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The Transmission of Exogenous Commodity and Oil Prices shocks to Latin America: A Panel VAR approach

Rocío Gondo* y Fernando Pérez*

* Banco Central de Reserva del Perú

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The Transmission of Exogenous Commodity and Oil Prices shocks to Latin America - A Panel VAR approach*

Rocío Gondo† Fernando J. Pérez Forero‡

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Abstract

During the last sixteen years, we have experienced an episode of commodity price boom and bust. Despite being exogenous to Latin America, commodity and oil price shocks are extremely relevant for explaining macroeconomic fluctuations. Thus, in this paper we assess the dynamic impact of these price fluctuations for relevant macroeconomic and financial variables for commodity exporting countries in the region (Chile, Colombia, Mexico and Peru) using a Bayesian Hierarchical Panel VAR with an exogenous block. This model is more flexible and less restrictive than a stylized DSGE model. We quantify the strong expansionary effect of these price shocks, and we discuss the connection with i) monetary and macro-prudential policy, ii) the financial sector and iii) the real economy. Furthermore, we observe some degree of heterogeneity across countries both in amplification and propagation patterns.

Resumen

Durante los últimos dieciséis años hemos experimentado un episodio de auge y caída de los precios de los *commodities*. A pesar de ser estas fluctuaciones exógenas para América Latina, los choques de precios de *commodities* y del petróleo son extremadamente relevantes para explicar las fluctuaciones macroeconómicas. Por lo tanto, en este documento evaluamos el impacto dinámico de fluctuaciones en estos precios para las variables macroeconómicas y financieras relevantes para los países exportadores de *commodities* en la región (Chile, Colombia, México y Perú) utilizando un modelo Panel VAR Jerárquico Bayesiano y con un bloque exógeno. Este modelo es más flexible y menos restrictivo que un modelo DSGE estilizado. Dado lo anterior, se cuantifica el fuerte efecto expansivo de estos choques de precios y discutimos la conexión con i) la política monetaria y macroprudencial, ii) el sector financiero y iii) la economía real. Además, observamos cierto grado de heterogeneidad entre países, tanto en los patrones de amplificación como de propagación.

C11, C23, E44, E52, Q02

Panel Vector Autoregressions, Exogenous Block, Bayesian Estimation, Commodity Prices

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†Research Department, Central Reserve Bank of Peru (BCRP), Jr. Santa Rosa 441, Lima 1, Perú; Email address: rocio.gondo@bcrp.gob.pe

‡Head of Monetary Programming Department, Central Reserve Bank of Peru (BCRP), Jr. Santa Rosa 441, Lima 1, Perú; Email address: fernando.perez@bcrp.gob.pe

1 Introduction

Commodity and oil price shocks are extremely relevant for explaining macroeconomic fluctuations in Latin America, especially for commodity exporters, though these variables are exogenous to these economies. The large commodity price boom and bust cycles observed in the recent decade (see Figure 1) have led policymakers to focus on the macroeconomic effects of these price swings in major commodity exporting economies. In a scenario of rising (falling) commodity prices and loosening (tightening) of financial and monetary conditions in advanced economies, commodity exporting countries are especially exposed to an exogenous improvement (a deterioration) in macroeconomic performance (see Figure 2).

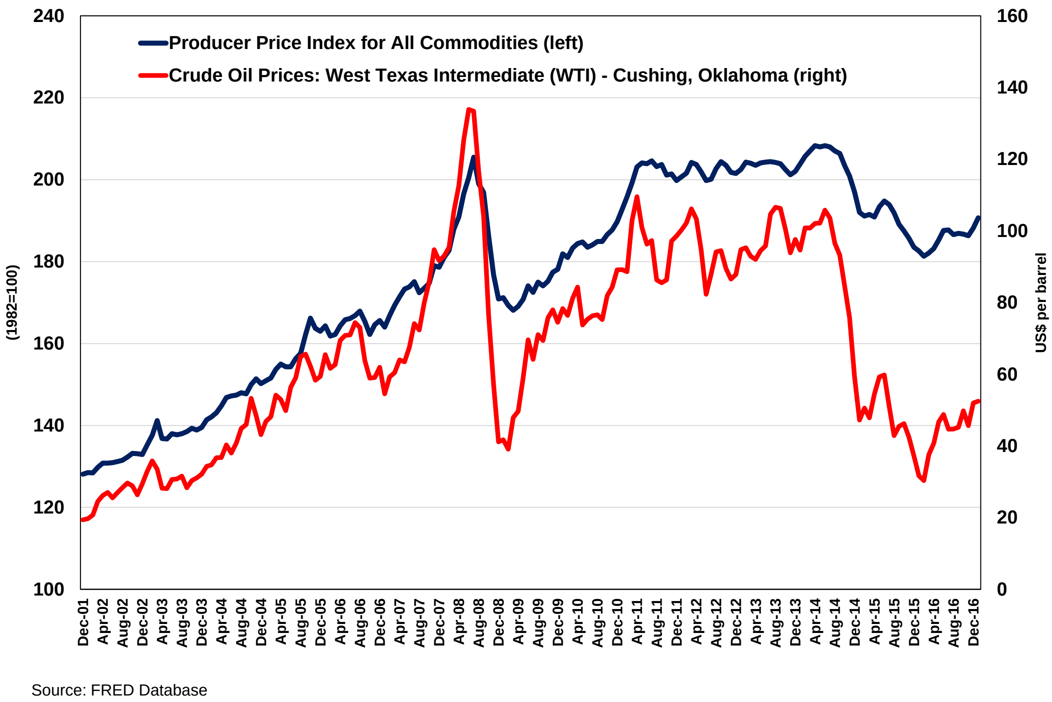

Figure 1: Evolution of Commodity and Oil Prices

Moreover, commodity booms attracted surges in capital inflows and higher access credit from international capital markets, partly due to improved growth prospects, creating an additional linkage with the financial cycle. In some cases, the effect of commodity price shocks also lead to policy tradeoffs. For instance, a reduction in commodity prices led to lower GDP growth but higher prices due to pass-through effects from the exchange rate depreciation. In 2015-2016 many countries in the region had to react by increasing the monetary policy rate in order to meet their inflation targets.

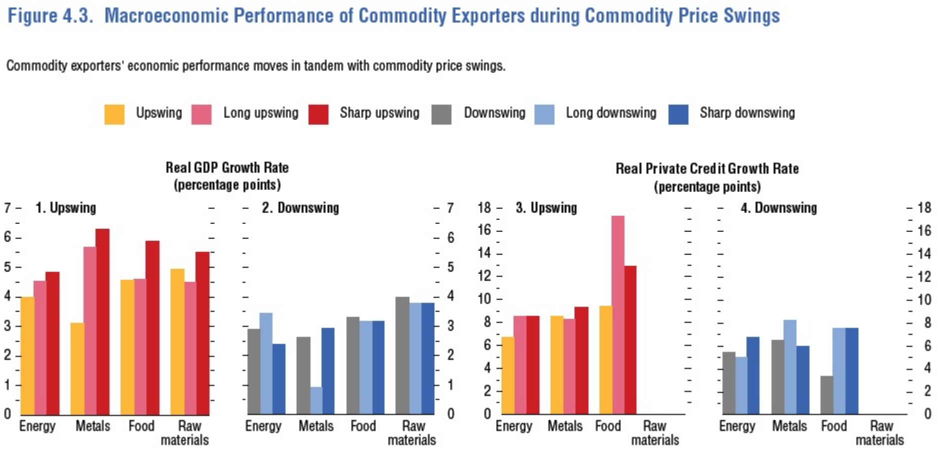

Figure 2: Source: IMF WEO Apr 2012

In this paper we assess the dynamic impact of these price fluctuations on a set of macroeconomic and financial variables for a sample of Latin American commodity exporting countries (Chile, Colombia, Mexico and Peru) and their effect on foreign and domestic financing conditions. We focus out attention in these economies since they share many structural characteristics, and for that reason we do not consider other commodity exporter countries that are more advanced in terms of economic development. For that purpose, we use a Bayesian Hierarchical Panel Vector Autorregressive model (Panel VAR) with an exogenous block as a common framework to make the results comparable between countries. The model is flexible and it is less restrictive than a stylized DSGE model such as García-Cicco *et al.* (2017) and complements this group of models by contrasting their implications with empirical data. We quantify the strong expansionary effect of these price shocks, and we discuss the connection with i) monetary and macro-prudential policy, ii) the financial sector and iii) the real economy. Furthermore, we observe a high degree of heterogeneity across countries both in amplification and propagation.

A key aspect of the sample of countries under study is their different degrees of exposure to commodity price fluctuations, both in terms of participation of commodities in total exports and in the composition and types of commodities for each country. All countries in the sample have an export structure that is highly concentrated on a particular product: oil in the case of Colombia and Mexico and metals in the case of Chile (copper) and Peru (copper and gold). Metal exporters have maintained a roughly constant share in total exports of goods since the 1990s, whereas in the case of oil exporters, Colombia's oil exports as a share of total exports nearly doubled whereas Mexico's share fell drastically to roughly 5 percent in the last years.

A second relevant feature of these countries is their exposure to external financial factors. The four economies depend heavily on capital flows and external financing conditions; at the same time, their domestic financial systems are less developed than those of advanced economies and highly concentrated in the banking sector. Therefore, the effect of external shocks on macroeconomic variables, in turn, is amplified by the higher access to international capital markets when growth prospects are revised upwards due to a commodity price upturn.

In this paper we estimate a Bayesian Hierarchical Panel VAR (see Ciccarelli and Rebucci (2006), Jarociński (2010), Canova and Pappa (2011) and Pérez Forero (2015)). We extend the standard approach in order to consider an exogenous block as in Canova (2005) that is common for all countries, and we identify structural shocks by imposing zero restrictions in the exogenous block, the one that will include the commodity prices[1]. It turns out that comparison across countries is fair, since we apply the same set of identification restrictions to the same set of variables in all these countries.

Our results show that a positive commodity price shock generates a large and persistent real exchange rate appreciation. On the external side this triggers higher exports, which improves the trade balance for all the analyzed countries. Moreover, in many countries exchange rate appreciation due to capital inflows is smoothed using FX intervention and precautionary accumulation of international reserves.

On the financial side, there is a fall in the country risk premium measured by the EMBI spread, which reduces the risk profile of the domestic economy, increases capital inflows and foreign financing for domestic banks and firms. Moreover, cheaper access to external funding by banks

[1] Another Panel VAR application for Emerging Economies and Commodity prices can be found in Kataryniuk and Martínez-Martín (2017). The abstract from the effects related to the financial sector of the economy.

increase their expected profitability, leading to an increase in bank valuation as measured by the bank price to book ratio. Cheaper access to external funding and higher availability of domestic bank deposits due to the positive temporary income shock contributes to higher credit growth. On the real side, as a consequence of this favorable scenario with easier access to credit and higher income in the economy, there is a large expansion in economic activity. However, even though higher GDP growth leads to higher inflationary pressures, the exchange rate appreciation works in the opposite direction. Therefore, we find some heterogeneity in the final effect, where countries with milder (larger) increases in GDP growth and larger (smaller) exchange rate appreciation show a reduction (increase) in inflation and monetary policy rates. Given that the financial sector works as an amplifier of the effect of commodity price shocks, there is room for macro-prudential policies to complement monetary policy and mitigate the effect of these shocks. Positive (negative) commodity price shocks increase (decrease) bank valuation due to higher (lower) expected profitability and reduce (increase) the perception of credit risk, which amplifies the reaction of credit growth. Macro-prudential policies can help smooth excessive credit boom bust cycles and reduce the impact of commodity price shocks on the domestic economy.

Secondly, small open economies are particularly vulnerable to the so called 'risk taking channel' of exchange rates. Moreover, in the case of countries with partially dollarized financial systems such as Peru, exchange rate fluctuations create an additional vulnerability and macro-prudential policies can mitigate the amplification and propagation of the identified external macroeconomic shocks in this context.

This work relates to the strand of the literature that focuses on the impact of external shocks on macroeconomic variables. Canova (2005) on impact of US shocks on Latin America, where he finds evidence of significant importance of the financial channel. In particular, several works have been made related to different aspects of the impact of commodity prices on business cycles in Latin America, such as Gruss (2014) who measures the effect on output growth using global VAR and Medina (2010) who considers its impact on fiscal positions. Céspedes and Velasco (2012) consider the effect of commodity price shocks on output and investment and find that these effects are larger when an economy faces more severe financial market imperfections.

However, none of these consider the interaction between commodity cycles and financial market imperfections in emerging economies.

Among studies that consider a theoretical model, Fornero *et al.* (2016) find that commodity price shocks generate spillovers on investment on related non-commodity sectors. García-Cicco *et al.* (2017) also analyze the effect of commodity price shocks in a model with imperfect domestic financial intermediation, which creates an amplification mechanism to the expansionary effects on output and investment, and find that macro-prudential policies can complement monetary policy to counteract the contraction in credit growth during commodity busts. We evaluate these effects from an empirical perspective, without restricting the model to the particular amplification mechanisms in the model.

Shousha (2016) present both a theoretical and empirical analysis of the effect of commodity prices on business cycles in both advanced and emerging economies, with an amplification mechanism created by working capital constraints which are more severe in emerging countries. Another work with both theoretical and empirical approaches is the one by Fernandez *et al.* (2015) who find the highly significant importance of commodity prices on business cycle fluctuations in emerging economies, and find that one of the key amplification mechanisms is through the negative correlation of commodity prices on the country's risk premium.

One possible limitation of our approach is the fact that we use an aggregated index of commodities. Therefore, our research agenda can also consider the approach of Fernández *et al.* (2017), where they include disaggregated terms of world prices such as metals and energy, and the also consider a small open economy approach by estimating VAR models for commodity exporter country using an exogenous block. In particular, they argue that there are significant gains from using disaggregated components for world prices.

The shock identification can also be improved. In particular, we have included a large information set for the external block as control variables. It is still an issue to discuss to what extent this external shock can be considered either as a demand or as a supply shocks[2]. For instance, Filardo and Lombardi (2014) find that the interpretation of external commodity price

[2] In this case, the alternative would be to use sign restrictions as in Arias *et al.* (2017) and Baumeister and Hamilton (2015).

shocks as supply ones could lead to suboptimal outcomes, since they could also be considered and global demand shocks. Therefore, this also has implications for the systematic reaction of monetary policy in emerging markets, and this idea is also reinforced through a global economy model by Filardo *et al.* (2018). Finally, the research agenda can also be extended to explore the explicit role of fiscal policy in this context of commodity prices shocks. For instance, Roch (2017) finds a significant reaction of fiscal variables for the case of Chile, Colombia and Peru by using individual VAR models, but he explores the impact of Terms of Trade indexes shocks for each country.

The document is organized as follows: section 2 describes the Panel VAR model used for the empirical analysis, section 3 presents the Gibbs Sampling algorithm for estimating the model, section 4 shows the data description, section 5 discusses the main results and section 6 concludes.

2 The model

We assume in this section that each economy can be modeled as an individual Vector Autoregressive (VAR) model with an exogenous block. Then we combine efficiently the information of these four economies in order to perform the estimation. A crucial point in this setup is the fact that the exogenous block is common to all the four economies, so that the dynamic effects derived from these external shocks will be easily comparable.

In this context, consider the set of countries $n = 1, \ldots, N$, where each country n is represented by a VAR model with exogenous variables:

$$y_{n,t} = \sum_{l=1}^{p} B'_{n,l} y_{n,t-l} + \sum_{l=0}^{p} B^{*\prime}_{n,l} y^*_{t-l} + \Delta_n z_t + u_{n,t} \tag{1}$$

where $y_{n,t}$ is a $M_1 \times 1$ vector of endogenous domestic variables, y^*_t is a $M_2 \times 1$ vector of endogenous domestic variables, z_t is a $W \times 1$ vector of exogenous variables common to all countries, $u_{n,t}$ is a $M_1 \times 1$ vector of reduced form shocks such that $u_{n,t} \sim N(\mathbf{0}, \Sigma_n)$, $E\left(u_{n,t} u'_{m,t}\right) = \mathbf{0}, n \neq m$ $\in \{1, \ldots, N\}$, p is the lag length and T_n is the sample size for each country $n \in \{1, \ldots, N\}$.

At the same time, there exists an exogenous block that evolves independently, such that

$$y_t^* = \sum_{l=1}^{p} \Phi_l^{*\prime} y_{t-l}^* + \Delta^* z_t + u_t^* \tag{2}$$

with $u_t^* \sim N(\mathbf{0}, \Sigma^*)$ and $E\left(u_t^* u_{n,t}'\right) = \mathbf{0}$.

The latter model can be expressed in a more compact form for each country $n \in \{1, \ldots, N\}$, so that:

$$\begin{aligned} \begin{bmatrix} \mathbf{I}_{M_1} & -B_{n,0}^{*\prime} \\ \mathbf{0} & \mathbf{I}_{M2} \end{bmatrix} \begin{bmatrix} y_{n,t} \\ y_t^* \end{bmatrix} = & \sum_{l=1}^{p} \begin{bmatrix} B_{n,l}' & B_{n,l}^{*\prime} \\ \mathbf{0} & \Phi_l^{*\prime} \end{bmatrix} \begin{bmatrix} y_{n,t} \\ y_t^* \end{bmatrix} + \\ & \begin{bmatrix} \Delta_n \\ \Delta^* \end{bmatrix} z_t + \begin{bmatrix} \Sigma_n & \mathbf{0} \\ \mathbf{0} & \Sigma^* \end{bmatrix} \begin{bmatrix} u_{n,t} \\ u_t^* \end{bmatrix} \end{aligned} \tag{3}$$

System (1) represents the small open economy in which its dynamics are influenced by the big economy block (2) through the parameters $B_{n,l}^{*\prime}$ and $\Phi_l^{*\prime}$. On the other hand, the big economy evolves independently, i.e. the small open economy cannot influence the dynamics of the big economy. Even though block (2) has effects over block (1), we assume that the block (2) is independent of block (1), and thus it will keep the same coefficients for each country model. This type of *Block Exogeneity* has been applied in the context of SVARs by Cushman and Zha (1997), Zha (1999) and Canova (2005), among others. Moreover, it turns out that this is a plausible strategy for representing small open economies such as the Latin American ones, since they are influenced by external shocks i.e. commodity prices fluctuations.

Reduced form estimation is performed by blocks as in Zha (1999), since they are independent. Assuming that we have a sample $t = 1, \ldots, T_n$ for each country $n \in \{1, \ldots, N\}$, the regression model for the domestic block can be re-expressed as

$$Y_n = X_n B_n + U_n \tag{4}$$

Where we have the data matrices $Y_n\,(T_n \times M_1)$, $X_n\,(T_n \times K)$, $U_n\,(T_n \times M_1)$, with $K = M_1 p + W$

and the corresponding parameter matrix $B_n\,(K \times M_1)$. In particular

$$B_n = \begin{bmatrix} B'_{n,1} & B'_{n,2} & \cdots & B'_{n,p} & B^{*\prime}_{n,1} & B^{*\prime}_{n,2} & \cdots & B^{*\prime}_{n,p} & \Delta'_n \end{bmatrix}'$$

The model in equation (4) can be re-written such that

$$\mathbf{y}_n = (I_{M_1} \otimes X_n)\,\beta_n + \mathbf{u}_n$$

where $\mathbf{y}_n = vec\,(Y_n)$, $\beta_n = vec\,(B_n)$ and $\mathbf{u}_n = vec\,(U_n)$ with

$$\mathbf{u}_n \sim N\,(0, \Sigma_n \otimes I_{T_n - p})$$

Under the normality assumption of the error terms, we have the likelihood function for each country

$$p\,(\mathbf{y}_n \mid \beta_n, \Sigma_n) = N\,((I_{M_1} \otimes X_n)\,\beta_n, \Sigma_n \otimes I_{T_n - p})$$

which is

$$\begin{aligned} p\,(\mathbf{y}_n \mid \beta_n, \Sigma_n) = (2\pi)^{-M_1(T_n-p)/2}\,|\Sigma_n \otimes I_{T_n-p}|^{-1/2} \times \\ \exp\left(-\frac{1}{2}\,(\mathbf{y}_n - (I_{M_1} \otimes X_n)\,\beta_n)'\,(\Sigma_n \otimes I_{T_n-p})^{-1}\,(\mathbf{y}_n - (I_{M_1} \otimes X_n)\,\beta_n)\right) \end{aligned} \tag{5}$$

where $n = 1, \ldots, N$.

On the other hand, the exogenous block estimation is as follows. First, rewrite equation (2) as a regression model

$$Y^* = X^* \Phi^* + U^*$$

Where we have the data matrices $Y^*\,(T^* \times M_2)$, $X^*\,(T^* \times K^*)$, $U^*\,(T^* \times M_2)$, with $K^* = M_2 p + W$ and the corresponding parameter matrix $\Phi^*\,(K^* \times M_2)$, and in particular:

$$\Phi^* = \begin{bmatrix} \Phi^{*\prime}_1 & \Phi^{*\prime}_2 & \cdots & \Phi^{*\prime}_p & \Delta^{*\prime} \end{bmatrix}'$$

The regression model can then be re-written such that

$$\mathbf{y}^* = (I_{M_2} \otimes X^*)\,\beta^* + \mathbf{u}^*$$

where $\mathbf{y}^* = vec\,(Y^*)$, $\beta^* = vec\,(\Phi^*)$ and $\mathbf{u}^* = vec\,(U^*)$ with

$$\mathbf{u}^* \sim N\,(0, \Sigma^* \otimes I_{T^*-p})$$

Under the normality assumption of the error terms, we have the likelihood function for the exogenous block:

$$p\,(\mathbf{y}^* \mid \beta^*, \Sigma^*) = N\,((I_{M_2} \otimes X^*)\,\beta^*, \Sigma^* \otimes I_{T^*-p})$$

which is

$$\begin{gathered} p\,(\mathbf{y}^* \mid \beta^*, \Sigma^*) = (2\pi)^{-M_2(T^*-p)/2}\,|\Sigma^* \otimes I_{T^*-p}|^{-1/2} \times \\ \exp\begin{pmatrix} -\frac{1}{2}\,(\mathbf{y}^* - (I_{M_2} \otimes X^*)\,\beta^*)'\,(\Sigma^* \otimes I_{T^*-p})^{-1} \times \\ (\mathbf{y}_n - (I_{M_2} \otimes X^*)\,\beta^*) \end{pmatrix} \end{gathered} \tag{6}$$

As a consequence of the previous analysis, the statistical model described above has a joint likelihood function. Denote $\Theta = \left\{\{\beta_n, \Sigma_n\}_{n=1}^N, \beta^*, \Sigma^*\right\}$ as the set of parameters, then the likelihood function is

$$p\,(\mathbf{y}, \mathbf{y}^* \mid \Theta) \propto |\Sigma^*|^{-T^*/2} \prod_{n=1}^N |\Sigma_n|^{-T_n/2} \times \exp\begin{pmatrix} -\frac{1}{2}\sum_{n=1}^N (\mathbf{y}_n - (I_{M_1} \otimes X_n)\,\beta_n)'\,(\Sigma_n \otimes I_{T_n-p})^{-1} \times \\ (\mathbf{y}_n - (I_{M_1} \otimes X_n)\,\beta_n) \\ -\frac{1}{2}\,(\mathbf{y}^* - (I_{M_2} \otimes X^*)\,\beta^*)'\,(\Sigma^* \otimes I_{T^*-p})^{-1} \times \\ (\mathbf{y}_n - (I_{M_2} \otimes X^*)\,\beta^*) \end{pmatrix} \tag{7}$$

2.1 Priors

Given the normality assumption of the error terms, it follows that each country coefficients vector is normally distributed. As a result, we assume a normal prior for them in order get a posterior distribution that is also normal, i.e. a conjugated prior:

$$p\left(\beta_n \mid \overline{\beta}, O_n, \tau\right) = N\left(\overline{\beta}, \tau O_n\right) \tag{8}$$

with $\overline{\beta}$ as the common mean and τ as the overall tightness parameter. The covariance matrix O_n takes the form of the typical Minnesota prior (Litterman, 1986), i.e. $O_n = diag\left(o_{ij,l}\right)$ such that

$$o_{ij,l} = \begin{cases} \frac{1}{l^{\phi_3}} & , i = j \\ \frac{\phi_1}{l^{\phi_3}}\left(\frac{\widehat{\sigma}_j^2}{\widehat{\sigma}_i^2}\right) & , i \neq j \\ \phi_2 & , exogenous \end{cases} \tag{9}$$

where

$$i, j \in \{1, \ldots, M_1\} \text{ and } l = 1, \ldots, p$$

and where $\widehat{\sigma}_j^2$ is the variance of the residuals from an estimated $AR(p)$ model for each variable $j \in \{1, \ldots, M_1\}$. In addition, we assume the non-informative priors:

$$p\left(\Sigma_n\right) \propto \left|\Sigma_n\right|^{-\frac{1}{2}(M_1+1)} \tag{10}$$

In a standard Bayesian context, $\overline{\beta}$ and τ would be hyper-parameters that are supposed to be calibrated. In turn, in a Hierarchical context (see e.g. Gelman *et al.* (2003)), it is possible to derive a posterior distribution for both and therefore estimate them. That is, we do not want to impose any particular tightness for the prior distribution of coefficients, we want to get it from the data. Following Gelman (2006) and Jarociński (2010)[3] we assume an inverse-gamma prior distribution for τ, so that

$$p\left(\tau\right) = IG\left(\frac{\upsilon}{2}, \frac{s}{2}\right) \propto \tau^{-\frac{\upsilon+2}{2}} \exp\left(-\frac{1}{2}\frac{s}{\tau}\right) \tag{11}$$

[3]See Pérez Forero (2015) for a similar application for Latin America.

Finally, we assume the non-informative prior:

$$p\left(\overline{\beta}\right) \propto 1 \tag{12}$$

In addition, coefficients of the exogenous block have a traditional Litterman prior with

$$p\left(\beta^{*}\right) = N\left(\overline{\beta^{*}}, \tau_X O_X\right) \tag{13}$$

where $\overline{\beta^{*}}$ assumes an AR(1) process for each variable and $O_X = diag\left(o^{*}_{ij,l}\right)$ such that

$$o^{*}_{ij,l} = \begin{cases} \frac{1}{l^{\phi_3^{*}}} & , i = j \\ \frac{\phi_1^{*}}{l^{\phi_3^{*}}}\left(\frac{\widehat{\sigma}_j^2}{\widehat{\sigma}_i^2}\right) & , i \neq j \\ \phi_2^{*} & , exogenous \end{cases} \tag{14}$$

where

$$i, j \in \{1, \ldots, M_2\} \text{ and } l = 1, \ldots, p$$

and similarly $\widehat{\sigma}_j^2$ is the variance of the residuals from an estimated $AR(p)$ model for each variable $j \in \{1, \ldots, M_2\}$. As in the domestic block, we assume the non-informative priors for the covariance matrix of error terms, so that:

$$p\left(\Sigma^{*}\right) \propto \left|\Sigma^{*}\right|^{-\frac{1}{2}(M_2+1)} \tag{15}$$

Moreover, since this is a hierarchical model, we also estimate the overall tightness parameter for the prior variance as in the domestic block, so that we again assume the inverse-gamma distribution:

$$p\left(\tau_X\right) = IG\left(\frac{\upsilon_X}{2}, \frac{s_X}{2}\right) \propto \tau_X^{-\frac{\upsilon_X+2}{2}} \exp\left(-\frac{1}{2}\frac{s_X}{\tau_X}\right) \tag{16}$$

As a result of the hierarchical structure, our statistical model presented has several parameter blocks. Denote the parameter set as Θ, such that:

$$\Theta = \left\{ \{\beta_n, \Sigma_n\}_{n=1}^{N}, \beta^*, \Sigma^*, \tau, \overline{\beta}, \tau_X \right\}$$

so that the joint prior is given by (8), (10), (11), (12), (13), (15) and (16):

$$\begin{aligned} p(\Theta) \propto \prod_{n=1}^{N} p(\Sigma_n)\, p\left(\beta_n \mid \overline{\beta}, O_n, \tau\right) p(\tau) \\ = \prod_{n=1}^{N} |\Sigma_n|^{-\frac{1}{2}(M_1+1)} \times \\ \tau^{-\frac{NM_1K}{2}} \exp\left(-\frac{1}{2} \sum_{n=1}^{N} \left(\beta_n - \overline{\beta}\right)' \left(\tau^{-1} O_n\right)^{-1} \left(\beta_n - \overline{\beta}\right) \right) \times \\ \tau^{-\frac{v+2}{2}} \exp\left(-\frac{1}{2} \frac{s}{\tau} \right) \times \\ |\Sigma^*|^{-\frac{1}{2}(M_2+1)} \times \\ \tau_X^{-\frac{M_2K^*}{2}} \exp\left(-\frac{1}{2} \left(\beta^* - \overline{\beta^*}\right)' \left(\tau_X^{-1} O_X\right)^{-1} \left(\beta^* - \overline{\beta^*}\right) \right) \times \\ \tau_X^{-\frac{v_X+2}{2}} \exp\left(-\frac{1}{2} \frac{s_X}{\tau_X} \right) \end{aligned} \tag{17}$$

3 Bayesian Estimation

Given the specified priors (17) and the joint likelihood function (2), we combine efficiently these two pieces of information in order to get the estimated parameters included in Θ. Using the Bayes' theorem we have that:

$$p(\Theta \mid Y) \propto p(Y \mid \Theta)\, p(\Theta) \tag{18}$$

Given (2) and (17), the posterior distribution (18) takes the form:

$$
\begin{aligned}
&p\left(\Theta \mid \mathbf{y}, \mathbf{y}^{*}\right) \propto \left|\Sigma^{*}\right|^{-\frac{T^{*}+M_{2}+1}{2}} \\
&\prod_{n=1}^{N}\left|\Sigma_{n}\right|^{-\frac{T_{n}+M_{1}+1}{2}} \times \exp\begin{pmatrix}
-\frac{1}{2} \sum_{n=1}^{N}\left(\mathbf{y}_{n}-\left(I_{M_{1}} \otimes X_{n}\right) \beta_{n}\right)'\left(\Sigma_{n} \otimes I_{T_{n}-p}\right)^{-1} \\
\left(\mathbf{y}_{n}-\left(I_{M_{1}} \otimes X_{n}\right) \beta_{n}\right) \\
-\frac{1}{2}\left(\mathbf{y}^{*}-\left(I_{M_{2}} \otimes X^{*}\right) \beta^{*}\right)'\left(\Sigma^{*} \otimes I_{T^{*}-p}\right)^{-1} \\
\left(\mathbf{y}_{n}-\left(I_{M_{2}} \otimes X^{*}\right) \beta^{*}\right)
\end{pmatrix} \times \\
&\tau^{-\frac{\left(N M_{1} K+v\right)}{2}} \exp \left(-\frac{1}{2}\left[\sum_{n=1}^{N}\left(\beta_{n}-\overline{\beta}\right)' O_{n}^{-1}\left(\beta_{n}-\overline{\beta}\right)+s\right] \frac{1}{\tau}\right) \times \\
&\tau_{X}^{-\frac{\left(M_{2} K^{*}+v_{X}\right)}{2}} \exp \left(-\frac{1}{2}\left[\left(\beta^{*}-\overline{\beta^{*}}\right)' O_{X}^{-1}\left(\beta^{*}-\overline{\beta^{*}}\right)+s_{X}\right] \frac{1}{\tau_{X}}\right)
\end{aligned} \tag{19}
$$

Our target is now to maximize the right-hand side of equation (19) in order to get Θ. The common practice in Bayesian Econometrics (see e.g. Koop (2003) and Canova (2007) among others) is to simulate the posterior distribution (19) in order to conduct statistical inference. This is since any object of interest that is also a function of Θ can be easily computed given the simulated posterior. In this section we describe a Markov Chain Monte Carlo (MCMC) routine that helps us to accomplish this task.

3.1 A Gibbs sampling routine

In general, in every Macro-econometric model it is difficult to sample from the posterior distribution $p(\Theta \mid Y)$. The latter is a consequence of the complex functional form that the likelihood function (or posterior distribution) might take, given the specified model. Typically, the Metropolis-Hasting algorithm is the canonical routine to do that. However, in this case we will show that there exists an analytical expression for the posterior distribution, therefore it is possible to implement a Gibbs Sampling routine, which is much simpler than the mentioned Metropolis-Hastings. In this process, it is useful to divide the parameter set into different blocks and factorize (19) appropriately.

Recall that $\Theta = \left\{\{\beta_n, \Sigma_n\}_{n=1}^{N}, \beta^*, \Sigma^*, \tau, \overline{\beta}, \tau_X\right\}$. Then, use the notation Θ/χ whenever we denote the parameter vector Θ without the parameter . Details about the form of each block can be found in Appendix B.

The routine starts here. Set $k = 1$ and denote K as the total number of draws. Then follow the steps below:

1. Draw $p\left(\beta^* \mid \Theta/\beta^*, \mathbf{y}^*, \mathbf{y}_n\right)$. If the candidate draw is stable keep it, otherwise discard it.

2. For $n = 1, \ldots, N$ draw $p\left(\beta_n \mid \Theta/\beta_n, \mathbf{y}^*, \mathbf{y}_n\right)$. If the candidate draw is stable keep it, otherwise discard it.

3. Draw $p\left(\Sigma^* \mid \Theta/\Sigma^*, \mathbf{y}^*, \mathbf{y}_n\right)$.

4. For $n = 1, \ldots, N$ draw $p\left(\Sigma_n \mid \Theta/\Sigma_n, \mathbf{y}^*, \mathbf{y}_n\right)$.

5. Draw $p\left(\tau_X \mid \Theta/\tau_X, Y\right)$.

6. Draw $p\left(\overline{\beta} \mid \Theta/\overline{\beta}, Y\right)$. If the candidate draw is stable keep it, otherwise discard it.

7. Draw $p\left(\tau \mid \Theta/\tau, Y\right)$.

8. If $k < K$ set $k = k + 1$ and return to Step 1. Otherwise stop.

A complete cycle of all these steps gives us a draw for the paramater set Θ.

3.2 Estimation setup

We run the Gibbs sampler for $K = 150,000$ and discard the first $100,000$ draws in order to minimize the effect of initial the values. Moreover, in order to reduce the serial correlation across draws, we set a thinning factor of 50, i.e. given the remaining $50,000$ draws, we take 1 every 50 and discard the remaining ones. As a result, we have $1,000$ draws for conducting inference. Specific details about how we conduct inference and assess convergence can be found in Appendix B respectively.

Following the recommendation of Gelman (2006) and Jarociński (2010), we assume a uniform prior for the standard deviation, which translates into a prior for the variance as

$$p\left(\tau\right) \propto \tau^{-1/2} \tag{20}$$

by setting $v = -1$ and $s = 0$ in (11).

Regarding the Minnesota-stye prior, we do not have any information about the value of the hyper-parameters. Thus, we set a conservative $\phi_1 = 0.5$, $\phi_2 = 1$ and $\phi_3 = 2$ in equation (9). More specifically, ϕ_1 is related with *a priori* difference between own lags and lags of other variables; ϕ_2 is related with the *a priori* heteroskedasticity coming from exogenous variables[4]; and $\phi_3 = 2$ means that the shrinking pattern of coefficients is quadratic. It is worth to mention that, in order to have symmetry, we set the same hyper-parameter values for the exogenous block, i.e. $\phi_1^* = 0.5$, $\phi_2^* = 1$ and $\phi_3^* = 2$ in equation (14). Finally, the exogenous block has a standard Minnesota Prior, and we set an autorregressive parameter of 0.9 for the prior mean of the first lag of the own variable in each VAR equation.

4 Data Description and Identification Strategy

4.1 Data Description

The model includes an exogenous block and a domestic block for each of the four Latin American countries. The data frequency is monthly and the final sample of the empirical exercise covers the period from December 2001 until January 2017 for each country and the exogenous block as well. Since commodity prices fluctuate in a daily basis, we consider the highest possible frequency for macroeconomic analysis, i.e. monthly data. In this regard, the use of either quarterly or annual data complicates the analysis, since we have mentioned that commodity and oil prices are high frequency variables. We also included a constant and a time trend for both the domestic and foreign block, and the the foreign block is common to all countries. It includes the following variables: (i) US industrial production index, (ii) US consumer price index, (iii) Fed funds rate, (iv) 10-year Treasury bond yield, (v) VIX, and (vi) all producers commodity price index and (vii) Oil Prices (WTI). All the variables for the exogenous block were obtained from the FRED database of the St. Louis' Fed. We did not include any variable related with China, since we have used only monthly data, which in case is unavailable. Moreover, our

[4]Since this is a VARX, i.e. a model that includes the lags of exogenous variables, we cannot set a very large value of this hyper parameter as in standard Minnesota prior applications.

assumption regarding the exogeneity of commodity prices rests on the fact that commodity exporters cannot control the prices by affecting the total supply, so that they take prices as given.

For the domestic block, each country's model includes the following set of variables: (i) industrial production index (IP), (ii) consumer price index (CPI), (iii) trade balance, (iv) bank credit, (v) monetary policy rate, (vi) net international reserves, (vii) real effective exchange rate, (viii) EMBI spread and (ix) bank price to book ratio. IP, CPI, bank credit, net international reserves and real exchange rates are obtained from the IMF IFS database. Bank credit is defined as credit from the survey to other depository corporations. The trade balance is obtained from each country's central bank web site. Policy rates are obtained from the BIS statistics, EMBI from Bloomberg and bank price to book ratios come from the equity index for banks from Datastream.

Economic activity variables were seasonally adjusted using TRAMO-SEATS and all variables were introduced in the model in logs, with the exception of the interest rates, the EMBI spread, the trade balance and the Bank price to book ratio. The mentioned variables were untransformed and rescaled (see the data plots in Appendix A).

4.2 Identification of structural shocks

We use sign restrictions in order to compute the impulse responses using the output of the Gibbs Sampling estimation of the Panel VAR model, taking into account that the exogenous block serves also as an extension of the information set for the econometrician, as it mitigates the risk associated with the presence of the omitted variable bias for the domestic block.

4.3 Identification assumptions

The identification of commodity shocks is fairly standard. We have two types of restrictions, as shown in Table 1. The first group is related with zero restrictions in the contemporaneous coefficients matrix, as in the old literature of Structural VARs, i.e. Sims (1980) and Sims (1986). The second group are the sign restrictions as in Canova and De Nicoló (2002) and Uhlig (2005), where we set a horizon of six months.

In this case we assume that the commodity shock produces i) a rise in commodity prices (Pcom) and ii) a rise in Oil Prices (WTI). We do not restrict the remaining variables, neither in the exogenous block, nor in the domestic block. Moreover, we do not consider the two shocks separately (commodities and oil prices), as they might be highly correlated.

Var / Shock	Name	Commodity shock
Domestic Block	$\mathbf{y}$	?
Consumer Price Index	CPI_{US}	?
Industrial Production	IP_{US}	?
Federal Funds Rate	FFR	?
10-year yield (TB)	$10y$	?
VIX	VIX	?
Oil prices	WTI	$\geqslant 0$
Commodity prices	$Pcom$	$\geqslant 0$

Table 1: Identifying Restrictions

The identification restrictions shown in Table 1 are only associated with a particular shock. As a result, the remaining shocks are unidentified. However, it turns out that this is not a econometric problem, since the literature of SVARs with sign restrictions explains that in order to conduct proper inference the model needs to be only partially identified (Rubio-Ramírez *et al.*, 2010).

4.3.1 The algorithm

In this stage we use as an input the estimation output from subsection 3.1, i.e. the posterior distribution of the reduced-form of the model. Then we take draws from this distribution as it is described in the following estimation algorithm[5]:

1. Set first $K = 2,000$ number of draws.

[5]See Uhlig (2005), among others.

2. Draw (β^*_k, Σ^*_k) from the posterior distribution (foreign block) and get $(\mathbf{A}^*_0)_k = (P^*)^{-1}$ from the *Cholesky* decomposition of $\Sigma^*_k = P^* (P^*)'$.

3. Draw $\mathbf{X}^* \sim N(0, I_{n^*})$ and get $\mathbf{Q}^*$ such that $\mathbf{Q}^*\mathbf{R}^* = \mathbf{X}^*$, i.e. an orthogonal matrix $\mathbf{Q}^*$ that satisfies the QR decomposition of $\mathbf{X}^*$. The random matrix $\mathbf{Q}^*$ has the uniform distribution with respect to the Haar measure on $O(n)$(Arias *et al.*, 2017).

4. Construct the matrix:

$$\overline{\mathbf{Q}^*} = \begin{bmatrix} \mathbf{I}_{k^*} & \mathbf{0}_{(k^* \times M_2 - k^*)} \\ \mathbf{0}_{(M_2 - k^* \times k^*)} & \mathbf{Q}^* \end{bmatrix}$$

That is, a subset of $k^* < n^*$ variables in $(\mathbf{y}^*)$ are going to be *slow* and therefore they do not rotate. This how we impose zero restrictions in this case.

5. Draw $(\beta_{n,k}, \Sigma_{n,k})$ from the posterior distribution (domestic block) and get $(A_{n,0})_k = (P_n)^{-1}$ from the *Cholesky* decomposition of $\Sigma_{n,k} = P_n (P_n)'$.

6. Draw $\mathbf{X} \sim N(0, I_{M_1})$ and get $\mathbf{Q}$ such that $\mathbf{QR} = \mathbf{X}$, i.e. an orthogonal matrix $\mathbf{Q}$ that satisfies the QR decomposition of $\mathbf{X}$. The random matrix $\mathbf{Q}$ has the uniform distribution with respect to the Haar measure on $O(n)$(Arias *et al.*, 2017).

7. Construct the matrix:

$$\overline{\mathbf{Q}} = \begin{bmatrix} \mathbf{I}_{k} & \mathbf{0}_{(k \times M_1 - k)} \\ \mathbf{0}_{(M_1 - k \times k)} & \mathbf{Q} \end{bmatrix}$$

That is, a subset of $k < n$ variables in $(\mathbf{y})$ are going to be *slow* and therefore they do not rotate. This how we impose zero restrictions in this case.

8. Compute the matrices $\overline{\mathbf{A}}_{n,0} = (\mathbf{A}_{n,0})_k \overline{\mathbf{Q}}$ and $\overline{\mathbf{A}^*}_0 = (\mathbf{A}^*_0)_k \overline{\mathbf{Q}^*}$, then recover the system (3) and compute the impulse responses.

9. If sign restrictions are satisfied, keep the draw and set $k = k + 1$. If not, discard the draw and go to Step 10.

10. If $k < K$, return to Step 2, otherwise stop.

5 Results

Figure 3 shows the results for an increase in commodity prices[6]. All countries show a real exchange rate appreciation, ranging on impact from Colombia to Peru. This effect is also quite persistent. The trade balance shows an improvement in all countries due to higher commodity exports. Also, in many countries, the exchange rate appreciation and surge in capital inflows lead to central bank intervention in the foreign exchange market and the consequent accumulation of international reserves. Capital inflows also lead to a reduction in the domestic risk premium measured with the EMBI spread, given higher demand for bonds from international investors. The EMBI spread reacts quite similar in all the countries in the sample, although this effect reverts later.

On the financial side, there is an expansion in credit growth in all countries with some lag, although this effect is more persistent in Colombia and Peru [7]. It is worth to remark that most countries show an increase in credit growth, despite very different reactions associated with differences in the implementation of macro-prudential policies affecting bank and financial system variables.

Also, cheaper access to financing from abroad, both due to higher expected income and lower risk perception, increase bank valuation, as shown in the reaction of bank price to book ratios. This in turn provides higher access to funding and increases credit growth.

On the real side, as a consequence of this favorable scenario associated to access to credit and higher income in the economy, there is a large expansion in economic activity. In this set of variables, we observe larger heterogeneity across countries. The increase in GDP growth is larger for Colombia and Peru relative to the milder increase in Chile and Mexico. In the case of Mexico, this result is highly expected, as oil exports only represent around 5 percent of total GDP. In the case of inflation, even though higher GDP growth leads to higher inflationary pressures, the exchange rate appreciation works in the opposite direction. Therefore, we find

[6] An increase in commodity prices includes a weighted average of different types of commodities, including oil prices. Error bands for impulse responses are shown in Appendix C.

[7] Other work as García-Cicco *et al.* (2017) shows that incorporating a financial sector in an estimated DSGE model and the amplification of real and financial shocks is more relevant for countries like Colombia and Mexico and less so for Chile and Peru.

some heterogeneity in the final effect, where countries with larger (milder) increases in GDP growth and smaller (larger) exchange rate appreciation show an increase (reduction) in inflation and monetary policy rates react to fulfill the inflation target.

Although this is somewhat expected in countries that apply an Inflation Targeting regime, there is still room for macro prudential policies that can help to mitigate the amplification effect of these shocks through credit growth and the financial sector. This is particularly relevant for economies that are vulnerable to exchange rate fluctuations because of a partially dollarized financial system. Therefore, our research agenda points towards the assessment of the effects of macro-prudential policies within this context, and their role to mitigate the amplification and propagation of the identified external macroeconomic shocks.

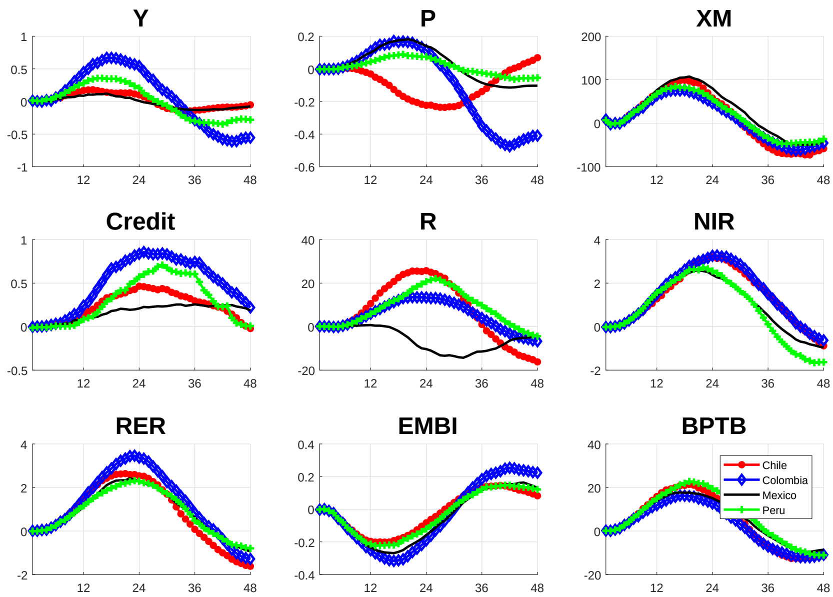

Figure 3: Commodity shocks comparison

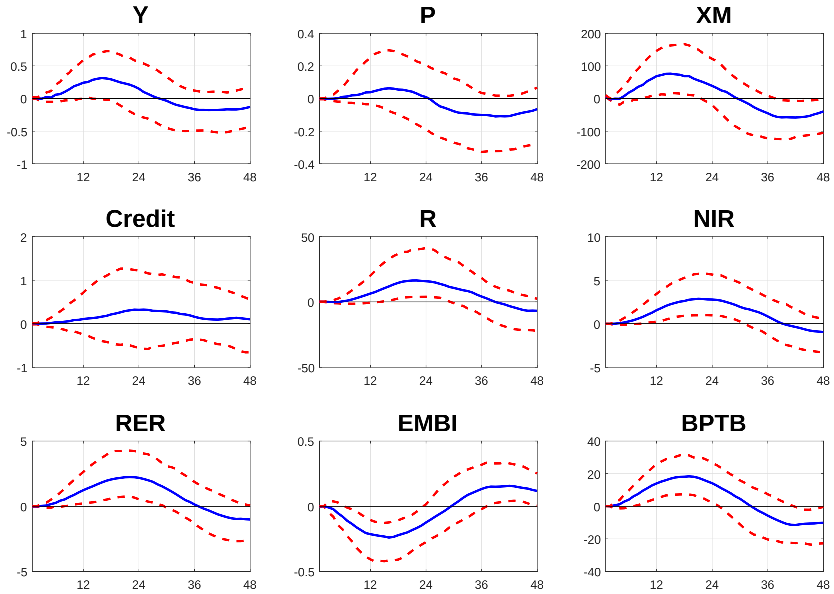

Figure 4: Commodity shocks: Average Effect in LATAM

6 Concluding Remarks

We have estimated a Bayesian Hierarchical Panel VAR (see Ciccarelli and Rebucci (2006), Jarociński (2010), Canova and Pappa (2011) and Pérez Forero (2015)), where we have extended the standard approach by including an exogenous block that is common for all countries, and we have identified structural shocks by imposing zero and sign restrictions. We extend the standard approach in order to consider an exogenous block as in Canova (2005) that is common for all countries. In this way, we provide a comparable framework across countries, since we apply the same set of identification restrictions to the same set of variables in all these countries. We use this framework to analyze the effect of external shocks such as commodity prices, which are highly relevant for commodity exporters such as the group of Latin American countries considered in this paper.

Our results show that a positive commodity price shock generates a large and persistent real exchange rate appreciation. On the external side this triggers higher exports, which improves the trade balance. Moreover, from the financial account, there is a surge in capital inflows due to improved profitability of domestic physical and financial assets. Monetary authorities smooth excessive fluctuations in the exchange rate by using FX intervention and precautionary accumulation of international reserves. In addition, external financial conditions improve, as shown by a fall in EMBI spreads, with increased foreign financing for domestic banks and firms. Lower external funding costs for banks increase bank valuation and contribute to higher credit growth.

On the real side, as a consequence of this favorable scenario with easier access to credit and higher in come in the economy, there is a large expansion in economic activity. However, even though higher GDP growth leads to higher inflationary pressures, the exchange rate appreciation works in the opposite direction. Therefore, we find some heterogeneity in the final effect, where countries with milder (larger) increases in GDP growth and larger (smaller) exchange rate appreciation show a reduction (increase) in inflation and monetary policy rates.

In terms of the model, one may argue that more work is needed in terms of the identification strategy, since we only imposed standard sign restrictions in order to pin down a commodity shock. However, it is important to mention that the included information set is useful enough to capture the presented transmission mechanism. Therefore, the use of more complicated assumptions such as hard-restrictions or a stylized structural modeling assumptions that can produce an over-identified model might be considered as robustness checks in the future. All in all, although simple, our empirical strategy is flexible and does not impose too many restrictions as Dynamic Stochastic General Equilibrium (DSGE) models that can be found in the literature related to this theme. More importantly, the results seem to be very much in line with the ones obtained in this group of models.

Also, future research could focus on quantifying the effect of macro-prudential policies as a financial policy to smooth excessive macroeconomic fluctuations. Particularly, it would be interesting to identify those macro-prudential policies that target excessive domestic credit growth and those that target cross-border flows separately.

A Data Description

In this section we present the plots from the data described in section 4, the one that coves the period 2001:12-2017:01. The variables in figures are already transformed, i.e. we show how they enter to the empirical model.

A.1 Domestic variables

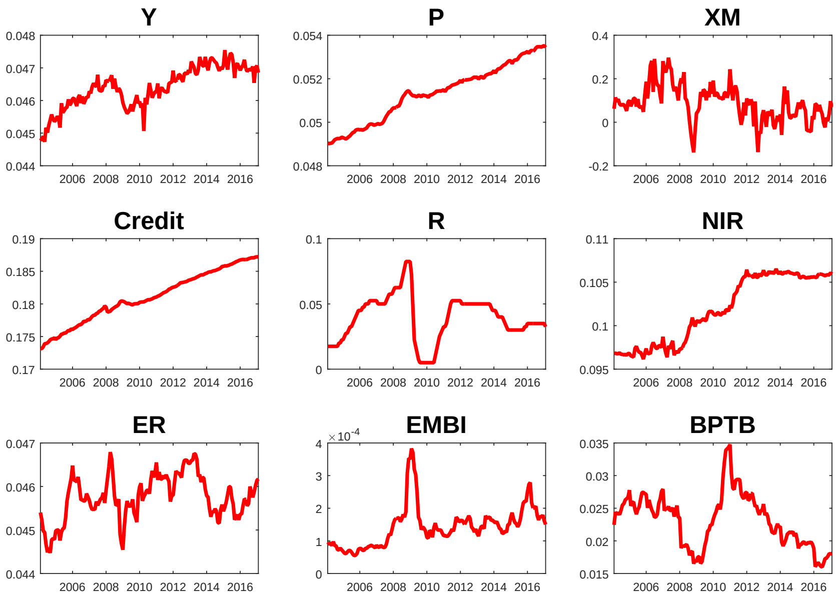

Figure 5: Chilean Data

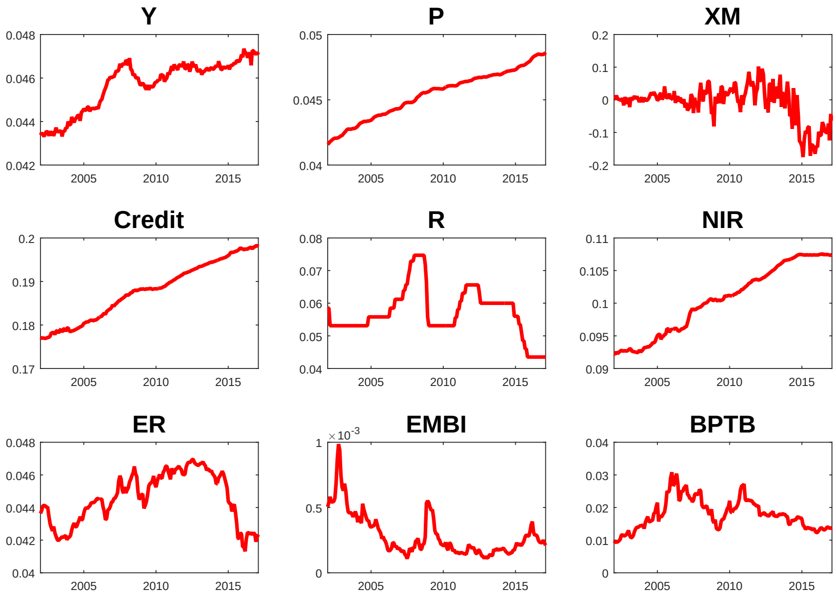

Figure 6: Colombian Data

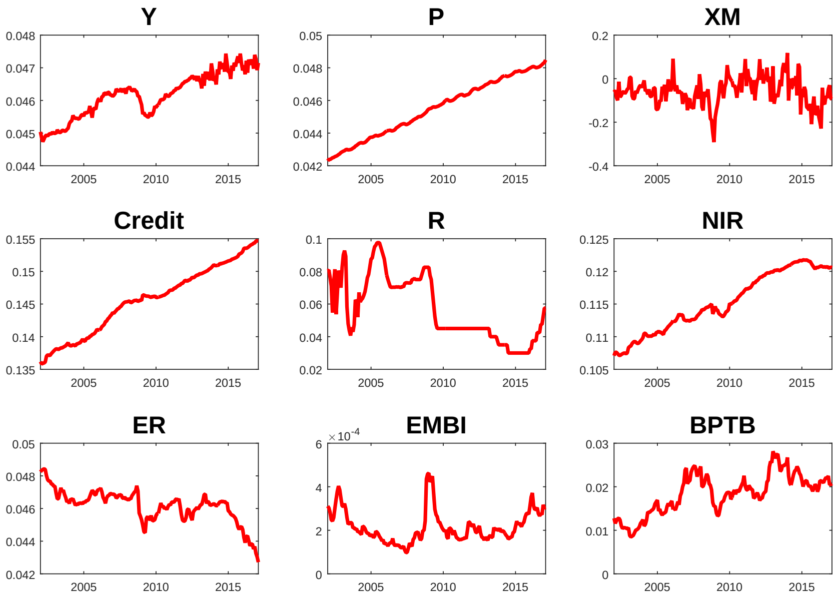

Figure 7: Mexican Data

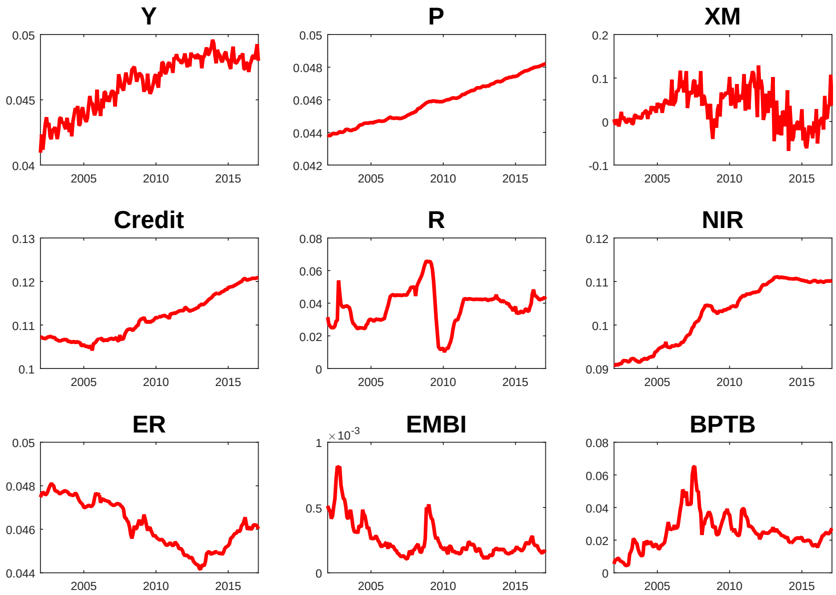

Figure 8: Peruvian Data

A.2 Exogenous variables

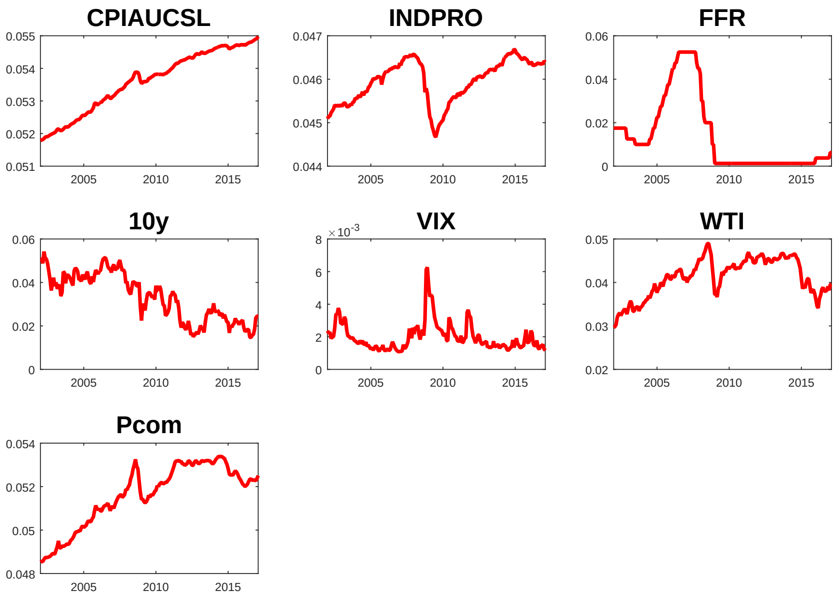

Figure 9: Exogenous Data

B Gibbs sampling details

The algorithm described in subsection 3.1 uses a set of conditional distributions for each parameter block. Here we provide specific details about the form that these distributions take and how they are constructed.

1. Block 1: $p(\beta^* \mid \Theta/\beta^*, \mathbf{y}^*)$: Given the likelihood (2) and the prior

$$p\left(\beta^* \mid \overline{\beta^*}, \tau\right) = N\left(\overline{\beta^*}, \tau_X O_X\right)$$

then the posterior is Normal

$$p\left(\beta^* \mid \Theta/\beta^*, \mathbf{y}^*\right) = N\left(\widetilde{\beta}^*, \widetilde{\Delta}^*\right)$$

with

$$\widetilde{\Delta}^* = \left(\Sigma^{*-1} \otimes X^{*\prime} X^* + \tau_X^{-1} O_X^{-1}\right)^{-1}$$

$$\widetilde{\beta}^* = \widetilde{\Delta}^* \left(\left(\Sigma^{*-1} \otimes X^{*\prime}\right)(\mathbf{y}^*) + \tau_X^{-1} O_X^{-1} \overline{\beta^*}\right)$$

2. Block 2: $p(\beta_n \mid \Theta/\beta_n, \mathbf{y}_n)$: Given the likelihood (2) and the prior

$$p\left(\beta_n \mid \overline{\beta}, \tau\right) = N\left(\overline{\beta}, \tau O_n\right)$$

then the posterior is Normal

$$p\left(\beta_n \mid \Theta/\beta_n, \mathbf{y}_n\right) = N\left(\widetilde{\beta}_n, \widetilde{\Delta}_n\right)$$

with

$$\widetilde{\Delta}_n = \left(\Sigma_n^{-1} \otimes X_n' X_n + \tau^{-1} O_n^{-1}\right)^{-1}$$

$$\widetilde{\beta}_n = \widetilde{\Delta}_n \left(\left(\Sigma_n^{-1} \otimes X_n'\right)(\mathbf{y}_n) + \tau^{-1} O_n^{-1} \overline{\beta}\right)$$

Block 3: $p\left(\Sigma^* \mid \Theta/\Sigma^*, \mathbf{y}^*\right)$: Given the likelihood (2) and the prior

$$p\left(\Sigma^*\right) \propto \left|\Sigma^*\right|^{-\frac{1}{2}(M_2+1)}$$

Denote the residuals

$$U^* = Y^* - X^* B^*$$

as in equation (4). Then the posterior variance term is Inverted-Wishart centered at the sum of squared residuals:

$$p\left(\Sigma^* \mid \Theta/\Sigma^*, \mathbf{y}^*\right) = IW\left(U^{*\prime}U^*, T^*\right)$$

Block 4: $p\left(\Sigma_n \mid \Theta/\Sigma_n, \mathbf{y}_n\right)$: Given the likelihood (2) and the prior

$$p\left(\Sigma_n\right) \propto \left|\Sigma_n\right|^{-\frac{1}{2}(M_1+1)}$$

Denote the residuals

$$U_n = Y_n - X_n B_n$$

as in equation (4). Then the posterior variance term is Inverted-Wishart centered at the sum of squared residuals:

$$p\left(\Sigma_n \mid \Theta/\Sigma_n, \mathbf{y}_n\right) = IW\left(U_n' U_n, T_n\right)$$

Block 5: $p\left(\tau_X \mid \Theta/\tau_X, Y\right)$: Given the priors

$$p\left(\tau_X\right) = IG\left(s, \upsilon\right) \propto \tau_X^{-\frac{\upsilon_X+2}{2}} \exp\left(-\frac{1}{2}\frac{s_X}{\tau_X}\right)$$

$$p\left(\beta_n \mid \overline{\beta}, O_n, \tau\right) = N\left(\overline{\beta}, \tau O_n\right)$$

then the posterior is

$$p(\tau_X \mid \Theta/\tau_X, Y) = IG\left(\frac{M_2K + \upsilon_X}{2}, \frac{\sum\limits_{n=1}^{N}\left(\beta_n - \overline{\beta}\right)' O_n^{-1}\left(\beta_n - \overline{\beta}\right) + s_X}{2}\right)$$

Block 6: $p\left(\overline{\beta} \mid \Theta/\overline{\beta}, Y\right)$: Given the prior

$$p\left(\beta_n \mid \overline{\beta}, O_n, \tau\right) = N\left(\overline{\beta}, \tau O_n\right)$$

by symmetry

$$p\left(\overline{\beta} \mid \beta_n, O_n, \tau\right) = N\left(\overline{\beta}, \tau O_n\right)$$

Then taking a weighted average across $n = 1, \ldots, N$:

$$p\left(\overline{\beta} \mid \{\beta_n\}_{n=1}^{N}, \tau\right) = N\left(\overline{\overline{\beta}}, \overline{\Delta}\right)$$

with

$$\overline{\Delta} = \left(\sum_{n=1}^{N} \tau^{-1} O_n^{-1}\right)^{-1}$$

$$\overline{\overline{\beta}} = \overline{\Delta}\left[\sum_{n=1}^{N} \tau^{-1} O_n^{-1} \beta_n\right]$$

Block 7: $p(\tau \mid \Theta/\tau, Y)$: Given the priors

$$p(\tau) = IG(s, \upsilon) \propto \tau^{-\frac{\upsilon+2}{2}} \exp\left(-\frac{1}{2}\frac{s}{\tau}\right)$$

$$p\left(\beta_n \mid \overline{\beta}, O_n, \tau\right) = N\left(\overline{\beta}, \tau O_n\right)$$

then the posterior is

$$p(\tau \mid \Theta/\tau, Y) = IG\left(\frac{NM_1K + \upsilon}{2}, \frac{\sum\limits_{n=1}^{N}\left(\beta_n - \overline{\beta}\right)' O_n^{-1}\left(\beta_n - \overline{\beta}\right) + s}{2}\right)$$

A complete cycle around these seven blocks produces a draw of Θ from $p(\Theta \mid Y)$.

C Impulse responses details

For each draw of Θ from the posterior distribution, we compute the companion form of the compact model as in equation (4). Then we compute the median value and the 68% credible interval for each impulse response. Results are shown below.

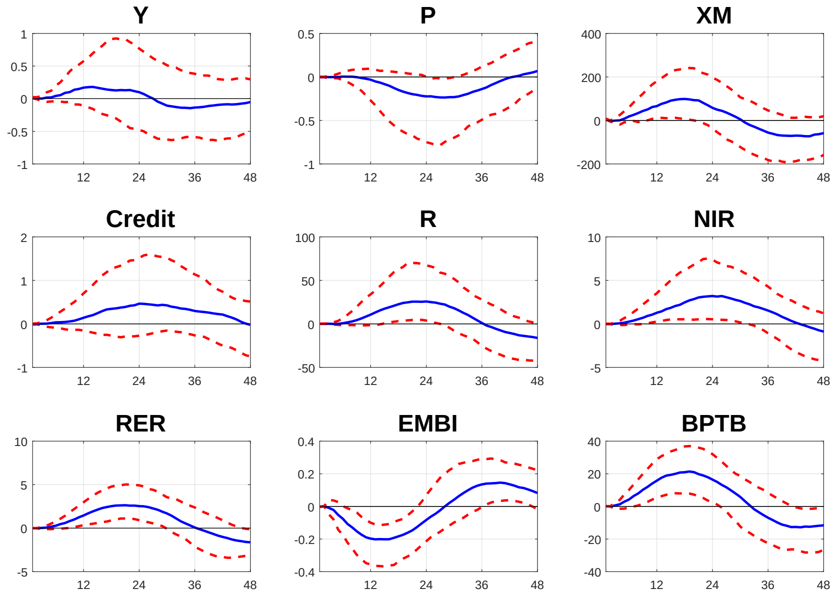

Figure 10: Commodity shocks effects in Chile, median value and 68% confidence interval

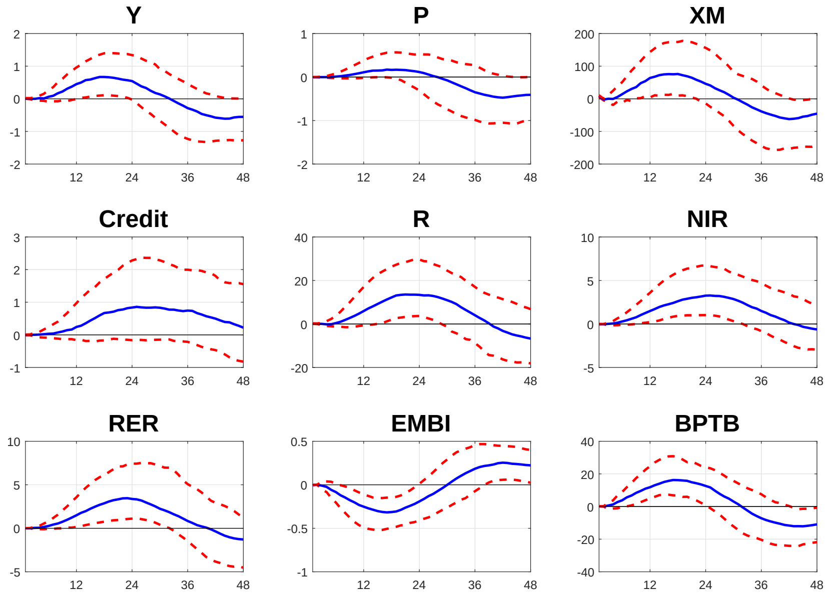

Figure 11: Commodity shocks effects in Colombia, median value and 68% confidence interval

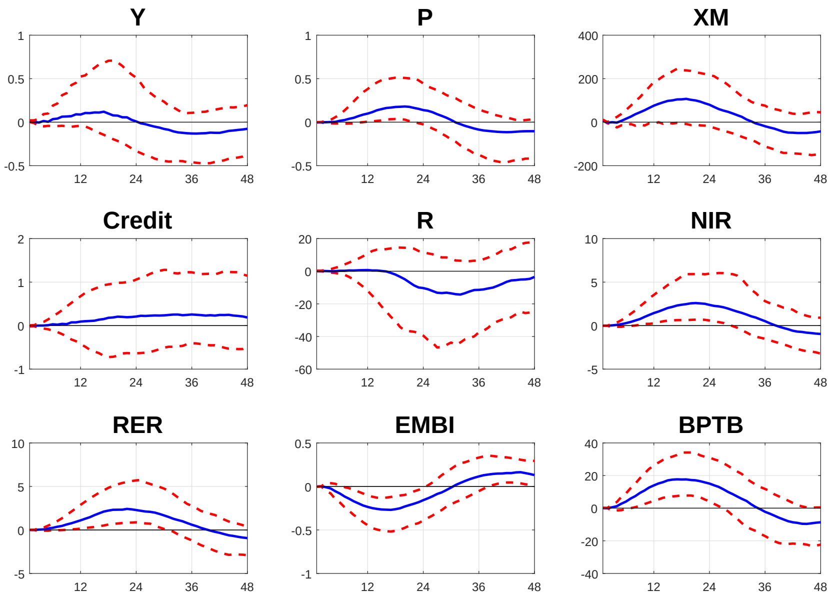

Figure 12: Commodity shocks effects in Mexico, median value and 68% confidence interval

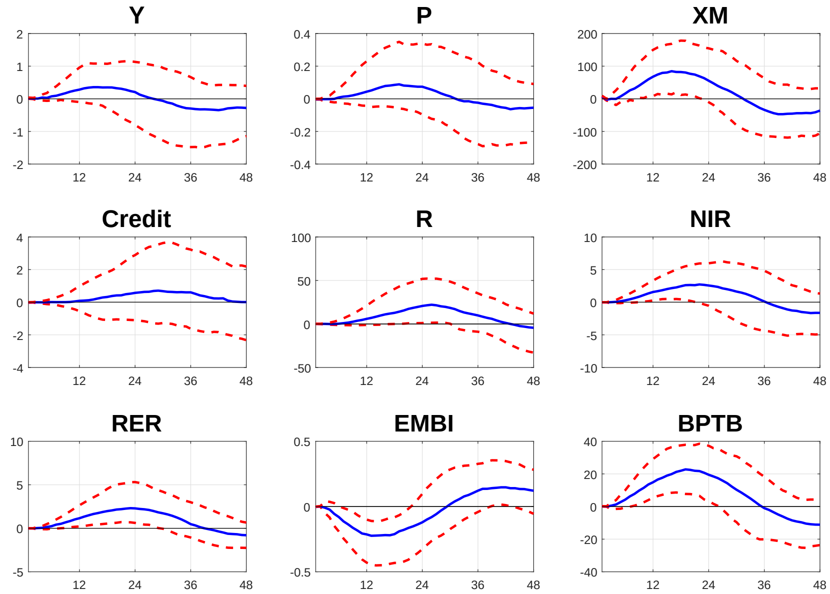

Figure 13: Commodity shocks effects in Peru, median value and 68% confidence interval

D Posterior distribution of hyperparameters

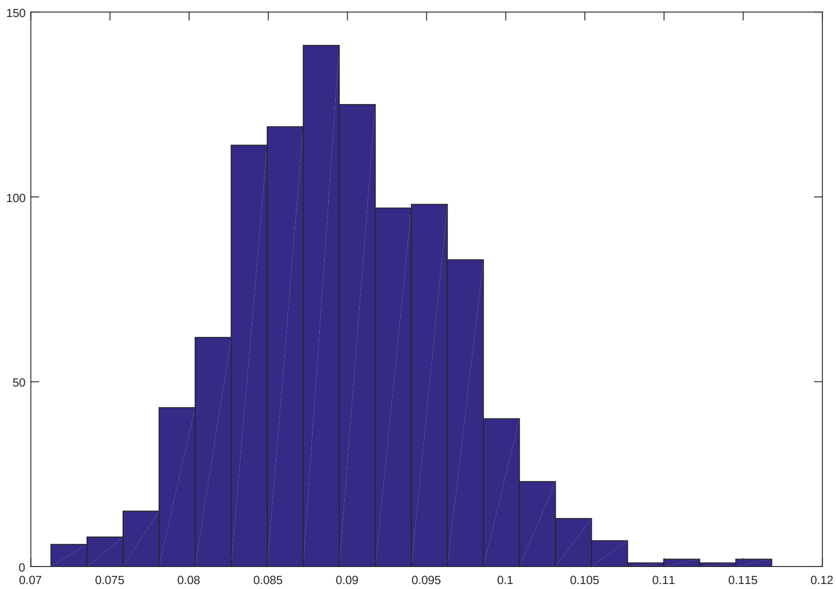

Figure 14: Posterior distribution of $\sqrt{\tau_X}$

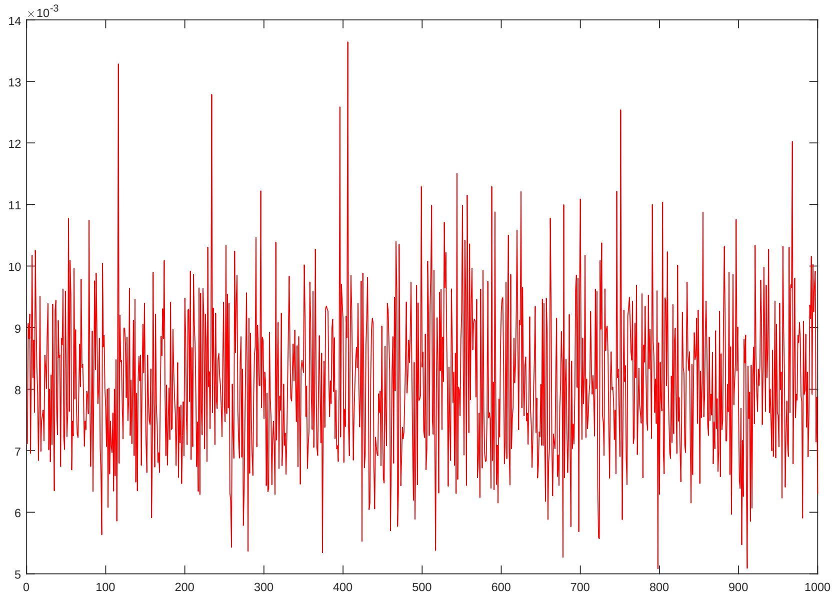

Figure 15: Posterior draws of τ_X

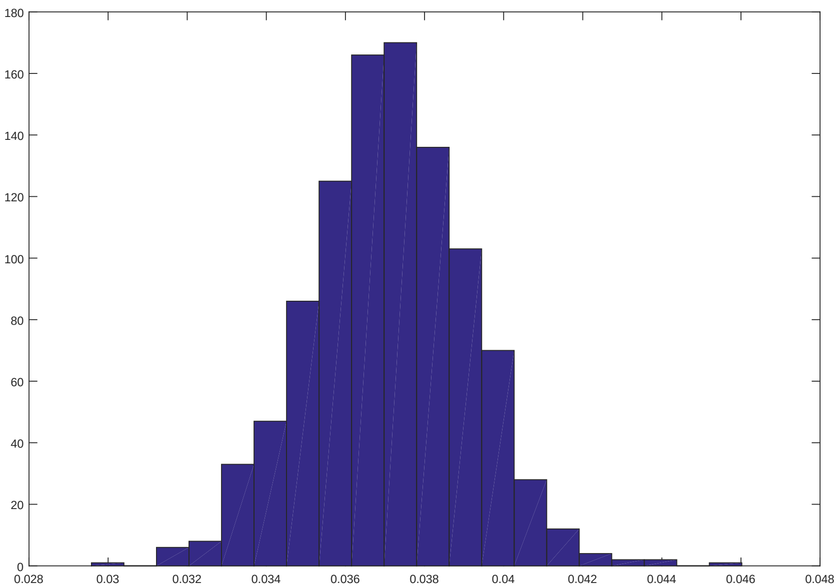

Figure 16: Posterior distribution of $\sqrt{\tau}$

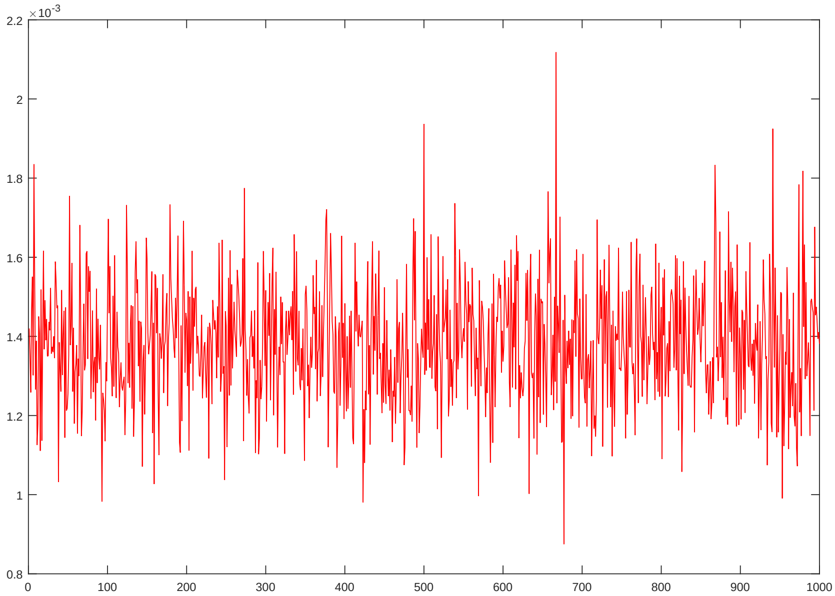

Figure 17: Posterior draws of τ

References

Arias, J., Rubio-Ramírez, J. F. and Waggoner, D. F. (2017). Inference based on svars identified with sign and zero restrictions: Theory and applications, econometrica (forthcoming).

Baumeister, C. and Hamilton, J. (2015). Sign restrictions, structural vector autoregressions and useful prior information. *Econometrica*, **83** (5), 1963–1999.

Canova, F. (2005). The transmission of us shocks to latin america. *Journal of Applied Econometrics*, **20** (2), 229–251.

— (2007). *Methods for Applied Macroeconomic Research*. Princeton University Press.

— and De Nicoló, G. (2002). Monetary disturbances matter for business fluctuations in the g-7. *Journal of Monetary Economics*, **49**, 11311159.

— and Pappa, E. (2011). Price differential in monetary unions: The role of fiscal shocks. *The Economic Journal*, **117**, 713–737.

Céspedes, L. F. and Velasco, A. (2012). Macroeconomic performance during commodity price booms and busts, nBER Working Paper No. 18569.

Ciccarelli, M. and Rebucci, A. (2006). Has the transmission mechanism of european monetary policy changed in the run-up to emu? *European Economic Review*, **50**, 737–776.

Cushman, D. and Zha, T. (1997). Identifying monetary policy in a small open economy under flexible exchange rates. *Journal of Monetary Economics*, **39** (3), 433–448.

Fernandez, A., Gonzalez, A. and Rodriguez, D. (2015). Sharing a ride on the commodities roller coaster: Common factors in business cycles of emerging economies, iMF Working Paper WP/15/280.

Fernández, A., Schmitt-Grohé, S. and Uribe, M. (2017). World shocks,world prices, and business cycles: An empirical investigation. *Journal of International Economics*, **108**, S2S14.

Filardo, A. and Lombardi, M. (2014). Has asian emerging market monetary policy been too procyclical when responding to swings in commodity prices?, published in Bank for International Settlements, 2014. "Globalisation, inflation and monetary policy in Asia and the Pacific," BIS Papers, Bank for International Settlements, number 77.

—, —, Montoro, C. and Ferrari, M. (2018). Monetary policy spillovers, global commodity prices and cooperation, bIS Working Papers No 696.

Fornero, J., Kirchner, M. and Yany, A. (2016). Terms of trade shocks and investment in commodity-exporting economies, documento de Trabajo No 773, Banco Central de Chile.

García-Cicco, J., Kirchner, M., Carrillo, J., Rodríguez, D., Pérez, F., Gondo, R., Montoro, C. and Chang, R. (2017). Financial and real shocks and the effectiveness of monetary and macroprudential policies in latin american countries, bank for International Settlements - Working Paper N668.

Gelman, A. (2006). Prior distributions for variance parameters in hierarchical models. *Bayesian Analysis*, **1** (3), 515–533.

—, Carlin, J., Stern, H., Dunson, A., D.B. Vehtari and Rubin, D. (2003). *Bayesian Data Analysis*. Chapman Hall/CRC Texts in Statistical Science, 3rd edition.

Gruss, B. (2014). After the boomcommodity prices and economic growth in latin america and the caribbean, iMF Working Paper WP/14/154.

Jarociński, M. (2010). Responses to monetary policy shocks in the east and the west of europe: A comparison. *Journal of Applied Econometrics*, **25**, 833–868.

Kataryniuk, I. and Martínez-Martín, J. (2017). Tfp growth and commodity prices in emerging economies, banco de España - Documentos de Trabajo No 1711.

Koop, G. (2003). *Bayesian Econometrics*. John Wiley and Sons Ltd.

Litterman, R. B. (1986). Forecasting with bayesian vector autoregressions-five years of experience. *Journal of Business Economic Statistics*, **4** (1), 25–38.

Medina, L. (2010). A commodity curse? the dynamic effects of commodity prices on fiscal performance in latin america, mPRA Paper No. 21690.

Pérez Forero, F. J. (2015). *Comparing the Transmission of Monetary Policy Shocks in Latin America: A Hierachical Panel VAR*. Premio de Banca Central Rodrigo Gómez / Central Banking Award "Rodrigo Gmez - Centro de Estudios Monetarios Latinoamericanos, CEMLA.

Roch, F. (2017). The adjustment to commodity price shocks in chile, colombia, and peru, iMF Working Papers WP/17/208.

Rubio-Ramírez, J., Waggoner, D. and Zha, T. (2010). Structural vector autoregressions: Theory of identification and algorithms for inference. *Review of Economic Studies*, **77**, 665696.

Shousha, S. (2016). Macroeconomic effects of commodity booms and busts: The role of financial frictions, mimeo Columbia University.

Sims, C. (1986). Are forecasting models usable for policy analysis? *Federal Reserve Bank of Minneapolis, Quarterly Review*, **1986**, 2–16.

Sims, C. A. (1980). Macroeconomics and reality. *Econometrica*, **48** (1), 1–48.

Uhlig, H. (2005). What are the effects of monetary policy on output? results from an agnostic identification procedure. *Journal of Monetary Economics*, **52**, 381–419.

Zha, T. (1999). Block recursion and structural vector autoregressions. *Journal of Econometrics*, **90** (2), 291–316.