What Do Latin American Inflation Targeters Care About? A Comparative Bayesian Estimation of Central Bank Preferences[*]

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Abstract

This paper uses Bayesian estimation techniques to uncover the central bank preferences of the big five Latin American inflation targeting countries: Brazil, Chile, Colombia, Mexico, and Peru. The target weights of each central bank's loss function are estimated using a medium-scale small open economy New Keynesian model with incomplete international asset markets and imperfect exchange-rate pass-through. Our results suggest that all central banks in the region place a high priority on stabilizing inflation and interest rate smoothing. While stabilizing the real exchange rate is a concern for all countries except Brazil, only Mexico is found to assign considerable weight to reducing real exchange rate fluctuations. Overall, Brazil, Colombia, and Peru show evidence of implementing a strict inflation targeting policy, whereas Chile and Mexico follow a more flexible policy by placing a sizeable weight to output gap stabilization. Finally, the posterior distributions for the central bank preference parameters are found to be strikingly different under complete asset markets. This highlights the sensitivity of Bayesian estimation, particularly when uncovering central bank preferences, to alternative international asset market structures.

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1 Introduction

For many central banks in both developed and developing countries, inflation targeting (IT) has become the operational monetary framework of choice to achieve price stability.[1] According to the International Monetary Fund (IMF) (see, e.g., Jahan, 2012), since the adoption of IT by New Zealand in December 1989, there are now 28 IT central banks worldwide, of which 6 currently originate from Latin America: Brazil, Chile and Colombia all adopted IT in 1999, shortly followed by Mexico (2001), Peru (2002), and Guatemala (2005) (see Table 1).[2] While there is some empirical evidence to suggest that IT has been successful in reducing inflation in developing countries (see, e.g., Batini and Laxton, 2007; Goncalves and Salles, 2008; Lin and Ye, 2009; Lee, 2011),[3] little is known about the policy preferences of central banks operating in these countries.[4] As discussed by Castelnuovo and Surico (2004) and Ilbas (2010, 2012), such information can help in evaluating the performance of central banks, as well as improving our understanding of monetary policy actions and its effects on the formation of expectations by private agents.

The aim of this paper is to use Bayesian estimation techniques to uncover and compare the central bank preferences of the big five Latin American inflation targeting (LAIT) countries.[5] Since the IT framework can be considered as "constrained discretion" (Bernanke and Mishkin, 1997), we assume that in each country monetary policy is conducted under discretion. Each central bank is assumed to optimally set the nominal interest rate by minimizing an intertemporal quadratic loss function that includes four specific policy objectives: price stability via control of inflation, stabilizing the output gap, reducing real exchange rate variability, and nominal interest rate smoothing. The weight attributed to each policy objective will depend on the institutional preferences of each central bank, which we can make inferences about using estimates of the respective Bayesian posterior distributions.

The structural model used to represent the LAIT economies is a dynamic medium-scale small open economy New Keynesian model. Following the modeling frameworks of Monacelli (2005), Kam et al. (2009) and Justiniano and Preston (2010), we allow for imperfect exchange-rate pass-through (ERPT) such that the law of one price fails to hold. As in Adolfson et al. (2007) and Justiniano and Preston (2010), international asset markets are assumed to be incomplete such that consumption risk sharing is not perfect. Both these features have been identified as empirically relevant, the more so for developing countries as in our sample. Using the popular Random-Walk Metropolis-Hastings Markov Chain Monte Carlo algorithm, we present posterior estimates and convergence diagnostics for both the structural parameters and the persistence and standard deviations of the shocks we consider as most important in

[1] See, for example, Mishkin and Schmidt-Hebbel (2001), Carare and Stone (2006), Roger (2010), Hammond (2011), Jahan (2012).

[2] As Table 1 reveals, there is significant heterogeneity across these 6 Latin American countries in terms of the inflation target set and their performance in steering actual inflation towards the target.

[3] Lee (2011) finds that IT has been particularly successful in reducing inflation in Colombia, while no significant reductions were found for Chile. In contrast to much of the earlier literature, Brito and Bystedt (2010) find no evidence that IT improves economic performance in developing countries, as measured by the behavior of inflation and output growth.

[4] There is also some evidence to suggest that IT has reduced the dispersion of long-run inflation expectations in developing countries. See Capistrán and Ramos-Francia (2010) for further details.

[5] We exclude Guatemala from the analysis due to the lack of reliable data.

emerging market economies (such as, among others, shocks to preferences, the risk premium, the terms of trade, and technology).

Our main findings are as follows. First, all five central banks are strongly concerned about stabilizing inflation and smoothing the nominal interest rate. In particular, relative to the weight of inflation stabilization, we find that Brazil and Peru place very high weights on interest rate smoothing. Second, there is significant heterogeneity amongst the five central banks concerning the priorities of output gap stabilization and real exchange rate stabilization. Brazil, Colombia, and Peru show little concern for the stabilization of the output gap, whereas Chile and Mexico assign sizeable weights. While Brazil and Colombia are not concerned about real exchange rate volatility, for the remaining three countries only Mexico is found to assign a sufficiently high weight to minimizing real exchange rate fluctuations. Overall, Brazil, Colombia and Peru show evidence of implementing a strict inflation targeting policy, whereas Chile and Mexico appear much more flexible in terms of their inflation targeting preferences.

In terms of the estimated key structural parameters influencing the endogenous propagation mechanism of the model, we find that these are statistically reliable, economically plausible, and broadly comparable to analogous estimates for other countries available in the literature using non-Bayesian econometric methods. For example, the estimated elasticity of substitution between home and foreign goods is within the typical range reported in Corsetti et al. (2008). Furthermore, we uncover some interesting differences across countries. For example, the degree of price stickiness in Brazil and Mexico is estimated to be over 3.5 times higher than in Peru.

In terms of the sources of exogenous fluctuations affecting the five LAIT economies, our results indicate a high volatility (measured by the posterior mean standard deviation) for the preference shock, the risk premium shock and the terms of trade shock. The most persistent shock appears to be the risk premium shock, yet it does not dominate the persistence of some among the other shocks in each individual country. Overall, the least persistent shock in all five LAIT economies is the preference shock.

There are few papers that have used Bayesian techniques to estimate central bank preferences in an open-economy setting.[6] Kam et al. (2009) estimate the central bank preferences for three developed IT countries, Australia, Canada, and New Zealand under optimal discretionary monetary policy. They find that the central banks of these countries all have very similar preferences: the highest priority is inflation stabilization, followed by interest rate smoothing, with no concern for stabilizing the output gap (with the exception of Australia) and the real exchange rate. Palma and Portugal (2014) estimate the model of Kam et al. (2009) using Brazilian data. They find that the major concern was inflation stabilization, followed by interest rate smoothing, real exchange rate stabilization and output gap stabilization.

While the estimation approach adopted in this paper is similar to Kam et al. (2009), one important difference relates to the asset market assumption used in the structural model. Kam et al. (2009) assume complete international asset markets which implies perfect risk sharing in consumption and a strong positive correlation between the real exchange rate and the marginal utilities of consumption across countries. However, there is clear empirical evidence for both de-

[6]In a closed-economy setting, Ilbas (2010, 2012) estimates the preferences of the European Central Bank and the US Federal Reserve under commitment using the structural models of Smets and Wouters (2003, 2007).

veloped and developing economies of low consumption risk sharing across countries.[7] Moreover, Rabanal and Tuesta (2010) show that the extent of international financial market integration affects both the Bayesian estimates of the parameters and the transmission mechanism of shocks. We therefore depart from Kam et al. (2009) and follow the modeling approach of Adolfson et al. (2007) and Justiniano and Preston (2010) in assuming incomplete international financial markets in addition to imperfect ERPT.[8] Furthermore, we test the sensitivity of our results to the alternative assumption of complete asset markets and show that a number of our key policy conclusions obtained under incomplete asset markets are now reversed. This is consistent with the Bayesian analysis of Rabanal and Tuesta (2010) who also find that the degree of model misspecification depends on the asset market structure.[9] Overall, our sensitivity analysis emphasizes the dangers of using the complete asset markets assumption in Bayesian estimation, particularly when uncovering central bank prevalences.

The paper is organized as follows. Section 2 outlines the theoretical model. Section 3 describes the data and explains the estimation strategy. Section 4 reports our main results. Finally, section 5 concludes.

2 The Model

This section outlines the model economy, which is based on the SOE frameworks of Monacelli (2005), Kam et al. (2009), and Justiniano and Preston (2010). The domestic economy is populated by infinitely-lived households of measure one, a continuum of domestic good producers, a continuum of retail firms who import foreign goods at competitive world prices, and a central bank. Both domestic and retail firms are assumed to operate under monopolistic competition and set prices in a staggered fashion according to Calvo (1983). Market power in the retail sector for imported goods results in incomplete ERPT and thus the law of one price fails to hold. International financial markets are assumed to be incomplete. Following Kam et al. (2009), the inflation-targeting central bank is assumed to minimize a quadratic loss function under discretion. In what follows, asterisks conventionally denote foreign variables, and subscripts H (F) denote variables of $Home$ $(Foreign)$ origin.

2.1 Households

Households consume a composite of domestic C_H and imported C_F goods:

$$C_t = \left[(1-\alpha)^{\frac{1}{n}} C_{H,t}^{\frac{\eta-1}{\eta}} + \alpha^{\frac{1}{n}} C_{F,t}^{\frac{\eta-1}{\eta}} \right], \tag{1}$$

$$C_{H,t} = \left[\int_0^1 C_{H,t}(i)^{\frac{\varepsilon-1}{\varepsilon}} di \right]^{\frac{\varepsilon}{\varepsilon-1}} \quad ; \quad C_{F,t} = \left[\int_0^1 C_{F,t}(j)^{\frac{\varepsilon-1}{\varepsilon}} dj \right]^{\frac{\varepsilon}{\varepsilon-1}}. \tag{2}$$

[7] See, e.g., Chari et al. (2002), Heathcote and Perri (2002), Corsetti et al. (2008), Rabanal and Tuesta (2010), Raffo (2010).

[8] Justiniano and Preston (2010) estimate the coefficients of a Taylor-type monetary policy rule for the same three advanced IT economies as in Kam et al. (2009).

[9] Using Bayesian estimation of a medium-scale two-country model for the US and the Euro Area, they find that incomplete asset markets fits the data better than complete asset markets, including matching the dynamics of the real exchange rate.

The parameter $\eta > 0$ measures the elasticity of substitution between home and foreign goods, $\alpha \in (0,1)$ is the share of foreign goods in the domestic consumption bundle, and $\varepsilon > 1$ measures the elasticity of substitution between the varieties of goods produced within H or F, where $i, j \in [0,1]$. The optimal allocation of expenditures between domestic and imported goods yields the following aggregate demand conditions:

$$C_{H,t} = (1-\alpha)\left(\frac{P_{H,t}}{P_t}\right)^{-\eta} C_t, \qquad C_{F,t} = \alpha \left(\frac{P_{F,t}}{P_t}\right)^{-\eta} C_t, \tag{3}$$

where the consumer price index P_t is given by:

$$P_t = \left[(1-\alpha)P_{H,t}^{1-\eta} + \alpha P_{F,t}^{1-\eta}\right]^{\frac{1}{1-\eta}}. \tag{4}$$

The (home) real exchange rate $\tilde{q}_t$ is defined by

$$\tilde{q}_t = \tilde{e}_t \frac{P_t^*}{P_t}, \tag{5}$$

where $\tilde{e}_t$ denotes the (home) nominal exchange rate. The relative price of foreign goods in terms of home goods, or the (home) terms of trade, S_t, is expressed as

$$S_t = \frac{P_{F,t}}{P_{H,t}}. \tag{6}$$

The representative household chooses consumption C_t and labor N_t to maximize expected discounted utility:

$$\max \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \epsilon_{g,t} U \left(\frac{(C_t - H_t)^{1-\sigma}}{1-\sigma} - \frac{N_t^{1+\varphi}}{1+\varphi}\right),$$

where the discount factor is $\beta \in (0,1)$, $\sigma, \varphi > 0$ are the inverse elasticities of intertemporal substitution and labor supply, respectively, $H_t \equiv hC_{t-1}$ is an external habit variable with $h \in (0,1)$, and $\epsilon_{g,t}$ is a preference shock.

The household during period t supplies labor to domestic firms receiving income from wages W_t and profits from the ownership of domestic and retail firms Π_t. As in Adolfson et al. (2007) and Justiniano and Preston (2010) – but differently from Monacelli (2005) and Kam et al. (2009) – the international asset market structure is assumed to be incomplete. Let B_{t-1} and B_{t-1}^* denote the holdings of home and foreign risk-free bonds that mature in period t with corresponding interest rates $\widetilde{r}_t$ and $\widetilde{r}_t^*$. Following Justiniano and Preston (2010), we assume that there is a debt-elastic interest rate premium $\omega_{t-1}(D_{t-1}, \epsilon_{q,t-1})$ given by:

$$\omega_{t-1} = \exp[-\chi(D_{t-1} + \epsilon_{q,t-1})], \qquad D_{t-1} \equiv \frac{\tilde{e}_{t-1}B_{t-1}^*}{Y_{ss}P_{t-1}}, \tag{7}$$

where $\epsilon_{q,t-1}$ is a risk premium shock and D_{t-1} is defined as the ratio of the real quantity of foreign bond holdings (expressed in terms of domestic currency) to steady state output Y_{ss}. If the household is a borrower ($D_t > 0$), it must pay a premium over the interest rate. This debt-elastic interest rate premium is sufficient to ensure that bond holdings are stationary.[10] Consequently, the period budget constraint of the domestic household can be expressed as:

$$P_tC_t + B_t + \tilde{e}_tB_t^* = B_{t-1}(1+\widetilde{r}_{t-1}) + \tilde{e}_tB_{t-1}^*(1+\widetilde{r}_{t-1}^*)\omega_{t-1}(D_{t-1}) + W_tN_t + \Pi_t. \quad (8)$$

The first-order conditions from the households maximization problem yield:

$$(C_t - H_t)^{\sigma}N_t^{\varphi} = \frac{W_t}{P_t}, \quad (9)$$

$$\beta(1+\widetilde{r}_t)\mathbb{E}_t\left\{\left(\frac{C_{t+1}-H_{t+1}}{C_t-H_t}\right)^{-\sigma}\left(\frac{P_t}{P_{t+1}}\right)\left(\frac{\epsilon_{g,t+1}}{\epsilon_{g,t}}\right)\right\} = 1, \quad (10)$$

$$\mathbb{E}_t\left\{\frac{\epsilon_{g,t+1}(C_{t+1}-H_{t+1})^{-\sigma}}{P_{t+1}}\left[(1+\widetilde{r}_t) - (1+\widetilde{r}_t^*)\left(\frac{\tilde{e}_{t+1}}{\tilde{e}_t}\right)\omega_t(D_t, \epsilon_{q,t})\right]\right\} = 0. \quad (11)$$

Equation (9) is the intratemporal labor supply condition, (10) is the intertemporal consumption Euler equation, and (11) is the interest rate parity condition.

2.2 Domestic Good Producers

The domestic goods market is comprised of a continuum of monopolistically competitive firms $i \in [0,1]$ that produce differentiated goods. Domestic firms hire labor N to produce output using a linear production technology

$$Y_{H,t}(i) = \epsilon_{a,t}N_t(i), \quad (12)$$

where $\epsilon_{a,t}$ is an exogenous domestic technology shock, and given competitive prices of labor, cost minimization yields

$$MC_t = \frac{W_t}{\epsilon_{a,t}P_{H,t}}, \quad (13)$$

where MC_t denotes real marginal cost.

Domestic firms set prices according to Calvo (1983), where in each period there is a constant probability $1-\theta_H$ that a firm will be randomly selected to adjust its price, while a fraction $0 < \theta_H < 1$ adjusts their prices according to the following indexation rule

$$P_{H,t}(i) = P_{H,t-1}(i)\left(\frac{P_{H,t-1}}{P_{H,t-2}}\right)^{\delta_H}, \quad (14)$$

where $\delta_H \in [0,1]$ measures the degree of inflation indexation. For simplicity, we assume that the export price of the domestic good is determined by the law of one price: $P_{H_t}^* = (1/S_t)P_{H,t}$. A domestic firm i, faced with changing its price at time t, has to choose $P_{H,t}(i)$ to maximize its expected discounted value of profits:

$$\max_{P_{H,t}(i)} \mathbb{E}_t\sum_{s=0}^{\infty}Q_{t,t+s}\theta_H^s\left[P_{H,t}(i)\left(\frac{P_{H,t+s-1}}{P_{H,t-1}}\right)^{\delta_H} - P_{H,t+s}MC_{t+s}\exp(\epsilon_{H,t+s})\right]Y_{H,t+s}(i),$$

[10] For an in-depth discussion of the stationarity problem of small open-economy models with incomplete asset markets, see Schmitt-Grohé and Uribe (2003).

where

$$Y_{H,t+s}(i) = \left(\frac{P_{H,t}(i)}{P_{H,t+s}} \left(\frac{P_{H,t+s-1}}{P_{H,t-1}} \right)^{\delta_H} \right)^{-\varepsilon} (C_{H,t+s} + C^*_{H,t+s}), \tag{15}$$

and $\epsilon_{H,t}$ is a cost-push shock. The first-order condition is:

$$\mathbb{E}_t \sum_{s=0}^{\infty} Q_{t,t+s}\theta_H^s Y_{H,t+s}(i) \left[\tilde{P}_{H,t} \left(\frac{P_{H,t+s-1}}{P_{H,t-1}} \right)^{\delta_H} - \left(\frac{\varepsilon}{\varepsilon - 1} \right) P_{H,t+s} MC_{t+s} \exp(\epsilon_{H,t+s}) \right] = 0. \tag{16}$$

The aggregate price level evolves according to:

$$P_{H,t} = \left[(1-\theta_H)(\tilde{P}_{H,t})^{1-\varepsilon} + \theta_H \left(P_{H,t-1} \left(\frac{P_{H,t-1}}{P_{H,t-2}} \right)^{\delta_H} \right)^{1-\varepsilon} \right]^{\frac{1}{1-\varepsilon}}. \tag{17}$$

2.3 Retail Firms

The retail market is comprised of a continuum of monopolistically competitive firms $j \in [0,1]$ that import differentiated goods from abroad. Similar to domestic firms, retail firms also set prices according to Calvo (1983) where in each period there is a constant probability $1 - \theta_F$ that a retail firm will be randomly selected to adjust its price.[11] Faced with changing its price at time t, a retail firm j importing a good at cost $\tilde{e}_t P^*_{F,t}(j)$ chooses $P_{F,t}(j)$ to maximize its expected discounted value of profits:

$$\max_{P_{F,t}(j)} \mathbb{E}_t \sum_{s=0}^{\infty} Q_{t,t+s}\theta_F^s \left[P_{F,t}(j) \left(\frac{P_{F,t+s-1}}{P_{F,t-1}} \right)^{\delta_F} - \tilde{e}_{t+s} P^*_{F,t+s}(j) \exp(\epsilon_{F,t+s}) \right] Y_{F,t+s}(j), \tag{18}$$

where

$$Y_{F,t+s}(j) = \left[\frac{P_{F,t}(j)}{P_{F,t+s}} \left(\frac{P_{F,t+s-1}}{P_{F,t-1}} \right)^{\delta_F} \right]^{-\varepsilon} C_{F,t+s}, \tag{19}$$

and $\epsilon_{F,t}$ is a cost-push shock to import retailers. The first-order condition is given by:

$$\mathbb{E}_t \sum_{s=0}^{\infty} Q_{t,t+s}\theta_F^s Y_{F,t+s}(j) \left[\tilde{P}_{F,t} \left(\frac{P_{F,t+s-1}}{P_{F,t-1}} \right)^{\delta_F} - \left(\frac{\varepsilon}{\varepsilon - 1} \right) \tilde{e}_{t+s} P^*_{F,t+s}(j) \exp(\epsilon_{F,t+s}) \right] = 0, \tag{20}$$

and the aggregate price index for imports:

$$P_{F,t} = \left[(1-\theta_F)(\tilde{P}_{F,t})^{1-\varepsilon} + \theta_F \left(P_{F,t-1} \left(\frac{P_{F,t-1}}{P_{F,t-2}} \right)^{\delta_F} \right)^{1-\varepsilon} \right]^{\frac{1}{1-\varepsilon}}. \tag{21}$$

[11] The parameter θ_F governs the degree of ERPT.

2.4 Market Clearing

Goods market clearing for domestic firms requires:

$$\begin{aligned} Y_{H,t}(i) = C_{H,t}(i) + C^*_{H,t}(i) = \left(\frac{P_{H,t}(i)}{P_{H,t}}\right)^{-\varepsilon} \left[C_{H,t} + C^*_{H,t}\right], \\ \Rightarrow Y_t \equiv \int_0^1 Y_{H,t}(i)di = C_{H,t} + C^*_{H,t}, \end{aligned} \tag{22}$$

where

$$C^*_{H,t} = \alpha \left(\frac{P^*_{H,t}}{P^*_t}\right)^{-\eta} C^*_t \quad \text{and} \quad Y^*_t = C^*_t.$$

Market clearing for domestic bonds requires:

$$B_t = 0. \tag{23}$$

2.5 The Log-Linearized Model

The model is log-linearized around a deterministic zero-inflation steady state where bond holdings are zero and the terms of trade are equal to $S_{ss} = 1$. Let lowercase letters denote the log-deviations of the respective variables from their steady-state values: i.e., $x_t = \ln(X_t/X_{ss})$. Log-linearizing the consumption Euler equation of the domestic household (10) yields:

$$c_t - hc_{t-1} = \mathbb{E}_t(c_{t+1} - hc_t) - \frac{1-h}{\sigma}(r_t - \mathbb{E}_t\pi_{t+1}) - \frac{1-h}{\sigma}(\mathbb{E}_t\epsilon_{g,t+1} - \epsilon_{g,t}). \tag{24}$$

Log-linearizing (16) and (17) gives the aggregate supply condition for domestic goods:

$$\pi_{H,t} - \delta_H\pi_{H,t-1} = \beta\mathbb{E}_t(\pi_{H,t+1} - \delta_H\pi_{H,t}) + \frac{(1-\beta\theta_H)(1-\theta_H)}{\theta_H}(mc_t + \epsilon_{H,t}), \tag{25}$$

where $\pi_{H,t} = p_{H,t} - p_{H,t-1}$ and

$$mc_t = \varphi y_t - (1+\varphi)\epsilon_{a,t} + \alpha s_t + \frac{\sigma}{1-h}(c_t - hc_{t-1}),$$

which is obtained after combining (9), the aggregate version of (12), (13) and noting that the log-linearized version of the CPI index (4) implies $p_t - p_{H,t} = \alpha s_t$ after using (6).

Log-linearizing (20) and (21) gives the aggregate supply condition for retail goods:

$$\pi_{F,t} - \delta_F\pi_{F,t-1} = \beta\mathbb{E}_t(\pi_{F,t+1} - \delta_F\pi_{F,t}) + \frac{(1-\beta\theta_F)(1-\theta_F)}{\theta_F}(\psi_{F,t} + \epsilon_{F,t}), \tag{26}$$

where $\pi_{F,t} = p_{F,t} - p_{F,t-1}$ and the law of one price gap $\psi_{F,t}$ is defined as:

$$\psi_{F,t} \equiv e_t + p^*_t - p_{F,t}.$$

Log linearizing equations (4)–(6) and using the above definition of $\psi_{F,t}$ yields the following relationship for the real exchange rate and the terms of trade:

$$q_t = e_t + p_t^* - p_t = \psi_{F,t} + (1-\alpha)s_t. \tag{27}$$

First-differencing the log-linearized version of equation (6) yields:

$$s_t - s_{t-1} = \pi_{F,t} - \pi_{H,t} + \epsilon_{s,t}, \tag{28}$$

where $\epsilon_{s,t}$ is an exogenous terms of trade shock, and first-differencing the log-linearized version of the CPI index (4) gives:

$$\pi_t = (1-\alpha)\pi_{H,t} + \alpha\pi_{F,t}, \tag{29}$$

where $\pi_t = p_t - p_{t-1}$.

The real interest rate parity condition is obtained by first-differencing (27) and combining with the log-linearized version of (11):

$$(r_t - \mathbb{E}_t\pi_{t+1}) - (r_t^* - \mathbb{E}_t\pi_{t+1}^*) = \mathbb{E}_t(q_{t+1} - q_t) - \chi(d_t + \epsilon_{q,t}). \tag{30}$$

The disturbance term $\epsilon_{q,t}$ captures time-varying deviations from real interest rate parity. Log-linearizing the budget constraint (8) implies:[12]

$$c_t + d_t = \frac{d_{t-1}}{\beta} - \alpha(s_t + \psi_{F,t}) + y_t, \tag{31}$$

where $d_t = \log(D_t) \equiv \log(\tilde{e}_t B_t^*/Y_{ss}P_t)$ is domestic-currency real foreign bond holdings (relative to steady state output). Finally, the goods market clearing condition (22) implies

$$y_t = (1-\alpha)c_t + \alpha\eta q_t + \alpha\eta s_t + \alpha y_t^*. \tag{32}$$

We assume that the stochastic processes for preferences, technology, the terms of trade, and risk-premium follow an independent AR(1) process:

$$\epsilon_{x,t} = \rho_x\epsilon_{x,t-1} + v_{x,t}, \text{ where } \rho_x \in (0,1), v_x \sim iid(0,\sigma_x^2) \tag{33}$$

for $x = g, a, s, q$, and the cost-push shocks in the domestic and retail sectors follow an i.i.d. process: $\epsilon_H \sim$ i.i.d.(0, σ_H) and $\epsilon_F \sim$ i.i.d. (0, σ_F). Following Kam et al. (2009), we further assume that the foreign country variables $\{\pi^*, y^*, r^*\}$ follow uncorrelated AR(1) processes:

$$\begin{pmatrix} \pi_t^* \\ y_t^* \\ r_t^* \end{pmatrix} = \begin{pmatrix} a_1 & 0 & 0 \\ 0 & b_2 & 0 \\ 0 & 0 & c_3 \end{pmatrix} \begin{pmatrix} \pi_{t-1}^* \\ y_{t-1}^* \\ r_{t-1}^* \end{pmatrix} + \begin{pmatrix} \sigma_{\pi^*} & 0 & 0 \\ 0 & \sigma_{y^*} & 0 \\ 0 & 0 & \sigma_{r^*} \end{pmatrix} \begin{pmatrix} v_{\pi^*,t} \\ v_{y^*,t} \\ v_{r^*,t} \end{pmatrix} \tag{34}$$

where $v_{\pi^*,t},\ v_{y^*,t},\ v_{r^*,t} \sim N(0, I_3)$.

[12] Similar to Justiniano and Preston (2010), in equilibrium household nominal income $W_tN_t + \Pi_t = P_{H,t}Y_t + (P_{F,t} - \tilde{e}_tP_t^*)C_{F,t}$.

Given the specification for monetary policy, the processes for $\{\epsilon_{a,t},\ \epsilon_{g,t},\ \epsilon_{q,t},\ \epsilon_{s,t}\}$ and $\{\pi_t^*,\ y_t^*,\ r_t^*\}$ described by (33) and (34), and the cost-push shocks $\{\epsilon_{H,t},\ \epsilon_{F,t}\}$, the system of equations (24)–(32) determines the following ten endogenous variables $\{c_t,\ y_t,\ d_t,\ q_t,\ s_t,\ \psi_{F,t},\ r_t,\ \pi_t,\ \pi_{H,t},\ \pi_{F,t}\}$.

2.6 Central Bank Preferences

As is standard in the literature, we assume that the central bank minimizes a one-period ad-hoc quadratic loss function where monetary policy targets inflation, the output gap, and interest rate smoothing.[13] In addition, following Kam et al. (2009) the central bank can also target the real exchange rate. Consequently, the loss function is given by:

$$L(\tilde{\pi}_t, y_t, q_t, r_t - r_{t-1}) = \frac{1}{2}\left[\tilde{\pi}_t^2 + \mu_y y_t^2 + \mu_q q_t^2 + \mu_r (r_t - r_{t-1})^2\right]. \tag{35}$$

The weight assigned to the annual inflation rate $\tilde{\pi}_t \equiv \sum_{i=0}^{3} \pi_{t-i}/4$ is normalized to one and the weights $\mu_y, \mu_q, \mu_r \in [0, +\infty)$ represent the relative importance assigned to output gap stabilization, RER stabilization and interest rate smoothing. The loss function specification given by (35) is consistent with flexible inflation targeting as described by Svensson (1999). Interest rate smoothing is included to capture monetary policy inertia.[14] As discussed by Svensson (2000), the RER plays a prominent role in the monetary policy transmission mechanism in SOEs.

We further assume that the central bank minimizes (35) subject to the structural equations (24)–(32) *under discretion*.[15] We employ the algorithm of Dennis (2007) to compute solutions to a linear-quadratic Markov perfect equilibrium (LQ-MPE) problem.[16] Following Kam et al. (2009), we add a noise term $\epsilon_{r,t} \sim N(0, \sigma_r^2)$ to the resulting optimal interest rate rule $r_t(\epsilon_t, z_{t-1})$ to capture imperfections in the setting of interest rates (i.e., an exogenous monetary policy shock).

3 Estimation

3.1 Data

To estimate the model we use quarterly data for each of the five LAIT countries. The *Foreign* economy is proxied by the United States (US). All data were downloaded from the IMF's *International Financial Statistics*, the OECD's *National Accounts*, and statistical tables published by the central bank of each country. Since the LAIT countries switched to IT at different periods during the late 1990s and early 2000s (see Table 1), the sample period differs for each country. To remove any country-specific noise in the data, the first few years of data after the adoption

[13] An alternative approach would be to derive an approximate welfare-based loss function using the preferences of the household. While this approach is theoretically appealing (see, e.g., Woodford, 2003; Galí, 2015), it does not carry over easily to open economies (see, e.g., Benigno and Benigno, 2003; Monacelli, 2005). In particular, an accurate quadratic approximation of household welfare can be obtained in open economy models only under special assumptions on household preferences and on the value of the trade elasticity parameter η.

[14] See Ilbas (2012) and McKnight and Mihailov (2015) for further discussion on the reasons behind its inclusion.

[15] When solving (35) under discretion, the central bank treats the problem as one of sequential optimization, whereas under commitment, the central bank credibly commits to a policy plan.

[16] For further details, see Dennis (2007) and Kam et al. (2009).

of IT were also omitted from the sample. Specifically, the sample period for each country used in the estimations is as follows: 2004:1 – 2014:4 for Brazil, 2002:1 – 2014:4 for Chile, 2003:1 – 2014:4 for Colombia, 2002:1 – 2014:4 for Mexico, and 2005:1 – 2014:4 for Peru.

As the model features 10 exogenous shock processes $\{\epsilon_{a,t},\ \epsilon_{g,t},\ \epsilon_{q,t},\ \epsilon_{r,t},\ \epsilon_{s,t},\ \epsilon_{H,t},\ \epsilon_{F,t},\ v_{\pi^*,t},\ v_{r^*,t},\ v_{y^*,t}\}$, 10 observable time series are needed to avoid stochastic singularity. Our data set contains the following 10 observable variables: imported goods inflation denominated in domestic currency ($\pi_{F,t}$), the terms of trade (measured as the price of imports to exports) (s_t), the real exchange rate (computed using the nominal exchange rate defined as national currency per 1 USD) (q_t), domestic real GDP (y_t), domestic CPI inflation (π_t), the nominal interest rate (r_t), US CPI inflation (π_t^*), US real output (y_t^*), and the US federal funds rate (r_t^*). Consistent with the definition in (7), the foreign bond holdings ratio (d_t) is proxied by the real value of international reserves of each country converted into national currency by the nominal exchange rate and divided by the Hodrick-Prescott trend in output. All variables are expressed in logs and detrended using the Hodrick–Prescott filter, except inflation rates and interest rates, which are expressed in quarterly percentage change. Since monetary policy in our framework is driven using an output gap methodology, for y_t we use the Hodrick–Prescott filter to construct an output gap of deviations from the trend. As is customary in the estimation of DSGE models analysis, all variables, including those in percentage terms, are demeaned to approximate theoretical deviations from the steady state.

In order to assess whether a positive or zero weight for the real exchange rate is more probable for the policy objectives of each central bank, we follow Kam et al. (2009) in estimating two versions of the model, $\mu_q > 0$ versus $\mu_q = 0$ in (35), to see which is more probable (given the same observables and shocks) via the comparison of Bayesian posterior odds.

3.2 Methodology and Prior Selection

The model M is estimated using Bayesian methods.[17] We update the *a priori* beliefs about the parameter vector θ, represented by the prior density $p(\theta|M)$ in view of the information contained in the observed sample Y.[18] According to Bayes Theorem (see, e.g., Herbst and Schorfheide, 2016),

$$p(\theta|Y,M) = \frac{p(Y|\theta,M)p(\theta|M)}{\int p(Y|\theta,M)p(\theta|M)d\theta}, \tag{36}$$

this updating generates a posterior distribution (or likelihood function) $p(\theta|Y,M)$. The denominator in (36) is commonly known as the marginal likelihood of the data (or marginal data density) associated with M. As discussed by Herbst and Schorfheide (2016), among others, Bayesian inference amounts to characterizing the properties of the posterior distribution $p(\theta|Y,M)$. Usually, posterior samplers are employed that generate sequences of draws θ^j, $j = 1,...,J$ from $p(\theta|Y,M)$. As is common in the literature, we apply the Random-Walk Metropolis-Hastings (RWMH) Markov Chain Monte Carlo (MCMC) algorithm to obtain draws from the posterior

[17]Bayesian methods are described in detail in Gelman et al. (2004) and Koop (2006), among others. Their application to DSGE models has been expanding rapidly and includes key references such as Smets and Wouters (2003, 2007), An and Schorfheide (2007), Fernández-Villaverde et al. (2010), DeJong and Dave (2011), Del Negro and Schorfheide (2011), Miao (2014), Herbst and Schorfheide (2016), Fernández-Villaverde et al. (in press).

[18]The parameter vector θ describes the preferences, technology, central bank policy weights, and exogenous shock processes of the model M.

distribution.[19] For each country, 2000000 RWMH-MCMC draws and 2500 Kalman filter iterations were obtained, where the first half of the draws was discarded (or burnt-in) in order to remove initial condition effects.

As the posterior density (36) is derived by combining the prior density $p(\theta|M)$ with the likelihood function $p(Y|\theta, M)$, the selection of priors for each parameter plays a fundamental role in Bayesian estimation. The priors used in our estimates are summarized in Table 2. As is customary, we conform to the established conventions in selecting the prior densities: we use the beta distribution for parameters in the interval $[0, 1]$, the inverse gamma distribution for the standard deviations of the stochastic innovations $[0, \infty)$, and the gamma distribution for the rest.

Due to limited information in the data set, some structural parameters cannot be estimated with sufficient precision, and were therefore calibrated prior to estimation. For each country, we calibrate the import share in domestic consumption, α, to values corresponding to the sample average share of imports of goods and services in consumption: we set α to 0.20 for Brazil; 0.51 for Chile; 0.29 for Colombia; 0.44 for Mexico; and 0.35 for Peru.[20] As is common in the literature, for all countries the discount factor (β) is fixed at 0.99 and the debt-elastic interest rate parameter (χ) is fixed at 0.05 consistent with the estimates of Selaive and Tuesta (2003 a, b).[21]

In order not to add any prior information into the estimation stage, we follow Kam et al. (2009) and assume that the prior distributions for the central bank preference parameters μ_y, μ_q and μ_r are exactly the same. Consequently, any resulting differences in the posterior distributions of these (as well as the other) parameters will be due to the data itself.

4 Results

For each of the five LAIT countries, Table 3 compares the characteristics of the Bayesian RWMH-MCMC estimation for the two model versions, $\mu_q > 0$ and $\mu_q = 0$. This table also presents a summary of the posterior mean and standard deviation estimates for the central bank preferences obtained under each model version. Figures 1–7 depict the estimated posterior distributions for each structural parameter and tables 4–10 report the posterior estimates and convergence diagnostics for both the structural parameters and the persistence and standard deviations of the shocks.

4.1 Model Comparison, Posterior Shapes and Convergence Diagnostics

For each country, Table 3 compares selected characteristics of the Bayesian RWMH-MCMC estimation for the two model versions, $\mu_q > 0$ and $\mu_q = 0$. Comparison of the marginal likelihoods reported in Table 3 suggests that the central banks of Chile, Mexico and Peru are

[19] The pseudo-code for this popular algorithm is detailed in Appendix B of Kam et al. (2009).

[20] These values correspond to the average quarterly share (in our whole sample, 1999:1-2014:4) of real imports of goods and services in real consumption by country. Since direct information for the latter ratio is usually not released in statistical publications, we obtained it indirectly, by the ratio of the average quarterly share of real imports of goods and services in GDP to the average quarterly share of real consumption in real GDP.

[21] Using GMM, Selaive and Tuesta (2003 a, b) estimate χ to be in the range of 0.004 and 0.071 for a sample of OECD countries.

explicitly concerned with stabilizing the real exchange rate. In the case of Brazil and Colombia, the results suggest that the model $\mu_q = 0$ is a better fit of the data in terms of having a higher marginal likelihood.[22] For the $\mu_q > 0$ model version, with the exception of Brazil where the acceptance rate is low, the acceptance rates obtained for Chile, Colombia, Mexico, and Peru all fall within conventional range.[23]

For each model version, Table 3 also summarizes the parameter estimates associated with the loss function of each central bank. By inspection, while there is significant heterogeneity in terms of the specific parameter weights estimated for each country, three main conclusions arise. First, there is significant evidence that all five central banks, with a possible exception of Colombia, are concerned about smoothing the nominal interest rate. In particular, relative to the weight of inflation stabilization (normalized at 1), we find that Brazil (0.8 or 1.06 depending on the model version) and Peru (1.6) place very high weights on interest rate smoothing. Colombia is found to have the lowest weight in both model versions (0.19 or 0.05). For the three countries concerned about real exchange rate volatility (i.e., where the $\mu_q > 0$ model version is preferable in terms of marginal likelihoods), the estimates for Chile (0.11), and Peru (0.06) yield low weights suggesting the low importance of real exchange rate stabilization in these countries. In stark contrast, Mexico assigns a significant weight to real exchange rate stabilization (nearly half of that to inflation stabilization, which is greater than interest rate smoothing (0.38) and output gap stabilization (0.3). Third, while Brazil, Colombia, and Peru show little concern for output gap stabilization, Chile (0.24) and Mexico (0.3) place a sizeable weight on it. Overall, there is evidence to suggest that the central banks of Brazil, Colombia, and Peru implement a strict inflation targeting policy with small weights assigned to real exchange rate stabilization and output gap stabilization, whereas the central banks of Chile and Mexico appear to be more flexible in terms of their inflation targeting preferences.

Figures 1–7 show both the assumed prior (dashed curve) and the estimated posterior (solid curve) distributions (also indicating the posterior mean by the vertical line) for each structural parameter for the five LAIT economies. For Brazil and Colombia we report both model versions in the figures that depict the posterior distributions, whereas for the remaining three LAIT countries we only report the $\mu_q > 0$ case.[24] By inspection, the posterior distributions are generally unimodal and nicely shaped in most cases.

Tables 4–10 report the posterior mean and standard deviation estimates, the 95% confidence sets for the posterior estimates, and selected diagnostic tests for MCMC convergence. As before, for Brazil and Colombia we report both model versions in the tables and for the remaining three LAIT countries we only report the $\mu_q > 0$ case.[25] In these tables, NSE stands for the numerical standard error, which approximates the true posterior moment as proposed by Geweke (1992). The NSE reported for each country uses an 8% autocovariance tapered estimate. The G p-values report the p-value associated with Geweke's (1992) chi-squared convergence

[22] This finding is further corroborated by the posterior densities reported in Table 3, which are very close to zero for Brazil and Colombia in the $\mu_q > 0$ version of the model.

[23] The Bayesian literature has no recommendation for a single value of an "optimal" acceptance rate. Common practice considers the range from about 20% to 50% to be the most reliable (see, e.g., Koop, 2006, or Herbst and Schorfheide, 2016).

[24] Results for all model versions for each country are available upon request.

[25] Results for the $\mu_q = 0$ model version are available upon request.

test.[26] The last column of tables 4–10 reports the Brooks-Gelman (B-G) univariate shrink factor proposed by Gelman and Rubin (1992) and extended by Brooks and Gelman (1998).[27] By inspection of tables 4–10, we can conclude that our estimation results are overall satisfactory. The reported convergence test statistics for the estimated parameters indicate that in general the latter converge to an invariant distribution.[28] There are few exceptions, however, where both the Geweke chi-squared convergence test and the univariate shrink B-G factor both agree on problems with convergence: δ_F for Brazil (for the $\mu_q = 0$ model); ρ_q for Chile; σ, θ_F, σ_F and σ_s for Colombia (for the $\mu_q = 0$ model); σ_H, σ_F, σ_a and σ_s for Mexico; σ_H and σ_{π^*} for Peru.[29]

4.2 Estimated Structural Parameters Influencing the Endogenous Propagation

We now check if the estimated key structural parameters are economically plausible and how they compare to analogous earlier estimates.

We obtain estimates for habit persistence that are broadly consistent with the literature: posterior mean of 0.87 for Brazil and 0.75 for Colombia for the preferred model version in the case of these two counties ($\mu_q = 0$), and of 0.77 for Chile and 0.92 for Mexico for the preferred $\mu_q > 0$ version. Peru, however, is an exception, with a low estimated habit persistence of 0.26. Our estimates of the coefficient of relative risk aversion vary considerably across the five LAIT countries: 0.97 in Chile and Mexico, 1.48 in Colombia and 2.7 in Brazil. Peru is again an exception, with extremely low degree of relative risk aversion, 0.07. The inverse of the Frish elasticity of labor supply is estimated within the usual range, 1.38 in Brazil, 1.46 in Chile, 1.36 in Colombia, 1.45 in Mexico, and 1.80 in Peru. With the exception of Peru, the estimated elasticity of substitution between home and foreign goods is (roughly) within the typical range of estimates found in the empirical literature of 0.1–2.0 (see, for example, Corsetti et al., 2008): 1.3 for Brazil, 1.72 for Chile, 2.89 for Colombia and 2.19 for Mexico.

The degree of domestic-output versus import price stickiness is estimated at 0.93 versus 0.68 for Brazil; 0.78 in both sectors for Chile; 0.80 versus 0.70 for Colombia; 0.97 versus 0.96 for Mexico; and 0.26 versus 0.63 for Peru; Kam et al. (2009) find for all their sample countries a consistently higher degree of price stickiness for domestic output relative to imported goods. We find little difference in the estimates of price stickiness across sectors for Chile and Mexico, whereas in Peru imported price stickiness is over twice higher in the imported sector. While there is evidence of a high degree of sticky prices in the domestic economy for Brazil, Colombia, Chile and Mexico, prices are much more flexible in Peru. The backward-lookingness of the NKPC for home prices versus that for import prices is estimated at 0.28 versus 0.34 in Brazil,

[26] If the Markov chain of draws has converged to a stable distribution, one would expect the means from the two halves of the generated sample to be statistically indistinguishable. The null hypothesis of the test is that the means are equal. Thus, a low p-value may indicate some evidence of problems in convergence.

[27] This test runs the chain two or more times from a widely-dispersed starting point to see if the Markov chain always converges to the same value. Commonly, a B-G factor below 1.1 is considered as little evidence of dispersion (see, e.g., Gelman et al., 2004).

[28] The MCMC diagnostics figures illustrating the convergence results discussed here are available upon request.

[29] Note that Kam et al. (2009) report similar problems for some of the parameters we enumerated for Australia, Canada, and New Zealand using a complete asset markets model.

0.24 versus 0.21 in Chile, 0.42 versus 0.15 in Colombia, 0.72 versus 0.38 in Peru, and 0.77 versus 0.65 in Mexico. For all five LAIT economies except Brazil, our findings reveal that backward-lookingness in the NKPC is higher in domestic versus imported price inflation.

4.3 Estimated Persistence and Volatility of the Latent Exogenous Shock Processes

We now turn to the sources of exogenous fluctuations in the LAIT economies, presenting our set of estimates for persistence and volatility of the structural shock processes we included in our model.[30]

As it was thus far, the estimated persistence of the exogenous shock processes reveals some common features as well as some country specificity. To start with the persistence parameter of the technology shock, its posterior mean is as follows: Brazil 0.98, Chile 0.46, Colombia 0.64, Mexico 0.51 and Peru 0.89. The estimated (posterior mean) persistence of the risk premium shock is, respectively: Brazil 0.80, Chile 0.86, Colombia 0.91, Mexico 0.99 and Peru 0.72; that of the ToT shock is: Brazil 0.99, Chile 0.96, Colombia 0.65, Mexico 0.24 and Peru 0.44; and the persistence of the preference shock is: Brazil 0.37, Chile 0.04, Colombia 0.04, Mexico 0.04, Peru 0.11. Across all five LAIT economies, the most persistent shock appears to be the risk premium shock, yet it does not dominate the persistence of some among the other shocks in each individual country. Overall, the least persistent shock in all five LAIT economies is the preference shock.

Looking now at the estimated volatility (measured by the posterior mean of the standard deviation) of the exogenous shock processes, we could summarize the following findings. The preference shock and the risk premium shock are among the most volatile in all countries except Peru, where the terms of trade shock is the most volatile. With the exception of Brazil, the terms of trade shock is of a comparable magnitude. Only in Mexico the technology shock plays a role in driving volatility (but is not precisely estimated in terms of the diagnostics). For Brazil, the cost-push shock in the imported goods sector appears to matter too, while for Peru it is the cost-push shock in the domestic economy and the foreign inflation shock that come out as sizable as well. Our estimates therefore show the importance in generating volatility of the preference shock, which we incorporated in the SOE model with incomplete asset markets as another realistic feature of the LAIT countries.

4.4 Complete versus Incomplete International Financial Markets

We now consider the role played by the assumption of incomplete international financial markets. More precisely, we investigate the sensitivity of our results for the estimated policy parameters to the asset market structure. We do this by reestimating the model under the alternative assumption of complete asset markets employed in the closely-related study of Kam et al. (2009).[31] Specifically, we follow Kam et al. (2009) in estimating the complete asset markets

[30] The prior and posterior figures per country, illustrating the results discussed in the present subsection, are available upon request.

[31] Palma and Portugal (2014) estimate the model of Kam et al. (2009) using Brazilian quarterly data (2000-2013).

model using 9 observables (as the variable d_t is now redundant) and 9 shocks (as the preference shock $\epsilon_{g,t}$ is now omitted). The key difference is that complete asset markets implies perfect risk sharing such that the real exchange rate equals the marginal rate of substitution in consumption across countries. Incomplete international financial markets breaks this link (see Chari et al., 2002, and Rabanal and Tuesta, 2010).

Table 11 summarizes the posterior mean and standard deviation for the central bank preferences for the two model versions, $\mu_q > 0$ and $\mu_q = 0$, estimated under complete asset markets.[32] In stark contrast with the results reported in Table 3 under incomplete asset markets, we find the following key differences. First, comparing the two model versions we now find for Peru the model $\mu_q = 0$ to be a better fit of the data (and not for Brazil and Colombia as under incomplete asset markets). Moreover, under complete asset markets both the central banks of Brazil[33] and Chile assign a sizeable weight (0.31) to real exchange rate stabilization, whereas Mexico, the only country to do so under incomplete asset markets, does not. Second, under complete asset markets Brazil (0.73) and Colombia (0.51) are concerned about output gap stabilization, instead of Chile and Mexico as under incomplete markets. Overall the central banks of Mexico and Peru appear to implement a strict inflation targeting policy if the model is estimated under complete asset markets, whereas Brazil, Colombia and Peru follow the same policy under incomplete asset markets. While the major similarity between the two asset market structures is the significant role of interest rate smoothing uncovered for all 5 LAIT central banks, the ranking of the central bank target weights changes, except for Colombia, according to the assumed structure of international financial markets.

The above exercise highlights the sensitivity of the Bayesian estimates for the central bank policy weights to the assumption of international asset markets (in)completeness. Similar to Rabanal and Tuesta (2010), we find that complete asset markets increase the degree of model misspecification and distort significantly the Bayesian estimates, in particular the reported policy parameters. For example, under incomplete asset markets we find no evidence that the central bank of Brazil engages in real exchange rate and output stabilization, whereas for the central bank of Mexico we find clear evidence of such concern. By contrast, under compete asset markets these conclusions are reversed. In so far that incomplete asset markets are a more realistic assumption, especially for developing countries, our analysis emphasizes the dangers of using the common complete asset markets assumption in Bayesian estimation, in our case focusing on uncovering central bank preferences.

5 Conclusions

The objective of this paper was to uncover and compare the central bank preferences of the big five LAIT countries using Bayesian estimation. We employed a medium-scale New Keynesian small open economy model that assumed incomplete international asset markets and imperfect ERPT. Optimal monetary policy was modeled under discretion, where the central

[32] The prior and posterior figures per country, illustrating the results discussed in the present subsection, are available upon request.

[33] This is consistent with the findings of Palma and Portugal (2014) for the same model but a different sample. However, they find a lower weight for output gap stabilization and a higher weight for interest rate inertia.

bank minimized an intertemporal quadratic loss function with four policy objectives: inflation control, output gap stabilization, real exchange rate volatility reduction and nominal interest rate smoothing. The weight attributed to each policy objective, which depends on the country-specific institutional preferences of each central bank, was represented in terms of Bayesian posterior distributions and convergence diagnostics.

The key insights of our analysis can be summarized as follows. The five LAIT economies we considered seem to fall broadly into two groups. The first group consists of Brazil, Colombia and Peru whose priority targets are to stabilize inflation with a significant degree of nominal interest rate smoothing, consistent with strict IT. The second group, Chile and Mexico, has broader policy objectives: both central banks additionally care about output gap stabilization, consistent with flexible IT; only Mexico further assigns a sizeable weight to reducing real exchange rate volatility. We also estimated the key structural parameters and the exogenous shocks. Our Bayesian estimates reveal that three of the ten considered shocks drive the fluctuations in the LAIT economies: the risk premium shock, the terms of trade shock and the preference shock.

We test the sensitivity of our results to the alternative assumption of complete asset markets and find that a number of our key policy conclusions are now reversed. This suggests that the degree of model misspecification depends on the asset market structure, highlighting the limitations of employing the assumption of complete asset markets in Bayesian estimation, especially when the aim is to uncover central bank preferences.

This work can be improved along various dimensions. It could be possible that preferences were not constant over this period. Personalities of governors, in addition to the legal mandate of the institution itself, could also shape the preferences of the central bank. For future research, it would therefore be interesting to study separate subperiods in the institutional history or in governor terms of office within a central bank, and compare across such subperiods the estimated policy preference parameters.

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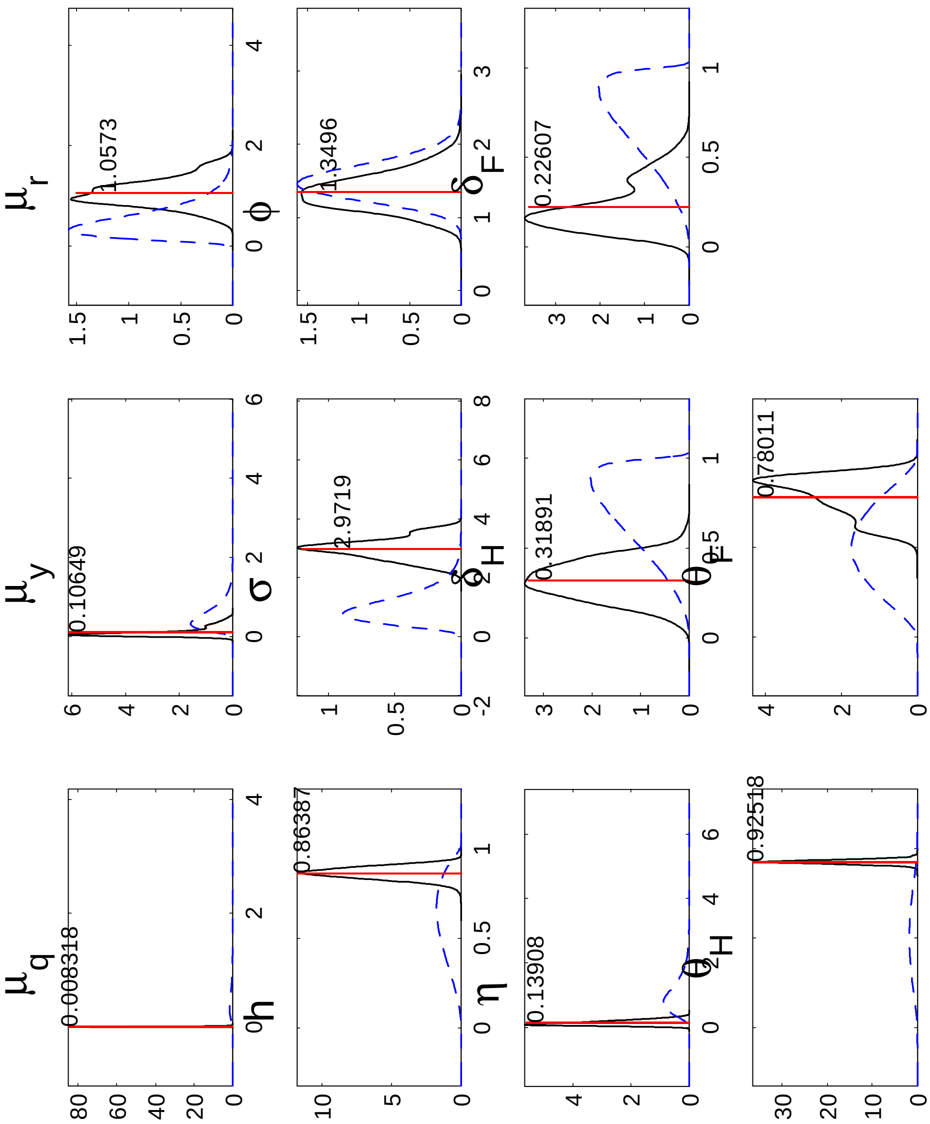

Figure 1: Brazil – Posterior Distributions of the Structural Parameters (Incomplete Markets, $\mu_q > 0$)

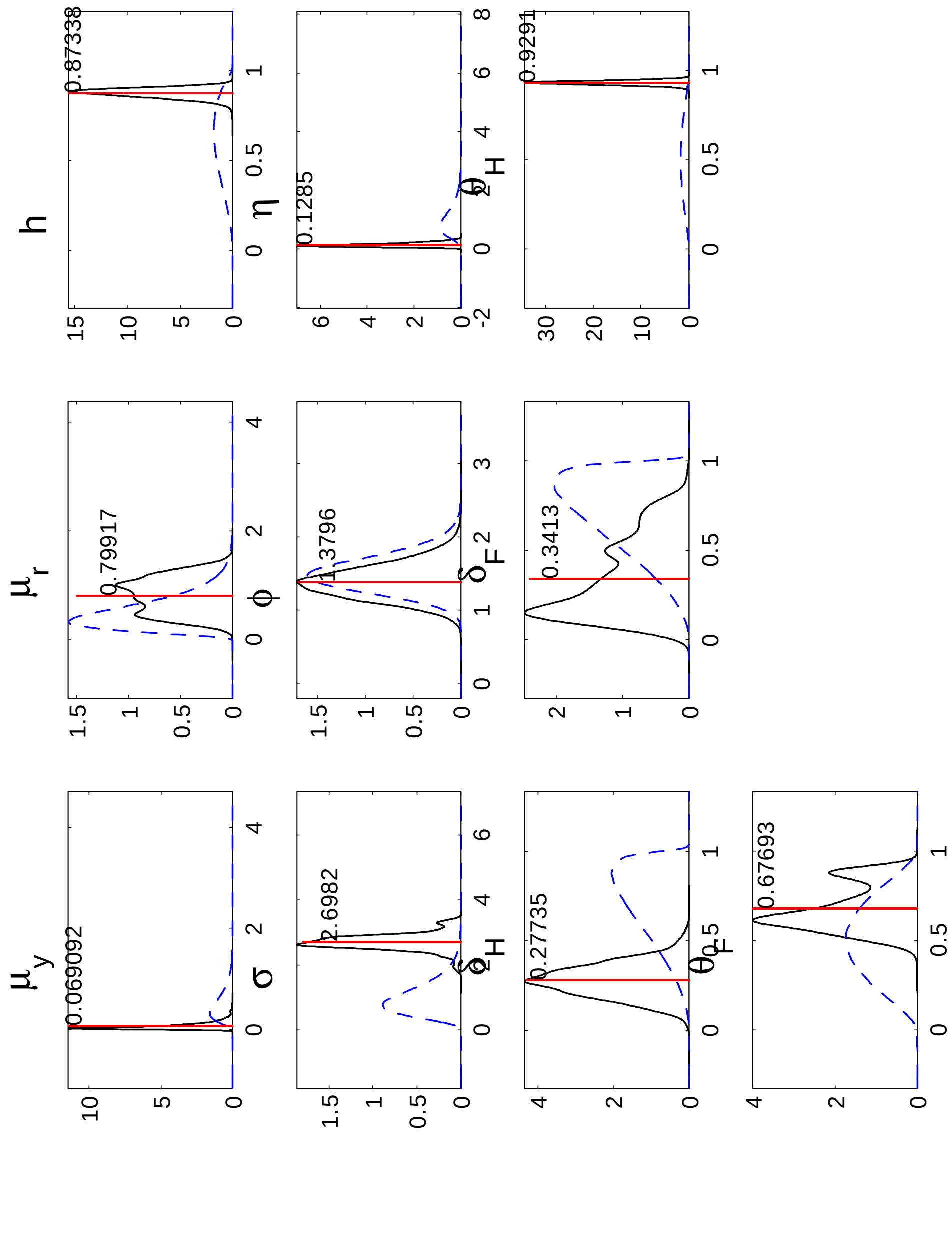

Figure 2: Brazil – Posterior Distributions of the Structural Parameters (Incomplete Markets, $\mu_q = 0$)

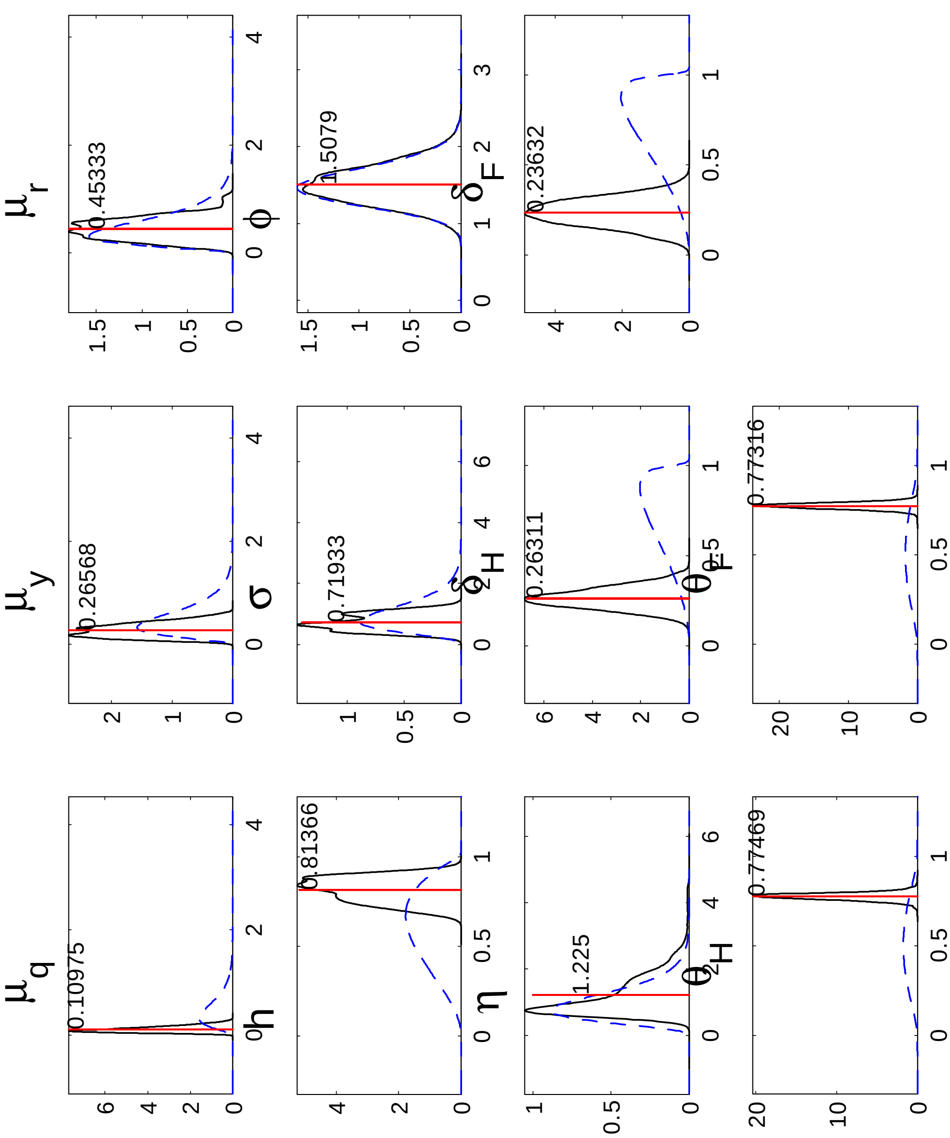

Figure 3: Chile – Posterior Distributions of the Structural Parameters (Incomplete Markets, $\mu_q > 0$)

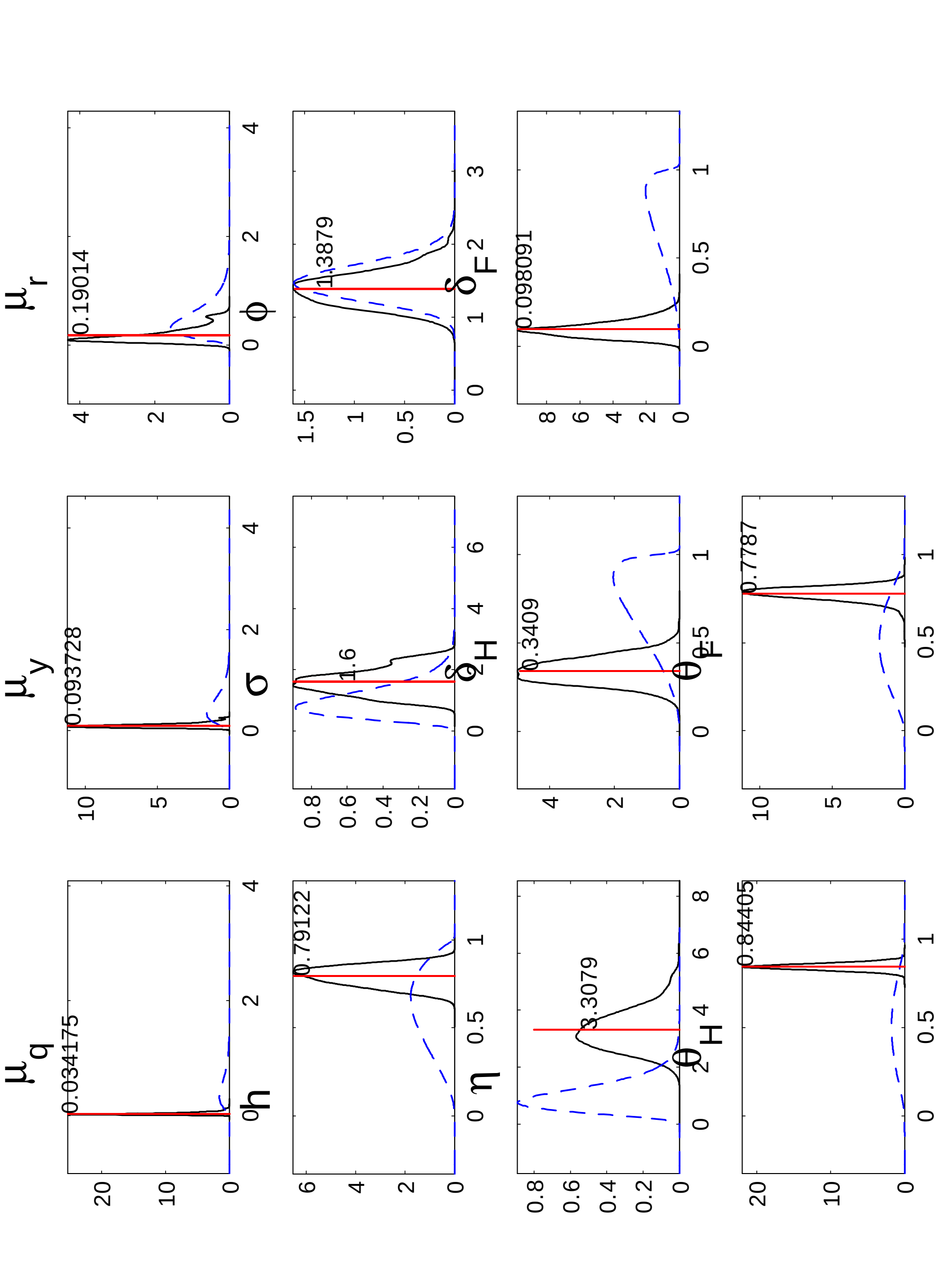

Figure 4: Colombia – Posterior Distributions of the Structural Parameters (Incomplete Markets, $\mu_q > 0$)

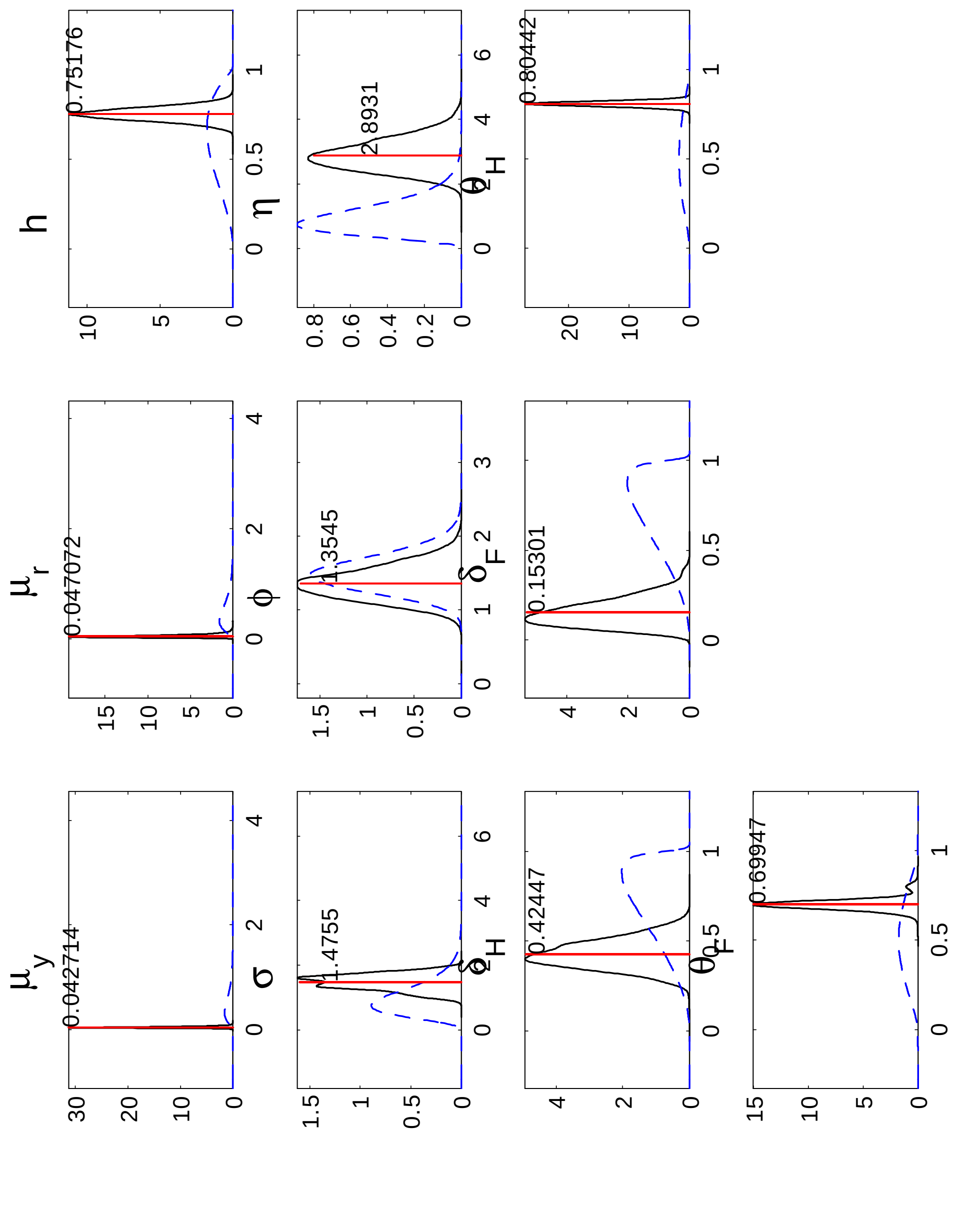

Figure 5: Colombia – Posterior Distributions of the Structural Parameters (Incomplete Markets, $\mu_q = 0$)

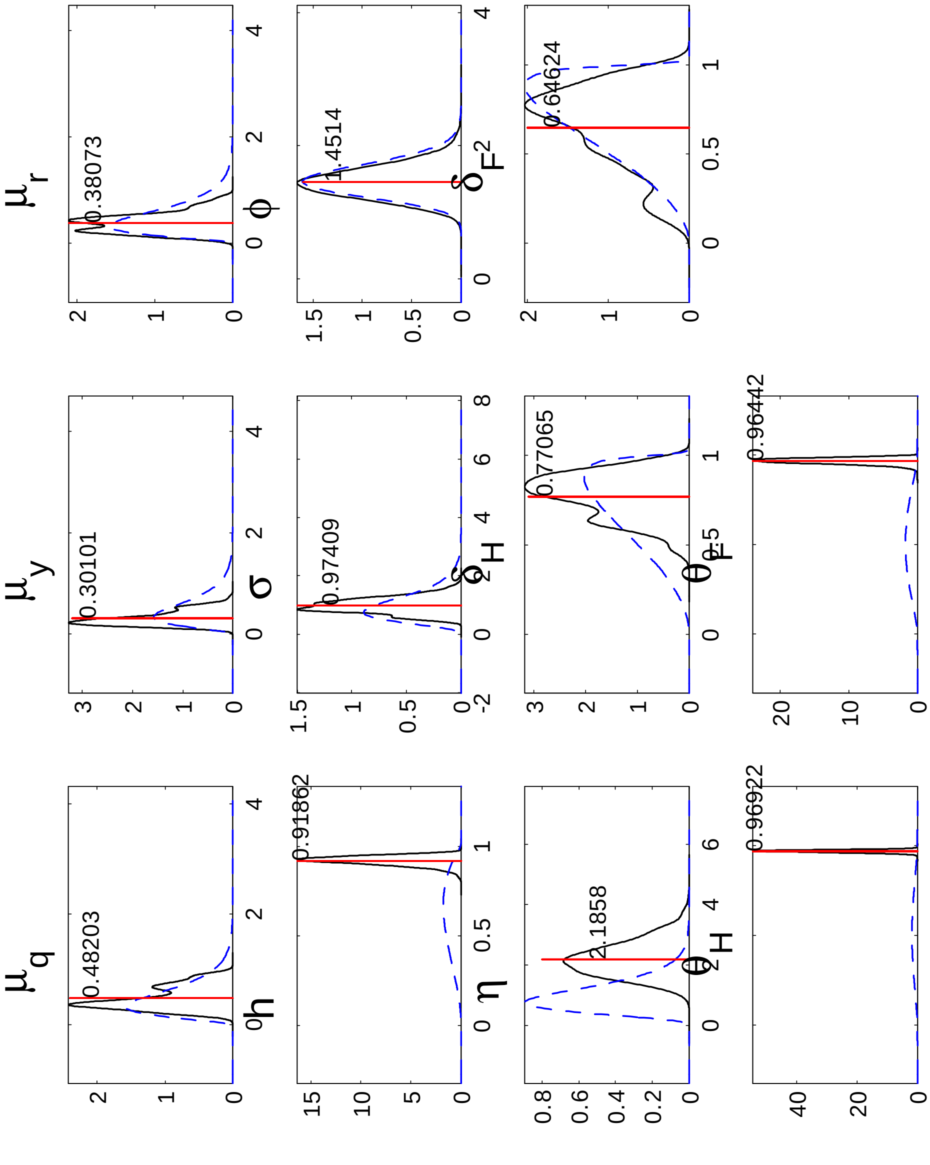

Figure 6: Mexico – Posterior Distributions of the Structural Parameters (Incomplete Markets, $\mu_q > 0$)

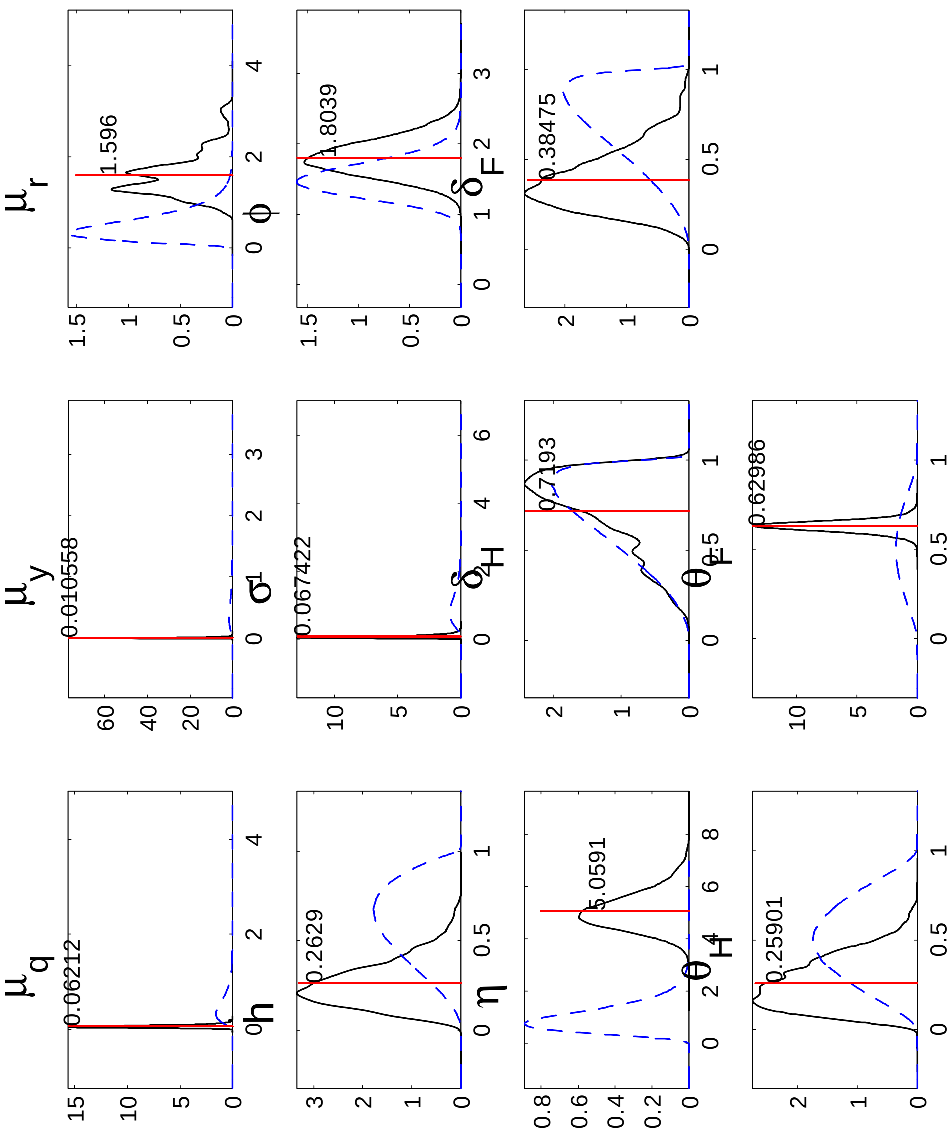

Figure 7: Peru – Posterior Distributions of the Structural Parameters (Incomplete Markets, $\mu_q > 0$)

Country	Exchange Rate Regime (2012)	Year of IT Adoption	Annual Inflation at IT Adoption	Average Quarterly Inlation since IT	Inflation Target (2012)
Brazil	Floating	1999	3.3	6.6	4.5 ± 1
Chile	Free floating	1999	3.2	3.2	3 ± 1
Colombia	Floating	1999	9.3	5.1	$2 - 4$
Guatemala	Stabilized arrangement	2005	9.2	5.4	5 ± 1
Mexico	Free floating	2001	9.0	4.3	3 ± 1
Peru	Floating	2002	-0.1	2.9	2 ± 1

Note: Annual inflation is expressed in % per annum. Average quarterly inflation is computed from the first quarter of the year following IT adoption through the last quarter of 2014, and is expressed as % per annum. The exchange-rate regime classification is taken from IMF (2012). The inflation target is extracted from Jahan (2012).

Table 1: The 6 Latin American Inflation Targeters: Some Basic Facts

Parameters	Description	Prior pdf	Prior Mean	Prior SD
	Structural Parameters			
h	degree of habit persistence	$\mathcal{B}$	0.6	0.2
σ	inverse of elasticity of intertemporal substitution in consumption	Γ	1.0	0.5
ϕ	inverse of Frish elasticity of intertemporal labor supply	Γ	1.5	0.25
η	elasticity of substitution between home and foreign goods	Γ	1.0	0.5
δ_H	degree of indexation in domestic-output markets	$\mathcal{B}$	0.7	0.2
δ_F	degree of indexation in imported-goods markets	$\mathcal{B}$	0.7	0.2
θ_H	degree of inflation persistence in domestic-output markets	$\mathcal{B}$	0.5	0.2
θ_F	degree of inflation persistence in imported-goods markets	$\mathcal{B}$	0.5	0.2
a_1	degree of persistence in foreign inflation	$\mathcal{B}$	0.5	0.2
b_2	degree of persistence in foreign output	$\mathcal{B}$	0.5	0.2
c_3	degree of persistence in foreign interest rate	$\mathcal{B}$	0.5	0.2
ρ_a	degree of persistence in technology shock	$\mathcal{B}$	0.5	0.2
ρ_q	degree of persistence in risk premium shock	$\mathcal{B}$	0.9	0.2
ρ_s	degree of persistence in terms of trade shock	$\mathcal{B}$	0.25	0.2
ρ_g	degree of persistence of preference shock	$\mathcal{B}$	0.5	0.2
	Relative Policy Target Weights			
μ_q	real exchange rate stabilization	Γ	0.5	0.3
μ_y	output gap stabilization	Γ	0.5	0.3
μ_r	interest rate smoothing	Γ	0.5	0.3
	Standard Deviation of Shock Innovations			
σ_H	domestic-output cost-push shock	Inverse Γ	0.5	0.25
σ_F	imported-goods cost-push shock	Inverse Γ	0.5	0.25
σ_a	technology shock	Inverse Γ	1.0	0.4
σ_q	risk premium shock	Inverse Γ	2.0	0.5
σ_s	terms of trade shock	Inverse Γ	1.0	0.4
σ_{π^*}	foreign inflation shock	Inverse Γ	1.0	0.4
σ_{y^*}	foreign output shock	Inverse Γ	1.0	0.4
σ_{r^*}	foreign interest rate shock	Inverse Γ	1.0	0.4
σ_r	interest rate shock	Inverse Γ	1.0	0.4
σ_g	preference shock	Inverse Γ	1.0	0.4

Note: Parameters calibrated to a value common to all countries: $\beta = 0.99$; $\chi = 0.05$. The parameter α is calibrated to a country-specific value. For the $\mu_q = 0$ model version, the prior and posterior distributions are degenerate at zero.

Table 2: Prior Distributions

	$\mu_q > 0$	$\mu_q = 0$
	Brazil (2004–2014)	
Marginal Likelihood	-1778.7	-1766.7
Acceptance Rate (%)	6.94	9.24
Indeterminacy Rate (%)	24.15	24.08
Invalid Likelihood Rate (%)	0.001208	0.000200
μ_q	0.01 (0.01)	0
μ_y	0.11 (0.11)	0.07 (0.07)
μ_r	1.06 (0.26)	0.80 (0.31)
	Chile (2002–2014)	
Marginal Likelihood	-1921.3	-1923.1
Acceptance Rate (%)	25.79	33.38
Indeterminacy Rate (%)	5.57	0.79
Invalid Likelihood Rate (%)	0.000273	0.001550
μ_q	0.11 (0.05)	0
μ_y	0.24 (0.10)	0.08 (0.04)
μ_r	0.51 (0.20)	0.25 (0.17)
	Colombia (2003–2014)	
Marginal Likelihood	-1799.3	-1793.0
Acceptance Rate (%)	20.31	5.54
Indeterminacy Rate (%)	5.10	7.03
Invalid Likelihood Rate (%)	0.000255	0.000350
μ_q	0.03 (0.02)	0
μ_y	0.09 (0.04)	0.043 (0.014)
μ_r	0.19 (0.14)	0.047 (0.027)
	Mexico (2002–2014)	
Marginal Likelihood	-1759.6	-1788.9
Acceptance Rate (%)	31.73	9.82
Indeterminacy Rate (%)	7.76	14.84
Invalid Likelihood Rate (%)	0.000388	0.007450
μ_q	0.48 (0.20)	0
μ_y	0.30 (0.14)	0.024 (0.01)
μ_r	0.38 (0.18)	1.50 (0.53)
	Peru (2005–2014)	
Marginal Likelihood	-1610.1	-1700.7
Acceptance Rate (%)	37.32	3.41
Indeterminacy Rate (%)	0.11	32.06
Invalid Likelihood Rate (%)	0.000006	0.000250
μ_q	0.06 (0.03)	0
μ_y	0.01 (0.01)	0.006 (0.005)
μ_r	1.60 (0.46)	0.037 (0.029)

Note: Estimates of the policy weights (relative to that of inflation, 1), μ_i with $i = \{q, y, r\}$, report the posterior mean, with the posterior standard deviation in parentheses.

Table 3: Model Comparison and Estimated Policy Weights

Parameters	Post Mean	Post SD	2.5%	97.5%	NSE (8%)	G p-value	B-GF
Structural Parameters							
h	0.86	0.03	0.19	0.93	0.01	0.11	1.03
σ	2.97	0.35	0.27	2.19	0.09	0.49	1.01
ϕ	1.35	0.24	1.05	2.03	0.02	0.07	1.01
η	0.14	0.07	0.27	2.19	0.01	0.09	1.03
δ_H	0.32	0.10	0.25	0.98	0.01	0.67	1.00
δ_F	0.23	0.13	0.25	0.98	0.02	0.12	1.04
θ_H	0.93	0.01	0.13	0.87	0.00	0.13	1.01
θ_F	0.78	0.10	0.13	0.87	0.02	0.46	1.01
a_1	0.63	0.12	0.19	0.96	0.01	0.86	1.00
b_2	0.82	0.06	0.19	0.96	0.00	0.08	1.00
c_3	0.99	0.01	0.19	0.96	0.00	0.37	1.00
ρ_a	0.99	0.01	0.13	0.87	0.00	0.72	1.00
ρ_q	0.89	0.21	0.23	1.00	0.05	0.29	1.02
ρ_s	0.99	0.01	0.01	0.72	0.00	0.51	1.00
ρ_g	0.38	0.05	0.13	0.87	0.00	0.64	1.00
Relative Policy Target Weights							
μ_q	0.01	0.01	0.09	1.24	0.00	0.01	1.05
μ_y	0.11	0.11	0.09	1.24	0.02	0.02	1.09
μ_r	1.06	0.26	0.09	1.24	0.06	0.45	1.01
Standard Deviation of Shock Innovations							
σ_H	2.07	1.31	0.91	7.33	0.31	0.00	1.25
σ_F	3.99	3.48	0.91	7.34	0.94	0.00	2.11
σ_a	5.18	1.87	0.52	2.66	0.45	0.04	1.07
σ_q	23.73	1.16	0.32	0.87	0.25	0.83	1.00
σ_s	13.39	2.88	0.52	2.66	0.76	0.25	1.02
σ_{π^*}	0.51	0.15	0.52	2.65	0.01	0.06	1.01
σ_{y^*}	5.83	1.01	0.52	2.65	0.19	0.71	1.00
σ_{r^*}	0.28	0.03	0.52	2.66	0.00	0.04	1.00
σ_r	0.58	0.07	0.52	2.66	0.00	0.74	1.00
σ_g	24.80	0.17	0.52	2.66	0.02	0.69	1.00

Note: Calibrated parameters: $\beta = 0.99$; $\chi = 0.05$; $\alpha = 0.20$.

Table 4: Brazil: Posterior Parameters and Convergence Diagnostics ($\mu_q > 0$)

Parameters	Post Mean	Post SD	2.5%	97.5%	NSE (8%)	G p-value	B-GF
			Structural Parameters				
h	0.87	0.03	0.19	0.93	0.00	0.46	1.00
σ	2.70	0.26	0.27	2.19	0.05	0.46	1.01
ϕ	1.38	0.23	1.05	2.03	0.01	0.16	1.00
η	0.13	0.07	0.27	2.20	0.01	0.99	1.00
δ_H	0.28	0.09	0.25	0.98	0.01	0.72	1.00
δ_F	0.34	0.21	0.25	0.98	0.05	0.00	1.14
θ_H	0.93	0.01	0.13	0.87	0.00	0.38	1.01
θ_F	0.68	0.12	0.13	0.87	0.03	0.15	1.03
a_1	0.66	0.12	0.19	0.96	0.01	0.15	1.00
b_2	0.82	0.06	0.19	0.96	0.00	0.78	1.00
c_3	0.98	0.01	0.19	0.96	0.00	0.43	1.00
ρ_a	0.98	0.01	0.13	0.87	0.00	0.02	1.02
ρ_q	0.80	0.28	0.23	1.00	0.06	0.32	1.02
ρ_s	0.99	0.01	0.01	0.73	0.00	0.00	1.03
ρ_g	0.37	0.05	0.13	0.87	0.00	0.45	1.00
			Relative Policy Target Weights				
μ_y	0.07	0.07	0.09	1.24	0.01	0.04	1.05
μ_r	0.80	0.31	0.10	1.23	0.07	0.79	1.00
			Standard Deviation of Shock Innovations				
σ_H	1.71	1.32	0.91	7.37	0.27	0.11	1.04
σ_F	16.91	1.97	0.92	7.37	0.49	0.09	1.05
σ_a	4.29	1.70	0.52	2.66	0.42	0.35	1.02
σ_q	23.47	1.21	0.32	0.88	0.24	0.00	1.26
σ_s	10.06	2.43	0.52	2.66	0.61	0.62	1.01
$\sigma_{\pi*}$	0.50	0.15	0.52	2.66	0.01	0.45	1.00
σ_{y*}	5.14	0.88	0.52	2.66	0.15	0.16	1.03
σ_{r*}	0.27	0.04	0.52	2.67	0.00	0.48	1.00
σ_r	0.58	0.07	0.52	2.66	0.00	0.52	1.00
σ_g	24.81	0.19	0.52	2.66	0.01	0.50	1.00

Note: Calibrated parameters: $\beta = 0.99$; $\chi = 0.05$; $\alpha = 0.20$.

Table 5: Brazil: Posterior Parameters and Convergence Diagnostics ($\mu_q = 0$)

Parameters	Post Mean	Post SD	2.5%	97.5%	NSE (8%)	G p-value	B-GF
			Structural Parameters				
h	0.77	0.05	0.19	0.93	0.01	0.08	1.04
σ	0.97	0.26	0.27	2.19	0.06	0.02	1.10
ϕ	1.46	0.24	1.05	2.03	0.02	0.75	1.00
η	1.72	0.66	0.27	2.19	0.12	0.82	1.00
δ_H	0.24	0.05	0.25	0.98	0.01	0.30	1.01
δ_F	0.21	0.06	0.25	0.98	0.01	0.52	1.00
θ_H	0.78	0.02	0.13	0.87	0.00	0.06	1.01
θ_F	0.78	0.02	0.13	0.87	0.00	0.30	1.00
a_1	0.50	0.11	0.19	0.96	0.00	0.58	1.00
b_2	0.85	0.07	0.19	0.96	0.00	0.11	1.00
c_3	0.99	0.02	0.19	0.96	0.00	0.14	1.00
ρ_a	0.46	0.14	0.13	0.87	0.02	0.93	1.00
ρ_q	0.86	0.25	0.24	1.00	0.06	0.01	1.14
ρ_s	0.96	0.03	0.01	0.72	0.00	0.69	1.00
ρ_g	0.04	0.04	0.13	0.87	0.00	0.67	1.00
			Relative Policy Target Weights				
μ_q	0.10	0.05	0.09	1.24	0.00	0.30	1.01
μ_y	0.20	0.08	0.09	1.24	0.01	0.05	1.04
μ_r	0.67	0.24	0.09	1.24	0.05	0.33	1.02
			Standard Deviation of Shock Innovations				
σ_H	1.35	0.68	0.91	7.33	0.11	0.36	1.01
σ_F	1.79	1.07	0.91	7.34	0.22	0.05	1.06
σ_a	1.34	0.69	0.52	2.66	0.09	0.42	1.01
σ_q	23.76	0.92	0.32	0.87	0.17	0.14	1.03
σ_s	24.60	0.35	0.52	2.65	0.04	0.43	1.01
σ_{π^*}	2.37	0.85	0.52	2.65	0.18	0.09	1.04
σ_{y^*}	12.60	1.27	0.52	2.66	0.28	0.95	1.00
σ_{r^*}	0.25	0.03	0.52	2.66	0.00	0.62	1.00
σ_r	0.62	0.07	0.52	2.66	0.00	0.35	1.00
σ_g	24.52	0.44	0.52	2.66	0.05	0.21	1.02

Note: Calibrated parameters: $\beta = 0.99$; $\chi = 0.05$; $\alpha = 0.51$.

Table 6: Chile: Posterior Parameters and Convergence Diagnostics ($\mu_q > 0$)

Parameters	Post Mean	Post SD	2.5%	97.5%	NSE (8%)	G p-value	B-GF
			Structural Parameters				
h	0.79	0.05	0.19	0.93	0.01	0.03	1.07
σ	1.60	0.41	0.27	2.19	0.10	0.00	1.24
ϕ	1.39	0.23	1.05	2.03	0.01	0.39	1.00
η	3.31	0.70	0.27	2.19	0.10	0.14	1.02
δ_H	0.34	0.07	0.25	0.98	0.01	0.65	1.00
δ_F	0.10	0.04	0.25	0.98	0.00	0.81	1.00
θ_H	0.84	0.02	0.13	0.87	0.00	0.12	1.02
θ_F	0.78	0.04	0.13	0.87	0.00	0.16	1.02
a_1	0.51	0.11	0.19	0.96	0.00	0.25	1.00
b_2	0.83	0.07	0.19	0.96	0.00	0.85	1.00
c_3	0.98	0.01	0.19	0.96	0.00	0.85	1.00
ρ_a	0.76	0.08	0.13	0.87	0.02	0.00	1.10
ρ_q	0.98	0.05	0.24	1.00	0.00	0.38	1.01
ρ_s	0.89	0.07	0.01	0.72	0.01	0.04	1.04
ρ_g	0.06	0.04	0.13	0.87	0.00	0.28	1.00
			Relative Policy Target Weights				
μ_q	0.03	0.02	0.09	1.23	0.00	0.00	1.08
μ_y	0.09	0.04	0.09	1.24	0.01	0.00	1.06
μ_r	0.19	0.14	0.09	1.24	0.02	0.00	1.19
			Standard Deviation of Shock Innovations				
σ_H	1.97	1.13	0.91	7.33	0.24	0.09	1.05
σ_F	3.08	1.75	0.91	7.34	0.42	0.00	1.18
σ_a	1.71	0.89	0.52	2.66	0.21	0.00	1.24
σ_q	24.20	0.66	0.32	0.87	0.11	0.22	1.02
σ_s	13.11	3.32	0.52	2.66	0.87	0.00	1.63
σ_{π^*}	0.69	0.25	0.52	2.65	0.03	0.34	1.01
σ_{y^*}	6.91	1.12	0.52	2.66	0.21	0.21	1.03
σ_{r^*}	0.25	0.03	0.52	2.65	0.00	0.61	1.00
σ_r	0.62	0.07	0.52	2.66	0.00	0.32	1.00
σ_g	24.21	0.79	0.52	2.65	0.16	0.08	1.05

Note: Calibrated parameters: $\beta = 0.99$; $\chi = 0.05$; $\alpha = 0.29$.

Table 7: Colombia: Posterior Parameters and Convergence Diagnostics ($\mu_q > 0$)

	Post Mean	Post SD	2.5%	97.5%	NSE (8%)	G p-value	B-GF
				Structural Parameters			
h	0.75	0.04	0.19	0.93	0.01	0.00	1.06
σ	1.48	0.23	0.27	2.20	0.05	0.00	1.15
ϕ	1.36	0.22	1.05	2.03	0.01	0.81	1.00
η	2.89	0.47	0.27	2.20	0.03	0.57	1.00
δ_H	0.42	0.08	0.25	0.98	0.01	0.02	1.05
δ_F	0.15	0.08	0.25	0.98	0.01	0.57	1.00
θ_H	0.80	0.02	0.13	0.87	0.00	0.14	1.01
θ_F	0.70	0.03	0.13	0.87	0.01	0.00	1.16
a_1	0.50	0.11	0.19	0.96	0.00	0.28	1.00
b_2	0.83	0.07	0.19	0.96	0.00	0.37	1.00
c_3	0.995	0.01	0.19	0.96	0.00	0.00	1.01
ρ_a	0.64	0.13	0.13	0.87	0.03	0.05	1.06
ρ_q	0.91	0.14	0.23	1.00	0.03	0.19	1.03
ρ_s	0.65	0.11	0.01	0.73	0.02	0.20	1.02
ρ_g	0.04	0.03	0.13	0.87	0.00	0.00	1.01
				Relative Policy Target Weights			
μ_y	0.04	0.01	0.10	1.24	0.00	0.15	1.00
μ_r	0.05	0.03	0.10	1.24	0.00	0.70	1.00
				Standard Deviation of Shock Innovations			
σ_H	2.01	1.39	0.91	7.36	0.29	0.998	1.00
σ_F	2.81	2.83	0.91	7.33	0.74	0.00	1.39
σ_a	1.22	0.47	0.52	2.65	0.06	0.98	1.00
σ_q	23.99	0.80	0.32	0.88	0.14	0.90	1.00
σ_s	19.73	3.54	0.52	2.66	0.95	0.00	1.33
$\sigma_{\pi *}$	0.53	0.16	0.52	2.65	0.01	0.35	1.00
$\sigma_{y *}$	4.37	0.80	0.52	2.65	0.17	0.37	1.01
$\sigma_{r *}$	0.25	0.03	0.52	2.65	0.00	0.87	1.00
σ_r	0.62	0.07	0.52	2.66	0.00	0.42	1.00
σ_g	24.44	0.54	0.52	2.66	0.08	0.07	1.04

Note: Calibrated parameters: $\beta = 0.99$; $\chi = 0.05$; $\alpha = 0.29$.

Table 8: Colombia: Posterior Parameters and Convergence Diagnostics ($\mu_q = 0$)

Parameters	Post Mean	Post SD	2.5%	97.5%	NSE (8%)	G p-value	B-GF
			Structural Parameters				
h	0.92	0.03	0.19	0.93	0.00	0.06	1.04
σ	0.97	0.26	0.27	2.19	0.06	0.02	1.09
ϕ	1.45	0.24	1.05	2.03	0.01	0.05	1.00
η	2.19	0.57	0.27	2.19	0.06	0.30	1.01
δ_H	0.77	0.12	0.25	0.98	0.02	0.04	1.05
δ_F	0.65	0.22	0.25	0.98	0.05	0.03	1.08
θ_H	0.97	0.01	0.13	0.87	0.00	0.58	1.00
θ_F	0.96	0.02	0.13	0.87	0.00	0.84	1.00
a_1	0.46	0.10	0.19	0.96	0.00	0.08	1.00
b_2	0.89	0.06	0.19	0.96	0.00	0.33	1.00
c_3	0.98	0.01	0.19	0.96	0.00	0.09	1.00
ρ_a	0.51	0.12	0.13	0.87	0.02	0.05	1.05
ρ_q	0.99	0.01	0.23	1.00	0.00	0.79	1.00
ρ_s	0.24	0.19	0.01	0.72	0.04	0.74	1.00
ρ_g	0.04	0.04	0.13	0.87	0.00	0.49	1.00
			Relative Policy Target Weights				
μ_q	0.48	0.20	0.09	1.24	0.05	0.46	1.01
μ_y	0.30	0.14	0.09	1.24	0.03	0.07	1.05
μ_r	0.38	0.18	0.09	1.23	0.04	0.13	1.04
			Standard Deviation of Shock Innovations				
σ_H	2.34	1.99	0.91	7.33	0.52	0.00	1.42
σ_F	2.94	2.61	0.91	7.32	0.71	0.00	1.86
σ_a	22.65	1.78	0.52	2.65	0.45	0.00	1.36
σ_q	18.50	1.16	0.32	0.88	0.24	0.13	1.04
σ_s	23.04	1.41	0.52	2.66	0.35	0.00	1.27
σ_{π^*}	3.08	0.45	0.52	2.66	0.05	0.30	1.01
σ_{y^*}	0.30	0.06	0.52	2.65	0.00	0.22	1.00
σ_{r^*}	0.25	0.03	0.52	2.65	0.00	0.22	1.00
σ_r	0.56	0.06	0.52	2.65	0.00	0.47	1.00
σ_g	23.27	1.40	0.52	2.66	0.32	0.19	1.03

Note: Calibrated parameters: $\beta = 0.99$; $\chi = 0.05$; $\alpha = 0.44$.

Table 9: Mexico: Posterior Parameters and Convergence Diagnostics ($\mu_q > 0$)

Parameters	Post Mean	Post SD	2.5%	97.5%	NSE (8%)	G p-value	B-GF
Structural Parameters							
h	0.26	0.13	0.07	0.56	0.01	0.68	1.00
σ	0.07	0.04	0.01	0.19	0.01	0.00	1.05
ϕ	1.80	0.26	1.33	2.34	0.02	0.00	1.03
η	5.06	0.70	3.84	6.63	0.05	0.29	1.01
δ_H	0.72	0.20	0.26	0.98	0.03	0.71	1.00
δ_F	0.38	0.16	0.13	0.74	0.02	0.27	1.02
θ_H	0.26	0.14	0.05	0.55	0.00	0.26	1.02
θ_F	0.63	0.03	0.56	0.69	0.00	0.32	1.00
a_1	0.57	0.13	0.32	0.81	0.00	0.37	1.00
b_2	0.86	0.07	0.71	1.01	0.00	0.52	1.00
c_3	0.99	0.02	0.95	1.02	0.00	0.18	1.04
ρ_a	0.89	0.06	0.75	0.98	0.02	0.00	1.00
ρ_q	0.72	0.11	0.50	0.96	0.03	0.996	1.00
ρ_s	0.44	0.18	0.11	0.88	0.00	0.998	1.00
ρ_g	0.11	0.07	0.02	0.28	0.00	0.01	1.00
Relative Policy Target Weights							
μ_q	0.06	0.03	0.02	0.13	0.00	0.00	1.00
μ_y	0.01	0.01	0.00	0.03	0.00	0.12	1.01
μ_r	1.60	0.46	0.92	2.93	0.11	0.00	1.00
Standard Deviation of Shock Innovations							
σ_H	18.85	4.40	7.71	24.65	1.14	0.00	1.47
σ_F	2.22	1.46	0.41	5.48	0.26	0.48	1.01
σ_a	3.66	0.71	2.40	5.17	0.06	0.48	1.00
σ_q	8.13	1.76	5.16	11.53	0.38	0.00	1.15
σ_s	24.32	0.67	22.35	24.98	0.09	0.09	1.05
σ_{π^*}	18.44	3.62	11.51	24.67	0.94	0.01	1.31
σ_{y^*}	5.85	0.75	4.44	7.38	0.10	0.78	1.00
σ_{r^*}	0.28	0.04	0.21	0.35	0.00	0.02	1.00
σ_r	0.60	0.07	0.48	0.76	0.00	0.47	1.00
σ_g	1.14	0.36	0.66	2.06	0.05	0.00	1.05

Note: Calibrated parameters: $\beta = 0.99$; $\chi = 0.05$; $\alpha = 0.35$.

Table 10: Peru: Posterior Parameters and Convergence Diagnostics ($\mu_q > 0$)

	$\mu_q > 0$	$\mu_q = 0$
	Brazil (2004–2014)	
Marginal Likelihood	-1431.5	-1469.8
Acceptance Rate (%)	12.93	8.24
Indeterminacy Rate (%)	4.72	18.16
Invalid Likelihood Rate (%)	2.247600	1.929550
μ_q	0.31 (0.12)	0
μ_y	0.73 (0.22)	0.78 (0.34)
μ_r	0.53 (0.30)	0.39 (0.18)
	Chile (2002–2014)	
Marginal Likelihood	-1591.9	-1606.4
Acceptance Rate (%)	20.50	20.28
Indeterminacy Rate (%)	9.13	14.36
Invalid Likelihood Rate (%)	1.534950	0.679450
μ_q	0.31 (0.11)	0
μ_y	0.17 (0.09)	0.06 (0.04)
μ_r	1.14 (0.39)	1.08 (0.17)
	Colombia (2003–2014)	
Marginal Likelihood	-1504.6	-1506.2
Acceptance Rate (%)	22.66	18.61
Indeterminacy Rate (%)	2.41	6.31
Invalid Likelihood Rate (%)	1.775850	1.896250
μ_q	0.13 (0.06)	0
μ_y	0.51 (0.22)	0.44 (0.19)
μ_r	0.54 (0.34)	0.47 (0.27)
	Mexico (2002–2014)	
Marginal Likelihood	-1527.6	-1560.7
Acceptance Rate (%)	20.86	22.69
Indeterminacy Rate (%)	4.19	5.12
Invalid Likelihood Rate (%)	1.600200	2.042550
μ_q	0.06 (0.04)	0
μ_y	0.10 (0.05)	0.01 (0.01)
μ_r	0.43 (0.30)	1.29 (0.33)
	Peru (2005–2014)	
Marginal Likelihood	-1437.1	-1399.8
Acceptance Rate (%)	29.18	15.97
Indeterminacy Rate (%)	6.67	1.14
Invalid Likelihood Rate (%)	1.018800	2.664300
μ_q	0.11 (0.07)	0
μ_y	0.62 (0.23)	0.03 (0.02)
μ_r	0.85 (0.26)	1.78 (0.48)

Note: Estimates of the policy weights (relative to that of inflation, 1), μ_i with $i = \{q, y, r\}$, report the posterior mean, with the posterior standard deviation in parentheses.

Table 11: Model Comparison under Complete Asset Markets