(0.0870)	(0.1268)	(0.2072)
N	1466	1466	1466	1466	1466

Notes: Standard errors in parentheses * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$; The omitted categories- basic education, tenure less than 1 year and South East Serbia.

TABLE A9.7: QR estimates for females – public sector in Serbia, 2003

	Females				
	θ=0.10	θ=0.25	θ=0.50	θ=0.75	θ=0.90
public	0.1785*	0.1850***	0.0874***	-0.0486	-0.1462
	(0.0923)	(0.0494)	(0.0328)	(0.0515)	(0.0930)
university	1.1518***	1.1522***	1.3055***	1.4300***	1.9437***
	(0.2197)	(0.3013)	(0.1887)	(0.2895)	(0.2399)
post-secondary	0.7911***	0.6314***	0.8395***	0.9548***	0.9862***
	(0.1629)	(0.0943)	(0.0614)	(0.0966)	(0.1850)
gymnasium	0.7340***	0.5566***	0.5849***	0.5070***	0.3889**
	(0.1827)	(0.1033)	(0.0656)	(0.1008)	(0.1951)
secondary (4 years)	0.5201**	0.2082	0.3037***	0.3156**	0.2286
	(0.2374)	(0.1271)	(0.0827)	(0.1252)	(0.2293)
secondary (1-2 years)	0.3606***	0.2386***	0.3084***	0.2856***	0.2748*
	(0.1355)	(0.0744)	(0.0475)	(0.0728)	(0.1410)
primary	0.1381	-0.2087	-0.0156	0.0500	0.0910
	(0.2668)	(0.1451)	(0.0974)	(0.1516)	(0.2836)
exp	0.0289**	0.0273***	0.0230***	0.0162**	0.0078
	(0.0129)	(0.0071)	(0.0046)	(0.0073)	(0.0129)
expsq	-0.0006***	-0.0005***	-0.0003***	-0.0002	-0.0000
	(0.0002)	(0.0001)	(0.0001)	(0.0001)	(0.0002)
1-2 years of tenure	0.1538	0.1530	0.1072	0.0855	-0.0102
	(0.2449)	(0.1346)	(0.0883)	(0.1344)	(0.2307)
3-5 years of tenure	0.2112	0.1695*	0.1467**	0.0807	0.1019
	(0.1712)	(0.0989)	(0.0628)	(0.0964)	(0.1844)
6-10 years of tenure	0.0931	0.0691	0.1399**	0.2148***	0.2778*
	(0.1552)	(0.0866)	(0.0545)	(0.0820)	(0.1596)
>10 years of tenure	-0.0768	0.0202	0.0507	0.0797	0.1065
	(0.1257)	(0.0706)	(0.0459)	(0.0701)	(0.1259)
married	0.0478	-0.0374	-0.0378	-0.0322	-0.1006
	(0.0910)	(0.0539)	(0.0349)	(0.0540)	(0.1028)
urban	0.1480	0.1075***	0.0832**	0.0183	-0.0282
	(0.0975)	(0.0541)	(0.0340)	(0.0524)	(0.1028)
Belgrade	0.3329**	0.2344***	0.2159***	0.1690**	0.1928
	(0.1421)	(0.0803)	(0.0516)	(0.0796)	(0.1505)
Vojvodina	0.2344*	0.1821**	0.1172**	0.1384*	0.0820
	(0.1325)	(0.0738)	(0.0471)	(0.0715)	(0.1346)
West Serbia	0.0573	-0.0540	-0.0402	-0.1351	-0.2319
	(0.1685)	(0.0945)	(0.0608)	(0.0919)	(0.1849)
Šumadija and Pomoravlje	0.1925	0.1112	0.0993**	0.0311	-0.0205
	(0.1414)	(0.0789)	(0.0503)	(0.0766)	(0.1477)
East Serbia	-0.1954	0.0910	0.0412	-0.0268	-0.1091
	(0.1679)	(0.0937)	(0.0601)	(0.0926)	(0.1758)
constant	2.3287***	2.8773***	3.1328***	3.6508***	4.1832***
	(0.2159)	(0.1204)	(0.0782)	(0.1207)	(0.2555)
N	984	984	984	984	984

Notes: Standard errors in parentheses * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$; The omitted categories- basic education, tenure less than 1 year and South East Serbia.

TABLE A9.8

QR estimates for males – public sector as a dummy variable, 2003 Russia

	Males				
	θ=0.10	θ=0.25	θ=0.50	θ=0.75	θ=0.90
public	0.0772***	0.0310**	-0.0119	-0.0451***	-0.0791***
	(0.0205)	(0.0123)	(0.0106)	(0.0119)	(0.0147)
university	0.5541***	0.4666***	0.3793***	0.3588***	0.3312***
	(0.0315)	(0.0192)	(0.0167)	(0.0187)	(0.0230)
college	0.4233***	0.3975***	0.3202***	0.2941***	0.2827***
	(0.0562)	(0.0342)	(0.0294)	(0.0327)	(0.0399)
secondary	0.2947***	0.2441***	0.1793***	0.1558***	0.1116***
	(0.0243)	(0.0148)	(0.0127)	(0.0143)	(0.0176)
primary	0.1426***	0.1421***	0.0858***	0.0552***	0.0161
	(0.0341)	(0.0207)	(0.0177)	(0.0198)	(0.0243)
exp	-0.0025	0.0008	0.0049***	0.0094***	0.0106***
	(0.0036)	(0.0022)	(0.0019)	(0.0021)	(0.0026)
expsq	-0.0001	-0.0001***	-0.0002***	-0.0003***	-0.0003***
	(0.0001)	(0.0000)	(0.0000)	(0.0000)	(0.0001)
managerial	0.2394***	0.1837***	0.1379***	0.1172***	0.1289***
	(0.0239)	(0.0147)	(0.0128)	(0.0145)	(0.0178)
married	0.0862***	0.1292***	0.1320***	0.1209***	0.1293***
	(0.0264)	(0.0159)	(0.0136)	(0.0153)	(0.0184)
1-2 years of tenure	0.1126***	0.1538***	0.1404***	0.1137***	0.0826***
	(0.0330)	(0.0201)	(0.0172)	(0.0192)	(0.0236)
3-5 years of tenure	0.2742***	0.2747***	0.2382***	0.1939***	0.1450***
	(0.0367)	(0.0223)	(0.0191)	(0.0214)	(0.0263)
6-10 years of tenure	0.3470***	0.3424***	0.2903***	0.2404***	0.2020***
	(0.0356)	(0.0215)	(0.0185)	(0.0207)	(0.0255)
>10 years of tenure	0.2756***	0.3237***	0.3238***	0.2838***	0.2360***
	(0.0333)	(0.0202)	(0.0172)	(0.0193)	(0.0238)
arrears	-0.6139***	-0.4895***	-0.3556***	-0.2939***	-0.2878***
	(0.0245)	(0.0150)	(0.0131)	(0.0146)	(0.0180)
Central region	0.2950***	0.2723***	0.2365***	0.2425***	0.2697***
	(0.0342)	(0.0207)	(0.0177)	(0.0198)	(0.0243)
North-West	0.6324***	0.6616***	0.6185***	0.6185***	0.6139***
	(0.0375)	(0.0227)	(0.0194)	(0.0217)	(0.0266)
Siberia	0.2861***	0.3376***	0.3606***	0.3742***	0.4274***
	(0.0379)	(0.0230)	(0.0197)	(0.0220)	(0.0270)
Far-East	0.7709***	0.7702***	0.7490***	0.7379***	0.7429***
	(0.0377)	(0.0229)	(0.0196)	(0.0218)	(0.0268)
Urals	0.4793***	0.4808***	0.5247***	0.6139***	0.8012***
	(0.0433)	(0.0262)	(0.0224)	(0.0250)	(0.0307)
Volga	0.1397***	0.1811***	0.1920***	0.1811***	0.1218***
	(0.0357)	(0.0217)	(0.0186)	(0.0208)	(0.0255)
constant	1.2077***	1.6390***	2.1356***	2.5448***	2.9461***
	(0.0481)	(0.0292)	(0.0250)	(0.0278)	(0.0337)
N	21874	21874	21874	21874	21874
Pseudo R^2	0.1541	0.1430	0.1389	0.1331	0.1339

Notes: Standard errors in parentheses * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$; The omitted categories- basic education, tenure less than 1 year and South region.

TABLE A9.9

QR estimates for females – public sector as a dummy variable, 2003 Russia

	Females				
	θ=0.10	θ=0.25	θ=0.50	θ=0.75	θ=0.90
public	-0.0113	-0.0932***	-0.1050***	-0.0785***	-0.0953***
	(0.0161)	(0.0117)	(0.0098)	(0.0129)	(0.0143)
university	0.5505***	0.5496***	0.5055***	0.4387***	0.4158***
	(0.0227)	(0.0165)	(0.0137)	(0.0182)	(0.0200)
college	0.4333***	0.4159***	0.4169***	0.3612***	0.3511***
	(0.0394)	(0.0287)	(0.0239)	(0.0314)	(0.0341)
secondary	0.2103***	0.2046***	0.1660***	0.1125***	0.0914***
	(0.0187)	(0.0136)	(0.0113)	(0.0150)	(0.0165)
primary	0.0792***	0.0942***	0.1044***	0.0738***	0.0654**
	(0.0301)	(0.0220)	(0.0182)	(0.0239)	(0.0260)
exp	0.0087***	0.0125***	0.0121***	0.0135***	0.0153***
	(0.0026)	(0.0019)	(0.0016)	(0.0020)	(0.0022)
expsq	-0.0002***	-0.0002***	-0.0002***	-0.0003***	-0.0003***
	(0.0001)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
managerial	0.2632***	0.2303***	0.1691***	0.1411***	0.1509***
	(0.0166)	(0.0122)	(0.0102)	(0.0137)	(0.0152)
married	-0.0394***	-0.0479***	-0.0433***	-0.0165	-0.0164
	(0.0150)	(0.0109)	(0.0090)	(0.0119)	(0.0130)
1-2 years of tenure	0.0740***	0.0854***	0.0867***	0.0677***	0.0623***
	(0.0258)	(0.0189)	(0.0156)	(0.0205)	(0.0224)
3-5 years of tenure	0.1663***	0.1455***	0.1453***	0.1336***	0.1079***
	(0.0289)	(0.0212)	(0.0176)	(0.0230)	(0.0251)
6-10 years of tenure	0.2319***	0.2109***	0.2181***	0.1844***	0.1509***
	(0.0278)	(0.0203)	(0.0168)	(0.0221)	(0.0240)
>10 years of tenure	0.2581***	0.2440***	0.2390***	0.1972***	0.1599***
	(0.0260)	(0.0190)	(0.0158)	(0.0209)	(0.0230)
arrears	-0.2556***	-0.2172***	-0.2142***	-0.2126***	-0.2164***
	(0.0200)	(0.0148)	(0.0124)	(0.0163)	(0.0178)
Central region	0.1400***	0.1520***	0.1786***	0.2274***	0.2836***
	(0.0247)	(0.0180)	(0.0149)	(0.0196)	(0.0213)
North-West	0.5008***	0.5544***	0.5859***	0.5884***	0.6070***
	(0.0272)	(0.0199)	(0.0164)	(0.0216)	(0.0234)
Siberia	0.2573***	0.2998***	0.3134***	0.3508***	0.4312***
	(0.0275)	(0.0201)	(0.0167)	(0.0219)	(0.0238)
Far-East	0.7035***	0.7404***	0.7694***	0.8015***	0.8314***
	(0.0276)	(0.0202)	(0.0167)	(0.0219)	(0.0237)
Urals	0.2518***	0.3181***	0.4097***	0.5402***	0.6810***
	(0.0315)	(0.0230)	(0.0191)	(0.0250)	(0.0272)
Volga	0.1019***	0.1171***	0.1223***	0.1449***	0.1503***
	(0.0260)	(0.0190)	(0.0158)	(0.0207)	(0.0225)
constant	1.0344***	1.4114***	1.8585***	2.2389***	2.5662***
	(0.0357)	(0.0266)	(0.0222)	(0.0294)	(0.0318)
N	24318	24318	24318	24318	24318
Pseudo R^2	0.1471	0.1539	0.1535	0.1488	0.1438

Notes: Standard errors in parentheses * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$; The omitted categories- basic education, tenure less than 1 year and South region.

TABLE A9.10

QR estimates for males – public sector in Tajikistan, 2003

	Males				
	θ=0.10	θ=0.25	θ=0.50	θ=0.75	θ=0.90
public	-0.0132	-0.0869*	-0.1991***	-0.3818***	-0.4338***
	(0.0683)	(0.0498)	(0.0426)	(0.0530)	(0.0695)
university	0.2844**	0.3565***	0.2856***	0.3872***	0.2891**
	(0.1293)	(0.0924)	(0.0790)	(0.0986)	(0.1287)
secondary technical	0.1540	0.1113	0.2032**	0.3508***	0.3857**
	(0.1518)	(0.1096)	(0.0933)	(0.1156)	(0.1505)
Secondary special	-0.0685	0.0824	0.0382	0.1452	0.0979
	(0.1446)	(0.1060)	(0.0895)	(0.1117)	(0.1461)
secondary general	-0.0373	0.0243	0.0012	0.1105	0.0914
	(0.1125)	(0.0816)	(0.0694)	(0.0858)	(0.1115)
primary	0.0823	0.0061	-0.1532	-0.3042*	-0.4963**
	(0.2492)	(0.1811)	(0.1501)	(0.1820)	(0.2291)
exp	0.0477***	0.0388***	0.0389***	0.0341***	0.0321***
	(0.0107)	(0.0076)	(0.0065)	(0.0081)	(0.0110)
expsq	-0.0009***	-0.0007***	-0.0007***	-0.0006***	-0.0006***
	(0.0002)	(0.0001)	(0.0001)	(0.0002)	(0.0002)
1-2 years of tenure	-0.0431	-0.0538	0.0012	0.1093	0.2758*
	(0.1429)	(0.1027)	(0.0874)	(0.1079)	(0.1427)
3-5 years of tenure	-0.2013	-0.0538	-0.1263	-0.0529	-0.0685
	(0.1341)	(0.0965)	(0.0819)	(0.1008)	(0.1337)
6-10 years of tenure	-0.1230	-0.0653	-0.1437*	-0.1377	-0.0563
	(0.1299)	(0.0925)	(0.0791)	(0.0980)	(0.1295)
>10 years of tenure	-0.2760**	-0.3334***	-0.4763***	-0.4807***	-0.3947***
	(0.1365)	(0.0963)	(0.0816)	(0.0999)	(0.1292)
married	0.0493	0.0939	0.0991	0.0638	-0.0126
	(0.1117)	(0.0781)	(0.0648)	(0.0810)	(0.1061)
urban	0.5066***	0.4502***	0.5215***	0.5363***	0.5098***
	(0.0813)	(0.0586)	(0.0501)	(0.0621)	(0.0791)
Dushanbe	-0.1255	-0.0072	-0.0375	0.0915	-0.1117
	(0.1235)	(0.0897)	(0.0767)	(0.0951)	(0.1247)
Khatlon	-0.1785	-0.1173	-0.2332***	-0.2177**	-0.4393***
	(0.1196)	(0.0881)	(0.0762)	(0.0968)	(0.1273)
Rrg region	0.3083**	0.2738**	0.1638*	0.2304**	-0.0013
	(0.1485)	(0.1077)	(0.0905)	(0.1101)	(0.1460)
Gbao region	0.0993	0.3410***	0.3591***	0.4503***	0.2978**
	(0.1417)	(0.1031)	(0.0880)	(0.1090)	(0.1449)
constant	1.9347***	2.4494***	3.1894***	3.7238***	4.6155***
	(0.1977)	(0.1458)	(0.1210)	(0.1507)	(0.1968)
N	3006	3006	3006	3006	3006
Pseudo R^2	0.1005	0.1078	0.1236	0.1271	0.1078

Notes: Standard errors in parentheses * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$; The omitted categories- basic education, tenure less than 1 year and Sogd region.

TABLE A9.11

QR estimates for females – public sector in Tajikistan, 2003

	Females				
	θ=0.10	θ=0.25	θ=0.50	θ=0.75	θ=0.90
public	-0.0625*	-0.1562***	-0.2424***	-0.2385***	-0.3144***
	(0.0372)	(0.0536)	(0.0578)	(0.0616)	(0.0970)
university	0.5280***	0.6678***	0.7903***	0.5911***	0.6515***
	(0.0771)	(0.1090)	(0.1113)	(0.1144)	(0.1813)
secondary technical	0.1995*	0.4660***	0.7111***	0.5876***	0.3939*
	(0.1168)	(0.1642)	(0.1666)	(0.1633)	(0.2332)
Secondary special	0.0377	0.1500	0.2655**	0.2545**	0.3181*
	(0.0786)	(0.1142)	(0.1157)	(0.1169)	(0.1832)
secondary general	0.0220	0.0435	0.1657**	0.1761**	0.2220*
	(0.0567)	(0.0799)	(0.0799)	(0.0796)	(0.1218)
primary	-0.2774***	-0.1930	-0.1316	0.0484	-0.0140
	(0.0985)	(0.1397)	(0.1437)	(0.1383)	(0.2035)
exp	0.0128**	0.0166**	0.0281***	0.0320***	0.0232**
	(0.0058)	(0.0080)	(0.0079)	(0.0079)	(0.0118)
expsq	-0.0002	-0.0002	-0.0004**	-0.0004**	-0.0002
	(0.0001)	(0.0002)	(0.0002)	(0.0002)	(0.0002)
1-2 years of tenure	0.0126	0.0914	-0.1269	-0.0064	0.0664
	(0.0869)	(0.1231)	(0.1260)	(0.1258)	(0.1889)
3-5 years of tenure	0.0711	0.1994*	0.0213	0.0656	0.0489
	(0.0812)	(0.1158)	(0.1176)	(0.1163)	(0.1742)
6-10 years of tenure	0.0303	0.0659	-0.0982	0.0257	0.1888
	(0.0780)	(0.1119)	(0.1147)	(0.1139)	(0.1718)
>10 years of tenure	-0.0735	-0.0791	-0.3332***	-0.2594**	-0.3499**
	(0.0761)	(0.1129)	(0.1168)	(0.1171)	(0.1741)
married	-0.1175***	-0.0922	-0.0493	-0.1372**	-0.0851
	(0.0439)	(0.0622)	(0.0614)	(0.0599)	(0.0924)
urban	0.2881***	0.3010***	0.3115***	0.4636***	0.4806***
	(0.0458)	(0.0665)	(0.0706)	(0.0715)	(0.1147)
Dushanbe	0.0813	0.2719***	0.3229***	0.0994	0.1006
	(0.0649)	(0.0920)	(0.0959)	(0.0973)	(0.1469)
Khatlon	-0.0267	-0.0420	0.0932	-0.1897*	-0.2674
	(0.0641)	(0.0919)	(0.0988)	(0.1031)	(0.1673)
Rrg region	0.3823***	0.6129***	0.8000***	0.5520***	0.4701***
	(0.0844)	(0.1143)	(0.1124)	(0.1101)	(0.1624)
Gbao region	-0.0853	0.0917	0.1106	0.1078	0.3413*
	(0.0772)	(0.1128)	(0.1151)	(0.1162)	(0.1785)
constant	1.6236***	1.8623***	2.2781***	2.8720***	3.4351***
	(0.1073)	(0.1464)	(0.1540)	(0.1550)	(0.2534)
N	1832	1832	1832	1832	1832
Pseudo R^2	0.0897	0.1111	0.1504	0.1877	0.1989

Notes: Standard errors in parentheses * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$; The omitted categories- basic education, tenure less than 1 year and Sogd region.

TABLE A9.12

Second stage QR estimates for the public and private sector earnings equations containing selection terms for males in Bulgaria, 2003

	Public sector earnings equation			Private sector earnings equation		
	θ=0.10	θ=0.50	θ=0.90	θ=0.10	θ=0.50	θ=0.90
university	0.7844*	1.4329***	0.9868	0.6430**	0.9374***	0.6259**
	(0.4055)	(0.3513)	(0.6573)	(0.3003)	(0.2437)	(0.2658)
college	0.3402	1.0179***	0.4792	0.3742	0.6716**	0.1033
	(0.4491)	(0.3607)	(0.7124)	(0.2861)	(0.2855)	(0.2873)
secondary	0.1554	0.2629**	0.1422	0.2042**	0.2143***	0.2123**
	(0.1327)	(0.1091)	(0.2003)	(0.0819)	(0.0733)	(0.0952)
exp	0.0214	0.0264**	0.0210	0.0077	0.0113	-0.0047
	(0.0142)	(0.0120)	(0.0220)	(0.0099)	(0.0085)	(0.0098)
expsq	-0.0005**	-0.0006***	-0.0004	-0.0001	-0.0002	0.0000
	(0.0002)	(0.0002)	(0.0004)	(0.0002)	(0.0002)	(0.0002)
1-2 years	0.0033	0.3039*	0.0462	0.1992**	0.1723**	-0.0426
	(0.1311)	(0.1559)	(0.2778)	(0.0880)	(0.0770)	(0.0893)
3-5 years	0.3161	1.0302***	0.4947	0.2167	0.4456**	-0.0811
	(0.2902)	(0.2784)	(0.4903)	(0.2211)	(0.1845)	(0.2017)
6-10 years	0.4279	1.6247***	0.8257	0.4770	0.9490***	0.0418
	(0.4470)	(0.4459)	(0.7993)	(0.3691)	(0.3140)	(0.3227)
>10 years	0.6371	2.3200***	1.2597	0.4781	1.1028**	0.0231
	(0.7148)	(0.6443)	(1.1894)	(0.5250)	(0.4482)	(0.4625)
married	0.0672	0.3512***	0.1569	-0.0079	0.0810	-0.0650
	(0.1274)	(0.1152)	(0.2167)	(0.0982)	(0.0821)	(0.0908)
managerial	0.1748	0.2309**	0.2892	0.4659**	0.3978**	0.1625
	(0.1218)	(0.1144)	(0.2027)	(0.1917)	(0.1680)	(0.1520)
urban	-0.0249	0.1701**	0.2795*	0.1493**	0.2343***	0.2252***
	(0.1117)	(0.0848)	(0.1565)	(0.0719)	(0.0631)	(0.0778)
Sofia	-0.2074	-0.3371*	-0.0843	0.0803	0.1708	0.5577***
	(0.2124)	(0.1833)	(0.3294)	(0.1388)	(0.1321)	(0.1475)
ps1	0.8196	2.9484***	1.8722	0.2739	0.7253	-0.9388
	(1.2636)	(1.0958)	(1.8714)	(1.0401)	(0.8470)	(0.8904)
ps2	-0.1930	-0.4214*	-0.3380	-0.2113	-0.4196	-0.4185
	(0.2831)	(0.2541)	(0.3864)	(0.5495)	(0.5373)	(0.5429)
constant	-0.2124	-3.2078**	-0.7205	0.5768***	0.9954***	1.4586***
	(1.7391)	(1.5292)	(2.8009)	(0.1461)	(0.1350)	(0.1665)
N	397	397	397	899	899	899

Standard errors in parentheses * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

TABLE A9.13

Second stage QR estimates for the public and private sector earnings equations containing selection terms for females in Bulgaria, 2003

	Public sector earnings equation			Private sector earnings equation		
	θ=0.10	θ=0.50	θ=0.90	θ=0.10	θ=0.50	θ=0.90
university	0.5745***	0.6527***	0.2095	0.7267***	0.6168***	0.7087***
	(0.1237)	(0.1126)	(0.2181)	(0.1814)	(0.1140)	(0.1384)
college	0.2718*	0.2602**	-0.1669	0.7208**	0.2646	-0.0086
	(0.1513)	(0.1301)	(0.2716)	(0.3138)	(0.1651)	(0.2434)
secondary	0.1344*	0.2594***	-0.2418*	0.4415***	0.1599**	0.1971**
	(0.0806)	(0.0664)	(0.1373)	(0.1098)	(0.0685)	(0.0782)
exp	0.0332***	0.0355***	0.0243*	0.0182*	0.0112*	0.0028
	(0.0071)	(0.0076)	(0.0144)	(0.0101)	(0.0068)	(0.0083)
expsq	-0.0007***	-0.0008***	-0.0005*	-0.0004*	-0.0002	0.0000
	(0.0001)	(0.0002)	(0.0003)	(0.0002)	(0.0002)	(0.0002)
1-2 years	-0.0298	0.0123	0.1518	0.0408	0.0953	0.0433
	(0.1132)	(0.1013)	(0.1734)	(0.1035)	(0.0697)	(0.0948)
3-5 years	0.2345	0.0008	0.4491	0.3038	0.3086**	0.1016
	(0.1611)	(0.1633)	(0.2959)	(0.2219)	(0.1382)	(0.1755)
6-10 years	0.1615	-0.0833	0.4992	0.3034	0.5269***	0.0667
	(0.2569)	(0.2330)	(0.4163)	(0.2969)	(0.1926)	(0.2542)
>10 years	0.2191	-0.0412	0.6107	0.4530	0.4215**	0.1162
	(0.2912)	(0.2687)	(0.4929)	(0.3193)	(0.1897)	(0.2223)
married	-0.0158	-0.0778*	-0.0445	0.0493	-0.0301	0.0067
	(0.0581)	(0.0470)	(0.0934)	(0.0831)	(0.0534)	(0.0624)
managerial	0.0636	0.0199	0.0634	0.7005***	0.5234***	0.2868
	(0.0842)	(0.0871)	(0.1996)	(0.2681)	(0.1721)	(0.2294)
urban	0.0079	0.0512	0.1308	0.0366	0.0654	0.1394*
	(0.0624)	(0.0546)	(0.0995)	(0.0880)	(0.0545)	(0.0711)
Sofia	-0.0308	0.0959	0.1492	0.0756	0.1297	-0.0924
	(0.0912)	(0.0814)	(0.1418)	(0.1420)	(0.1203)	(0.1610)
ps1	0.0010	-0.1789	0.0791	0.1892	0.3154	0.2340
	(0.4309)	(0.3833)	(0.7401)	(0.6560)	(0.4513)	(0.5650)
ps2	-0.0019	-0.0094	0.0359	-0.1563	0.0677	0.3854
	(0.1409)	(0.1302)	(0.2661)	(0.5074)	(0.3509)	(0.4319)
constant	0.4123	1.0029**	1.3218*	0.2161	0.9217***	1.3534***
	(0.4563)	(0.4118)	(0.7724)	(0.1551)	(0.1119)	(0.1483)
N	462	462	462	724	724	724

Standard errors in parentheses * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

TABLE A9.14

Second stage QR estimates for the public and private sector earnings equations containing selection terms for males in Serbia, 2003

	Public sector earnings equation			Private sector earnings equation		
	θ=0.10	θ=0.50	θ=0.10	θ=0.50	θ=0.10	θ=0.50
university	0.9957***	1.0854***	0.7145*	0.8024***	0.1387	-0.4223
	(0.2656)	(0.3097)	(0.4076)	(0.1930)	(0.3418)	(0.3338)
post-secondary	0.6174	0.6402***	0.9263	0.2567	0.5798***	0.7785
	(0.4462)	(0.2236)	(0.5828)	(0.2637)	(0.1998)	(0.5647)
gymnasium	0.5476**	0.4809***	0.6338**	0.3255*	0.3184**	0.2750
	(0.2402)	(0.1235)	(0.3054)	(0.1976)	(0.1304)	(0.3045)
secondary (4 years)	-0.0880	0.2604	-0.0733	0.2245	0.4601***	0.8416**
	(0.3300)	(0.1891)	(0.4522)	(0.1911)	(0.1655)	(0.4068)
secondary (1-2 years)	-0.0172	0.2080**	0.2953	0.1465	0.1679*	0.3559
	(0.1975)	(0.0997)	(0.2427)	(0.1298)	(0.0961)	(0.2389)
primary	0.0029	0.1177	0.0685	0.1451	0.0997	0.1130
	(0.1415)	(0.0832)	(0.2349)	(0.1065)	(0.0728)	(0.1628)
exp	0.0218	0.0084	0.0230	0.0066	0.0121	0.0180
	(0.0259)	(0.0125)	(0.0330)	(0.0103)	(0.0079)	(0.0195)
expsq	-0.0004	-0.0001	-0.0006	-0.0001	-0.0002	-0.0001
	(0.0005)	(0.0002)	(0.0007)	(0.0002)	(0.0001)	(0.0004)
1-2 years of tenure	0.3530	0.1639	1.8629***	-0.4730	-0.1230	-0.1048
	(0.5511)	(0.2804)	(0.6857)	(0.2952)	(0.2213)	(0.5741)
3-5 years of tenure	-0.2021	0.0034	0.5426	-0.2868	0.0139	-0.1479
	(0.5931)	(0.2980)	(0.7873)	(0.3486)	(0.2418)	(0.6966)
6-10 years of tenure	-0.1117	0.0962	0.8050	-0.5299	-0.1448	-0.5136
	(0.7044)	(0.3444)	(0.9448)	(0.4041)	(0.2941)	(0.8621)
>10 years of tenure	-0.4607	-0.0443	1.1441	-0.6709	-0.1552	-0.3181
	(1.0114)	(0.5014)	(1.3692)	(0.5762)	(0.4085)	(1.1921)
married	0.0079	0.0119	-0.0173	0.0687	0.0832	0.0489
	(0.1256)	(0.0712)	(0.1773)	(0.0994)	(0.0698)	(0.1590)
urban	0.0254	0.0156	-0.0071	0.1378*	0.0491	-0.1716
	(0.1157)	(0.0649)	(0.1641)	(0.0815)	(0.0595)	(0.1480)
Belgrade	0.4141	0.3872***	0.0688	0.7458***	0.3497***	0.4343
	(0.2586)	(0.1269)	(0.3168)	(0.1666)	(0.1167)	(0.3032)
Vojvodina	0.0536	0.1378	-0.1130	0.5760***	0.1140	0.1583
	(0.2774)	(0.1332)	(0.3575)	(0.1481)	(0.1103)	(0.3068)
West Serbia	-0.0215	0.1367	-0.2237	0.5797***	0.0567	0.2176
	(0.3381)	(0.1607)	(0.4482)	(0.1878)	(0.1358)	(0.3855)
Šumadija and Pomoravlje	0.0699	0.1592*	0.0706	0.4685***	0.1810**	0.3058
	(0.1647)	(0.0863)	(0.2130)	(0.1222)	(0.0869)	(0.2054)
East Serbia	-0.1256	0.0584	-0.2261	0.4419***	0.0664	-0.0227
	(0.2025)	(0.1041)	(0.2638)	(0.1534)	(0.1034)	(0.2467)
ps1	-0.1071	-0.4298	-0.4085	-0.3001	1.1723	-1.2442
	(1.8254)	(0.9428)	(2.5084)	(1.4410)	(0.9576)	(2.5357)
ps2	-0.1036	0.0779	1.0147	0.9899	1.0393*	-0.4540
	(0.5994)	(0.2785)	(0.6597)	(0.7870)	(0.5618)	(1.3924)
constant	3.4444*	3.8852***	3.0324	2.1726***	3.9672***	3.8778***
	(1.8107)	(0.9031)	(2.4690)	(0.6563)	(0.4402)	(1.2473)
N	670	670	670	796	796	796

Standard errors in parentheses * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

TABLE A9.15

Second stage QR estimates for the public and private sector earnings equations containing selection terms for females in Serbia, 2003

	Public sector earnings equation			Private sector earnings equation		
	θ=0.10	θ=0.50	θ=0.10	θ=0.50	θ=0.10	θ=0.50
university	0.9023***	1.1896***	1.8276***	2.5658***	1.9037***	1.1523**
	(0.3448)	(0.1656)	(0.4244)	(0.3404)	(0.1410)	(0.4593)
post-secondary	0.2730	0.6438***	1.1892***	0.2026	0.8703***	1.3229***
	(0.2828)	(0.0795)	(0.4159)	(0.2955)	(0.1443)	(0.4260)
gymnasium	0.5568**	0.5211***	0.5343	0.6052**	0.5354***	0.3628
	(0.2742)	(0.0723)	(0.3539)	(0.2565)	(0.1367)	(0.4120)
secondary (4 years)	0.2869	0.3551***	0.2143	0.2966	0.2826**	0.1151
	(0.3282)	(0.0964)	(0.4484)	(0.2554)	(0.1397)	(0.4274)
secondary (1-2 years)	0.2032	0.3281***	0.1873	0.4185**	0.3010***	0.4189*
	(0.2123)	(0.0538)	(0.2955)	(0.1915)	(0.0802)	(0.2425)
primary	0.0041	0.1882***	-0.0127	0.2653	0.0913	0.1712
	(0.2089)	(0.0614)	(0.3371)	(0.2123)	(0.0914)	(0.2806)
exp	-0.0096	-0.0075	0.0164	0.0265	0.0351***	0.0103
	(0.0356)	(0.0097)	(0.0415)	(0.0189)	(0.0082)	(0.0248)
expsq	0.0002	0.0001	-0.0003	-0.0005	-0.0005***	-0.0000
	(0.0007)	(0.0002)	(0.0008)	(0.0003)	(0.0001)	(0.0004)
1-2 years of tenure	-0.6588	-0.3109**	0.4949	-0.1216	0.2598	-0.1633
	(0.5954)	(0.1303)	(0.4940)	(0.3843)	(0.1879)	(0.5508)
3-5 years of tenure	-0.4923	-0.1005	0.6424	-0.0137	0.2311	-0.2113
	(0.5534)	(0.1094)	(0.5092)	(0.3789)	(0.1650)	(0.4654)
6-10 years of tenure	-0.6343	-0.1450	0.6414	-0.3472	0.1103	0.1506
	(0.4825)	(0.1075)	(0.5951)	(0.3356)	(0.1618)	(0.5164)
>10 years of tenure	-0.9432*	-0.2130*	0.7624	-0.3117	0.1502	-0.1066
	(0.5202)	(0.1155)	(0.6161)	(0.3608)	(0.1782)	(0.5749)
married	-0.0509	-0.0396	-0.0587	-0.0648	-0.0829	-0.0098
	(0.1580)	(0.0388)	(0.2015)	(0.1223)	(0.0627)	(0.1870)
urban	0.0695	0.0657*	0.0356	0.1895	0.1200*	0.0455
	(0.1598)	(0.0375)	(0.1817)	(0.1462)	(0.0621)	(0.2213)
Belgrade	0.5245**	0.3299***	0.2054	0.2833	0.2471**	0.3039
	(0.2237)	(0.0590)	(0.3012)	(0.2250)	(0.1033)	(0.3276)
Vojvodina	0.4789**	0.2407***	-0.0680	0.3909*	0.0824	0.1778
	(0.2248)	(0.0585)	(0.2994)	(0.2194)	(0.0946)	(0.3010)
West Serbia	0.2434	0.1493**	-0.2231	-0.0440	-0.1646	-0.3433
	(0.2620)	(0.0643)	(0.3487)	(0.2567)	(0.1158)	(0.3490)
Šumadija and Pomoravlje	0.2188	0.1790***	0.0538	0.1811	0.0682	0.2200
	(0.2232)	(0.0532)	(0.2628)	(0.1965)	(0.0955)	(0.2887)
East Serbia	0.1367	0.2059***	-0.0805	-0.3495	-0.1043	-0.0645
	(0.2764)	(0.0677)	(0.3298)	(0.2342)	(0.1122)	(0.3728)
ps1	-2.0892**	-0.4289**	0.3416	-0.5050	0.4949	-0.2732
	(0.8308)	(0.2172)	(1.0097)	(1.2798)	(0.5766)	(1.9890)
ps2	0.5569	0.0253	0.0783	0.0748	0.2861	-0.2246
	(0.4521)	(0.1057)	(0.4288)	(0.8931)	(0.4235)	(1.4249)
constant	4.9557***	4.0516***	3.2224***	2.1322***	3.1710***	3.9271***
	(0.8569)	(0.2168)	(1.1194)	(0.3883)	(0.1799)	(0.5860)
N	498	498	498	486	486	486

Standard errors in parentheses * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

TABLE A9.16

Second stage QR estimates for the public and private sector earnings equations containing selection terms for males in Russia, 2003

	Public sector earnings equation			Private sector earnings equation		
	θ=0.10	θ=0.50	θ=0.10	θ=0.50	θ=0.10	θ=0.50
university	0.4964***	0.3450***	0.2949***	0.4431***	0.4009***	0.3897***
	(0.0351)	(0.0201)	(0.0283)	(0.0541)	(0.0309)	(0.0393)
college	0.2593***	0.2320***	0.2265***	0.4535***	0.3685***	0.3505***
	(0.0646)	(0.0377)	(0.0521)	(0.0879)	(0.0492)	(0.0616)
secondary	0.2482***	0.1681***	0.0782***	0.2446***	0.1743***	0.1441***
	(0.0284)	(0.0162)	(0.0230)	(0.0383)	(0.0215)	(0.0273)
primary	0.0797**	0.0408*	-0.0139	0.1373***	0.1127***	0.0455
	(0.0403)	(0.0230)	(0.0323)	(0.0528)	(0.0291)	(0.0365)
exp	-0.0049	0.0075***	0.0125***	-0.0039	0.0101***	0.0069*
	(0.0042)	(0.0024)	(0.0034)	(0.0058)	(0.0031)	(0.0039)
expsq	-0.0000	-0.0003***	-0.0003***	0.0000	-0.0003***	-0.0002**
	(0.0001)	(0.0000)	(0.0001)	(0.0001)	(0.0001)	(0.0001)
managerial	0.1746***	0.0996***	0.0992***	0.3392***	0.2260***	0.1710***
	(0.0264)	(0.0154)	(0.0220)	(0.0422)	(0.0240)	(0.0303)
married	0.1177***	0.1266***	0.1305***	0.0857**	0.1353***	0.1599***
	(0.0313)	(0.0175)	(0.0244)	(0.0406)	(0.0225)	(0.0278)
1-2 years of tenure	0.1403***	0.1474***	0.0940***	0.1248***	0.0964***	0.0656**
	(0.0425)	(0.0244)	(0.0344)	(0.0465)	(0.0259)	(0.0320)
3-5 years of tenure	0.2762***	0.2263***	0.1826***	0.2583***	0.1839***	0.1140***
	(0.0450)	(0.0259)	(0.0367)	(0.0555)	(0.0304)	(0.0379)
6-10 years of tenure	0.3341***	0.2312***	0.1951***	0.1722***	0.1940***	0.2218***
	(0.0452)	(0.0251)	(0.0353)	(0.0571)	(0.0313)	(0.0386)
>10 years of tenure	0.2881***	0.2744***	0.2464***	0.0907*	0.1948***	0.2304***
	(0.0429)	(0.0240)	(0.0337)	(0.0515)	(0.0283)	(0.0355)
arrears	-0.3164***	-0.2814***	-0.2768***	-0.7190***	-0.4329***	-0.2900***
	(0.0281)	(0.0168)	(0.0236)	(0.0383)	(0.0222)	(0.0283)
Central region	0.2255***	0.3074***	0.3699***	0.3021***	0.1247***	0.1155***
	(0.0391)	(0.0222)	(0.0313)	(0.0550)	(0.0304)	(0.0380)
North-West	0.5834***	0.6320***	0.6683***	0.5793***	0.5583***	0.5149***
	(0.0422)	(0.0241)	(0.0340)	(0.0618)	(0.0341)	(0.0427)
Siberia	0.2805***	0.3593***	0.4463***	0.2715***	0.2982***	0.4044***
	(0.0436)	(0.0249)	(0.0350)	(0.0608)	(0.0336)	(0.0418)
Far-East	0.7503***	0.7967***	0.7643***	0.6742***	0.6209***	0.7450***
	(0.0412)	(0.0236)	(0.0332)	(0.0656)	(0.0365)	(0.0452)
Urals	0.3177***	0.5251***	0.8542***	0.5587***	0.4515***	0.6600***
	(0.0506)	(0.0288)	(0.0405)	(0.0678)	(0.0374)	(0.0464)
Volga	0.1166***	0.2171***	0.1872***	0.1670***	0.1152***	0.0053
	(0.0407)	(0.0233)	(0.0329)	(0.0580)	(0.0321)	(0.0399)
ps1	-0.5765***	0.1017*	0.5784***	-12.7486***	-9.5304***	-7.8672***
	(0.1083)	(0.0565)	(0.0728)	(1.1738)	(0.6573)	(0.9042)
ps2	0.0734	-0.3551***	-0.5592***	-7.5415***	-5.4560***	-4.6119***
	(0.0849)	(0.0454)	(0.0578)	(0.7334)	(0.4115)	(0.5719)
constant	1.4666***	2.1551***	2.7806***	-3.7306***	-1.7887***	-0.1630
	(0.0598)	(0.0336)	(0.0456)	(0.4395)	(0.2441)	(0.3285)
N	13020	13020	13020	8854	8854	8854

Standard errors in parentheses * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

TABLE A9.17

QR estimates for second stage estimates for the public and private sector earnings equations containing selection terms for males in Tajikistan, 2003

	Public sector earnings equation			Private sector earnings equation		
	θ=0.10	θ=0.50	θ=0.10	θ=0.50	θ=0.10	θ=0.50
university	0.4529***	0.3078***	0.1980	-0.0881	0.0062	0.1711
	(0.1639)	(0.0868)	(0.1953)	(0.1707)	(0.1148)	(0.2167)
secondary technical	0.2267	0.2222**	0.4671**	-0.0137	0.0160	0.0927
	(0.1791)	(0.0968)	(0.2142)	(0.1614)	(0.1118)	(0.2137)
secondary special	-0.0574	0.0859	0.2429	0.1286	-0.1750	0.0667
	(0.1810)	(0.0947)	(0.2137)	(0.1542)	(0.1065)	(0.1997)
secondary general	0.0958	0.0923	0.3135*	0.1046	-0.0194	0.1045
	(0.1481)	(0.0781)	(0.1747)	(0.1090)	(0.0780)	(0.1423)
primary	-0.1977	-0.0039	-0.0756	0.4091*	-0.2001	-0.4808*
	(0.2609)	(0.1663)	(0.3979)	(0.2379)	(0.1640)	(0.2789)
exp	0.0218	0.0262***	0.0288**	0.0471***	0.0285***	0.0166
	(0.0136)	(0.0067)	(0.0145)	(0.0099)	(0.0075)	(0.0156)
expsq	-0.0004	-0.0006***	-0.0006**	-0.0010***	-0.0006***	-0.0003
	(0.0003)	(0.0001)	(0.0003)	(0.0002)	(0.0001)	(0.0003)
1-2 years of tenure	-0.1325	-0.1016	0.4219*	-0.0158	0.0983	0.3286*
	(0.2131)	(0.1113)	(0.2469)	(0.1329)	(0.0919)	(0.1742)
3-5 years of tenure	0.0420	-0.1017	-0.0476	-0.1326	-0.1772**	-0.0191
	(0.1848)	(0.0978)	(0.2155)	(0.1338)	(0.0893)	(0.1745)
6-10 years of tenure	-0.0824	-0.1249	-0.2263	-0.2231	-0.2279**	-0.1787
	(0.1772)	(0.0933)	(0.1989)	(0.1402)	(0.0922)	(0.1857)
>10 years of tenure	0.1331	-0.1750*	-0.5306***	-0.6207***	-0.6846***	-0.3667
	(0.1878)	(0.0985)	(0.2028)	(0.1628)	(0.1105)	(0.2245)
married	0.1229	-0.0158	-0.2871*	-0.0238	0.1759**	0.0618
	(0.1407)	(0.0711)	(0.1623)	(0.1075)	(0.0751)	(0.1389)
urban	0.2180**	0.3165***	0.4103***	0.7021***	0.6630***	0.6587***
	(0.0918)	(0.0507)	(0.1107)	(0.0898)	(0.0606)	(0.1201)
Dushanbe	0.0002	0.1338*	-0.1221	-0.4557**	-0.0112	0.3023
	(0.1361)	(0.0709)	(0.1529)	(0.1911)	(0.1256)	(0.2384)
Khalton	0.0411	0.1736**	-0.1538	-0.3533*	-0.2257*	-0.0017
	(0.1303)	(0.0710)	(0.1556)	(0.1925)	(0.1303)	(0.2593)
Rrg region	0.4008***	0.4354***	0.3396*	0.0763	-0.0109	-0.3668
	(0.1525)	(0.0779)	(0.1812)	(0.2140)	(0.1385)	(0.2546)
Gbao region	-0.0340	0.5064***	0.9380***	0.0560	0.2107	0.3159
	(0.1497)	(0.0808)	(0.1639)	(0.2020)	(0.1347)	(0.2534)
ps1	-0.2940	0.5874***	0.7703*	0.2632	0.9592	-0.1961
	(0.3747)	(0.1882)	(0.4295)	(1.0104)	(0.7136)	(1.3748)
ps2	0.1886	-0.4642***	-0.6724***	1.2816*	1.8105***	1.0133
	(0.2153)	(0.1096)	(0.2568)	(0.7077)	(0.4980)	(0.9122)
constant	1.8959***	2.7859***	4.1621***	1.8434***	3.1891***	3.8763***
	(0.3008)	(0.1503)	(0.3115)	(0.3895)	(0.2848)	(0.5645)
N	1397	1397	1397	1609	1609	1609

Standard errors in parentheses * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

TABLE A9.18

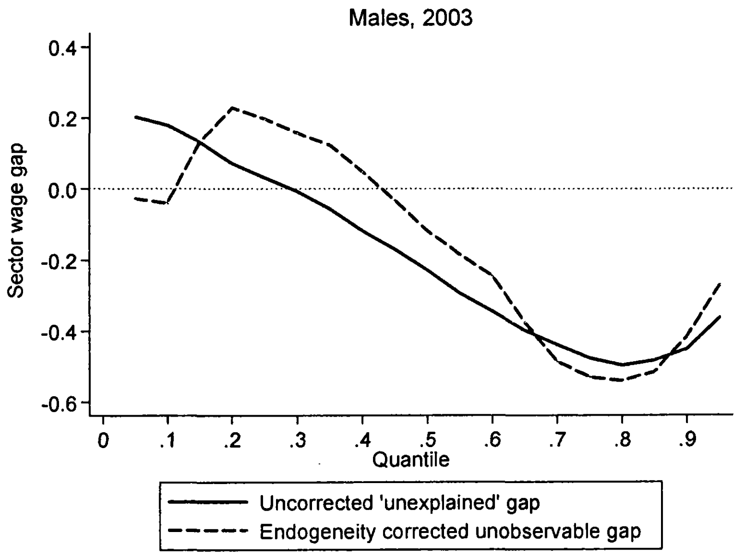
QR estimates for second stage estimates for the public and private sector earnings equations containing selection terms for females in Tajikistan, 2003

	Public sector earnings equation			Private sector earnings equation		
	θ=0.10	θ=0.50	θ=0.10	θ=0.50	θ=0.10	θ=0.50
university	0.7362***	0.8116***	0.7736***	-0.1387***	0.6445***	0.6659
	(0.2452)	(0.1184)	(0.1597)	(0.0430)	(0.2040)	(0.4335)
secondary technical	0.4278	0.4074***	0.6304***	0.0697	0.4037	0.0021
	(0.3201)	(0.1500)	(0.1992)	(0.0604)	(0.2559)	(0.2924)
secondary special	0.1448	0.2329*	0.3850**	0.1051**	0.5333***	0.3128
	(0.2529)	(0.1215)	(0.1572)	(0.0482)	(0.1966)	(0.3978)
secondary general	0.1730	0.0901	0.1189	0.0188	0.1284	0.1586
	(0.2056)	(0.0969)	(0.1182)	(0.0221)	(0.0906)	(0.1666)
primary	-0.0628	-0.1410	0.1013	-0.2859***	-0.1917	-0.4616*
	(0.3535)	(0.1721)	(0.1881)	(0.0472)	(0.1660)	(0.2799)
exp	0.0146	0.0236***	0.0328***	0.0010	0.0212**	0.0087
	(0.0169)	(0.0089)	(0.0116)	(0.0030)	(0.0100)	(0.0176)
expsq	-0.0003	-0.0005***	-0.0006**	-0.0000	-0.0002	0.0001
	(0.0004)	(0.0002)	(0.0003)	(0.0001)	(0.0002)	(0.0003)
1-2 years of tenure	-0.5217*	-0.1450	-0.0738	0.2721***	-0.0326	0.0946
	(0.2717)	(0.1438)	(0.1891)	(0.0408)	(0.1514)	(0.3047)
3-5 years of tenure	0.0213	-0.0205	0.2196	0.1891***	0.1365	0.1132
	(0.2381)	(0.1243)	(0.1580)	(0.0370)	(0.1393)	(0.2669)
6-10 years of tenure	-0.1129	-0.1592	0.0694	0.1492***	0.1400	0.3563
	(0.2227)	(0.1206)	(0.1596)	(0.0378)	(0.1414)	(0.2632)
>10 years of tenure	-0.0810	0.0224	0.0689	0.1171***	0.0527	-0.2736
	(0.2396)	(0.1251)	(0.1610)	(0.0426)	(0.1644)	(0.3213)
married	-0.1417	-0.0999*	-0.2148***	0.0005	-0.0383	0.0751
	(0.1198)	(0.0581)	(0.0787)	(0.0213)	(0.0840)	(0.1520)
urban	-0.1019	0.0154	0.2395***	0.6993***	0.5044***	0.6083***
	(0.1219)	(0.0625)	(0.0865)	(0.0283)	(0.0988)	(0.2256)
Dushanbe	0.0185	0.5184***	0.2591**	-0.9655***	-0.4881***	-0.2720
	(0.1520)	(0.0778)	(0.1054)	(0.0406)	(0.1658)	(0.3084)
Khalton	0.0489	0.3512***	0.0191	-0.8772***	-0.7508***	-0.7833**
	(0.1729)	(0.0893)	(0.1330)	(0.0485)	(0.1890)	(0.3652)
Rrg region	0.4742**	0.8405***	0.7975***	-0.3819***	0.0859	-0.0637
	(0.1872)	(0.0938)	(0.1119)	(0.0463)	(0.1802)	(0.3319)
Gbao region	0.0216	0.2614***	0.6310***	-0.9926***	-0.6989***	-0.1977
	(0.1879)	(0.0959)	(0.1361)	(0.0472)	(0.1883)	(0.3615)
ps1	-0.3952	0.2724	0.9975***	0.0500	1.1380	1.4732
	(0.4687)	(0.2149)	(0.2710)	(0.3651)	(1.2326)	(2.2706)
ps2	0.3792	-0.0047	-0.3895**	0.5678**	1.2764	1.8406
	(0.3062)	(0.1399)	(0.1820)	(0.2358)	(0.8314)	(1.6000)
constant	1.6365***	1.9013***	2.3315***	2.2105***	3.1366***	4.1153***
	(0.3287)	(0.1722)	(0.2154)	(0.1382)	(0.4802)	(0.9220)
N	812	812	812	1020	1020	1020

Standard errors in parentheses * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

FIGURE A 9.1

Comparison between uncorrected 'unexplained' gap and endogeneity corrected unobservable gap for males in Tajikistan, 2003



CHAPTER TEN

CONCLUSIONS AND POLICY IMPLICATIONS

10.1. Introduction

This thesis undertook country specific empirical research on a range of key labour market indicators. The empirical findings have contributed in filling the gap in the existing literature on returns to education, gender wage gap and public-private wage differential in a number of post-transition countries. The current Chapter highlights and summarises the key findings from each of the preceding empirical essays and develops overall conclusions. Broader examination of the contribution of the thesis provides a more informative point to consider the possible policy implications and directions for future research.

Each chapter in the present thesis examined our main objectives' providing empirical evidence to implement and support this. Several contributions have been made. First, the literature has been enhanced by providing comprehensive analysis on the returns to education, gender wage gap and public-private sector wage differential in four transition countries – Bulgaria, Serbia, Russia and Tajikistan. Relatively few studies were found which had covered together a number of countries. To the best of our knowledge, this is the first empirical study that estimates evolution of returns to education in Bulgaria over the transition period from the early 1986 to 2003. The main novelty in respect of the first empirical study lays on the estimation of the rate of returns to education controlling for endogeneity and selectivity bias in a quantile regression framework. Second, contributions have been made to the econometrics literature by extending the popular Machado and Mata (2005) decomposition technique and considering the 'index' number problem. The proposed technique has not been applied before in the context of gender earnings gap and sector wage differential in the selected transition countries. The third contribution has been to further extend the technique by considering the endogenous sector choice. The idea was to use the technique to simulate other counterfactual distributions for public and private sector workers by using sector selection adjusted coefficients and being sampled from the whole sample of workers. The method for accounting for endogenous selection also allowed decomposition of the selection gap into a component due to observables and component due to unobservable characteristics. Finally, the technique has been further modified by adapting a time-wise decomposition framework in order to identify the main factors contributing to the changes in the ratio of the public to private sector earnings for two periods identified as being of interest. Again the study for Bulgaria is the first one which examined the differences in the sector wage differential across the earrings distribution in early 1986. As far as it is known, the study on public-private sector gap is the first one that both controls for endogeneous sector choice and analyses the public sector pay gap across the distribution in the selected transition countries. The results will be of interest to both policy makers and practitioners interested in what has happened to the relative pay of public and private sector workers over time.

10.2. Summary of main findings

Returns to education over time

The first empirical analysis permitted us to account for the variation in the rate of returns to education in Bulgaria over time, looking at the rates of returns to education in pre-1990s period. The study estimated returns to education before and during transition from centrally planned to market economy for a country that has received little attention in the literature. The main finding was that returns to education in Bulgaria increased over time, which confirmed the statement that increase in human capital would raise productivity and thus would increase individual earnings. While the average returns to education were positive and quite low during the communist regime – in the range of 1.1% for males and 2.1% for females in 1986, it increased to 5.8% for males and 6.7% for females in 2003, when the wage constraints imposed by the 'old regime' were relaxed. In line with Flabbi *et al.* (2008) the largest increase in the rate of returns to education took place in the early transition period, when it increased from 1.1% in 1986 to 4.5% in 1995 for males and from 2.1% to 7.0% for females. Women also showed higher returns to education than men. The most prominent increase in the rate of returns to education over the period of 1986 to 2003 occurred at the top of the earnings distribution, where the rate of returns to education, in particular for females, rose from a negative and insignificant value in 1986 to nearly 7% in 2003.

The study also addressed the selectivity and endogeneity biases by adopting Heckman's two-step, semi-parametric two-step approach, and instrumental variable method. The sample selection correction has not made a significant difference in the coefficients for both males and females in Bulgaria. In almost all specifications, the Mills ratio and the power series approximation terms were statistically insignificant. Only for males in 1995, the two approximation terms were significant at 5% level. The similarity of the parametric and semi-parametric two-step estimates indicated that inclusion of the two approximation terms essentially performed the same task as the inclusion of the Mills ratio. The endogeneity issue in the education-earnings relationship was examined by instrumenting years of schooling with mother's education and variation in the average regional schooling. Assumption in applying the regional average differences in schooling as an instrument was based on the data district's educational composition, which provided evidence on diverse regional human capital concentration. The corresponding IV results revealed a similar increasing pattern over time. By using mother education as an instrument for schooling, the IV estimates for males rose from 2.2% in 1986 to 9.9% in 2003. Similarly, for females the returns to education increased from 4.2% in 1986 to 10.5% in 2003. Overall, the endogeneity adjusted returns to education estimates increased by approximately 42% as compared to the OLS estimates, and this is broadly in line with the literature. The estimates have not changed significantly in the over-identified specifications where both mother's education and regional average schooling were used as instruments for schooling.

An explanation for the steady increase in the rate of returns to education was that, as a transition country, Bulgaria moved from a system of compressed wage scale to a market-oriented system. The 1980s was a decade of considerable economic change. The situation changed with transition, when the structure of incentives was progressively altered in favour of college and university education. The expansion of skill-intensive services increased the demand for more educated employees in Bulgaria. However, it was not the intention to evaluate the systematic link between education, economic reforms and technological changes. Following Flabbi *et al.*

(2008) it was found that returns to education in Bulgaria evolved differently across the distribution and we indicated that non-competitive forces might have determined the wage levels. The results demonstrated that the returns to education for both males and females in Bulgaria were driven by an increasing trend in the upper end of the conditional wage distribution.

Heterogeneity in returns to education

In the second empirical study we examined dispersion of the returns to education in Bulgaria, Serbia, Russia and Tajikistan, attempting to test for evidence of heterogeneity for individuals in the upper part of the wage distribution as compared to the individuals in the lower part of the distribution. It was advantageous to have data for each country for the same year, as this allowed comparison of the results between countries. The estimates suggested a monotonic increase in the rate of returns to education across the conditional earnings distribution for both males and females in Bulgaria and for females in Tajikistan. Returns to education in Bulgaria increased between the 10^{th} and 90^{th} percentile levels from 3.9% to 6.2% for males and from 4.7% to 7.2% for females. For females in Tajikistan, the return at the 10^{th} percentile was no greater than 6% and it reached 12% at the 90^{th} percentile. In contrast, for both males and females in Russia, the estimated returns to education were higher at the lower part of the earnings distribution, suggesting that education had higher impact on the 'less able' individuals. A little heterogeneity in the estimated returns to education was found in Serbia.

The selectivity bias was also addressed by applying Buchinsky (1998) semi-parametric approach. The results for both males and females in Russia and for males in Tajikistan were sensitive to the inclusion of the power series approximation terms. In Russia there was compelling evidence of strong sample selection effect at the bottom of the distribution where the estimated coefficients increased from 11.7% to 14.6% for males and from 12.1% to 19.4% for females. Correcting for selection had a drastic effect on the returns to schooling at the 10^{th} percentiles for males in Tajikistan, changing the coefficient from insignificant to significantly different from zero. However, the two power-series correction terms were not significant for either males or females in Bulgaria and Serbia. The test of whether the selectivity corrected returns to education differ across each quanitle levels was significantly different for female workers in Bulgaria and Tajikistan and for males in Russia.

The issue of endogeneity in the quantile regression framework was considered by applying Lee's (2007) control function approach. For all four countries, the DWH test undertaken at the mean led to strong rejection of the null hypothesis of exogeneity in education, therefore correcting for endogeneity was a necessary step in the analysis. Overall, the endogeneity adjusted QR estimates were higher with the unadjusted estimates. Finally, by considering both endogeneity and sample selection biases, the quantile regression point estimates for Russia suggested considerable homogeneity across conditional wage distribution. The negative earnings premium, found in the unadjusted QR estimates, was no more evident.

Gender wage differential in transition

The third empirical study considered the gender wage gap across entire earnings distribution by examining the same countries of interest. Applying the extended quantile regression decomposition to data from Bulgaria, Serbia, Russia and Tajikistan, positive coefficient effect was found across the whole earnings distribution, indicating that in all these countries otherwise identical men and women receive different returns to their characteristics. The coefficient effect was larger at the top of the distribution than at the bottom for both years in Bulgaria, Serbia and Tajikistan. The observed positive and significant differential, found in Russia, was larger at the middle of the distribution. The decomposition results indicated that the gender wage differential in Bulgaria, Russia and Serbia was mainly attributed to differences in return, that is, the 'male advantage' and 'female disadvantage' and to a lesser extent to the employee's endowment component. Hence, the conclusion is that the observed gender differences in human capital characteristics appeared to be too small (and often in favour of women) to explain the large wage discrepancies between men and women.

A further interesting consideration concerned the changes in the composition of the samples over time. The study focused on two time periods in order to observe whether the differences in payment between men and women have changed over time. Unfortunately, the data availability did not permit us to explore the changes over a longer time span and to examine the position before 1990 in all four countries. We found that both coefficient and covariate components of the gap contributed to the actual evolution of the gender wage differential over time. The largest fraction of the gap was attributed to the differences in the coefficient component. Overall, there was a clear trend in narrowing of the earnings differential over the examined period for both males and females in Bulgaria, Serbia and Russia. In 2007, the overall gender wage gap decreased in Serbia, but the coefficient effect of the gap was pronounced especially at the higher end of the wage distribution. In contrast, the estimated differential for Tajikistan has risen over the observed period and this was mainly due to the increased effect of the endowment component. The relatively high unexplained wage differential in Russia is in line with Newell and Reilly (2001) findings. In Bulgaria and Serbia, women showed more favourable observed characteristics than men.

Public-private wage differential

The third essay provided a comprehensive empirical study of the wage differential between the public and private sector in Bulgaria, Serbia, Russia and Tajikistan. Given differences in the sector wage setting procedures, the study sought to ascertain whether identical individuals working in the same job in the public and in the private sector would earn the same or a different amount. Public sector men and women in Bulgaria were found to be the ones who derive the greatest financial benefit from such employment. The advantage enjoyed by public sector workers in Bulgaria tended to be higher for males. Although public sector employees in Russia have individual characteristics that were more conducive to a higher remuneration, they faced a wage penalty, especially at the top of the distribution. A negative sector wage differential across the entire wage distribution was found for females in Russia and for males in Serbia. In Serbia, the advantage enjoyed by public sector females tended to be higher than those found for males. At the top of the conditional earnings distribution, however, Serbian women, like men, tended to face a wage penalty from being employed in the public sector. In Tajikistan, a positive wage premium was found at the lower part of the distribution, and significant wage penalty at the upper part of the distribution, a result similar to a number of international studies (see for example Blackaby *et al.*, 1999, Poterba and Rueben, 1998).

The selection process between both sectors was also considered. As a first step estimation for the endogenous selection correction, an estimation was made of the probability of working in the public sector conditional on a set of characteristics in a probit equation. Since probit equation depends for consistency heavily on the distributional assumption, the sector choice equation was also estimated using single index model (Ichimura, 1993). For the identification of the selection effect a variable which allowed for a possible regional variation in the public sector employment was constructed. With some exceptions, the coefficients of the probit and the single index estimates were not fundamentally different. The standard errors of the probit estimates were generally slightly lower than those of the single index model. Overall, in three of the countries, the probability to work in the public sector increased with higher education and the worker's tenure. As the intercept in the wage equations is not separately identified in the quantile regression models and it is conflated with the constant term of the higher order power series, Anndrews and Scahfgan (1998) identification at unity was followed.

The results of the sector wage gap show that controlling for the endogenous sector choice has an important impact on the estimated public-private wage gap. Public sector employees located at the lower end of the wage distribution were positively selected for this sector. At the end of the distribution, those with the highest wage potential chose the private sector. These findings are largely consistent with the rest of the related literature (Heitmueller (2004), Barbosa and Filho (2008)). In particular, our results indicate that when the endogeneity is considered, in most cases, the sector wage gap tended to be lower compared against the uncorrected results obtained by adopting the modified Machado and Mata procedure. However, the sector wage gap for females in Bulgaria widened, which suggests that Bulgarian women were positively selected into the public sector.

The study also explored what had happened to the public sector hourly wage differential at different points in the conditional earnings distribution over time by providing further evidence for Bulgaria and Russia. We found that the sector wage differential widened over time for both males and females in Bulgaria. The increase in the public earnings tended to be higher at the top of the conditional earnings distribution. Differences in the observable characteristics gained importance over time in accounting for the general increasing trend in the public sector earnings distribution in Bulgaria. The sector wage differential in Russia tended to increase over time for males. In contrast, the differential changed substantially over the observed period for females. Further investigation of what lay behind this change suggests that both changes in the wage advantages/disadvantages of public and private sector in Russia and changes in the characteristics of workers at which the advantages/disadvantages was measured, contributed to the overall coefficient effect.

10.3. Policy implications and recommendations

In concluding the present thesis, the policy implications of our research will be considered. This is a difficult task due to the fact that establishing a link between research results and policy initiative is not always straightforward. The main policy implications and recommendations that can be drawn from the empirical evidence can be summarised in several points.

The economic return to education is a fundamental parameter of interest in many different areas of economics and public policy. It allows researchers to perceive the significance of the role played by education in terms of enhancing an individual's wage perspectives. The importance of the returns to education is seen in the OECD adoption as a key indicator in their annual *Education at Glance* series. The observed increasing trend in the rate of returns to education between 1986 and 2003 suggests that the liberalised Bulgarian economy has responded to the market forces by providing larger returns for human capital investment. These results may reflect transformation of the labour market to a more flexible and less rigid wage structure. Increased returns to education may be due to the need of wages to respond to market forces after transition.

The finding of education-ability complementarities has potentially strong implications. The fact that returns to education can be heterogeneous across individuals has implications for the inequality-reducing role of education. If returns to education for the less able are lower, it follows that a policy of raising the level of schooling for everyone will generally increase the inequality of earnings. Even the policies which are targeted at the less able may be a poor use of resources if marginal returns are very low (Denny and O'Sullivan, 2007). On the other hand if education is benefiting only the higher-earners, then education policy needs to be targeted towards improving educational quality and outcomes for the lower income individuals.

The gender wage decomposition findings have a very important message for the policy makers, oriented in providing equal opportunities for women. Gender inequality constraints economic growth, increases poverty and has negative effect on the well-being of individuals. A recent study '*Global Employment Trends for Women*' published by the ILO in 2009 provides information about the global gender pay gap. The report reconfirms that gender inequality remains an issue within the labour market globally. However, the report does not look at the gap in different points of the wage distribution. This study attests that a more detailed explanation of the wage gap, utilizing quantile regression technique, is required.

Despite the remarkable improvement of female labour marker characteristics, we found that men are paid more than women. The estimated gender wage differential was found to be high in all four countries. Males in Tajikistan for example, were found to be paid on average nearly 24% more compared to their female counterparts. At the bottom of the distribution the estimated coefficient effect of the gap was lower, suggesting that minimum wage legislations in Tajikistan helped to reduce the wage gap and led to a relatively equal distribution for low-paid workers. The wage differential was found to be mainly attributed to differences in returns and to a lesser extent to the workers observable characteristics. In other words, the pricing of workers' characteristics drives gender wage differentials in all these countries. Although the Tajikistan data show women are still less educated than men (23% of employed men in 2007 have completed university degree compared to 15% for women), further improvements in the human capital alone are unlikely to be sufficient as long as gender differences remain in returns.

One further implication of our results is the necessity to improve the research in transition countries and to investigate the nature of the 'unexplained' component. Concrete policy measures need to take into account the evidence on the various factors underlying the gender wage differential. This step requires not only clear sex-disaggregated data from baseline household surveys but also attention in those surveys to the causes of gender inequality, for example barriers in accessing certain productive inputs. Furthermore, a pro-active policies and measures to enhance payment equity in all examined countries are needed, especially in Tajikistan, where a significant gender pay differential was found. Clearly, labour market policies in that particular country should be aimed at removing barriers to female career progression.

Existence of a sizable public sector wage differential in the examined countries can make it difficult for the public sector to retain and recruit 'high ability' workers. The positive premium for public sector workers, found in Bulgaria, has obvious implications for the country's labour market and can create 'queues' for public sector jobs given they are comparatively well-paid across the entire wage distribution. Higher public sector wages in Bulgaria may have negative consequence on the fiscal position of the government. It will increase the wage bill and place a strain on the fiscal position of the public sector. In this context, the recent government's decision in Bulgaria to freeze the public sector wages is an obvious solution to halt the differences in the relative pay of public and private sector workers.

On the other hand, underpayment of the public sector workers relative to the private sector at the top of the distribution, found in Serbia, Russia and Tajikistan, may induce negative selection into the public sector. A recent call by the Prime Minister of Russia to raise the public sector wages by 6.5% was clearly meant to address the underpaid public sector workers. The existence of a positive public sector wage differential in the lower tail of the wage distribution compared to the top tail, found for females in Serbia, males in Russia and for both males and females in Tajikistan, implied that skilled workers in the public sectors in these countries were much less well paid than their private sector counterparts. This may result in the public sector being characterised as a sector that has difficulty in retaining high-skilled workers. In order to attract high skill workers and to improve the quality of public service, public wage setting policy in these countries should introduce more flexible pay structures in the public sector, such as increasing the role of competition and linking pay with performance. It is vital for these countries, in both sectors, to pursue wage policy that aims to increase medium-run productivity and achieve moderate inflation.

10.4. Limitations and avenues for future research

Limitations

The analysis in this thesis is not without limitations. These limitations could also help to open possibilities for further empirical research. The first issue to be pointed out is data limitation and availability. One problem was the lack of reliable data, or sometimes any data in the countries of interest. It was impossible to obtain a longer time span of data for each country, as required by the analysis over time. Poorer countries, such as Bulgaria, Serbia, Russia or Tajikistan, have been largely ignored by statistical authorities. Examination of the individual data from the transition period in Eastern Europe and Central Asia suggests that availability of data in Britain for example, compares favourably with that in these countries. The position in the Central European countries, such as Czech Republic, Hungary and Poland differs as well. In these countries there has been a long tradition of data collection.

The concern about quality of data on household incomes in transition countries extends beyond the constraints of this thesis. In the literature a common expressed view is that Eastern European data cover only 'official' income and that 'hidden' economy income is not reported. As discussed in Chapter Nine, illegal employment and tax evasion have been problematic during the early transition period. Because these problems occurred in the private sector, actual labour income in the private sector might be underestimated both by official wage statistics and probably by reported wages in the surveys.

Despite the availability of detailed information in the data files, it was not possible to fully capture all workers specific characteristics. In the case of Bulgaria the Sargan test of instrumental validity was passed, lending support to the case for the instruments. Due to the lack of information on the parent education in the other samples, however, we were unable to test the over-identifying restriction on the returns to education for Serbia, Russia and Tajikistan. Moreover, due to issues of comparability in using national cross-sectional data it is recommended that World wide data will be a potentially useful source of information from which to provide up to date estimations on different indicators on the labour market outcomes.

Although we considered human capital and work experience as major determinants of the gender pay gap, another possible answer to that question might be that the degree of gender specific policies and antidiscrimination laws differ from country to country. Unfortunately, it was difficult to capture the effect of these policies with the models utilised in this thesis. Better data that brings information on actual experience, firm size and female decisions could improve analysis of the evolution of the gender wage gap since it will bring more aspects that were not captured here.

Finally, the workplace characteristics such as presence of performance-related payment, company pension schemes and family employment practices are important in explaining the sector wage differential. Unfortunately, the data in this study had no information on work efforts and non-wage benefits, and it was not possible to include more job related characteristics than those used in the models. Therefore, the sector wage differential examined in essay three may not reflect the total compensation differentials between the two sectors.

Empirical evidence shows the importance of accounting for endogeneity and selectivity biases in the estimates of the gender wage differential. It was surprising, that apart from Albrecht *et al.* (2009), Picchino and Mussida (2010), Badel and Pena (2010) and Chzhen and Mumford (2011), there were no studies that correct the gender wage gap for sample selection across the conditional wage distribution. Given the data, however, we were not able to cope with all of the methodological issues. No attempt was made to account for the potential endogeneity problem in occupational decision as there was a lack of appropriate instruments in the samples. Using the presence or total number of young children as instrumental variables in the selection equation is a common practise in the literature on wage modelling. However, the variable was found to be not valid for men in the estimates of this research. No significant selection effects were detected in the case of Bulgaria and Serbia. Due to these difficulties, the influence of the endogeneity issue of occupational choice and selectivity on the estimates of gender wage gap was left for future research. The proposed method that accounts for endogenous sector selection, applied in essay three, could successfully be utilised to account for selection in the gender wage gap. This is proposed as a subject for further research.

The most important point to note was that the applied decomposition method allowed differences across the conditional wage distribution to be taken into account and the index number problem was overcome by choosing the pooled sample of males and females as a non-discriminatory wage structure. Although the decomposition method of Brown *et al.* (1980) was appealing as it focuses on occupational segregation, the method was not applicable across the earnings distribution.

Directions for future research

There are a number of specific aspects that would provide an avenue for future research. First, a future updates on the returns to education based on more recent data could not only provide more robust estimates, but also provide evidence on whether the performance of transition's labour market are improving over time. Second, as the evidence presented in this thesis has all been based on cross sectional data, an alternative identifying strategy relying on time variation in individual wages and work histories could be extended to analyse the gender wage gap and account for unobserved characteristics.

An obvious agenda for future research in the public-private sector wage gap literature would be to investigate the extent to which the sector wage differential in these countries changes by controlling for both non-labour force participation and endogeneous sector choice biases. Another suggestion for future research would be to examine whether the differences in the public-private sector wage differential can be explained by compensating factors such as fringe benefits and working conditions.

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