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Partial dollarization and financial frictions in emerging economies

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Abstract

How do financial frictions affect macroeconomic volatility and monetary policy in emerging market economies? This article assesses the empirical relevance of such frictions by estimating a two-bloc emerging market/rest-of-the-world model containing two key features of emerging economies: partial transaction and liability dollarization, and financial frictions where capital financing is partially or totally in foreign currency. Our estimation employs the "one-step approach" which allows us to be "agnostic" regarding nonstationarity in the data and simultaneously estimate structural and trend parameters. Using data for Peru and the US, we find substantial empirical support for both the financial accelerator and partial dollarization mechanisms. The data fit of the baseline model improves with the addition of each of these frictions, exogenous shocks are significantly amplified in their presence and our preferred model captures several important stylized facts of a small emerging open economy.

KEYWORDS

dollarization, emerging economies, financial frictions

JEL CLASSIFICATION

E52, E37, E58

1 | INTRODUCTION

The financial crisis of 2007–2008 and the ensuing Great Recession have highlighted the need to understand how large external shocks are propagated in small open economies. This is particularly relevant in emerging market countries, since these economies face additional vulnerabilities that make them very different from advanced economies and aggravate their responses to global shocks. Indeed, these economies usually display weak fiscal, monetary, and financial institutional frameworks, and have imperfect access to capital markets. These frailties can lead to sudden and sharp reversals of capital inflows (the “sudden stops” highlighted in Calvo, 1998), the inability of firms to borrow in domestic currency only, or the presence of significant monitoring costs in credit markets, thus exacerbating finance premiums faced by borrowers (the “financial accelerator” mechanism).

These features may substantially amplify the effects of large external disturbances to the domestic economy. Indeed, as Table 1 shows, many emerging economies experienced episodes of high inflation that have seen economic agents resorting to a foreign currency as a partial replacement for the domestic currency, a phenomenon known as dollarization.[1]

There is a broad consensus that dollarization is a potential cause for balance of payments and financial crises, thus endangering macroeconomic and financial stability. A currency depreciation will deteriorate the balance sheets of borrowers relying on foreign currency denominated debt and increase the external finance premium accrued at the international interest rate. The ensuing fall in the demand for capital reduces the value of the borrowers' existing capital stock and corresponding net worth, further amplifying the increase in the costs of borrowing and the swings in investment and production.

While there is a substantial body of literature devoted to understanding business cycle dynamics (and financial frictions) in developed economies, research focusing on emerging economies is relatively sparser. In addition, developing countries are usually small open economies, which requires careful modeling of a mechanism that allows for the transmission of external shocks (see, e.g., Edwards, 2010).

A series of papers have separately addressed some of these issues. Cespedes et al. (2004) and Gertler et al. (2007) introduce a “financial accelerator” mechanism into a small open economy model and find that fixing exchange rates may lead to unwanted outcomes due to “counter-cyclical' monetary policy decisions. On the other hand, earlier currency substitution models emphasize the fact that monetary aggregates become more sensitive to changes in devaluation expectations, with a more recent view stressing the higher pass-through and a weaker monetary transmission (see Rennhack & Nozaki, 2006).

Batini et al. (2010) analyze the consequences of (partial) dollarization for exchange rate stabilization and the conduct of monetary policy (see also Castillo et al., 2013; Felices & Tuesta, 2013). In fact, when dollarization is present, monetary transmission is different than in the case of no dollarization, given that changes in the exchange rate have a stronger impact either on inflation expectations, or on real activity, or both. This poses an extra challenge when setting monetary conditions in response to shocks.

There are various other forms of financial frictions in the business cycle literature that acknowledge powerful transmission mechanisms of shocks in advanced and emerging economies with financial sectors. The Gertler et al. (2007)-type mechanism has been developed by Gertler and Kiyotaki (2010) and Gertler and Karadi (2011) in a number of respects by introducing a banking sector, the costs of enforcing contracts and allowing for interbank frictions. Another simple

TABLE 1 Degree of deposit and loan dollarization across emerging economies

	Foreign currency deposits/loans (in % of total deposits/loans)										
Year	**Bolivia**	**Costa Rica**	**Kazakhstan**	**Morocco**	**Nigeria**	**Pakistan**	**Paraguay**	**Peru**	**Sri Lanka**	**Tunisia**	**Turkey**
2002	80.6/85.2	49.5/41.1	–	7.7/0	9.6/0	14.4/0	69/22	67.4/77.7	0/0	37/0	–
2003	81.4/75.9	46.5/40.6	–	3.7/0	8.8/0	12.3/0	62.2/26.3	65.9/74.1	0/0	43.8/0	75.1/97.6
2004	78.3/68.4	50.2/41	16.6/92.2	4/0	5.4/0	11.3/0	51/23.3	61.4/72	0/0	57.1/0	84.4/86.8
2005	62.3/64.2	41.1/40.3	8.8/95.7	3.3/0	3.6/0	10/0	51.4/19.2	55/68.5	0/0	48.8/0	82/72.8
2006	42.7/98.2	40.3/38.6	11.5/95.6	3.5/0	2/0	8.5/0	49.9/17.4	54.3/61.7	0/0	71.2/0	78.3/87.2
2007	37/96.6	35.1/34.7	40.9/43.3	3.7/0	0/0	8.1/0	39.7/13.7	49.6/56.6	0/0	81.8/0	78/1/100
2008	37.1/92.8	32.9/44.9	38.8/47.4	4/0	5.4/0	13/0	51/23.3	42.5/52.4	0/0	73.3/0	71.9/18.2
2009	50.1/79.1	47.6/42.8	28/0	5/0	0.3/0	14.7/0	37.8/15	45.4/47.6	0/0	74.6/0	65.4/12.1
2010	37.1/49.4	50/16	15.5/0	7.8/0	0.2/0	13.2/0	36.4/24	39.7/44.2	0/0	68.9/0	68.2/1.34
2011	28/35.1	47.8/4.8	14.8/6.2	9.4/0	0.2/0	12.5/0	41.3/12.4	38.3/44.3	10/29.6	80.7/0	56/7.5
2012	31.3/21.7	47.5/5.3	33.3/18.5	10.8/0	0.1/0	13.1/0	42.9/12.7	33.6/43.5	15.2/51.3	84.5/0	82.7/19.4
2013	53.7/19.7	50.4/2.7	32.8/10.7	15.9/0	0.1/0.3	14.7/0	52/3.3	31.8/41.9	10/32.3	79.3/0	90/27.7
2014	50.8/26.7	50.2/1	41.2/0	28.8/0	5.4/0.5	12.7/0	56.5/4.3	33.3/39.1	16.5/0.1	66.2/0	94/32.1
2015	44.5/21.9	46.4/0.4	47/9/0	26.2/0	0.8/0.4	10.2/0	60.2/5.7	36.3/33.8	16.4/0.1	60.2/0	92.5/25.5
2016	43.3/17.9	44.2/0	58.2/0	28.3/0	0/0.9	7.6/0	59.3/6.2	36.8/29.2	18.5/0.1	51.7/0	90.2/23.3
2017	56.6/14.5	49.5/0	46.2/0	27.2/0	0/1.4	8.1/0	54.3/6	33.3/28.6	18.1/0.1	50.2/0	85.6/30.4
2018	46.2/12	48.5/0	43.7/0	22.4/0	0.2/4.4	11.4/0	50.4/5.3	31.9/28.4	16.2/0.1	51.1/0	85/34.5

Note: Data sources—The dollarization data are constructed from the IMF's Standardized Report Form (SRF) and BCRPData at the Banco Central de Reserva del Peru. As not all countries report their foreign currency deposits and loans via the SRF, in most cases data is available only since 2002. We have removed some countries from the sample as their deposits and loans are either fully dollarized (e.g. Ecuador and El Salvador) or not denominated in foreign currency (e.g. Brazil and Colombia).

type of friction arising between the firm-entrepreneurs relation takes the form of a collateral constraint in the spirit of Kiyotaki and Moore (2007), which imposes credit limits endogenously. The differences between the types of models are in the location of the financial frictions, their nature, the role of net worth and the distribution of risk-taking, and therefore these types of financial frictions have different roles in accentuating business cycle fluctuations.[2]

Another strand of literature estimates emerging and small open economy models to understand the contributions of various shocks and financial frictions in conjunction with the effects from nonstationary productivity shocks (Aguiar & Gopinath, 2007; Chang & Fernandez, 2013; Garcia-Cicco et al., 2010). Extending an RBC model with trend technology shocks and departing from the frictionless credit market assumption with a simple form of risk premium that depends on the level of external debt similar to our assumption, these papers show some successful performance in capturing several important stylized facts of emerging economies (such as countercyclicality of current account balance and excessive volatility in consumption).

Neumeyer and Perri (2005) and Fernandez and Gulan (2015) are the two other important papers that study emerging economies business cycles and develop small open economy models with financial frictions that generate cyclical movements consistent with the data. Both of these papers focus on real interest rate shocks and aim to explain empirically the observed countercyclicality of country-specific interest rates in emerging economies. The latter focuses on a financial accelerator, similar to our assumption, that provides a microfoundation for linking interest rate spreads to corporate leverage. More recently, Caballero et al. (2019) provide empirical evidence about foreign financing of firms and study the real effects of external financial shocks. However, all these papers motivating the use of financial frictions in models for emerging economies have not considered denomination of liabilities and the transmission of the balance sheet effects for understanding business cycles and stabilization policy in these economies.

Thus, building on these papers, this study aims at gauging the empirical relevance of financial frictions and how they affect macroeconomic volatility and monetary policy. We follow Batini et al. (2010) and employ a two-bloc emerging-market/rest-of-the-world dynamic stochastic general equilibrium (DSGE) model, with the emerging market bloc displaying a strong link between changes in the exchange rate and financial distress of households and firms. More precisely, the model features: (a) financial frictions in the form of a financial accelerator, since firms are obliged to finance at least part of their capital requirements in foreign currency (see Gertler et al., 2007; Gilchrist, 2014); (b) domestic households hold both local and foreign currency money balances for transaction purposes; (c) the relative demand of foreign currency is endogenous to the extent of exchange rate stabilization by the central bank.

Features (a)–(c), directly linking the cost of external borrowing and domestic firms' net worth, play a central role in our empirical analysis and are novel in the context of an estimated two-bloc model. While there are well-known shortcomings of the Gertler et al. (2007)-type mechanism, it is suitable to conceptualize our analysis and provides the simplest framework to incorporate and study the role of dollarization associated with borrowers' net worth, which is accumulated by risk-neutral entrepreneurs and a financial intermediary.

We focus on the case of Peru, taking the US as the rest of the world. Although Peru has not officially dollarized, de facto dollarization, involving both monetary substitution and asset substitution is widespread. The ratio of per capita holdings of US currency and domestic currency is over 1:1, suggesting that Peru is in fact a highly dollarized economy. Also, estimates of monetary and asset substitution indices indicate that asset substitution dominates currency substitution.[3]

Liability dollarization has not been a significant impediment in stabilizing inflation, as shown in Reinhart et al. (2003). These authors study a sample of partially dollarized economies (including Peru) and find that dollarization does not necessarily prevent monetary policy from bringing inflation under control. On the other hand, using the exchange rate as a flexible anchor has proved successful. Indeed, it is well known that liability dollarization has remained at high levels despite the sharp decline in inflation in dollarized countries such as Peru.

Thus, by bringing our model to the data, we should be able to throw some light on the mechanisms by which the financial frictions interact with shocks hitting the Peruvian economy and provide an empirical assessment of the implications of the different types of financial imperfections considered here, for which there is scant empirical evidence in the literature. This is important, as the different types of dollarization, coupled with the "financial accelerator" mechanism, affect the economy through different channels, which is crucial for the design of stabilization policies.

To do so, we employ Bayesian system estimation techniques to estimate distinct variants of the model, but do so considering several novel aspects. First, an important innovation of this article is our "trend agnostic" stance regarding the nature of nonstationarity in the data. Indeed, on the estimation front we believe that the role of trends in the data requires careful treatment. For example, inflation in Peru ranges around 10%–40% in 1991–1993, just before the beginning of our sample period. GDP growth ranged from −1% to 12% from 1994 to 2006. It is difficult to argue that, with a log-linearized model around a steady state, undergoing data transformation/filtering before estimation to account for nonstationarity is a reasonable approach especially when dealing with emerging economies.[4]

Therefore, in this article, we address this issue by adopting the alternative "one-step approach" of Ferroni (2011) in which the raw data is used to simultaneously estimate the structural parameters in a hybrid framework, in which the cyclical fluctuations of the data are seized by the solution of the DSGE model and the noncyclical fluctuations are captured by a flexible reduced form representation. This contrasts with the common "two-step" practice of first filtering the data and then estimating structural parameters. However, the usually arbitrary choice of a particular statistical filter is likely to affect the estimation, as estimates of structural parameters will inevitably "compensate" for potential trend misspecifications (see Cogley, 2001; Gorodnichenko & Ng, 2010). We believe this is an important contribution of this article, dealing with small open economy modeling with a more transparent treatment of nonstationary series.

In addition, another distinctive aspect of our analysis is our assumption regarding the US policy stance, both in the theoretical model and the estimation. Indeed, standard DSGE models fail to account properly for the behavior of key (US) macroeconomic variables around and following the Great Recession. Given that our sample period encompasses the Great Recession, and in line with recent evidence on policy shifts in the Federal Reserve's approach to inflation management, we depart from the restrictive assumption that the Federal Reserve has used the same policy rule with the constant predetermined target inflation throughout the sample period. We follow Del Negro and Eusepi (2011) and Del Negro et al. (2015) in including financial frictions and allowing for time variation in the inflation target in the Taylor rule. In implementing this feature, we employ data on long-run inflation expectations as an additional observable, thus expanding our information set to include the forecasts of changes in the target inflation. Indeed, long run inflation expectations essentially determine the level of the target inflation rate at each point in time.

Indeed, in our empirical analysis, we find substantial evidence in favor of a model variant incorporating both the financial accelerator and partial dollarization mechanisms, thus suggesting that this type of financial frictions may be pivotal in understanding business cycles in developing open economies.

The rest of this article is organized as follows. Section 2 outlines the model. Section 3 describes the data and estimation methodology. Section 4 sets out the results, while Sections 4.2 and 4.3 include model comparisons and further validation. Sections 5 and 6 contain the analysis of the dynamic properties of the key variables under the different frictions. Section 7 considers additional robustness checks while Section 8 provides concluding remarks.[5]

2 | THE MODEL

As in Batini et al. (2010), the model features two asymmetric and unequally sized blocs, each one with different household preferences and technologies, with the small open economy as the limit when the relative size of the larger bloc tends to infinity. Households work, save and consume tradable goods produced both at home and abroad. At home, there are three types of firms: wholesale, retail, and capital producers. As in Gertler et al. (2007), wholesale firms purchase capital and employ household labor to produce wholesale goods that are sold to the retail sector. Wholesalers' demand for capital in turn depends on their financial position, which varies inversely with wholesalers' net worth. In the monopolistically competitive retail sector, firms differentiate wholesale goods and sell the differentiated (repackaged) goods to households and capital producers. The capital goods sector is competitive and converts the final good into capital.

There are four departures from the standard open-economy model that lead to interesting results. First, money enters utility in a non-separable way and results in a direct impact of the interest rate on the supply side.[6] Second, in the emerging market bloc, households derive utility from holding both domestic and foreign money (dollars) balances as in Felices and Tuesta (2013). Third, along the lines of Gilchrist (2014) (see also Cespedes et al., 2004), firms face an external finance premium that increases with leverage and part of the debt of wholesale firms is financed in foreign currency (dollars), because it is impossible for firms to borrow 100% in domestic currency owing to "original sin" type constraints. Finally, there are frictions in the world financial markets facing households as in Benigno (2009).

Departures two and three add an additional dimension to openness itself. Indeed, domestic agents not only hold foreign bonds and derive utility from consuming foreign-produced goods, as in standard open-economy models, but also borrow in foreign currency from domestic agents and derive utility from holding foreign money balances. The main elements of the model are as follows (see Batini et al., 2010 for details).[7]

2.1 | Households

Normalizing the total population to be unity, there are ν households in the "home', emerging economy bloc and $(1-\nu)$ households in the "foreign" bloc. A representative household h in the home bloc maximizes the non-separable utility function

$$E_t \sum_{t=0}^{\infty} \beta^t \left(\frac{\left[\Phi_t(h)^{1-\varrho}(1-L_t(h))^{\varrho}\right]^{1-\sigma}}{1-\sigma} \right), \tag{1}$$

where

$$\Phi_t(h) \equiv \left[b(C_t(h) - h_C C_{t-1})^{\frac{\theta-1}{\theta}} + (1-b) Z_t(h)^{\frac{\theta-1}{\theta}} \right]^{\frac{\theta}{\theta-1}},$$

$$Z_t(h) \equiv \left[a \left(\frac{M_{H,t}(h)}{P_t} \right)^{\frac{\chi_M - 1}{\chi_M}} + (1-a) \left(\frac{S_t M_{F,t}(h)}{P_t} \right)^{\frac{\chi_M - 1}{\chi_M}} \right]^{\frac{\chi_M}{\chi_M - 1}}.$$

$C_t(h)$ is a Dixit–Stiglitz index of consumption defined below in (4), $M_{H,t}(h)$ and $M_{F,t}(h)$ are end-of-period nominal domestic and foreign currency balances respectively, P_t is a Dixit–Stiglitz price index defined in (6) below, S_t is the nominal exchange rate and $L_t(h)$ are hours worked, measured as a proportion of a day, normalized at unity.[8] Note that consumption and money holdings together, and leisure (equal to $1 - L_t(h)$) are substitutes. We augment the utility function with a preference shock to the marginal utility of consumption, $\varepsilon_{C,t}$, and a shock to labor supply, $\varepsilon_{L,t}$. The degree of transaction dollarization is captured by $1 - a \in [0, 1]$.

Financial frictions facing households are incorporated as in Benigno (2009). There are two risk-free one-period bonds denominated in the currencies of each bloc with payments in period t, $B_{H,t}$ and $B_{F,t}$, respectively, in (per capita) aggregate. The prices of these bonds are given by

$$P_{B,t} = \frac{1}{1 + R_{n,t}}; \qquad P^*_{B,t} = \frac{1}{\left(1 + R^*_{n,t}\right) \phi \left(\frac{S_t B_{F,t}}{P_t} \right)},$$

where $\phi(\cdot)$ captures the cost in the form of a risk premium for home households to hold foreign bonds. We assume $\phi(0) = 1$ and $\phi' < 0$. $R_{n,t}$ and $R^*_{n,t}$ denote the nominal interest rate over the interval $[t, t+1]$. For analytical convenience, the home households can hold foreign bonds, but foreign households cannot hold home bonds. Then the net and gross foreign assets in the home bloc are equal. The representative household h must obey a budget constraint

$$\begin{aligned} &P_t C_t(h) + P_{B,t} B_{H,t}(h) + P^*_{B,t} S_t B_{F,t}(h) + M_{H,t}(h) + S_t M_{F,t}(h) + TF_t \\ &\quad = W_t(h) L_t(h) + B_{H,t-1}(h) + S_t B_{F,t-1}(h) + M_{H,t-1}(h) + S_t M_{F,t-1}(h) + \Gamma_t(h), \end{aligned} \tag{2}$$

where $W_t(h)$ is the wage rate, $\Gamma_t(h)$ are dividends from ownership of firms, and TF_t is net flat-rate transfers received per household. In addition, if we assume that households' labor supply is differentiated with elasticity of supply η, then (as we shall see below) the demand for each consumer's labor supplied by ν identical households is given by

$$L_t(h) = \left(\frac{W_t(h)}{W_t} \right)^{-\eta} L_t, \tag{3}$$

where $W_t = \left[\frac{1}{\nu} \sum_{r=1}^{\nu} W_t(h)^{1-\eta} \right]^{\frac{1}{1-\eta}}$ and $L_t = \left[\left(\frac{1}{\nu} \right) \sum_{r=1}^{\nu} L_t(h)^{\frac{\eta-1}{\eta}} \right]^{\frac{\eta}{\eta-1}}$ are the average wage index and average employment, respectively.[9]

Let the number of differentiated goods produced in the home and foreign blocs be n and $(1 - n)$, respectively, again normalizing the total number of goods in the world at unity. We also assume that the ratio of households to firms are the same in each bloc. It follows that n and $(1 - n)$

(or ν and $(1-\nu)$) are measures of size. The per capita consumption index in the home bloc is given by

$$C_t(h) = \left[\mathrm{w}^{\frac{1}{\mu}} C_{H,t}(h)^{\frac{\mu-1}{\mu}} + (1-\mathrm{w})^{\frac{1}{\mu}} C_{F,t}(h)^{\frac{\mu-1}{\mu}} \right]^{\frac{\mu}{\mu-1}}, \tag{4}$$

where μ is the elasticity of substitution between home and foreign goods, with $\zeta > 1$ denoting the elasticity of substitution between varieties in each bloc aggregated in the usual way.[10]

Weights in the consumption baskets in the two blocs are defined by

$$\mathrm{w} = 1 - (1-n)(1-\omega)\ ; \quad \mathrm{w}^* = 1 - n(1-\omega^*), \tag{5}$$

where $\omega,\ \omega^* \in [0,1]$ are parameters that capture the degree of "bias" in the two blocs. If $\omega = \omega^* = 1$, we have autarky, while $\omega = \omega^* = 0$ gives us the case of perfect integration. In the limit, as the home country becomes small, $n \to 0$ and $\nu \to 0$. Hence, $\mathrm{w} \to \omega$ and $\mathrm{w}^* \to 1$. Thus, the foreign bloc becomes closed, but as long as there is a degree of home bias and $\omega > 0$, the home bloc continues to consume foreign-produced consumption goods.

Standard results from the households optimal intra-temporal decisions and the usual Dixit–Stiglitz aggregation yields the domestic consumer price index

$$P_t = \left[\mathrm{w}(P_{H,t})^{1-\mu} + (1-\mathrm{w})(P_{F,t})^{1-\mu} \right]^{\frac{1}{1-\mu}}. \tag{6}$$

Letting S_t be the nominal exchange rate, the law of one price applies to differentiated goods, so that $\frac{S_t P^*_{F,t}}{P_{F,t}} = \frac{S_t P^*_{H,t}}{P_{H,t}} = 1$. Then it follows that the real exchange rate $RER_t = \frac{S_t P^*_t}{P_t}$ and the terms of trade, defined as the domestic currency relative price of imports to exports, $\mathcal{T}_t = \frac{P_{F,t}}{P_{H,t}}$, are related by the relationship

$$RER_t \equiv \frac{S_t P^*_t}{P_t} = \frac{\left[\mathrm{w}^* + (1-\mathrm{w}^*)\mathcal{T}_t^{\mu^*-1} \right]^{\frac{1}{1-\mu^*}}}{\left[1 - \mathrm{w} + \mathrm{w}\mathcal{T}_t^{\mu-1} \right]^{\frac{1}{1-\mu}}}.$$

Thus, if $\mu = \mu^*$, then $RER_t = 1$ and the law of one price applies to the aggregate price indices *iff* $\mathrm{w}^* = 1 - \mathrm{w}$. The latter condition holds if there is no home bias. If there is home bias, the real exchange rate appreciates (RER_t falls) as the terms of trade deteriorates.

Maximizing (1) subject to (2) and (3), treating habit as exogenous, and imposing symmetry on households (so that $C_t(h) = C_t$, etc.) yields standard results, from which we arrive at the *modified UIP condition*

$$\frac{P_{B,t}}{P^*_{B,t}} = \frac{E_t \left[U_{C,t+1} \frac{P_t}{P_{t+1}} \right]}{E_t \left[U_{C,t+1} \frac{S_{t+1} P_t}{S_t P_{t+1}} \right]},$$

where $U_{C,t}$ is the marginal utility of consumption. Moreover, assuming flexible wages, we obtain

$$\frac{W_t}{P_t} = -\frac{\eta}{(\eta-1)} \frac{U_{L,t}}{U_{C,t}}, \tag{7}$$

where $U_{L,t}$ is the marginal disutility of work, noting that the real disposable wage is proportional to the marginal rate of substitution between consumption and leisure, $-\frac{U_{L,t}}{U_{C,t}}$, and the constant of proportionality reflects the market power of households that arises from their monopolistic supply of a differentiated factor input with elasticity η.

2.1.1 | Rule of thumb households

We assume there are two groups of household, a proportion λ without credit constraints and the remaining proportion $1-\lambda$ who consume out of wage income. Let $C_{1,t}(h)$, $W_{1,t}(h)$, and $L_{1,t}(h)$ be the per capita consumption, wage rate, and labor supply, respectively, for this latter group. Then the optimizing households are denoted as before with $C_t(h)$, $W_t(h)$, and $L_t(h)$ replaced with $C_{2,t}(h)$, $W_{2,t}(h)$, and $L_{2,t}(h)$. We then have the budget constraint of the rule of thumb (RT) consumers

$$P_t C_{1,t}(h) = W_{1,t}(h) L_{1,t}(r) + TF_{1,t},$$

where $TF_{1,t}$ is net flat-rate transfers received per credit-constrained household. Following Erceg et al. (2005), we further assume that RT households set their wage to be the average of the optimizing households. Then, since RT households face the same demand schedule as the optimizing ones, they also work the same number of hours. Hence, in a symmetric equilibrium of identical households of each type, the wage rate is given by $W_{1,t}(r) = W_{1,t} = W_{2,t}(r) = W_{2,t} = W_t$ and hours worked per household is $L_{1,t}(h) = L_{2,t}(h) = L_t$. The only difference between the income of the two groups of households is that optimizing households as owners receive the profits from the markup of domestic monopolistic firms. Average consumption per household over the two groups is given by $C_t = \lambda C_{1,t} + (1-\lambda) C_{2,t}$.[11]

2.2 | Firms

2.2.1 | Wholesale firms

Wholesale goods are homogeneous and produced by entrepreneurs who combine differentiated labor and capital with a technology $Y_t^W = K_t^{\alpha}(A_t L_t)^{1-\alpha}$, where K_t is beginning-of-period t capital stock.

$$L_t = \left[\left(\frac{1}{\nu} \right)^{\frac{1}{\eta}} \sum_{r=1}^{\nu} L_t(h)^{(\eta-1)/\eta} \right]^{\eta/(\eta-1)},$$

where we recall that $L_t(h)$ is the labor input of type h, and A_t is an exogenous labor-augmenting technology shock. Equating the marginal product and cost of aggregate labor gives

$$W_t = P_{H,t}^W (1-\alpha) \frac{Y_t^W}{L_t}. \tag{8}$$

Let Q_t be the real market price of capital in units of total household consumption. Then noting that profits per period are $P_{H,t}^W Y_t^W - W_t L_t = \alpha P_{H,t}^W Y_t^W$, using (8), the expected return on capital over the period, acquired at the beginning of period t, is given by

$$E_t(1+R_t^k) = \frac{\frac{P_{H,t}^W}{P_t}\alpha\frac{Y_t^W}{K_t} + (1-\delta)E_t[Q_{t+1}]}{Q_t},$$

where δ is the depreciation rate of capital. This expected return must be equated with the expected cost of funds over $[t, t+1]$, taking into account credit market frictions. Wholesale firms borrow in both home and foreign currency, with proportion of the former given by $\varphi \in [0,1]$, so that this expected cost is

$$\begin{aligned}(1+\Theta_t)\varphi E_t\left[(1+R_{n,t})\frac{P_t}{P_{t+1}}\right] &+ (1+\Theta_t)(1-\varphi)E_t\left[(1+R^*_{n,t})\frac{P^*_t}{P^*_{t+1}}\frac{RER_{t+1}}{RER_t}\right] \\ = (1+\Theta_t)&\left[\varphi E_t\left[(1+R_t)\right] + (1-\varphi)E_t\left[(1+R^*_t)\frac{RER_{t+1}}{RER_t}\right]\right]. \end{aligned} \tag{9}$$

If $\varphi = 1$ or if UIP holds this becomes $(1+\Theta_t)E_t\,[1+R_t]$. The presence of dollarized liabilities is measured by the fraction $1-\varphi$ which, along with the transaction dollarization parameter a, are the key parameter configurations used by our model variants defined in Section 2.5. In (9), $R_{t-1} \equiv \left[(1+R_{n,t-1})\frac{P_{t-1}}{P_t}\right] - 1$ is the ex post real interest rate over $[t-1, t]$ and $\Theta_t \geq 0$ is the external finance premium given by

$$\Theta_t = \Theta\left(\frac{B_t}{N_t}\right) \;\; ; \;\; \Theta'(\cdot) > 0, \;\; \Theta(0) = 0, \;\; \Theta(\infty) = \infty,$$

where $B_t = Q_tK_t - N_t$ is bond-financed acquisition of capital in period t and N_t is the beginning-of-period t entrepreneurial net worth, the equity of the firm. Note that the ex post return at the beginning of period t, R^k_{t-1}, is given by

$$1 + R^k_{t-1} = \frac{\frac{P^W_{H,t-1}}{P_{t-1}}\alpha\frac{Y_{t-1}}{K_{t-1}} + (1-\delta)Q_t}{Q_{t-1}}$$

and this can deviate from the *ex ante* return on capital.

Assuming that entrepreneurs exit with a given probability $1-\xi_e$, net worth accumulates according to $N_t = \xi_e V_t$, where V_t, the net value carried over from the previous period, is given by

$$\begin{aligned} V_t = \Big[&(1+R^k_{t-1})Q_{t-1}K_{t-1} \\ &- (1+\Theta_{t-1})\left(\varphi(1+R_{t-1}) + (1-\varphi)(1+R^*_{t-1})\frac{RER_t}{RER_{t-1}}\right)(Q_{t-1}K_{t-1} - N_{t-1})\Big]. \end{aligned} \tag{10}$$

Note that in (10), $(1+R^k_{t-1})$ is the ex post return on capital acquired at the beginning of period $t-1$, $(1+R_{t-1})$ is the ex post real cost of borrowing in home currency and $(1+R^*_{t-1})\frac{RER_t}{RER_{t-1}}$ is the ex post real cost of borrowing in foreign currency. Also note that net worth N_t at the beginning of period t is a non-predetermined variable since the ex post return depends on the current market value Q_t, itself a non-predetermined variable.

2.2.2 | Retail firms

Retail firms are monopolistically competitive, buying wholesale goods and differentiating the product at a fixed resource cost ϕ_F. In a free-entry equilibrium profits are driven to zero. Retail output for firm f is then $Y_t(f) = Y_t^W(f) - \phi_F$, where Y_t^W is produced according to the above production technology. Retail firms set prices of differentiated goods according to the following. Assume that there is a probability of $1 - \xi_H$ at each period that the price of each good f is set optimally to $\hat{P}_{H,t}(f)$. If the price is not reoptimized, then it is held constant.[12] For each producer f the objective is at time t to choose $\hat{P}_{H,t}(f)$ to maximize discounted profits

$$E_t \sum_{k=0}^{\infty} \xi_H^k D_{t,t+k} Y_{t+k}(f) \left[\hat{P}_{H,t}(f) - P_{H,t+k} \mathrm{MC}_{t+k}\right],$$

where $D_{t,t+k}$ is the discount factor over the interval $[t, t+k]$, subject to a common downward sloping demand from domestic consumers and foreign importers of elasticity ζ and $MC_t = \frac{P_{H,t}^W}{P_{H,t}}$ are marginal costs.[13] The solution to this is

$$E_t \sum_{k=0}^{\infty} \xi_H^k D_{t,t+k} Y_{t+k}(f) \left[\hat{P}_{Ht}(f) - \frac{\zeta}{(\zeta - 1)} P_{H,t+k} \mathrm{MC}_{t+k}\right] = 0$$

and by the law of large numbers the evolution of the price index is given by

$$P_{H,t+1}^{1-\zeta} = \xi_H \left(P_{H,t}\right)^{1-\zeta} + (1 - \xi_H)(\hat{P}_{H,t+1}(f))^{1-\zeta}.$$

2.2.3 | Capital producers

Capital producers combine existing capital, K_t, leased from the entrepreneurs to transform an input, I_t, gross investment, into new capital according to

$$K_{t+1} = (1 - \delta)K_t + (1 - S(\tilde{X}_t))I_t \ ; \quad S',\ S'' \geq 0 \ ; \quad S(1) = S'(1) = 0,$$

where $\tilde{X}_t \equiv \frac{I_t}{I_{t-1}}$, so that adjustment costs are associated with *changes* rather than *levels* of investment. Gross investment consists of domestic and foreign final goods

$$I_t = \left[\mathrm{w}_I^{\frac{1}{\rho_I}} I_{H,t}^{\frac{\rho_I - 1}{\rho_I}} + (1 - \mathrm{w}_I)^{\frac{1}{\rho_I}} I_{F,t}^{\frac{\rho_I - 1}{\rho_I}}\right]^{\frac{\rho_I}{1-\rho_I}},$$

where weights in investment are defined as in the consumption baskets, with investment price given by

$$P_{I,t} = \left[\mathrm{w}_I (P_{H,t})^{1-\rho_I} + (1 - \mathrm{w}_I)(P_{F,t})^{1-\rho_I}\right]^{\frac{1}{1-\rho_I}}.$$

The capital producing firm at time 0 then maximizes expected discounted profits

$$E_t \sum_{t=0}^{\infty} D_{0,t} \left[Q_t (1 - S(\tilde{X}_t)) I_t - \frac{P_{I,t} I_t}{P_t}\right],$$

which, with $X_t \equiv \frac{I_t}{I_{t-1}}$, results in the first-order condition

$$Q_t(1 - S(\tilde{X}_t) - \tilde{X}_t S'(\tilde{X}_t)) + E_t \left[\frac{1}{(1+R_{t+1})} \; Q_{t+1} S'(\tilde{X}_t) \frac{I_{t+1}^2}{I_t^2} \right] = \frac{P_{I,t}}{P_t}.$$

2.3 | The equilibrium, fiscal policy, and foreign asset accumulation

In equilibrium, goods markets, money markets, and the bond market all clear. Equating the supply and demand of the home consumer goods and assuming that government expenditure, taken as exogenous, goes exclusively on home goods, we obtain[14]

$$Y_t = C_{H,t} + C_{H,t}^e + I_{H,t} + \frac{1-\nu}{\nu} \left(C_{H,t}^* + C_{H,t}^{e\,*} + I_{H,t}^* \right) + G_t.$$

Fiscal policy is rudimentary: a balanced government budget constraint given by

$$P_{H,t} G_t = T_t + M_{H,t} - M_{H,t-1}. \tag{11}$$

Adjustments to taxes, T_t, in response to shocks to government spending away from the steady state are assumed to be non-distortionary.

Let $\sum_{h=1}^{\nu} B_{F,t}(h) = \nu B_{F,t}$ be the net holdings by the household sector of foreign bonds. Summing over the household budget constraints (including entrepreneurs and capital producers), noting that net holdings of domestic bonds are zero (since home bonds are not held by foreign households) and subtracting (11), we arrive at the accumulation of net foreign assets

$$\begin{aligned} P_{B,t}^* S_t B_{F,t} + S_t M_{F,t} &= S_t B_{F,t-1} + S_t M_{F,t-1} + W_t L_t + \Gamma_t + (1-\xi_e) P_t V_t + P_t Q_t (1 - S(X_t)) I_t \\ &\quad - P_t C_t - P_t C_t^e - P_{I,t} I_t - P_{H,t} G_t \\ &\equiv S_t B_{F,t-1} + S_t M_{F,t-1} + TB_t, \end{aligned}$$

where the trade balance, TB_t, is given by the national accounting identity

$$P_{H,t} Y_t = P_t C_t + P_t C_t^e + P_{I,t} I_t + P_{H,t} G_t + TB_t.$$

This completes the model. Given nominal interest rates $R_{n,t}$ and $R_{n,t}^*$, the money supply is fixed by the central banks to accommodate money demand. By Walras' law, we can dispense with the bond market equilibrium condition. Then the equilibrium is defined at $t = 0$ as stochastic sequences C_t, C_t^e, $C_{H,t}$, $C_{F,t}$, $P_{H,t}$, $P_{F,t}$, P_t, $M_{H,t}$, $M_{F,t}$, $B_{H,t}$, $B_{F,t}$, W_t, Y_t, L_t, $P_{H,t}^0$, P_t^I, K_t, I_t, Q_t, V_t, foreign counterparts C_t^*, and so forth, RER_t, and S_t, given the monetary instruments $R_{n,t}$, $R_{n,t}^*$ and exogenous processes.

Finally, we follow Felices and Tuesta (2013) in modeling the small open economy (SOE) by letting its relative size in the world economy $n \to 0$, while retaining its linkages with the rest of the world (ROW). In particular, the demand for exports is modeled in a consistent way that retains its dependence on shocks to the home and ROW economies. From (5), we have that $\mathrm{w} \to \omega$ and $\mathrm{w}^* \to 1$ as $n \to 0$. Similarly, for investment we have $\mathrm{w}_I \to \omega_I$ and $\mathrm{w}_I^* \to 1$ as $n \to 0$. From the viewpoint of the ROW, our SOE becomes invisible, but not vice versa. Exports to and imports from the ROW are modeled explicitly in a way that captures all the interactions between shocks in the ROW and the transmission to the SOE.[15]

2.4 | Monetary policy and the interest rate rule

We assume that the central bank in the emerging market sets the interest rate to respond to deviations of domestic or CPI inflation from a predetermined target (inflation targeting under fully flexible exchange rates). This takes the form of Taylor rule with CPI (or domestic) inflation and output growth targets

$$r_{n,t} = \rho r_{n,t-1} + \theta_\pi(1-\rho)E_t\pi_t + \theta_y(1-\rho)\Delta y_t + \nu_{R,t}, \tag{12}$$

where $\rho \in [0,1]$ is an interest rate smoothing parameter.

For the central bank of the ROW, in order to capture the variations in the interest rates and inflation, we augment the interest rate feedback rule by a time-varying inflation target and incorporate long-run inflation expectations as an observable in the estimation of the US closed-economy model. As specified in Aruoba and Schorfheide (2011), Del Negro and Eusepi (2011), and Del Negro et al. (2015), the interest rate rule is modified as follows

$$r^*_{n,t} = \rho^* r^*_{n,t-1} + \theta^*_\pi(1-\rho^*)(\pi^*_t - \pi^{tar}_t) + \theta^*_y(1-\rho^*)\Delta y^*_t + \nu^*_{R,t} \tag{13}$$

with the time-varying inflation target π^{tar}_t evolving as

$$\pi^{tar}_t = \rho_{\pi^{tar}}\pi^{tar}_{t-1} + \sigma_{\pi^{tar}}\nu_{\pi^{tar},t}.$$

Data on inflation expectations can be used to pin down π^{tar}_t, with the corresponding measurement equation given by

$$\pi^{obs,K}_t = \pi^{tar}_t + 100E_t\left[\frac{1}{K}\sum_{k=1}^{K}\pi^*_{t+k}\right].$$

In our application, we present results for $K = 4$, using data from the Survey of Professional Forecasters (SPF) and the Blue Chip Economic Indicators (BCEI) survey, additional estimations with 10-year expectations data (i.e. $K = 40$) yield similar results.

2.5 | Liability dollarization and financial accelerator: Model variants

The DSGE model makes explicit the operating mechanism of (unofficial) dollarization. According to Elekdag et al. (2005), first, emerging market economies can typically borrow only or partly in foreign currency denominations (liability dollarization). This feature increases the susceptibility to external shocks, since a potential depreciation can substantially inflate debt service costs due to currency mismatches and thus increase rollover risk. In other words, borrowers in these economies may find that both interest and exchange rate fluctuations have large effects on their real net worth positions (recall Equation 10), and so, through balance sheet constraints that affect investment spending, have much more serious macroeconomic consequences than for richer developed economies. Second, these countries usually have imperfect access to capital markets. Foreign credit (or entrepreneurial debt) is typically associated with a risk premium above and beyond the international lending rate, thus borrowers' demand for capital in turn depends on their respective financial positions (e.g. ratio of debt to net worth)—this underpins what is known as the financial accelerator. Moreover, this premium typically moves in tandem with the business cycles of the borrowing country, which implies a higher risk premium when it is experiencing a recession.[16]

We parameterize the model according to four alternatives, ordered by increasing degrees of frictions and dollarization:

- Model 1 (Baseline): This is a fairly standard SOE model similar to many in the NK open-economy literature, with the only nonstandard features being a non-separable utility function in money balances, consumption, and leisure consistent with a BGP and a fully articulated ROW bloc with a policy reaction function augmented by a time-varying inflation target.
- Model 2: Baseline with transaction dollarization (TD) only where the degree of TD is captured by $1 - a \in [0, 1]$.
- Model 3: Financial accelerator (FA) with liability dollarization (LD), assuming that firms borrow a fraction $1 - \varphi \in [0, 1]$ of their financing requirements in dollars.
- Model 4: FA, LD, and TD.

As noted, it is not uncommon to assume that entrepreneurial debt is denominated in foreign currency. Some authors refer to the combination of the financial accelerator channel and liability dollarization as collateral constraints or simply just financial accelerator (e.g. Elekdag et al., 2005). Their respective empirical relevance and properties are tested and studied below.

3 | DATA AND METHODOLOGY

Ten observable variables at quarterly frequency for Peru and the US are used to estimate the model parameters and shocks: real GDP, real private consumption, real private investment, annualized inflation, annualized interest rate (the discount rate), and the real CPI-based exchanged rate. The US GDP, inflation and interest rate are also used in the estimation including our measures of 1- and 10-year ahead inflation expectations based on the SPF and the BCEI survey. The data are obtained from the FRED Database available through the Federal Reserve Bank of St. Louis and the Banco Central de Reserva del Peru (BCRP) Statistics Department. The inflation rate is defined as the annualized log percentage change in the CPI index. The real exchange rate represents key relative prices in an open economy and corresponds to the multilateral real exchange rate index. The sample runs from the first quarter of 1994 to the fourth quarter of 2019 (in Section 7.2, we consider an inflation-targeting-only period starting in 2002).

As noted, we do not make these variables stationary in the data so all the real and nominal variables enter in levels but all the data are demeaned prior to estimation.[17] In addition, another possible caveat of our estimated models is that we assume a conventional monetary policy rule for the entire sample period for the US so that we do not consider the fact that the behavior of capital flows has differed through time. The zero lower bound may invalidate the conventional policy analysis when the key interest rates are near zero. However, this latter challenge can be addressed by the use of the Wu-Xia Shadow Federal Funds Rate (Wu & Xia, 2016) as a measure of the macroeconomic effects of quantitative easing. Therefore, in this estimation, we extend the empirical sample by replacing the US Federal Funds rate with the Wu-Xia Shadow rate from the first quarter of 2009 to cover the zero lower bound period. The vector of observable variables in deviation form that enters in the measurement equation (14) below consists of

$$Y^T = \left\{ y_t^{obs}, c_t^{obs}, i_t^{obs}, r_{n,t}^{obs}, \pi_t^{obs}, rer_t^{obs}, y_t^{*,obs}, r_{n,t}^{*,obs}, \pi_t^{*,obs}, \pi_t^{obs,K} \right\}'.$$

We estimate the model variants by Bayesian methods which entail obtaining the posterior distribution of the model's parameters.[18] Each variant has an associated set of unknown parameters $\theta \in \Theta$ for which we want to characterize the posterior distribution. Model 4, which contains all the frictions, is augmented with 16 orthogonal structural shocks capturing changes in technology, preferences, government expenditure, external finance premium, possible deviations of the UIP conditions, cost-push factors, foreign factors, and the time-varying inflation target that follow a stationary AR(1), such that $\vec{\epsilon}_t = \vec{\rho}\vec{\epsilon}_{t-1} + \vec{v}_t$, with $\vec{\epsilon}_t = \left\{ A_t, G_t, \epsilon_{C,t}, \epsilon_{L,t}, \epsilon_{PM,t}, \epsilon_{P,t}, \epsilon_{UIP,t}, A_t^*, G_t^*, \epsilon_{C,t}^*, \epsilon_{L,t}^*, \epsilon_{PM,t}^*, \epsilon_{P,t}^* \; \pi_t^{tar} \right\}$, and two i.i.d. monetary policy shocks $v_{R,t}, v_{R,t}^*$ for (12) and (13), respectively.

The estimation strategy that follows takes two stages: (i) first we estimate the ROW model using US data. This is a closed economy, so estimation can be done separately. We examine the baseline model with FA only for the US; (ii) then we estimate the parameters and shocks in the Peru-ROW model using the information obtained from stage (i), from which the estimated US parameters are used for calibrating stage (ii). Note that, overall, there are more shocks to be estimated than observable variables. To aid in the identification of the open economy effects and the propagation mechanism of foreign disturbances to the domestic economy, we follow Adolfson et al. (2007) and also use US observables in the second stage.

3.1 | One-step approach with a hybrid framework

To bridge the models and the raw data, we implement a hybrid framework where the cyclical fluctuations of the data are seized by the solution of the DSGE model and the noncyclical fluctuations are captured by a reduced form representation. The approach thus links the observables to the model counterparts via a flexible specification which does not require the cyclical component to be solely located at business cycle frequencies, see Canova (2014), and allows the noncyclical component to take various time series patterns, see Ferroni (2011) and Cantore et al. (2015). Now, we briefly describe the general framework in Ferroni (2011) that we use to set out our measurement equations.

We assume that a vector of times series $y = \{y_t\}_{t=1}^T$ is available and postulate that the data consist of a noncyclical component y_t^τ and a cyclical component y_t^c, so that

$$y_t = y_t^\tau + y_t^c, \tag{14}$$

$$y_t^c = S y_t^\dagger, \tag{15}$$

$$y_{t+1}^\dagger = \Phi(\theta) y_t^\dagger + \Psi(\theta) v_{t+1}, \tag{16}$$

where $y^{\dagger}_{t+1}$ represents the variables in the DSGE model and S is a selection matrix that selects the observable variables; Φ and Ψ are matrices which are functions of the structural parameters of the model, θ; v_{t+1} are mutually uncorrelated innovations of the structural model.

Furthermore, we assume that y^{τ}_{t} is represented by

$$y^{\tau}_{t+1} = y^{\tau}_{t} + \mu_t + \epsilon^{(1)}_{t+1}\ ; \quad \mu_{t+1} = \mu_t + \epsilon^{(2)}_{t+1}, \tag{17}$$

where $\epsilon^{(j)}_{t+1} \sim N(0, \Sigma_j)$, and Σ_j is diagonal for $j = 1, 2$. The setup is very flexible and able to capture various low frequency movement of the data. If $\Sigma_j = 0$ for $j = 1, 2$, $y^{\tau}_{0} = 0$ and $\mu_{t+1} = \mu_t = \mu$ for all t, then the specification becomes a deterministic time trend, that is, $y^{\tau}_{t} = \mu t$. If $\Sigma_2 = 0$ and $\mu_{t+1} = \mu_t = \mu$ for all t, then y^{τ}_{t+1} has a unit root with drift, that is, $y^{\tau}_{t+1} = \mu + y^{\tau}_{t} + \epsilon^{(1)}_{t+1}$. If we leave parameters unrestricted, the trend displays a smooth stochastic trend. Hence, (17) nests, as special cases, the structures which are typically thought to motivate the use of filters.

Given (14)–(17), we let the data endogenously select the specification for the noncyclical component which is more appropriate for each series and this is done jointly with the estimation of the structural parameters θ. There are several advantages of such an approach. First, there is no need to build noncyclical components directly into a DSGE model nor to worry about their exact time series features. Second, the procedure eliminates by construction the first source of measurement error and considerably reduces the second because the spectrum of the data is endogenously split into cyclical and noncyclical parts.

It is important to stress that the traditional approaches are problematic, since they typically create a mismatch between the model and the data. If we perform some transformation in the data to account for nonstationarity, it would be appropriate to have the same transformation also imposed in the model. According to Ferroni (2011), the widespread practice of specifying a unit root process with a drift in technology and considering co-integrated variables in estimation is unsatisfactory for at least two reasons. First, while "great ratios" are stationary in the model, the stationarity of nominal or real great ratios in the data is dubious (see Canova, 2014). Second, a model-based transformation requires the exact knowledge of the shocks that drive noncyclical fluctuations of the data. Absent such knowledge, misspecifications of all sorts may be present, which may bias estimates. While the other feasible route is to prefilter the data, as is currently standard, the choice of the statistical filter is arbitrary and does affect the structural estimation.

3.2 | Calibrated parameters and priors

In order to implement Bayesian estimation, prior distributions must be defined for the parameters θ, structural shocks ϵ_t and the exogenous shocks associated with the trend, $\epsilon^{(j)}_{t+1}$. First, a number of the structural parameters from both the closed and open economy models are kept fixed in the estimation procedure. These parameters can often be related to the steady state values of the observed variables in the model and are, therefore, calibrated so as to match their sample mean. Table 2 reports the calibrated parameters along with the key implied steady state relationships based on previous studies for the Peruvian (and the US) economies.[19]

For the remaining parameters, we use normal distributions as priors for the unbounded parameters when more informative priors seem to be necessary, and beta distributions are used for all parameters bounded between 0 and 1, that is, fractions or probabilities. We use

TABLE 2 Calibrated parameters

Calibrated parameter	SOE		ROW	
Discount factor	β	0.99	β^*	0.99
Depreciation rate	δ	0.025	δ^*	0.025
World growth rate	g	0.03		
Capital share in production	α	0.5	α^*	0.33
Substitution elasticity (H/F goods)			μ^*	1.5
Working day	L	0.5	L^*	0.4
Fixed cost	ϕ_F	1.2	ϕ_F^*	1.2
Trade shares				
Imported investment share	is_{import}	0.15		
Imported consumption share	cs_{import}	0.1		
Exported investment share	is_{export}	0.02		
Exported consumption share	cs_{export}	0.23		
Implied steady state relationships				
Consumption-output ratio	c_y	0.7	c_y^*	0.6
Investment-output ratio	i_y	0.15	i_y^*	0.2
Government expenditure-output ratio	g_y	0.15	g_y^*	0.2
Export/import-output ratio		0.25		

Note: Trade shares are derived as follows: Total exports and imports are around 25% for Peru so $0.25 = cs_{imports} + is_{imports} = cs_{exports} + is_{exports}$ for balanced trade. Data on consumption and capital goods exports show $\frac{is_{imports}}{cs_{imports}} = 1.6$ and $\frac{is_{exports}}{cs_{exports}} = 0.1$.

inverse gamma distributions as priors when non-negativity constraints are necessary. All priors are assumed to be the same across specifications. As far as possible, parameter values are chosen based on quarterly data for Peru. Elsewhere the parameters and prior standard deviations reflect broad characteristics of emerging economies. A variety of sources are used: for Peru we draw upon Castillo et al. (2013) (CMT). For emerging economies more generally and for parameters related to the financial accelerator we use Gertler et al. (2007) (GGN) and Bernanke et al. (1999) (BGG). The rest of the world is represented by US data. Here we draw upon Smets and Wouters (2007) (SW).

For estimating the US closed economy model, the priors for the nonfinancial parameters are chosen in line with those in SW (see Table S1 in Online Appendix E). Our prior choices for the parameters, Θ^*, χ_θ^*, n_k^*, ξ_e^*, are derived from earlier studies conducted for the financial accelerator, notably by BGG. Turning to the prior setting for the Peru-ROW model, as mentioned before, we have little information in addition to the Peruvian data, hence we impose more prior uncertainty to the parameters and center the prior distribution to the values suggested by CMT, GGN, and the BCRP. The whole ROW bloc is calibrated based on the US estimation from the previous stage. Table 3 shows the assumptions for the prior distribution of the estimated structural parameters and shocks in the Peru-ROW model. The location of the prior distribution corresponds, to a large extent, to a number of theoretical and empirical studies for both closed and open economies.[20] Details of the priors are also appended to the article in Online Appendix D.

TABLE 3 Prior and posterior distributions: Peru-ROW 2-bloc model

		Prior distribution			Posterior distribution			
Parameter	**Notation**	**Density**	**Mean**	**SD/df**	**Model 1**	**Model 2**	**Model 3**	**Model 4**
Investment adjustment	$S''(1+g)$	N	4.00	1.50	3.1228 [1.0144:4.9983]	3.7527 [1.6502:5.7933]	0.8319 [0.4652:1.1964]	1.7101 [0.8727:2.6532]
Risk aversion	σ	N	1.50	0.50	1.7047 [0.6900:2.6515]	1.9059 [1.1873:2.5570]	1.7222 [0.9220:2.3842]	1.2004 [0.7308:1.6503]
Consumption habit	h_C	B	0.50	0.20	0.9463 [0.8749:0.9985]	0.9042 [0.8083:0.9942]	0.8020 [0.6601:0.9316]	0.4909 [0.2472:0.7223]
Calvo prices	ξ_H	B	0.50	0.10	0.7250 [0.5908:0.8748]	0.7898 [0.6543:0.9299]	0.5334 [0.3570:0.7536]	0.6876 [0.5461:0.8087]
Substitution elasticity (balances in H/F currencies)	χ_m	N	4.00	1.50	–	3.8766 [1.5374:6.1336]	–	4.6791 [2.6604:6.6106]
Fraction of real balances held in H currency	a	B	0.50	0.20	–	0.4232 [0.1408:0.7350]	–	0.2155 [0.0825:0.3428]
Fraction of borrowing in H currency	φ	B	0.50	0.20	–	–	0.8767 [0.7698:0.9850]	0.7775 [0.5364:0.9838]
External finance premium	Θ	IG	0.01	4.00	–	–	0.0060 [0.0024:0.0097]	0.0053 [0.0023:0.0081]
External finance premium elasticity	χ_θ	IG	0.05	4.00	–	–	0.0234 [0.0134:0.0336]	0.0257 [0.0140:0.0367]
Inverse of leverage	n_k	B	0.40	0.10	–	–	0.2818 [0.1640:0.3965]	0.3839 [0.2463:0.5315]
Entrepreneurs survival rate	ξ_e	B	0.90	0.05	–	–	0.9591 [0.9385:0.9803]	0.9777 [0.9605:0.9937]
Proportion of RT consumption	λ_{C_1}	B	0.25	0.10	0.2189 [0.1553:0.2716]	0.1888 [0.1260:0.2400]	0.3017 [0.1880:0.4382]	0.1936 [0.1107:0.2564]

TABLE 3 (Continued)

		Prior distribution			Posterior distribution			
Parameter	**Notation**	**Density**	**Mean**	**SD/df**	**Model 1**	**Model 2**	**Model 3**	**Model 4**
Policy rule: inflation	θ_π	N	1.50	0.50	1.6896 [1.3381:2.0618]	1.7913 [0.1194:0.3643]	2.9887 [2.4215:3.5116]	2.6773 [2.1773:3.2122]
Policy rule: output	θ_y	N	0.125	0.10	0.2282 [0.1095:0.3495]	0.2440 [1.5048:2.1079]	0.2626 [0.1427:0.3821]	0.2944 [0.1764:0.4014]
Policy rule: interest rate smoothing	ρ	B	0.50	0.20	0.6631 [0.5208:0.8155]	0.7088 [0.6142:0.8008]	0.4172 [0.2477:0.5837]	0.4560 [0.3387:0.5936]
AR(1) coefficient								
Technology	ρ_A	B	0.50	0.20	0.9518 [0.9197:0.9880]	0.9350 [0.8842:0.9961]	0.5715 [0.3690:0.7708]	0.5758 [0.2761:0.9979]
Preferences	ρ_C	B	0.50	0.20	0.5294 [0.2027:0.8629]	0.4845 [0.1468:0.8124]	0.4978 [0.1689:0.8225]	0.3569 [0.1073:0.5890]
Government spending	ρ_G	B	0.50	0.20	0.4808 [0.1421:0.7924]	0.5001 [0.1863:0.8280]	0.4768 [0.1520:0.8104]	0.4841 [0.1320:0.7960]
Labor supply	ρ_L	B	0.50	0.20	0.5315 [0.1999:0.8754]	0.5025 [0.1682:0.8196]	0.5011 [0.1716:0.8227]	0.4953 [0.1739:0.8161]
External finance premium	ρ_P	B	0.50	0.20	–	–	0.3422 [0.1171:0.5476]	0.4051 [0.1080:0.7942]
Price markup	ρ_{PM}	B	0.50	0.20	0.6893 [0.5717:0.8182]	0.6271 [0.4066:0.8070]	0.5965 [0.3326:0.8051]	0.6805 [0.5623:0.8027]
UIP	ρ_{UIP}	B	0.50	0.20	0.6193 [0.3742:0.8631]	0.5960 [0.3338:0.8521]	0.6278 [0.3435:0.9481]	0.4950 [0.1662:0.8019]
Standard deviation of AR(1) innovations/i.i.d. shocks								
Technology	σ_{v_A}	IG	0.40	2.00	2.0803 [1.5396:2.5797]	2.0915 [1.6059:2.5984]	1.0235 [0.5763:1.6199]	0.5026 [0.1100:1.1638]

TABLE 3 (Continued)

		Prior distribution			Posterior distribution			
Parameter	**Notation**	**Density**	**Mean**	**SD/df**	**Model 1**	**Model 2**	**Model 3**	**Model 4**
Preferences	σ_{v_C}	IG	0.40	2.00	0.3293 [0.0969:0.6391]	0.3257 [0.0993:0.5613]	0.3041 [0.1011:0.5250]	1.2722 [0.1599:1.9879]
Government spending	σ_{v_G}	IG	0.40	2.00	0.2795 [0.1091:0.4578]	0.3447 [0.1106:0.6483]	1.1763 [0.0902:5.0812]	0.3548 [0.1024:0.6582]
Labor supply	σ_{v_L}	IG	0.40	2.00	0.2828 [0.0980:0.4971]	0.3141 [0.0974:0.5993]	0.3716 [0.1013:0.7365]	0.3752 [0.0978:0.6529]
External finance premium	σ_{v_P}	IG	0.40	2.00	–	–	2.0731 [0.8960:3.1762]	3.6718 [1.0567:6.9341]
Price markup	$\sigma_{v_{PM}}$	IG	0.40	2.00	1.5824 [0.5878:2.5046]	1.2270 [0.5265:2.1153]	3.5030 [1.0619:5.7529]	1.8569 [0.7384:3.2547]
Monetary policy	σ_{v_R}	IG	0.40	2.00	0.2083 [0.1025:0.3040]	0.1929 [0.1126:0.2768]	1.0285 [0.7190:1.3222]	0.9206 [0.7042:1.1346]
UIP	$\sigma_{v_{UIP}}$	IG	0.40	2.00	0.5593 [0.2580:0.8781]	0.5922 [0.2419:0.9601]	0.4148 [0.1520:0.6537]	0.5408 [0.2517:0.8274]
Degree of TD	$1-a$				–	0.5768 [0.2650:0.8592]	–	0.7845 [0.6572:0.9175]
Degree of LD	$1-\varphi$				–	–	0.1233 [0.0150:0.2302]	0.2225 [0.0162:0.4636]
RT consumers	λ				0.4793	0.4592	0.5345	0.4624
Price contract length	$\frac{1}{1-\xi_H}$				3.6364	4.7574	2.1432	3.2010
Log-likelihood					−1154.6284	−1155.1447	−1124.3356	−1113.8972
Probability					0	0	0	1

Note: λ is derived as follows: $\frac{C_2}{C} = \frac{C_2^*}{C^*} = 1.5$, then $\lambda_{C_1} = \lambda \frac{C_1}{C} = 1 - (1-\lambda)\frac{C_2}{C}$.
Abbreviations: B, beta; IG, inverse gamma; N, normal.

4 | ESTIMATION AND RESULTS

Table S1 in Online Appendix E and Table 3 report the parameter estimates, summarizing posterior means of the estimated parameters and 90% high posterior density (HPD) intervals for all the model specifications, as well as the posterior marginal likelihoods. Overall, parameter estimates are plausible and reasonably robust across specifications. In what follows, we first discuss the parameter estimates. Second, we evaluate the model's empirical fit using the estimated log-likelihoods. Third, we compute the impulse response functions and carry out the moment analysis in order to assess the impact of the key structural shocks and model-implied dynamics. As noted, the ROW estimation is done separately and is reported in Online Appendix E.[21]

4.1 | Peru-ROW 2-bloc open economy

We now turn to the two-country model estimation results with data for Peru and using the US estimates discussed above as calibration. First, we note that the estimated value of the key parameter, χ_θ, the elasticity of the external finance premium with respect to leverage is 0.03 [0.01:0.04], which is consistent with the value range often used to calibrate this parameter (e.g. Bernanke et al., 1999). Note that the estimates across the two specifications are statistically away from zero, suggesting that the financial accelerator is playing an important role in the Peruvian economy. The estimated financial accelerator risk premium is around 1% per quarter, implying an annualized endogenous risk premium of 4%, which is again a plausible estimate and is in line with the data (Gertler et al., 2007). Our models estimate the leverage ratio, measured by n_k, at around 0.28–0.38. A relatively high debt-equity ratio (2.6–3.6) indicates that the financial markets in Peru are somewhat less developed compared with those of developed economies. The entrepreneurs' exit rate is 0.02–0.04 which is very close to the value suggested by Gertler et al. (2007).

The estimation results show that the degree of transaction dollarization is much higher than the degree of liability dollarization regardless of the model assumptions. Estimates of TD are very significant, although the posterior means are somewhat high. Note, however, that the HPD intervals suggest a range between 0.26 and 0.91—that is, well within the range of those reported in Castillo et al. (2013) (which does not model financial dollarization, though). Moreover, while the overall degree of dollarization has recently decreased, from Table 1 we can see that, over that period, the average share of foreign currency in deposits and loans in Peru is around 50%, so in this context, the estimates are not wholly unreasonable.[22]

Our models suggest that the Peruvian domestic firms are required to finance about 12%–22% of capital requirements in dollars. The estimation is robust to alternative modeling assumptions.[23] Interestingly, the model with only TD (Model 2) delivers a slightly larger inflation response (1.79) compared with Model 1 (1.67). Models 3 and 4 with LD suggest a much more aggressive inflation response compared with the other models. This implies the reaction of the central bank to offset the effects of dollarization on the volatility of consumption.

We find that the estimated BCRP reaction function shows strong responses to inflation and the responses are more aggressive when financial frictions and dollarization are included. Evidence from all four models shows moderate feedback of the central bank to the growth rate of output and this can be rationalized. This may be explained by the fact that the Peruvian government has implemented a monetary regime from which an inflation targeting policy has been pursued. Comparing to our prior, the estimates of ρ show a higher degree of interest rate smoothing across the first two model variants (0.66–0.71). As suggested by Christensen and Dib (2008), which estimates a New Keynesian model incorporating the financial accelerator for the US, the monetary authority should respond more aggressively to changes in output to stabilize the economy than it would if there were no financial accelerator, because the presence of the financial accelerator leads to an amplification and propagation of the impacts of the shocks on output, since investment is more volatile. Indeed, our results show some small and gradual increases from the estimated reactions to output growth from Models 1 to 4. The increase, ranging from 0.23 to 0.29, is relatively small compared with the inflation responses. This, again, can be explained by the fact that the BCRP is inflation averse.

Moving to the preference parameters, the estimates of risk-aversion (σ) indicate that the intertemporal elasticity of substitution in consumption (proportional to $1/\sigma$) is less than one although our estimates are slightly below those usually found in earlier empirical studies. The posterior means of the estimated price stickiness parameter are relatively high in comparison to those estimated for the developed economies. These values imply that the average price contract length is over 2–4 quarters, which is broadly consistent with the findings from recent empirical research conducted for emerging market economies (e.g. Castillo et al., 2013; Caputo et al., 2007). The estimated values of ξ_H tend to be lower in the models that incorporate liability dollarization—a finding that is consistent with that in Castillo et al. (2013).

Our estimation delivers, based on the posterior estimates of λ_{C_1}, high shares of liquidity constrained consumers in the Peruvian economy relative to the ROW. This is not surprising as it shows that 46%–53% of the households is liquidity constrained, which is consistent with our prior assumption.[24] These households do not trade on asset markets and consume entirely their disposable income each period. This result potentially has important implications for fiscal policymaking in Peru.[25]

In terms of the persistence of the exogenous shocks, the estimates of the AR(1) coefficients show that most shocks are moderately inertial, the markup shock (ρ_{PM}) being the most persistent shock in all models. The technology shock is far more persistent in the models without the financial accelerator. Moreover, according to their estimated standard deviations, the markup and technology shocks stand out as being the most volatile shocks in the home economy in the absence of the FA mechanism. As for the model with the financial accelerator and financial dollarization, the external finance premium shock is by far the most volatile shock, followed by the shock to domestic price markup.

Overall, most of the posterior estimates are statistically significant and reasonable. More importantly, the Peru-ROW estimation seems to be able to deliver plausible estimates of the financial accelerator risk premium and elasticity of the external finance premium which are tightly estimated away from zero. Thus, our estimates support the inclusion of the financial accelerator mechanism.

4.2 | Bayes factor comparison

One of the great advantages of adopting a Bayesian approach is that it facilitates a formal comparison of different models through their posterior marginal likelihoods, computed using the Geweke (1999) modified harmonic-mean estimator. To compare models, we calculate the posterior odds ratio, which is the ratio of their posterior model probabilities (or Bayes Factor when the prior odds ratio is set to unity) in terms of the log-likelihood. Such comparisons are important in the assessment of rival models, as the model which attains the highest odds outperforms its rivals and is therefore favored. Another great advantage, according to Ferroni (2011), is that we can directly compare which noncyclical component specification best fits the data by using the posterior odds ratio.

The central question of interest for the empirical analysis now is which mechanism(s) help in improving the model fit to the data. Table 3 reports the posterior log marginal likelihood (LL) of the four competing model variants. Model 4 clearly wins the likelihood ranking. In other words, it indicates that the models with both financial accelerator and partial dollarization clearly appear to outperform their rivals, implying that both frictions are empirically relevant and strongly improve the fit to the Peruvian data. In Model 4, LD is included through the financial accelerator assuming that domestic firms use foreign currency to finance working capital. In the absence of the financial accelerator (Models 1 and 2), the data do not favor the model with some degree of transaction dollarization over the one with no dollarization. This comparison result using our model assuming RT households that is estimated directly based on a much longer sample with nonstationary time series does not deliver the findings in Castillo et al. (2013) that the transaction form of dollarization is important to explain Peruvian data.[26]

The differences in log marginal likelihoods or the posterior odds ratio (Bayes factors) are important as decisive evidence in favoring one model over others. The LL difference between Model 1 and Model 2 is 0.5163. As suggested by Kass and Raftery (1995), in order to choose Model 1 over Model 2, it needs a prior probability over Model 1 1.6758 ($\approx e^{0.5163}$) times larger than the prior probability over Model 2.[27] Thus, we cannot confirm that Model 1 statistically outperforms Model 2 in terms of the ability to explain the data. The goodness of fit of the baseline model improves significantly by adding the remaining frictions (FA and LD), eventually resulting in a large LL difference of 40.7312 between Models 1 and 4, although assuming no TD deteriorates the marginal likelihood slightly as is evident from Model 3. The posterior probabilities also confirm our results.

Moreover, we estimate and compare the three filtering configurations with real data in addition to our parameter estimation using the flexible specification as described in Section 3.1. In particular, we examine separately the following filters incorporated into the state-space representation of the preferred Model 4 for the trend component (Ferroni, 2011) and apply the same estimation procedure set out above for each of them

$$\begin{aligned}
\text{Linear trend (LT)}: &\quad y^{\tau}_{t+1} = A + Bt + \epsilon^{lt}_{t+1}, \\
\text{First differences (FD)}: &\quad y^{\tau}_{t+1} = \gamma + \Gamma y^{\tau}_{t} + \epsilon^{fd}_{t+1}, \\
\text{Hodrick-Prescott (HP)}: &\quad y^{\tau}_{t+1} = y^{\tau}_{t} + \mu_t \;; \quad \mu_{t+1} = \mu_t + \epsilon^{hp}_{t+1}.
\end{aligned}$$

As expected, using different filters prior to estimation induces large differences in the estimated posterior likelihoods: LT: −2179.8885; FD: −1127.0345; HP: −1257.9273, and indeed each data transformation procedure determines statistically different estimates. The FD filter, extracting the noisy trend components, is preferred by the posterior density of data and the LT specification that adjusts the level data by a deterministic linear trend results in the lowest probability. The important message here is that, as Table 3 reveals, the use of the flexible specification that can freely choose the individual trend component clearly has the great advantage over any of the three specific data transformations. Our finding that the trend agnostic procedure produces a significantly better fit than the traditional approaches supports the use of Ferroni's (2011) approach for estimating SOE DSGE models.

4.3 | Assessing model fit

The next exercise of our quantitative analysis is to further assess the contribution of the models' frictions by computing the theoretical moments implied by the estimated models and comparing them with the empirical ones in the actual data, which we observe as the target variables in the estimation.[28] This RBC-type moment-validating exercise conducted in Table 4 is not only useful to gain intuition behind the LL comparison results, but also important to address how well the model does in explaining some stylized facts of the emerging economy.

Models 3 and 4 seem to fit better the data in terms of implied volatilities of real investment, the real exchange rate and inflation, getting closer to the data in this dimension. Our model variants with FA can successfully replicate the stylized fact that investment is more volatile than output whereas Models 1 and 2 fall considerably short of capturing that. By providing a feedback mechanism from the shocks to net worth, the financial accelerator increases the volatilities of investment, lending support to the findings about the importance of credit market imperfections when modeling emerging economies. Inflation volatility is practically lower owing to the lower responsiveness (and volatility) of the interest rate. This is consistent with our estimated Taylor rules. It is interesting to note that all the models predict excessive volatility in consumption, a result that replicates the finding in Garcia-Cicco et al. (2010) and Chang and Fernandez (2013) for emerging open economies.

To further evaluate the relative merits of alternative models, Table 4 reports the cross-correlations of the observable variables vis-à-vis output and Figure 1 displays the autocorrelation functions. Our models perform poorly in capturing the countercyclicality of inflation, generating the wrong sign. The countercyclicality of the exchange rate and interest rate is consistent with the data for Peru. When it comes to matching the autocorrelograms, the evidence is now much clearer, with the encompassing Model 4 generally fitting better the dynamics seen

TABLE 4 Data and model moments

Observable	$dlGDP_t$	$dlCON_t$	$dlINV_t$	RER_t	INT_t	$dlCPI_t$
Standard deviation (in %)						
Data	2.8916	3.0592	5.7438	1.9833	1.8099	0.7956
Model 1	2.8651	3.0968	2.8249	3.2598	1.6938	1.2131
Model 2	2.8437	2.9646	2.8389	3.3646	1.6754	1.2243
Model 3	2.5079	3.2615	5.9611	1.5198	1.0998	0.7752
Model 4	2.6658	2.9808	6.7785	1.7734	1.1493	0.7618
Cross-correlation with $dlGDP_t$						
Data	1.0000	0.8565	0.7751	−0.4068	−0.2413	−0.2960
Model 1	1.0000	0.7841	0.6881	−0.1555	−0.2959	0.2096
Model 2	1.0000	0.7632	0.7013	−0.1890	−0.4207	0.2344
Model 3	1.0000	0.9080	0.5574	−0.0176	0.0132	0.0714
Model 4	1.0000	0.9397	0.5141	−0.0610	−0.0083	0.0718

Note: The sample is 1994:1–2019:4. Theoretical moments implied by the estimation are compared with those in the actual data. These model implied statistics are computed by simulating the models at the posterior means using the parameters reported in Table 3. For the observables in this comparison exercise, we adjust the raw data, take natural logarithms and first differences of output, consumption, investment and the CPI index, respectively.

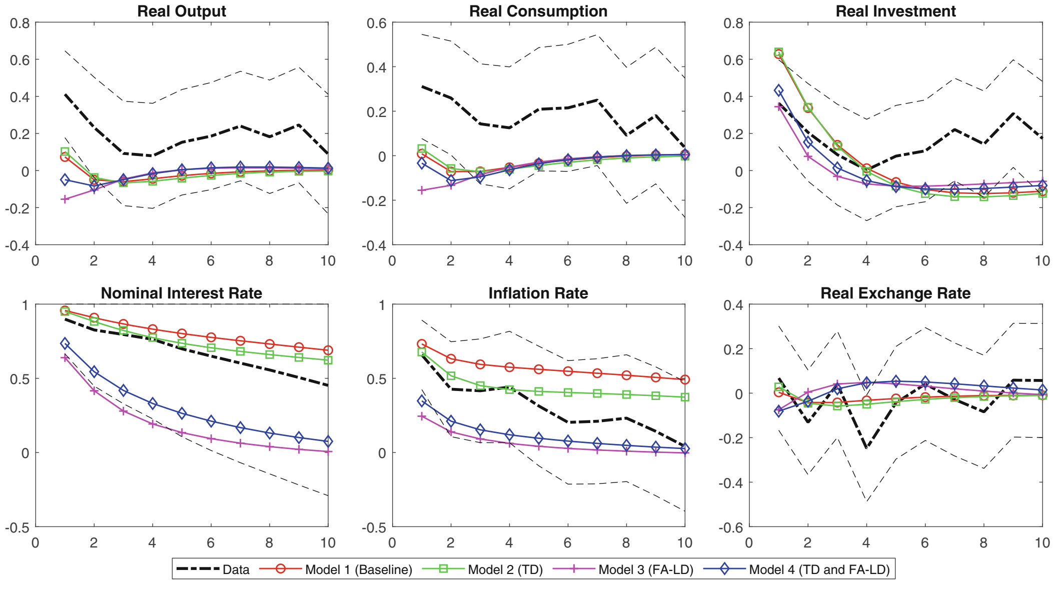


FIGURE 1 Autocorrelations of observables in the actual data and in the estimated models. The approximate 95% confidence bands are constructed using the large-lag standard errors (see Anderson, 1976). [Colour figure can be viewed at wileyonlinelibrary.com]

in the data, especially for higher order autocorrelations. Of particular interest is that the implied autocorrelograms produced by Model 4 match well the observed autocorrelation of investment and inflation while the baseline model generates too much sluggishness and is less able to match the inflation autocorrelation observed from the second lag onwards. Most of our estimated models capture the empirical autocovariance structure reasonably well, with most of the autocorrelograms lying inside the 95% uncertainty bands.[29]

5 | POSTERIOR IMPULSE RESPONSE ANALYSIS

In this section, we study the estimated posterior impulse responses (henceforth IRFs) for four selected shocks: a domestic technology shock ($v_{A,t}$), a shock to the country's external risk premium ($v_{UIP,t}$), and domestic monetary and fiscal policy shocks (v_R and v_G, respectively). The aim of this exercise is two-fold. First, we are interested in comparing the transmission of the key internal and external shocks when the accelerator mechanism is "turned on" and "turned off." This way, we assess the impact of imposing the financial frictions on different model dynamics. Second, we investigate the importance of shocks to the endogenous variables of interests in order to gain a better understanding of the innovation and forecasting uncertainties and, thus, the model uncertainties faced by policymakers.

To focus the presentation, this exercise is only performed for Model 4 and Model 1, that is, the preferred encompassing model vis-à-vis the baseline model. Figures 2 and 3 depict the mean responses corresponding to a positive one standard deviation of the shocks' innovations. Each response is for a 20 period (5 years) horizon and is level deviation of a variable from its steady-state value. We also report the IRFs of two counterfactual models in which the setup of RT households is taken out of the estimated models to examine the liquidity effect in details (Figure S1 in Online Appendix F) and the degree of LD is switched off in the net worth position (Section 5.3).

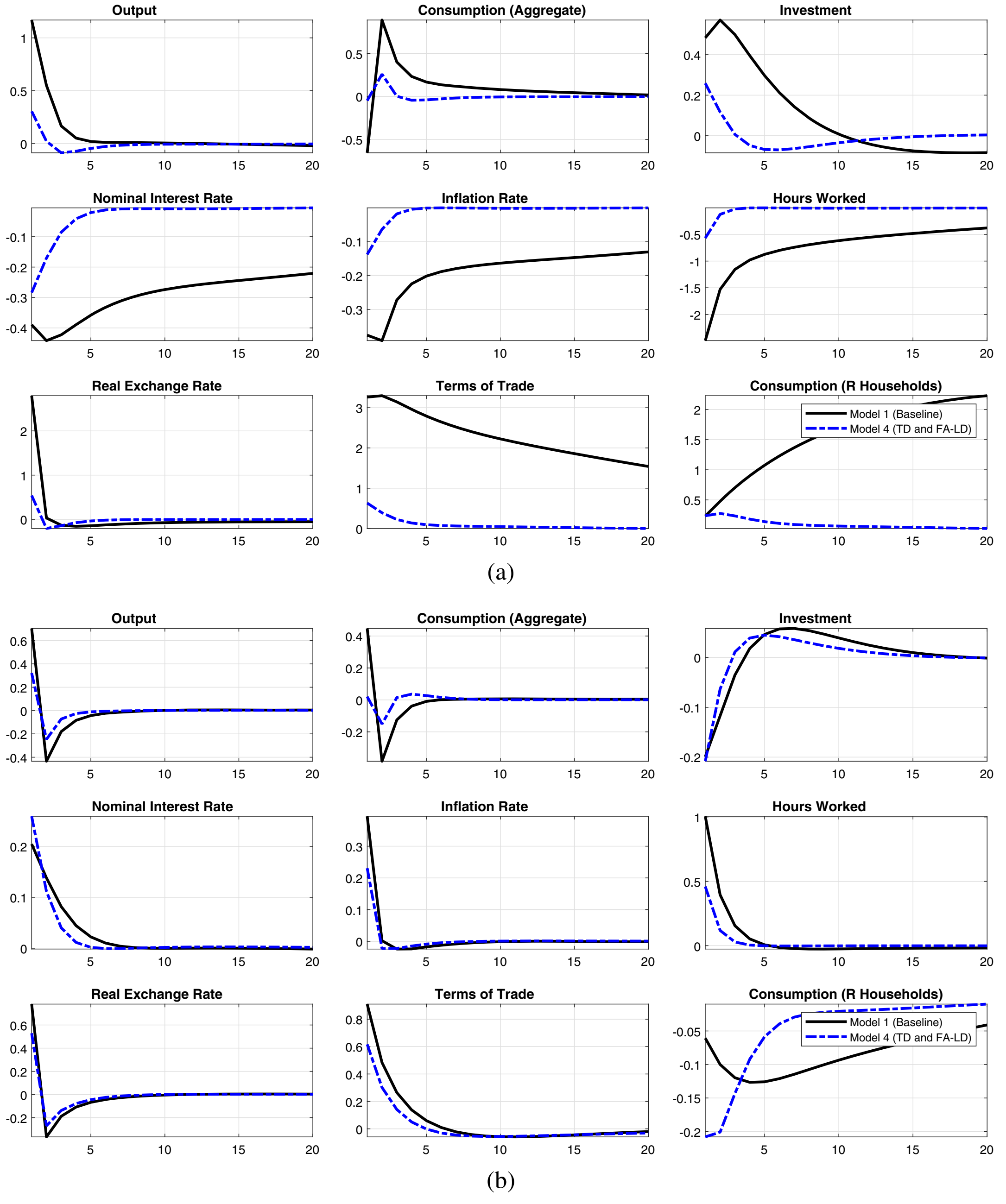
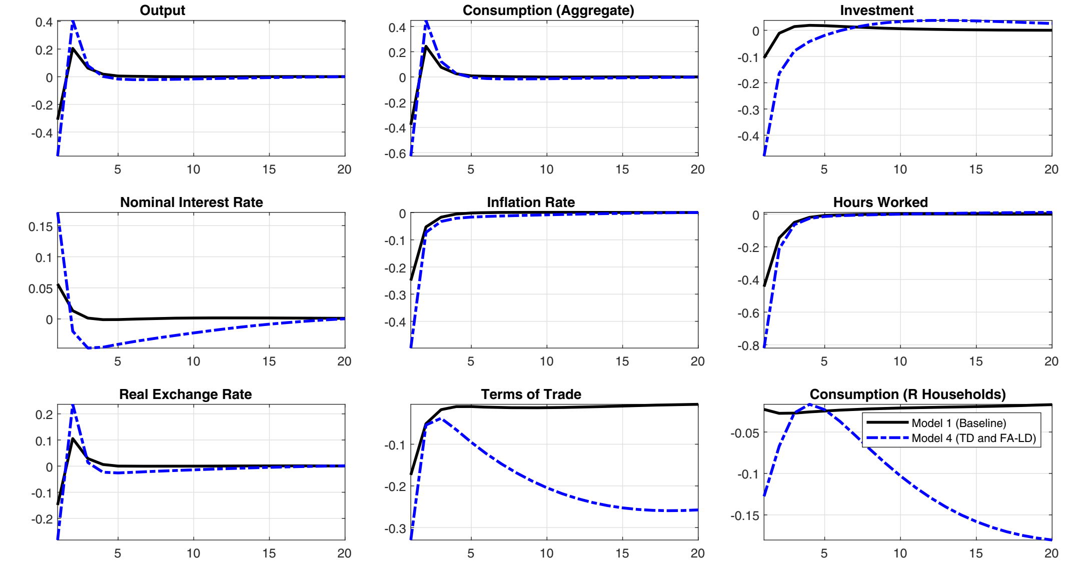


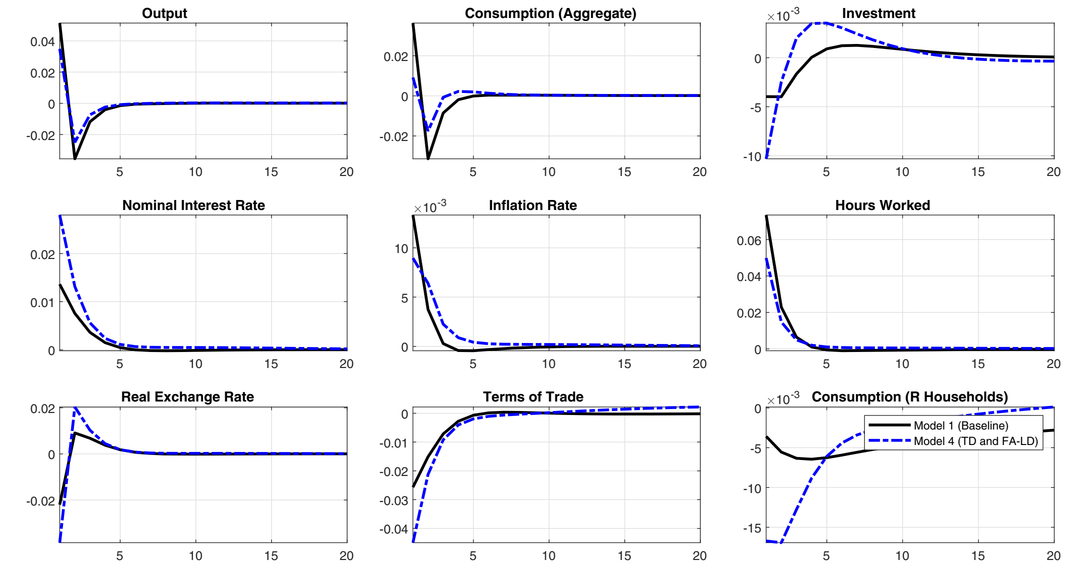
FIGURE 2 Impulse response functions: (A) Technology shock and (B) risk premium shock [Colour figure can be viewed at wileyonlinelibrary.com]

5.1 | Technology and risk premium shocks

For Model 4, a positive technology shock has the usual positive impact on output, consumption and investment and implies an immediate fall in inflation and the nominal interest rate. Nevertheless, the effect dies out relatively rapidly (about 1 year) when affecting most of the variables. This shock becomes fairly persistent when affecting investment. Output and consumption



(a)



(b)

FIGURE 3 Impulse response functions: (A) Monetary policy shock and (B) government spending shock [Colour figure can be viewed at wileyonlinelibrary.com]

expand following the shock, with the latter jumping upwards in response to the lower expected real interest rate, so demand for consumption rises in line with supply.

A positive technology shock in Peru acts as a labor demand shifter: higher productivity shrinks labor demand, pushing marginal cost down, lowers prices and interest rates, but its effect on prices is not persistent as inflation returns to preshock levels after less than one year (about 2 quarters). When all firms experience a decline in their marginal cost as a result of a shock in technology, they will adjust prices downwards only partially in the short run. In addition, a technology shock also depreciates the currency in response to the lowering of the interest rate, thereby increasing the terms of trade, which improves goods' competitiveness.

However, there are marked differences between (frictionless) Model 1 and Model 4 with financial frictions. In Figure 2, we also see an immediate fall in (aggregate) consumption following the technology shock from Model 1. This can be explained by the labor market and the behavior of the real wage with this less formal economy. To show how this works, we simulate our estimated models by assuming that the technology shock has a unit standard deviation and fixed AR(1) parameter, and conduct a counterfactual exercise with only the Ricardian consumers (assuming $\lambda = 0$) in Figure S1 in Online Appendix F. This analysis, by creating a "level playing field" for assessing the effect from a technology shock, is also part of a robustness check reported in Online Appendix F.

With sticky prices, the increase in demand for output is less than the 1% technology increase, so the demand for labor falls shifting the demand curve in (real-wage, quantity) space inwards. This puts an initial downward pressure on the real wage. With liquidity constrained households (with a relatively larger fraction of population: 48% in Model 1) consuming out of their wage income in these models, this means that consumption and output are initially depressed (see Figure S1 in Online Appendix F), shifting the demand curve further inwards due to this liquidity effect, even though the shock boosts consumption and output for the Ricardian households.

On the supply-side, we see that the reduction in hours (that subsequently raises the real wage) is very short-lived as the liquidity effect dominates. The model where dollarization is absent displays a greater reaction on these two aggregate variables (Model 1) but this initial downward shift disappears when $\lambda = 0$. The ensuing negative effect on investment (due to the initial demand deficiency) can be significantly magnified and more persistent because of the FA mechanism.[30]

In Figure 2, we evaluate the responses from a one standard deviation increase in the domestic country external risk premium (UIP). As before, most responses are consistent with the findings of Bernanke et al. (1999) and Batini et al. (2010), using US-calibrated models. In particular, and in contrast with the responses from the technology shock, the models now predict that a positive risk premium shock depreciates the real exchange rate. As expected, there is an immediate drop in investment because the increase in the cost of capital drives Tobin's Q down. The nominal interest rate goes up on impact. Indeed, it is intuitive to observe an interest rate increase when there is an increase in the external risk premium.

Almost all the responses are short-lived, as suggested by the parameter's estimate. It produces only a relatively modest impact on output following a UIP and technology shock and this contradicts the findings from Gertler et al.'s (2007) model simulation for the US. Nonetheless, the main results are generally in line with the findings in Gabriel et al. (2016). The key messages are that (i) there is a long-stabilizing effect as investment and output are "accelerated" faster back to potential (this explains a further balance sheet effect that brings investment and hence output faster back to equilibrium); (ii) the mild output response suggests that there may be off-setting effects among the individual output components, coming from various frictions (e.g. a delayed response of consumption as is evident in Figure 2).

5.2 | Monetary and fiscal shocks

Figure 3a reports the impact of an unanticipated increase in the nominal interest rate. Modeling financial frictions in Peru does lead to a much stronger (and more persistent) response of real and nominal variables (Model 4). There are clear amplifications of all the responses in the presence of the financial accelerator and dollarization. The amplification seems to be far more substantial for the investment response (over 5 times stronger).

Further, the persistence of the real effects is substantially greater in response to domestic contractionary policy in Model 4. When there is an unanticipated increase in the cost of capital, the demand for capital is depressed, which in turn reduces investment and the price of capital. The decline in asset prices pushes net worth down, forcing up the external finance premium and this, as a result, helps to further lower private investment and consequently the demand for output. The negative impact on employment becomes larger too. The strong linkages modeled between exchange rate changes and financial distress of firms suggest that these effects are likely to be more substantial given the firm's external borrowing requirements and domestic currency appreciation in response to the shock. Hence, our estimation predicts that monetary policy would be far more effective if FA and dollarization are absent and this also depends on the liquidity effect (λ).

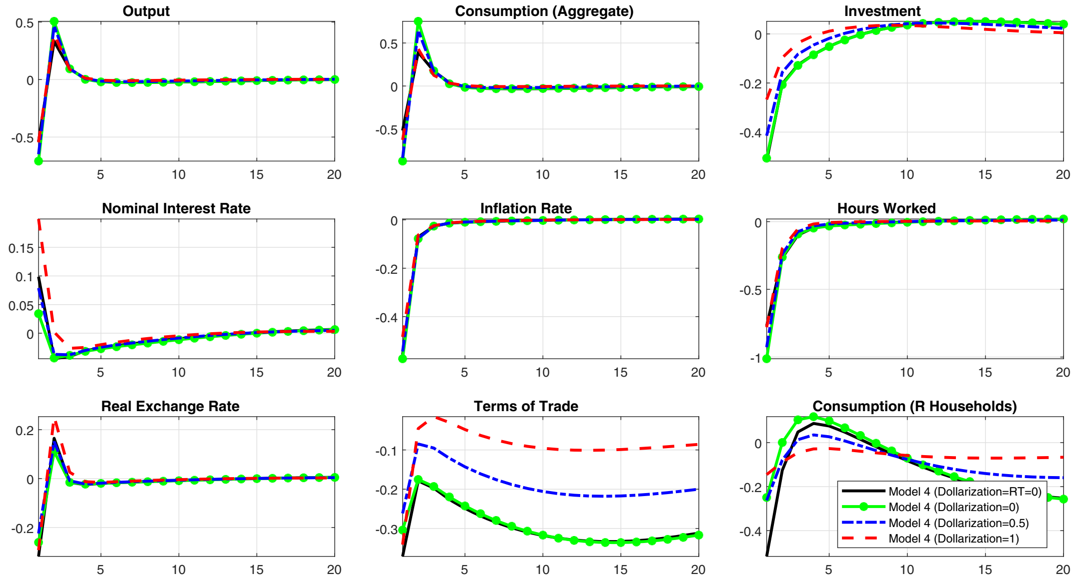
The effect on output of a positive shock to government expenditure is reported in the second panel (Figure 3). It shows a similar impact on the real activities that our financial frictions magnify and propagate the demand shift. It is interesting to observe that, again, the impact and propagation effects on investment are stronger. In other words, it is evident that investment crowding-out is mitigated by the investment boom associated with the rise in asset prices (the price level) and the demand shift, which raises net worth and thus reduces the external finance premium. A reason for explaining the magnitude and persistence of the effects is because the shock has a much smaller effect on the interest rate and the real exchange rate in Model 4.

As before, there are now marked differences between the baseline model and Model 4 and this depends on the liquidity effect. Consider the labor market again: the increase in demand for output leads to a corresponding increase in labor supply and demand, putting an upward pressure on the real wage. The liquidity effect means that this increases the aggregate consumption further and the increase in hours is very short-lived. As before, the initial fiscal stimulus would result in a sharp fall in consumption of Ricardian consumers. It is worth noting that fiscal policy would again be more effective in stabilizing the real economy if these frictions are absent.

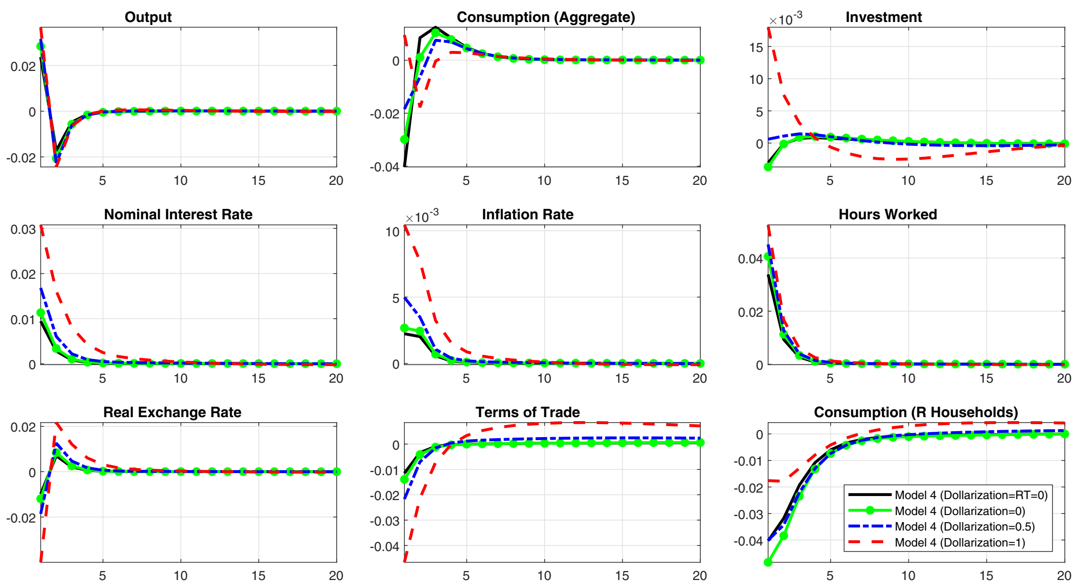
5.3 | Counterfactual experiments: Degrees of dollarization

Finally, and as mentioned above, because entrepreneurial debt can often be denominated in foreign currency, one may refer to the combination of the financial accelerator channel and liability dollarization as collateral constraints. Because in this article, we mainly consider and study inseparably the effects of the two mechanisms (FA-LD), we want to see, in Figures 4 and 5, how the IRFs change when we switch off LD in the net worth position and then increase its degrees gradually so that $(1-\varphi) \in [0, 0.5, 1]$. The focus is on assessing, using this second counterfactual exercise, how eliminating dollarization can affect our model responses reported above.

Indeed, for most cases, the preferred model without dollarization displays much more moderate responses that often die out more rapidly. The impulse responses show the larger (or very different) effects on impact that often last longer over time when the borrowers are in a partially dollarized environment, amplifying the effects of financial stress (especially in the case of responding to the government spending and technology shocks). Once again, this is particularly noticeable in the investment IRFs. Consumption and investment demand is most significantly affected by the entrepreneurs' varying degree of dollarization in our model. This clearly shows evidence of a direct supply-side effect of dollarization through the real marginal costs of firms. For example, as noted, the ensuing negative response on investment (and consumption) following a



(a)

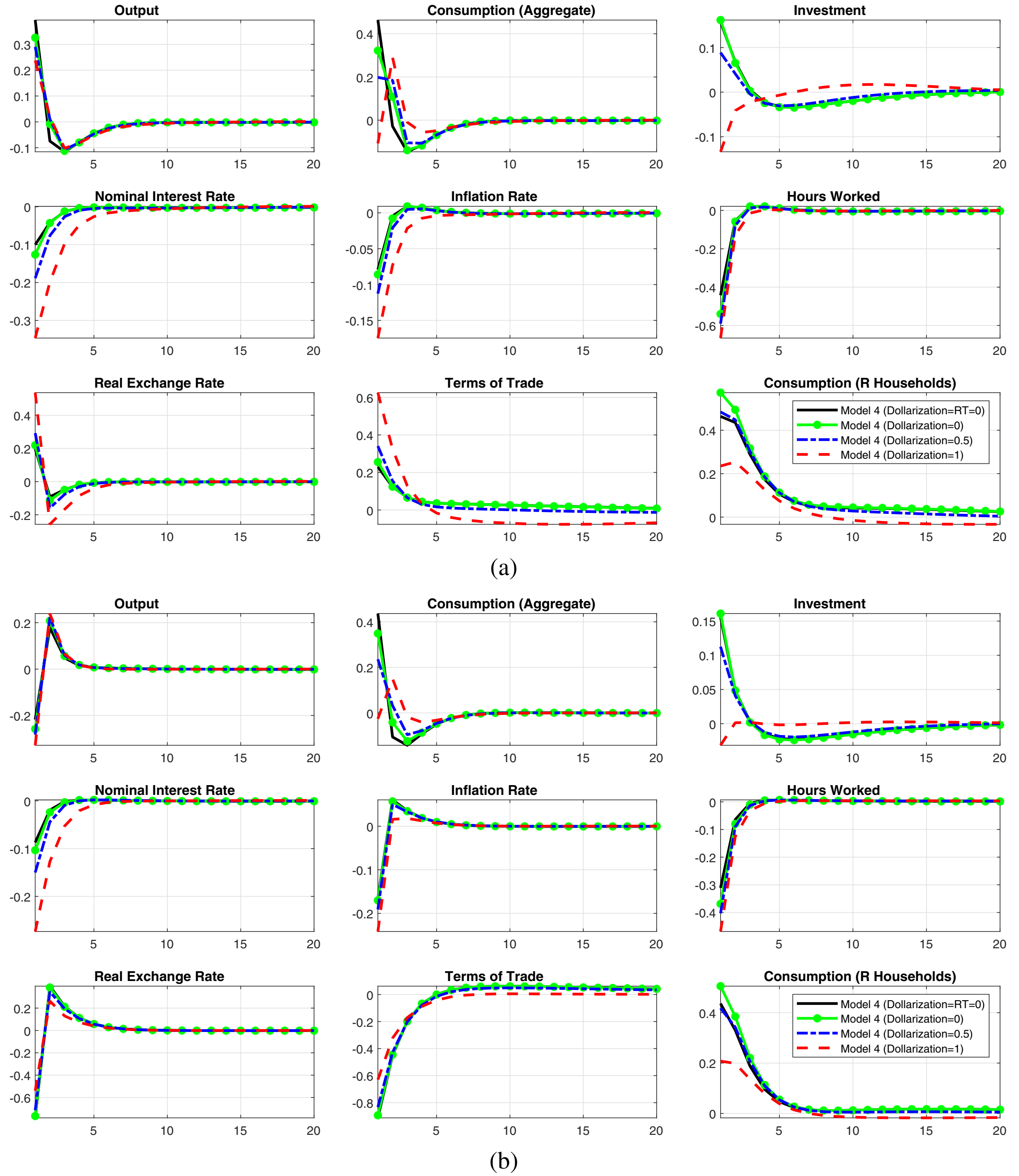


(b)

FIGURE 4 Impulse response functions (Model 4): (A) Monetary policy shock and (B) government spending shock. Each panel plots the mean response based on the estimated Model 4 with varying degree of dollarization $\in [0, 0.5, 1]$. [Colour figure can be viewed at wileyonlinelibrary.com]

technology shock in Model 4, due to the initial demand deficiency, can be significantly magnified on impact because of the combined FA-LD mechanism (Figure 5a).

On the other hand, when looking at a demand shock, the effect of increasing the degree of dollarization seems to be much weaker in response to domestic contractionary monetary policy, at least on the real variables and employment. This is very informative for stabilization policies since the result suggests that the size of the policy transmission effect is mainly driven by the



(a)

(b)

FIGURE 5 Impulse response functions (Model 4): (A) Technology shock and (B) risk premium shock. Each panel plots the mean response based on the estimated Model 4 with varying degree of dollarization $\in [0, 0.5, 1]$. [Colour figure can be viewed at wileyonlinelibrary.com]

financial accelerator mechanism only, which captures the balance sheet effects when there is an unanticipated increase in the cost of capital, irrespective of the degree of dollarization.

Overall, the results from the estimated posterior IRFs confirm our key findings discussed above, that is, there is substantial evidence in the data to support the presence of a financial accelerator mechanism that depends on dollarization. More precisely, the inclusion of the accelerator along with the dollarized borrowers and consumers affects the transmission mechanism of monetary and fiscal policy and significantly magnifies or alters the real effects of shocks. The effectiveness of policy also depends on the presence of a large number of credit constrained households.

In the IRF analysis, the response of consumption is generally stronger than that of GDP so it appears that consumption variability is generally larger than that of GDP. This is again consistent with our findings in Section 4.3. The ability of these additional frictions in explaining the overall persistence and variability we observe in the data is the reason why our preferred model can significantly outperform its rivals in the likelihood race. Our empirical analysis helps us understand whether our shocks generate distinct IRFs, aiming to provide an explanation on how the macroeconomic implications of identified shocks may differ depending on our key modeling features relevant to an emerging SOE.

6 | VARIANCE DECOMPOSITION OF BUSINESS CYCLES

This section investigates the contribution of each of the structural shocks to the forecast error variance of the observable variables in the model. The results are based on the models' posterior distributions reported in Table 3. Table 5 provides a summary of variance decomposition of all the domestic shocks in the estimated models for all the observed domestic variables in the long run.

The most noteworthy finding is the significant role played by the price markup shock and the external financing premium shock in explaining mostly the variance of the real variables in Peru. Movements in real GDP and real consumption are primarily driven by the exogenous markup shock (ν_{PM}) with the dominant influence of over 60%. This shock dominates, accounting for the biggest part of the unexpected output and consumption fluctuations in all four models, and its effects are amplified by financial frictions, particularly when explaining the variance of consumption. It is clearly evident that this shock plays a significantly role in generating the excess consumption volatility reported in Section 4.3. In contrast, the shocks to preferences and fiscal spending shock have little impact on output and consumption variability.

Not surprisingly, the exogenous price markup shock also dominates the investment variance, but the contribution to its variability becomes smaller but still nontrivial (about 40%) when financial frictions are assumed in Models 3 and 4. Instead, the role of external financing premium shock (ν_P) in accounting for the variance of investment increases within the estimated dollarized models with financial frictions. Over time, ν_P dominates, accounting for the biggest part of the investment variance under these models, while the technology shock (ν_A) is a moderate, but significant, factor behind long-run movements in Peruvian investment growth predicted by the models without LD. ν_A is important in explaining its real effects with Models 1 and 2. Over 15% of the variance decomposition are due to an impact from ν_A. This interesting result highlights the systematic importance of the external factors in recent decades that have persistent growth spillovers effects via bilateral trade on an emerging economy. Remarkably, external financing disturbances account for over 50% of investment fluctuations over time and the effects (real and nominal) of the shock are amplified by the encompassing model (Model 4).

Under the estimated interest rate rule we find that the external financing premium shock is by far the most determinant influence to the nominal interest rate in Peru (Model 4) while ν_A dominates the effect on the nominal interest rate with the baseline Model 1 and extended Model 2. Interestingly, the encompassing Model 4 suggests that there are only limited effects on inflation from the productivity shocks and various demand shocks. This finding seems to be plausible,

TABLE 5 Forecast error variance decomposition (in %)

Observable	Shocks of the estimated models							
	v_A	v_C	v_G	v_L	v_P	v_{PM}	v_R	v_{UIP}
Model 1								
$dlGDP_t$	21.27	0.03	0.05	0.03	–	67.97	1.74	8.85
$dlCON_t$	16.53	0.20	0.02	0.24	–	76.99	2.19	3.81
$dlINV_t$	15.84	0.01	0.00	0.02	–	83.05	0.16	0.92
RER_t	75.39	0.07	0.01	0.07	–	16.56	0.32	7.50
INT_t	89.71	0.07	0.01	0.07	–	7.55	0.12	2.46
$dlCPI_t$	71.97	0.13	0.01	0.16	–	12.52	4.45	10.69
Model 2								
$dlGDP_t$	17.19	0.04	0.09	0.07	–	69.22	1.88	11.45
$dlCON_t$	24.96	0.28	0.05	0.41	–	66.35	1.77	6.16
$dlINV_t$	15.05	0.01	0.00	0.01	–	84.19	0.33	0.41
RER_t	67.32	0.05	0.01	0.06	–	24.07	0.67	7.74
INT_t	85.32	0.05	0.01	0.07	–	12.10	0.30	2.15
$dlCPI_t$	63.14	0.08	0.02	0.13	–	24.46	2.97	9.17
Model 3								
$dlGDP_t$	12.59	0.06	0.27	0.16	2.84	78.97	3.16	1.92
$dlCON_t$	2.32	0.01	0.03	0.04	1.01	92.47	3.86	0.26
$dlINV_t$	6.00	0.03	0.03	0.07	50.12	41.90	0.64	1.20
RER_t	59.76	0.44	0.75	0.82	8.23	14.83	0.40	14.46
INT_t	37.01	0.26	0.79	0.58	33.18	20.42	1.54	6.16
$dlCPI_t$	11.63	0.07	0.15	0.20	14.09	9.09	57.66	6.99
Model 4								
$dlGDP_t$	1.55	0.07	0.03	0.07	3.96	84.78	7.11	2.40
$dlCON_t$	0.84	4.89	0.00	0.07	1.36	85.47	6.90	0.29
$dlINV_t$	0.23	0.14	0.00	0.01	56.50	42.38	0.62	0.12
RER_t	11.50	14.25	0.06	0.68	7.99	47.97	4.46	12.08
INT_t	9.09	13.55	0.08	0.59	52.42	14.64	3.23	6.16
$dlCPI_t$	4.13	5.18	0.02	0.29	23.70	12.95	44.19	9.39

Note: All the variance decomposition is computed from the model solutions (order of approximation = 1). The results are based on the model's posterior means. In this exercise, we only examine the role of the structural shocks, which are: v_A—technology, v_C—preferences, v_G—government expenditure, v_L—labor supply, v_P—external finance premium, v_{PM}—price markup, v_R—monetary policy, v_{UIP}—deviations of the UIP conditions.

given that the estimated slope of the Phillips curve is small ($(1 - \beta\xi_H)(1 - \xi_H)/\xi_H$), so that only large and persistent changes in the marginal cost will have an impact on inflation.

This exercise can also help address a concern about the model fit where we find that, from all the four models in Figure 1, the first-order autocorrelation of output and consumption lies outside the confidence interval of the data counterpart. One possible explanation for the poor performance in fitting these moments in the data is that the main driving forces behind these short-run movements are the models' responses to the price markup shock which is indeed responsible for most of the long-run variations in output and consumption. This is in line with the estimated standard deviation of the shock—Model 3 has by far the largest $\sigma_{v_{PM}}$. Overall, these results confirm the view that our estimated Model 4 provides a reasonable approximation to the Peruvian data where fluctuations are amplified in the presence of dollarized liabilities and financial frictions.

7 | ROBUSTNESS CHECKS

In this section, we carry out robustness checks on key aspects of the model and data, respectively. In particular, we investigate the sensitivity of our results to an alternative model setup where we add real wage rigidity to capture frictions in the Peruvian labor market. As an additional robustness exercise, we re-estimate our model based on a truncated sample from 2002:1 to 2019:4. While the first robustness check is implemented to gauge the sensitivity of our results to an alternative assumption that departs from flexible wages and the standard efficiency condition for the labor market, the second check is aimed at exploring the results when we shorten our data sample to start from 2002 in which the BCRP officially adopted inflation targeting (IT). To be consistent with our empirical analysis, these exercises are performed for Model 4—the preferred encompassing model.

7.1 | Persistence of real wages

Following Blanchard and Gali (2007) and Castillo et al. (2013), we add a simple real friction in the labor market to capture the possible persistence when real wages adjust slowly to changes in the marginal rate of substitution between consumption and leisure. For Model 4, we depart from the standard flexible wages assumption set out in (7) by assuming an additional endogenous persistence mechanism through wage indexation. The first-order condition that determines the supply of labor is now given by

$$\frac{W_t}{P_t} = \left(\frac{W_{t-1}}{P_{t-1}}\right)^{\lambda_w} \left(-\frac{\eta}{(\eta - 1)} \frac{U_{L,t}}{U_{C,t}}\right)^{1-\lambda_w},$$

where λ_w measures the degree of wage rigidity or the speed of adjustment of real wages. We estimate the model again using the same sample period that covers from 1994:1 to 2019:4, the same number of MCMC-MH draws, and assuming that the prior for λ_w is beta distributed, centered at 0.5 with a standard deviation of 0.1, which is the same prior assumption used for the probability of price setting, ξ_H.

The results summarizing posterior means of the estimated parameters and 90% HPD intervals are reported in the first column of Table 6. The parameter λ_w is estimated to be 0.62, suggesting a substantial degree of wage inertia in the Peruvian data. However, we notice that the LL is almost

TABLE 6 Prior and posterior distributions: Robustness

Parameter	Notation	Posterior distribution Model 4 (with wage inertia)	Model 4 (data from 2002)
Investment adjustment	$S''(1+g)$	2.0198 [0.8506:3.5912]	2.6667 [1.6252:3.6993]
Risk aversion	σ	1.0049 [0.5043:1.5085]	0.6825 [0.2360:1.1274]
Consumption habit	h_C	0.5587 [0.1310:0.9096]	0.5330 [0.2640:0.7956]
Calvo prices	ξ_H	0.7715 [0.7029:0.8466]	0.8367 [0.7621:0.9121]
Degree of wage rigidity	λ_w	0.6194 [0.4235:0.8488]	–
Substitution elasticity (balances in H/F currencies)	χ_m	4.8320 [2.8311:6.6269]	3.9785 [1.5723:6.4514]
Fraction of real balances held in H currency	a	0.2209 [0.0856:0.3516]	0.6000 [0.3004:0.9075]
Fraction of borrowing in H currency	φ	0.7504 [0.5349:0.9719]	0.5497 [0.2383:0.8878]
External finance premium	Θ	0.0059 [0.0026:0.0099]	0.0055 [0.0024:0.0088]
External finance premium elasticity	χ_θ	0.0264 [0.0135:0.0375]	0.0239 [0.0121:0.0352]
Inverse of leverage	n_k	0.3899 [0.2483:0.5244]	0.5798 [0.4327:0.7294]
Entrepreneurs survival rate	ξ_e	0.9815 [0.9686:0.9954]	0.9922 [0.9863:0.9988]
Proportion of RT consumption	λ_{C_1}	0.3645 [0.1942:0.5366]	0.1906 [0.1311:0.2444]
Taylor rule: inflation	θ_π	2.4334 [1.9641:2.9753]	1.6552 [1.1122:2.1406]
Taylor rule: output	θ_y	0.3160 [0.2140:0.4424]	0.2228 [0.1301:0.3161]
Taylor rule: interest rate smoothing	ρ	0.4938 [0.3676:0.6164]	0.8530 [0.7863:0.9249]
AR(1) coefficient			
Technology	ρ_A	0.4086 [0.1199:0.7283]	0.5068 [0.1531:0.8661]
Preferences	ρ_C	0.4109 [0.0827:0.7275]	0.5807 [0.2874:0.8614]
Government spending	ρ_G	0.5018 [0.1872:0.8441]	0.4973 [0.1764:0.8309]
Labor supply	ρ_L	0.6395 [0.2703:0.9974]	0.5041 [0.1577:0.8234]
External finance premium	ρ_P	0.4247 [0.4247:0.6838]	0.3421 [0.0720:0.6018]
Price markup	ρ_{PM}	0.7393 [0.5861:0.9126]	0.6867 [0.5634:0.8118]
UIP	ρ_{UIP}	0.4689 [0.1645:0.7697]	0.4581 [0.1622:0.7569]
Standard deviation of AR(1) innovations/i.i.d. shocks			
Technology	σ_{v_A}	0.6831 [0.1363:1.0778]	0.3518 [0.0997:0.6402]
Preferences	σ_{v_C}	0.7200 [0.0980:1.7060]	0.9714 [0.1422:1.5701]
Government spending	σ_{v_G}	0.3880 [0.1019:0.7351]	0.3223 [0.1012:0.5892]
Labor supply	σ_{v_L}	1.3046 [0.1456:3.7677]	0.3894 [0.0903:0.7421]

TABLE 6 (Continued)

Parameter	Notation	Posterior distribution Model 4 (with wage inertia)	 Model 4 (data from 2002)
External finance premium	σ_{v_P}	3.1819 [2.0398:4.9597]	5.2886 [1.1275:8.7451]
Price markup	$\sigma_{v_{PM}}$	0.8277 [0.5512:1.1705]	0.7153 [0.3885:1.0429]
Monetary policy	σ_{v_R}	0.7910 [0.6519:0.9224]	0.1148 [0.0788:0.1509]
UIP	$\sigma_{v_{UIP}}$	0.4680 [0.1478:0.7306]	0.2313 [0.1018:0.3564]
Degree of TD	$1-\alpha$	0.7791 [0.6484:0.9144]	0.4000 [0.0925:0.6996]
Degree of LD	$1-\varphi$	0.2496 [0.0281:0.4651]	0.4503 [0.1122:0.7617]
RT consumers	λ	0.5763	0.4604
Price contract length	$\frac{1}{1-\xi_H}$	4.3764	6.1237
Log-likelihood		−1115.26	N/A

Note: λ is derived as follows: $\frac{C_2}{C} = \frac{C_2^*}{C^*} = 1.5$, then $\lambda_{C_1} = \lambda\frac{C_1}{C} = 1-(1-\lambda)\frac{C_2}{C}$.

unchanged (−1115.26); the LL difference is −1.36. As explained in Section 4.2, based on this small difference in log marginal likelihoods, we cannot confirm that one model statistically outperforms the other in terms of its ability to explain the data.

In line with our findings under the original Model 4, the estimates of the structural parameters do no change significantly. In particular, the estimated parameters and their confidence intervals capturing the key features in our model—TD, FD, and FA—are very similar across the two variants of the model. However, the main differences with respect to the original preferred Model 4 are the parameter measuring the exogenous probability of changing prices, ξ_H, and the proportion of RT consumers, λ_{C_1}. The former shows that the price contract length changes from 3.20 to 4.38, implying slightly stronger price stickiness, and the latter suggests that the share of liquidity constrained consumers, λ, increases from 0.46 to 0.58. These values are in the range reported by studies using similar models for emerging market economies (see Bhattarai et al., 2021; Caputo et al., 2007; Castillo et al., 2013, among others) although the estimated λ in this case is higher than the assumed prior mean of 0.50 and the typical value that one probably finds in some of the recent SOE studies (see, for example, Gabriel et al., 2016; Omotosho, 2019).

7.2 | Shorter sample period (2002–2019)

In this section, we retain our benchmark model assumption of flexible wages and perform our second check: we shorten the sample period (2002–2019), disregarding 32 observations (1994–2001), and compare the results with respect to the previously preferred model. First, as mentioned in the introduction, both volatilities of nominal variables before the first year of the sample and volatilities of real variables, such as GDP growth, during the first several years of the sample are relatively high. Including this post-stabilization period in our sample has motivated us to adopt the flexible trend agnostic procedure in our estimations. Second, while our estimation procedure is robust to potential trend misspecifications in the data, there has been a monetary regime change in the full sample period of 1994–2019. Since 2002, the BCRP has adopted a fully fledged IT regime as part of the dollarization risk control framework. In light of this change of monetary policy actions, we also briefly discuss how the estimation results may be different based on the reduced sample period.

Now, we compare the estimates in the second column of Table 6 with those presented in the rightmost column of Table 3. First, the parameters on the reaction of the Taylor rule to inflation and output are estimated to be smaller, with the estimates depending on the inflation regime (from 2.68 to 1.66 and from 0.29 to 0.22, respectively). The less aggressive responses are consistent with the relatively low degree of TD (0.40 in this case) which can provide some intuition behind the results. This suggests a milder reaction of the BCRP to offset the effects of dollarization on the volatility of consumption in the reduced sample. Our estimation using the shorter sample period also shows a much higher degree of interest rate smoothing—the value almost doubles. The stronger weight attached to interest rate smoothing in the Taylor rule is expected to lead to a less aggressive reaction toward inflation.

Second, given that we observe reduced volatilities in the shorter sample period, it is not surprising to find that the estimated volatilities of structural shocks are much reduced—with one exception, the shock to external finance premium, ν_P, which has the largest standard deviation (around 5.28%). Interestingly, we observe a significant reduction in the estimated standard deviations of monetary policy shock (from 0.92 to 0.11). With respect to the structural parameters, while the FA parameters are largely unchanged compared with the estimation using the full sample, some differences can be found from the investment adjustment and risk aversion parameter estimates. Notably, the average price contract length has further increased to 6.12 quarters in this case. The DSGE literature has reported similar values for the US and the Euro area.

Finally, we uphold our main finding that Model 4, which contains all the frictions, significantly outperforms Model 1, the baseline model, in terms of their empirical relevance in explaining the data. As we carry out the estimations for both models in this section, we uphold this result with both the full sample and the shorter sample period during which the country becomes an official inflation targeter. Importantly, our likelihood comparison exercise confirms our earlier results that there is a significant difference in the estimated posterior likelihoods: −574.5546 (Model 4) and −632.4708 (Model 1), providing yet again decisive evidence in favoring the former. Our overall conclusion in this section regarding the choice of models is robust to a structure with a much reduced sample period and regime changes in the volatility of time series observations.

8 | CONCLUSIONS AND FUTURE RESEARCH

Are the financial accelerator and partial dollarization empirically relevant in emerging market economies? In this article, we tackled this question by developing and estimating a two-bloc emerging-market/rest-of-the-world DSGE model, where in the emerging market bloc there is a strong link between changes in the exchange rate and financial distress of household and firms. In other words, we assumed that the emerging market bloc incorporates partial transaction and liability dollarization, as well as financial frictions, including a financial accelerator mechanism, where capital financing is partly or totally in foreign currency as in Gertler et al. (2007) and Gilchrist (2014).

Our modeling framework is general enough to be relevant for the study of propagation of shocks in open (developing) economies featuring dollarization and various other forms of credit frictions, while highlighting the importance of using flexible and robust estimation procedures that address the considerable uncertainty about the nature of trends in emerging economies. Our empirical strategy is novel and comprehensive in terms of estimating a SOE-DSGE model incorporated with a fully fledged ROW economy. Several variants of the model were then estimated by Bayesian maximum likelihood using data for Peru and the US. We obtained plausible and tight posterior estimates for both the financial accelerator and partial dollarization parameters, thus confirming their empirical relevance.

Applying the marginal likelihood criterion and common validation methods, we found strong evidence in the data (for both Peru and the US) to support the presence of a financial accelerator and strong evidence for the existence of transaction and liability dollarization. In fact, Model 4, a dollarized model incorporating the financial accelerator and presence of credit constraint consumers, significantly outperforms its rivals, suggesting that both frictions play important roles in improving the conformity of the model with the actual data. The impulse response analysis further highlights the empirical importance of the accelerator mechanism. It magnifies the effects of productivity and monetary policy shocks on private investment. The article also discusses the linkages and interaction between the key features of the model for the propagation of policy shocks to the real activity. The results and analysis provide important insights into the design of stabilization policy necessary for sustaining economic growth in emerging open economies.

An issue for further research is the form of rational expectations (RE) solution employed for solving and estimating the model. In this article, we adopted the standard assumption in the literature that agents have perfect information despite the number of shocks being greater than the number of observables. The RE solution is therefore not invertible and neither agents nor the econometrician can recover the shocks using data. The perfect information assumption assumes that this exists as an endowment for the agents in the model. However, Levine et al. (2019) show how this extreme (but standard) assumption can be relaxed in favor of imperfect information. In particular, by giving agents the same information as the econometrician (the data), this will introduce informational learning into the model via a Kalman filter and the added persistence may well improve the data fit of the model (see Levine et al., 2020 for an example of this).

Another issue that certainly deserves further attention is the way the policy rule is modeled, in particular for emerging economies. Indeed, it is possible that exchange rates play a role in how monetary policy is conducted. In normal times, central banks may not wish to focus on the exchange rate, as it can lead to a weakening of the inflation target as a nominal anchor (see, e.g., Berganza & Broto, 2012; Mishkin & Savastano, 2001). However, in a sudden-stops world, when capital abruptly stops to flow into a country, interest rate interventions may not work and monetary authorities may occasionally resort to foreign exchange rate interventions instead. Note, however, that the log-linear approach employed here may not be able to capture the nonlinearities implied by this phenomenon.

In terms of the modeling framework, a transactions-based framework, rather than a money-in-the-utility-function approach, could be pursued, but we feel that the latter adds tractability and is general enough to capture different possibilities. The model could also be enriched by incorporating additional endogenous persistence, for example through price indexation, which could refine our parameter estimations. Also, while in our framework firms are assumed to apply producer currency pricing, recent studies suggest that firms set prices in very few currencies (with the dollar being the most frequently used currency) and do not change prices often (Goldberg & Tille, 2010; Gopinath, 2015; Gopinath et al., 2020). This new Dominant Currency Pricing paradigm implies that changes in nominal exchange rate will only weakly impact the terms of trade, while the driver of prices and quantity of imported goods will be the country's currency value *vis-à-vis* the dollar, and therefore its implications for monetary policy is likely to be substantially different. This should be taken further by proper consideration of price rigidities in imported goods, introducing an additional channel for the propagation of foreign shocks, as in Christiano et al. (2011). These features, however, are beyond the scope of the present article and we leave this for future research.

ACKNOWLEDGMENTS

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DATA AVAILABILITY STATEMENT

The data that support the findings of this study are available from the corresponding author upon reasonable request.

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ENDNOTES

1. The term is often used interchangeably to indicate that dollars (or a foreign currency, more generally) serve as a unit of account (real dollarization or price dollarization), as a store of value (financial dollarization), or as a medium of exchange (transaction dollarization, also known as currency substitution).
2. Since the 2008–2009 financial crisis the literature on credit frictions and macroeconomic fluctuations has expanded considerably and focused both on improving the existing approaches and on developing new theories to model the interactions between the financial distress and the real economy. The main focus has been on how conventional and unconventional monetary policy should react to financial crises and how various forms of financial frictions can be used for guiding macro-prudential policy (e.g. Brzoza-Brezina, 2014; Quint & Rabanal, 2014). See, for a more comprehensive survey of literature, Brunnermeier et al. (2012) and Duncan and Nolan (2017).
3. Although there have been moderate declines in foreign currency deposits as a share of total deposits, the extent of dollarization still remains high, with this share close to 20% in the past 20 years.
4. Andrle (2008) argues that assumptions on trending behavior should be explicitly modeled, rather than side-stepped by means of an ad-hoc prefiltering procedure, while recent developments by Canova and Ferroni (2011) and Ferroni (2011) allow for formal statistical comparisons among different de-trending procedures.
5. Online appendices provide details of the model, including the list of equations in the linearized system, as well as discussing the priors, US model estimation and an additional robustness exercise.
6. See Woodford (2003), a "cost channel" as in Ravenna and Walsh (2006) has a similar supply-side effect on the Phillips curve.
7. For simplicity, analogous "foreign" variables are largely omitted in the exposition.
8. A BGP requires that the real wage, real money balances and consumption grow at the same rate at the steady state with labor supply steady. It is straightforward to show that (1) has these properties.
9. In Section 7.1, we further consider the introduction of real wage rigidities as in Blanchard and Gali (2007).
10. We impose $\zeta = \zeta^*$ for reasons that become apparent in Section 2.2.2. Consistently, we adopt a notation where subscript H or F refers to goods H or F produced in the home and foreign bloc, respectively. The presence (for the foreign bloc) or the absence (for the home bloc) of a superscript "$*$' indicates where the good is consumed or used as an input. Thus $C^*_{H,t}$ refers to the consumption of the home good by households in the foreign bloc. Parameters w and w* refer to the home and foreign bloc, respectively, and so forth.
11. Note that, for estimation purposes, we define λ_{C_1}, which measures the share of consumption consumed by the liquidity constrained consumers, such that $\lambda_{C_1} = \lambda \frac{C_1}{C} = 1 - (1-\lambda)\frac{C_2}{C}$.

12. Thus we can interpret $\frac{1}{1-\xi_H}$ as the average duration for which prices are left unchanged.
13. Recall that we have imposed a symmetry condition $\zeta = \zeta^*$ at this point; that is, the elasticity of substitution between differentiated goods produced in any one bloc is the same for consumers in both blocs.
14. Note that all aggregates, Y_t, $C_{H,t}$, and so forth are expressed in per capita (household) terms.
15. Details of this can be found in a supplementary appendix available from the authors. Details of the associated linearization can be found in Online Appendix A.
16. See Gertler et al. (2007), Cespedes et al. (2004), Elekdag et al. (2005), and Devereux et al. (2006).
17. The real variables are seasonally adjusted with ARIMA X-12. For the nominal variables, especially the domestic CPI inflation and policy interest rate, any trend in their deterministic component throughout the sample period is not prespecified and should be picked up by the agnostic procedure we use for the measurement equations.
18. The posterior mode and corresponding Hessian are obtained and then the posterior density is approximated by using the Monte Carlo Markov chain Metropolis–Hastings (MCMC-MH) algorithm, starting from the posterior kernel mode, with two parallel chains with 500,000 draws. We burn-in the first 25% of the chain.
19. A number of parameters can be further calibrated or derived—see Online Appendices B and C for details.
20. See Castillo et al. (2013), Bernanke et al. (1999), and Gertler et al. (2007) for further details.
21. The last two columns of Table S1 in Appendix E report results when 1-year and 10-year expectations data are employed and, as can be seen, estimates are very similar. Thus, to save space, we only report second-stage SOE-ROW results for the 1-year inflation expectations specification.
22. Regardless, in Section 5.3, we conduct further sensitivity analysis by employing a range of values for the degree of dollarization.
23. Our posterior mode obtained for the both dollarization parameters is robust to assuming even looser prior on them (e.g. we consider an alternative prior specification with standard deviation of 0.25).
24. Importantly, we consider again alternatively prior specifications (e.g. priors with a looser precision) to ensure robustness of our posterior mode before sampling for the posterior draws.
25. To further validate our estimates, a recent study by Demirguc-Kunt et al. (2018) finds that 57.4% and 7% of households in Peru and US do not have access to financial institutions, which are very close to our estimates of λ for Peru and the US, respectively.
26. As a robutsness/counterfactual check, we remove all the financial frictions including the nonoptimizing group of RT households (so that $\lambda = 0$) in the first two models. We can observe some improvement in goodness of fit from the model incorporating TD.
27. Equivalently, in a Bayesian model comparison, a posterior Bayes factor needs to be at least 3 for there to be decisive evidence favoring Model 1.
28. For this comparison among the four model variants, the relevant data and model variables are transformed using a first-difference filter since our focus is on validating our results against data properties obtained based on the different underlying modeling assumptions.
29. It is not surprising to see that Model 4 does not offer the best fit for some moments in the data. Perhaps the main message to emerge from this validity exercise is that it can be misleading to assess model fit using a selective choice of second moment comparisons. Likelihood comparisons provide the most comprehensive form of assessment that can still leave trade-offs in terms of fitting some second moments well at the expense of others.
30. From Figure S1 in Online Appendix F, the several responses based on Model 1 with and without the RT households (i.e. with black and red lines) and Model 4 (blue and green broken lines) show clear predictions of such effects.

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