The Evolution of External Shocks and Macrofiscal Outcomes in the Andes

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Abstract

The abundance of natural resources can adversely affect the macroeconomic stability of countries. Developing economic institutions that support the proper management of extractive industry resources is one of the fundamental factors for maintaining macroeconomic stability and generating long-term economic growth. This paper evaluates the impact of terms-of-trade shocks on fiscal performance and economic growth in the countries of the Andean region. To adequately capture the variation in both the magnitude of shocks and the sensitivity of economic and fiscal variables to those shocks, this paper makes use of a TV-SVAR model. This method allows us to evaluate whether the presence of fiscal institutions change their sensitivity to external disturbances. The results presented in this paper show a time-varying dynamic of both (1) terms-of-trade shocks and (2) the sensitivity of the Andean region's economies to such shocks, although with marked heterogeneity across countries. This heterogeneity coincides with the uneven development of the economic institutions in the countries of the region. The results highlight the importance of developing adequate economic policy frameworks in order to adequately mitigate the volatility in the terms-of-trade.

Keywords: Fiscal Policy, Natural Resource Management, Time-Varying Bayesian SVAR.

JEL classification: E62, D72, F3, H60.

1   Intro and Motivation

The availability of abundant natural resources can generate adverse effects on an economy, giving rise to the phenomenon known as the "natural resource curse". Among these effects are: (1) risks to macroeconomic stability arising from the high volatility of commodity prices, (2) restraints on economic growth and productive diversification due to specialization in natural resource activities, (3) social conflicts over natural resource extraction, and (4) the proliferation of weak institutions and corruption (Frankel, 2010).

Concerning macroeconomic stability, terms-of-trade volatility has been found to be an important source of macroeconomic fluctuations (Kose , 2002; Mendoza, 1995), particularly in resource-rich countries. Negative terms-of-trade shocks, for instance, can generate larger current account deficits and significant exchange rate variability (Harberger, 1950; Laursen and Metzler, 1950; Ostry and Reinhart, 1992). In emerging markets, terms-of-trade shocks are an *essential* source of cyclical fluctuations. Fernández et al. (2018) find that fluctuations in commodity prices explain an important part of the business cycle behavior in emerging countries. In terms of economic growth, Fernández et al. (2020) report that global shocks affecting commodity prices are a relevant (although not dominant) source of variation in the economic growth of countries worldwide, and Céspedes and Velasco (2012) show that commodity price shocks have a significant impact on the dynamics of production and investment.

Despite abundant evidence on the effects of commodity price volatility in the economic activity of commodity-rich countries, the discussion about the evolution of its magnitude and transmission channels is far from being settled, particularly in a global and changing environment, where the sources of commodity price volatility are multiple. In this paper, we focus on studying the time-varying transmission of commodity price shocks to economic activity through the fiscal channel,[1] paying special attention to the procyclicality bias of fiscal policy found in commodity-dependent economies.

Standard economic theory predicts that fiscal policy should be countercyclical. The neo classical tax smoothing model of Barro (1979) says that because taxes are distorting to economic activity, the optimal fiscal response by economic authorities would be to smooth public spending and taxes throughout the economic cycle, generating budget surpluses in good times and temporary deficits in bad times. A similar result occurs in the standard neo-Keynesian framework (Mankiw and Romer, 1991). In this case, the reasons are related to market failures that can spur a recession, which in turn pushes a government to run budget deficits during such times in order to return the economy to a path of growth.

In practice, however, the empirical literature has found that countries present persistent deficits throughout the economic cycle. Such procyclicality is most evident in Latin America (Ilzetzki and Vegh, 2008) and in countries where public finances rely in a significant way on the revenues (taxes, royalties, profits, and credit) coming from the exploitation of natural

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[1]Villafuerte and Lopez-Murphy (2009), Erbil (2011), and Levy et al. (2020) show that fluctuations in commodity prices can generate significant impacts on countries' fiscal variables, especially in countries with emerging economies.

resources (Alberola and Sousa, 2017). In this sense, fiscal policy presents a procyclical tendency or a procyclical bias.

One of the reasons why such procyclicality occurs is credit constraints.The implication of high commodity price volatility is that, when international prices of natural resources are low, governments in resource-rich nations are credit constrained, and when these prices are high, governments are credit relaxed. This condition, imposed by commodity price volatility, results in a procyclical bias on fiscal policy. Thus, the loss (or the lack) of access to financial markets during economic downturns may make it impossible for countries to follow a countercyclical policy (Caballero and Krishnamurthy, 2004; Gavin and Perotti, 1997). This effect that can be reinforced by incomplete financial markets, which affect the financing options available (Riascos and Vegh, 2003).

Nevertheless, as Céspedes and Velasco (2014) claim, the fact that a government can borrow during a commodity-price boom does not mean that it will find it *desirable* to borrow (or increase spending) in such a period. Thus, credit restrictions that change according to the economic cycle do not fully explain the fiscal procyclicality in countries rich in natural resources. For this reason, additional explanations of the procyclicality of fiscal policy have come from the field of political economy.

Based on the observation that, during booms, government expenditures increase more proportionally than revenues, some political economy theories have been proposed under different political arrangements and institutional settings. One, the "voracity effect" (Lane and Tornell, 1996; Tornell and Velasco, 1992; Tornell and Lane , 1999), posits that in political systems in which power is diffused among several agents (each of them pushing for their share of the windfall), fiscal policy will be more procyclical, relative to a more centralized system.

Another theory focuses on the interaction of political agency and information asymmetries (Alesina et al., 2008). The setting is straight forward: while voters observe the state of the economy during a boom (and therefore demand higher utility for themselves), they cannot observe government behavior; thus, they end up facing corrupt governments that appropriate a portion of tax revenues for unproductive public consumption or rent-seeking, which results in a procyclical bias in fiscal policy. Moreover, when voters are not well informed and there exists a high degree of polarization, politicians may have incentives to accumulate public debt and pursue a procyclical fiscal policy (Aguiar and Amador, 2011; Ilzetzky, 2011).

The undesired economic and social outcomes of fiscal procyclicality are multiple. Among them are an unsustainable level of public debt, regressive and intergenerational distributional impacts, and the occurrence of sovereign debt defaults and other economic crises, with the consequential increases in poverty, inequality, migration, social conflict, and political turmoil.

Given the importance of resources derived from commodity exports in the Andean region and the heterogeneous development of fiscal institutions in some of the region's countries in recent years (Kehoe & Nicolini, 2022), this paper studies the time-varying impact of terms-of-trade shocks on economic growth and fiscal performance in Bolivia, Colombia, Ecuador, Peru, and Venezuela. To accurately capture the variation on both, (i)the magnitude of shocks and (ii)the response of economic and fiscal variables to them, this paper makes use of a TV-SVAR model (Primiceri, 2005). This methodology allows us to evaluate whether

improvements in economic (fiscal) institutions generated a lower sensitivity to terms-of-trade fluctuations during the recent decades and what the fiscal and economy-wide time varying responses to these type of shocks have been.

There could be at least two sources of variation in the relationship between terms-of-trade shocks and macroeconomic and fiscal outcomes through time. The first is related to changes in the heteroskedasticity of the shocks and the second consists of changes in the transmission mechanism, that is, how macroeconomic and fiscal variables respond to shocks. If terms-of-trade shocks vary over time, their transmission mechanism could be directly impacted and, if agents are rational and forward-looking, fiscal policy responses to those changes will be incorporated into the agents economic decisions, introducing additional modifications to the transmission mechanism. This is key for the Andean region where, due to both shocks and policies, agent responses have varied significantly over time.

The results found in this paper show a time-varying dynamic for both terms-of-trade shocks and the sensitivity of the Andean region's economies to these shocks. In addition, the results reveal heterogeneity in the impacts of terms-of-trade shocks in the different countries. Thus, countries such as Venezuela, Bolivia and Ecuador exhibit a marked procyclical effect in their fiscal policy, while Peru and Colombia stand out for having a more countercyclical stance. This heterogeneity coincides with the uneven institutional development in the countries of the region. These results highlight the importance of developing adequate economic policy frameworks to mitigate the volatility of the terms-of-trade in the region.

The remainder of the paper is structured as follows: Section 2 includes a short literature review. Section 3 introduces the empirical methodology. Section 4 describes the data and our main results. Section 5 offers some concluding remarks.

2   Literature Review

Different studies have shown that variations in the terms-of-trade can generate significant impacts on the main macroeconomic variables, especially in the case of emerging countries that are highly dependent on the exportation of natural resources. Employing a general equilibrium model developed for 23 developing countries, Mendoza (1995) finds that variations in the terms-of-trade explain approximately half of the variation observed in GDP in the countries evaluated. These results are supported by the findings of Kose (2002) who constructs a more detailed production structure than the original framework developed by Mendoza (1995). In the case of Latin America and the Caribbean (LAC), Izquierdo et al. (2008) show that economic growth in industrialized countries, U.S. high-yield-bond spreads, and terms-of-trade are the external factors that most affect the economic dynamics of LAC countries. In addition, these authors argue that growth in the terms-of-trade is associated with long-term growth in the GDPs of the countries in the region.

Concerning the influence of the terms-of-trade on fiscal variables, Villafuerte and Lopez-Murphy (2009) analyze the fiscal responses of oil-producing countries from 2003 to 2008, finding that the countries present deterioration in their non oil primary balances that is explained by an increase in primary expenditure. In addition, the authors find that most

of the countries in their sample presented a procyclical fiscal policy and a deterioration in the long-term fiscal position. In the same vein, Erbil (2011) studies the cyclical behavior of fiscal policy in 28 oil-producing developing countries from 1990 to 2009, finding a high procyclicality of different fiscal variables. At the same time, this author presents evidence of heterogeneous effects across countries according to different income levels, with spending being procyclical in low- and middle-income countries and countercyclical in high-income countries.

Different studies have found procyclical behavior in LAC (Gavin et al., 1996; Gavin and Perotti, 1997; Stein et al., 1999); Camacho and Perez-Quiros (2014) analyze the behavior of output in Argentina, Brazil, Chile, Colombia, Mexico, Peru, and Venezuela in the face of changes in commodity prices. The authors report that each country's response to commodity price shocks depends on the moment in time, the magnitude of the shock, and whether the shock represents a reduction or an increase in prices. Melo-Becerra et al. (2020) find that the effects of changes in the price of oil on the Colombian economy have changed over time and depend on the direction of the variation in the commodity's price. In this sense, significant increases in the price of oil did not seem to affect public debt, while reductions in oil prices increased the level of indebtedness.

The literature has highlighted the development of adequate fiscal institutions as one of the factors thgat contributes to the reduction of fiscal policy procyclicality. Medina (2016) studies the effects of commodity price shocks on fiscal revenues and expenditures in LAC countries in the period 2005—2013, finding that fiscal aggregates respond to commodity price shocks, although with heterogeneities across countries: Medina (2016) finds that Venezuela has the highest fiscal sensitivity to commodity price shocks, whereas Chile's fiscal sensitivity is low. This result may be linked to the establishment of fiscal rules that reduce the sensitivity of fiscal variables to fluctuations in international prices.

Céspedes and Velasco (2012) argue that the implementation of credible fiscal rules can reduce the impact of external shocks on output and reduce the sensitivity of the exchange rate. Additionally, the type of monetary regime and the existence of stable political systems can reduce the sensitivity of macroeconomic variables to commodity price shocks. Frankel et al. (2013) find that many developing countries have managed to escape fiscal procyclicality in recent years and have even implemented countercyclical fiscal policies as developed countries do. According to these authors, a key driver of this effect is the development of effective institutions in economic, legal, and political terms. Similar results are found by Céspedes and Velasco (2014) and Cespedes et al. (2014). However, recent works such as Bjørnland and Thorsrud (2019) have found that even in countries such as Norway, the establishment of fiscal rules may not have contributed much to the reduction of the procyclicality of fiscal policy. These authors highlight that the time-varying dynamics of shocks are crucial elements to consider in order to accurately model shocks generated by changes in international commodity prices.

The countries of the Andean region have made heterogeneous progress in recent years in terms of the development of fiscal institutions (Andrian et al. (2022)). Colombia, Ecuador, and Peru have instituted fiscal rules during the last 20 years. Colombia and Peru are the only countries with independent fiscal councils and moreover Peru has a solid medium-term

fiscal framework in place. Meanwhile, as part of the fiscal institutional framework defined in Colombia, the Ministry of Finance and Public Credit (MHCP) publishes an annual Medium-Term Fiscal Framework (MFMP), which presents the results and sets forth the purposes of fiscal policy. In addition to an MFMP, Colombia has a Medium-Term Expenditure Framework (MGMP). Although Ecuador does not have a MFMP, the country has multiannual budgetary programming. Finally, Bolivia has not shown any significant development in its fiscal institutions in recent years compared to the rest of the countries in the region.

3 Methodology

3.1 Time-varying VAR

There could be at least two sources of variation in the relationship between external shocks and macroeconomic and fiscal outcomes through time. The first is related to changes in the heteroskedasticity of the shocks, which are, if anything, volatile. The second are changes in the transmission mechanism, that is, how macroeconomic and fiscal variables respond to shocks. If shocks vary over time, this could directly impact their transmission mechanism and, if agents are rational and forward-looking, the fiscal policy response to those changes will be incorporated into their economic decisions, introducing additional modifications to the transmission mechanism. For an econometric model to capture these dynamics, it must have two features: (1) a multiple equation model, in order to capture how changes in fiscal policy affect the rest of the economy, and (2) time-varying parameters, in order to measure and understand how volatility and policy changes affect agents' behavior. Therefore, to estimate changes in macrofiscal outcomes through time, as in Primiceri (2005), Nakajima et al. (2011), Del Negro and Primiceri (2013), and Melo-Becerra et al. (2020), we use the following time-varying structural vector auteregression model with stochastic volatility:

$$y_t = c_t + B_{1,t}y_{t-1} + ... + B_{k,t}y_{t-k} + u_t, \qquad (1)$$

where y_t is a $n \times 1$ vector of observed endogenous variables; c_t is a $n \times 1$ vector of time-varying coefficients that multiply constant terms; $B_{i,t}$, $i = 1, ..., k$, are $n \times n$ matrices of time-varying coefficients; and u_t are heteroskedastic unobservable shocks with variance-covariance matrix Ω_t. For simplicity, consider the triangular decomposition of Ω_t:

$$A_t \Omega_t A_t^T = \Sigma_t \Sigma_t^T, \qquad (2)$$

where A_t is the lower triangular matrix:

$$A_t = \begin{bmatrix} 1 & 0 & \cdots & 0 \\ a_{12,t} & 1 & \ddots & \vdots \\ \vdots & \ddots & \ddots & 0 \\ a_{n1,t} & \cdots & a_{nn-1,t} & 1 \end{bmatrix},$$

and Σ_t is the diagonal matrix:

$$\Sigma_t = \begin{bmatrix} \sigma_{1,t} & 0 & \cdots & 0 \\ 0 & \sigma_{2,t} & \ddots & \vdots \\ \vdots & \ddots & \ddots & 0 \\ 0 & \cdots & 0 & \sigma_{n,t} \end{bmatrix}.$$

This implies that

$$\begin{aligned} y_t &= c_t + B_{1,t}y_{t-1} + \ldots + B_{k,t}y_{t-k} + A_t^{-1}\Sigma_t\varepsilon_t \\ \beta_t &= \{c_t, B_{1,t}, \ldots, B_{k,t}\} \\ V(\varepsilon_t) &= I_n. \end{aligned} \tag{3}$$

The modeling strategy consists of estimating the coefficient processes in equation (3). This model has two sets of time-varying coefficients, β_t and $a_{ij,t}$, and a stochastic volatility model for the diagonal elements, $\sigma_{i,t}$. The dynamics of the model's time-varying parameters are specified as follows:

$$\beta_t = \beta_{t-1} + \eta_t, V(\eta_t) = Q \tag{4}$$

$$a_t = a_{t-1} + u_t, V(\zeta_t) = S \tag{5}$$

$$\ln \sigma_t = \ln \sigma_{t-1} + \nu_t, V(\nu_t) = W, \tag{6}$$

where the elements of the vector β_t are modeled as random walks as well as the free elements of the matrix A_t. The standard deviations (σ_t) are assumed to evolve as geometric random walks, which belong to the class of models known as stochastic volatility (see Blake and Mumtaz (2017)). All of the model's innovations are assumed to be jointly normally distributed with the following assumptions on the variance covariance matrix:

$$V = Var\left(\begin{bmatrix} \varepsilon_t \\ \eta_t \\ \zeta_t \\ \nu_t \end{bmatrix}\right) = \begin{bmatrix} I_n & 0 & 0 & 0 \\ 0 & Q & 0 & 0 \\ 0 & 0 & S & 0 \\ 0 & 0 & 0 & W \end{bmatrix}, \tag{7}$$

where I_n is a n-dimensional identity matrix and Q, S and W are positive definite matrices. Primiceri (2005) states these assumptions are justified, for at least two reasons: (1) the high number of parameters in the model (adding all the off-diagonal elements of V would require the specification of a sensible prior that is able to prevent cases of ill-determined

parameters), and (2) allowing for a completely generic correlation structure among different sources of uncertainty would preclude any structural interpretation of the innovations.

Finally, S is assumed to be block diagonal, with the blocks corresponding to parameters belonging to separate equations. The intuition is that the contemporaneous relationship among variables are assumed to evolve independently in each equation. More formally, from equations (2) and (3), consider the following relationship:

$$A_t u_t = \varepsilon_t. \tag{8}$$

In the case of a three-variable VAR, this relationship implies the following set of equations:

$$\begin{pmatrix} 1 & 0 & 0 \\ a_{12,t} & 1 & 0 \\ a_{13,t} & a_{23,t} & 1 \end{pmatrix} = \begin{pmatrix} u_{1,t} \\ u_{2,t} \\ u_{3,t} \end{pmatrix} = \begin{pmatrix} \varepsilon_{1,t} \\ \varepsilon_{2,t} \\ \varepsilon_{3,t} \end{pmatrix}, \tag{9}$$

which can be rearranged and translated into the following equation forms:

$$u_{1,t} = \varepsilon_{1,t}$$
$$u_{2,t} = -a_{12,t} u_{1,t} + \varepsilon_{2,t}$$
$$u_{3,t} = -a_{13,t} u_{1,t} + -a_{23,t} u_{2,t} + \varepsilon_{3,t},$$

where $V(\varepsilon_t) = \Sigma_t$ and

$$a_{12,t} = a_{12,t-1} + \zeta_{1,t}, V(u_{1,t}) = S_1 \tag{10}$$

$$\begin{pmatrix} a_{13,t} \\ a_{23,t} \end{pmatrix} = \begin{pmatrix} a_{13,t-1} \\ a_{23,t-1} \end{pmatrix} + \begin{pmatrix} \zeta_{2,t} \\ \zeta_{3,t} \end{pmatrix}, V(\begin{pmatrix} \zeta_{2,t} \\ \zeta_{3,t} \end{pmatrix}) = S_2. \tag{11}$$

Therefore, the $a_{ij,t}$ are time-varying coefficients on regressions involving the VAR residuals.

3.1.1 Estimation procedure

Because the model includes a number of parameters that change over time, the econometric procedure for estimating the model is the Bayesian approach. In these types of models, parameters are considered latent variables so that they can be estimated with state-space models. The unobservable states correspond to the history of coefficients (β^T, A^T) and the history of volatilities (Σ_t). A Bayesian approach reduces uncertainty about the parameters and evaluates their posterior distribution and that of the hyperparameters (V).

Gibbs sampling, a Monte Carlo with Markov Chains (MCMC) simulation method, is used for the posterior numerical evaluation of the parameters of interest. This procedure sets

the prior joint distribution of the parameters function and then updates it with the data likelihood function, which generates the posterior probability function of the parameters. Because this function is analytically intractable, Gibbs sampling generates samples of the parameters from their full conditional distributions, which are smaller than their joint distribution. It is an iterative algorithm that builds sequence-dependent values of the parameters with distributions converging toward their posteriors after a number of iterations.

Priors   Following Blake and Mumtaz (2017), the priors for the initial states of the time-varying coefficients, simultaneous relations, and log standard errors $p(\beta_0)$, $p(a_0)$, and $p(\ln \sigma_0)$ are assumed to be normally distributed. These assumptions together with the autoregressive nature of time-varying parameters imply Gaussian priors on the entire sequences of the β's, a's, and $\ln \sigma$'s (conditional on Q, W, and S). The priors for the distributions of the hyperparameter Q is assumed to be distributed as inverse Wishart, because the parameter W is inverse gamma and the blocks of S are inverse gamma for S_1 and inverse Wishart for S_2. More formally, these prior distributions have the following form:

$$\beta_0 \sim N(\hat{\beta}, V(\hat{\beta}))$$

$$A_0 \sim N(\hat{A}, V(\hat{A}))$$

$$\ln \sigma_0 \sim N(\ln \hat{\sigma}, I_n)$$

$$Q \sim IW(Q_0, T_0)$$

$$W \sim IG(W_0, T_0)$$

$$S_1 \sim IG(S_{10}, T_0)$$

$$S_2 \sim IW(S_{20}, T_0)$$

The joint distribution of β^T, A^T, Σ^T, and V is given by

$$[\Sigma^T, \beta^T, A^T, V|y^T]\alpha[y^T|\Sigma^T, \beta^T, A^T, V][\Sigma^T, \beta^T, A^T, V],$$

where the rectangular parenthesis is interpreted as the probability distribution and the symbol α expresses proportionality.

Simulation method   A detailed explanation of the Gibbs sampling methods used to estimate the model is provided in section A.1.3 of the appendix. A summary is provided here. Following Del Negro and Primiceri (2013) and Blake and Mumtaz (2017) the Gibbs sampling algorithm is used to generate a sample from the joint posterior of (β^T, A^T, Σ^T, and V). The algorithm is carried out in four steps, drawing (1) time-varying coefficients (β^T), (2) simultaneous relations (A^T), (3) volatilities (Σ^T), and (4) hyperparameters (V), conditional on the observed data, y_t, and the rest of the parameters.

Conditional on A^T and Σ^T, the state-space form given by equations (3) and (4) is linear and Gaussian. Therefore, the conditional posterior of β^T is a product of Gaussian densities and can be drawn using the Carter and Kohn (1994) algorithm. By the same logic, the posterior of A^T conditional on β^T and Σ^T is also a product of normal distributions and A^T can thus be drawn in the same way. Drawing Σ^T instead is more complicated; this can be done using the independence Metropolis-Hastings algorithm, as presented by Blake and Mumtaz (2017). Simulating the conditional posterior of V is standard, because it is the product of independent inverse Wishart distributions.

Identification   As long as an exact identification scheme —like the triangular matrix described above— for the additive shock is available, a structural vector autoregression (VAR) can be easily estimated in two steps. Consider the following structural VAR:

$$y_t = X'_t + \Xi_t \varepsilon_t, \tag{12}$$

where Ξ_t, $t = 1, ..., T$, are $n \times n$ matrices[2] and for any t, Ξ_t contains at least $\frac{n(n-1)}{2}$ restrictions that guarantee identification. The first step consists of estimating the reduced form VAR, following the methodology illustrated in the previous sections. This first step delivers the posterior of the β's and the Ω's at every point in time. The second step is to obtain the numerical evaluation of the posterior of the $\Xi's$; it suffices to solve the system of equations given by

$$\Xi_t \Xi'_t = \Omega_t, t = 1, ..., T \tag{13}$$

for every draw of Ω_t. If the identification is based on a triangular scheme, the solution to (12) is given simply by $\Xi_t = A_t^{-1}\Sigma_t$, the same as in equation (3).

Sign restrictions.   In order to asses the time change in the degree of procyclicality of fiscal policy to terms-of-trade shocks, we identify the shock by assuming that a *positive* shock has the following contemporaneous macrofiscal effects:

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[2]These do not need to be triangular. See Primiceri (2005) for a discussion.

$$A_{0,t} = \begin{pmatrix} 1 & 0 & 0 & 0 \\ a_{12,t} & 1 & 0 & 0 \\ a_{13,t} & a_{23,t} & 1 & 0 \\ a_{14,t} & a_{24,t} & a_{34,t} & 1 \end{pmatrix}.$$

The second, third, and fourth rows of this matrix correspond to the contemporaneous response of the real exchange rate (rer), the primary balance (pb), and real GDP ($rgdp_t$), respectively, to a terms-of-trade shock. Restricting these contemporaneous responses the following way $a_{12,t} < 0$ (Real Exchange Rate [RER] appreciation); $a_{13,t} < 0$, $a_{23,t} < 0$; and $a_{14,t} > 0$, $a_{24,t} > 0$, $a_{23,t} > 0$ implies that a contemporaneous shock ε_{tot_t} leads to (1) a contemporaneous RER appreciation, (2) a fall in pb_t (e.g., an increase in primary expenditures that is not proportional to the increase in revenues) and (3) a contemporaneous increase in $rgdp_t$ (e.g., as higher oil income increases domestic demand, both public and private.).

To answer the question of whether there is an increase in fiscal procyclicality over time, one has to compare the dynamic behavior of the contemporaneous fiscal response to oil price impulses. Accordingly, we compare the contemporaneous response of the primary balance, $a_{13,t}$, over time. If $a_{13,t} < a_{123,t-s}$ (where $t > s$) during the period considered, fiscal procyclicality deteriorates. Conversely, if $a_{13,t} > a_{13,t-s}$, fiscal procyclicality improves.

The same criterion is applied to assess the direct multiplier effect of terms-of-trade shocks on the economy as well as their indirect multiplier effect (through fiscal policy). If $a_{14,t} > a_{14,t-s}$ and $a_{34,t} > a_{34,t-s}$, then terms-of-trade shocks have a higher multiplier effect on the economy, whereas if $a_{14,t} < a_{14,t-s}$ and $a_{34,t} < a_{34,t-s}$ the opposite occurs: the multiplier effect of terms-of-trade shocks on the economy is lower. The procedure to drawing matrix $A_{0,t}$ is detailed in Appendix B.

4 Data and stylized facts

This paper evaluates the impact of terms-of-trade shocks on fiscal performance and economic growth in the countries of the Andean region. Thus, in addition to a variable that captures the terms-of-trade dynamics, the VAR model for each country includes the following variables: Economic Growth, Primary Balance (Fiscal Measure), and Real Exchange Rate (Monetary Measure). This framework enables the capture of the main transmission channels, which is necessary to explain the impacts generated by terms-of-trade shocks. The data's periodicity is annual and the length of the periodicity depends on the availability of data for each country. Table 1 describes the variables and sources by country.

4.1 Bolivia

Following Kehoe et al. (2022), the recent economic history of Bolivia can be divided into five distinct periods. The first, 1960—1977, was characterized by a high rate of economic

growth. In contrast, the period 1977—1986 was characterized by a debt crisis and hyperinflation in the country. Subsequently, between 1986 and 1998, Bolivia experienced a period of 10 slow economic recovery, the dynamics of which were reversed in the period 1998—2002, when the country experienced a financial crisis. Finally, the period 2002—2017 was one of dynamic growth, driven by the commodity super cycle.

Bolivia was characterized by consistent positive GDP growth from the late 1980s up to 2019. Over the last 35 years, the Bolivian economy has experienced some short episodes of negative growth, one of which was the period after the Bolivian economic crisis of the mid-1980s. During this period, Bolivia was affected by reduced availability of foreign external credit, a political crisis, and economic decisions such as indexing the minimum wage to past inflation and printing money to finance the fiscal deficit. All these factors negatively affected the economy to such an extent that hyperinflation set in: the CPI reached the astonishing level of 11,000 percent (Kehoe et al. (2022)) according to the available data series, the most singular behavior in terms of RER variation happened in the mid-1980s (figure 1c), when it went from its highest to its lowest value in consecutive years (72 and -71 percent in in 1985 and 1986, respectively).

The most recent year with negative growth was 2020, when the COVID-19 pandemic brought the global economy to a near standstill for weeks. Even though Bolivia did not show negative growth in the late 1990s crisis, it is important to note its low GDP growth in 1999; this rate, a sign of the effects of the financial crisis according to Jemio (2006), was a constant until 2002. In 2003, gas exports from Bolivia to Brazil increased and Bolivian exporters increased their prices as well, which resulted in higher foreign income and an export boom (Kehoe et al. (2022)).

The export boom that began around 2004 helped revive the Bolivian economy after the late 1990s crisis. This is especially evident in the data available for the primary balance, which are presented in figure 1b: beginning in 2003, the primary deficit started to decline and the primary balance was positive from 2005 through the remainder of the 2000s. In addition, there was a primary surplus until 2013, the year when the exports boom came to an end as commodities prices fell. Secondly, the country saw its worst-ever primary balance deficit, almost 9 percent of GDP, in 2020, when the pandemic hit.

In terms of fiscal institutions, unlike other countries in the region Bolivia does not currently have a fiscal rule for the execution of the public budget that allows for maintaining fiscal balances and guarantees the sustainability of public sector finances.

4.2   Colombia

Based on an analysis of the dynamics of the main monetary and fiscal variables of Colombia in recent decades, (?Perez-Reyna & ?Osorio-Rodriguez (2022)) identify three distinct periods. The first period, 1960—1970, was one of low inflation and a low fiscal deficit. In contrast, in the second, 1971—1990, the country experienced high and persistent inflation in conjunction with high fiscal deficits. During this period, the use of monetary issuance was relatively frequent in order to finance government expenditures. Thus, fiscal dominance and

the nonexistence of an independent central bank characterized the first two periods. Finally, the 1991—2019 period was one of monetary dominance as a result of the independence of the central bank since 1991. During this third period, the country recorded the highest average fiscal deficit and saw its worst economic recession (except for the crisis generated by the COVID-19 pandemic) in 1999. Despite this crisis, inflation continued downward, reaching single-digit levels in 2010. This crisis management took place in the context of an independent central bank, greater flexibility in foreign exchange markets, and fiscal financing with a greater preponderance of the domestic market.

In GDP-growth terms, Colombia boasts one of the more consistent economies in Latin America for the last 60 years. This country has posted only two years of GDP contraction: 1999 and 2020. In 2020, the pandemic affected the GDP negatively, causing it to decrease by 6.8 percent; in 1999 the economy was noticeably affected when borrowers had trouble paying credits with interest rates indexed to the Constant Acquisitive Power Unit (UPAC in Spanish). The UPAC unit ascended, contrary to what was predicted by Banco de la República (the central bank), and installments that had to be paid rose, which led to the difficulty for people in meeting their obligations. This generated a financial crisis. According to Kalmanovitz (2017), some private banks and most of the public banks went through a process of being liquidated or rescued by the state. In this year, GDP fell 4.2 percent according to our data.

Colombia adopted a new constitution, in 1991, replacing the one that had been in place since 1886. Among many social and economic reforms, Perez-Reyna & Osorio-Rodriguez (2021) highlight two important changes, the first, regarding new resource transfers from central to regional governments and the second, regarding the Banco de la República, which made it independent from the government and responsible for controlling inflation and maintaining the value of the Colombian peso. The first reform resulted in an increase in government expenditure on social security. Because of this increase, the size of the government doubled in the 1990s, which, in tandem with tax revenue remaining flat, brought Colombia to the dire situation in which its economy ended the 20th century with a primary deficit that reached 3.6 percent of GDP. This was the second-largest deficit in 60 years, with the largest one being that which arose during the COVID-19 pandemic (almost 5 percent of GDP).

According to the available data, the terms-of-trade and the GDP growth seem to have had a certain common tendency. This is more evident from the early 1990s onwards, when the Economic Opening (Apertura Económica) policy went into effect under the presidency of César Gaviria, and Colombia opened to international trade, lowering tariffs and attracting foreign capital (Perez-Reyna & Osorio-Rodriguez (2021)). During this period as any of the indicators increase or decrease, so does the other one. This may be an indication of the high correlation of the economy to the terms-of-trade.

In terms of fiscal arrangements, in 2000 Colombia introduced an expenditure rule that regulates the growth of current spending by the central government. However, it was in 2011 that the biggest recent change in the country's fiscal institutional framework was instituted with the creation of the fiscal rule by Law 1473 of 2011. This rule applied to the Central National Government (CNG) due to the high representativeness of the CNG fiscal accounts on the Nonfinancial Public Sector (NFPS) Balance and the behavior of the public debt. The

fiscal rule enacted by this law established targets on the structural balance while allowing for deviations in its targets derived from changes in the output gap and the energy cycle. In addition, the rule allowed for additional countercyclical spending under special circumstances in order to effect countercyclical policies in cases of pronounced economic slowdown. Furthermore, the fiscal rule included an escape clause that made it possible to suspend compliance with the targets in the event of extraordinary events that compromised macroeconomic stability (Barreix & Corrales, 2019). The economic authorities used this escape clause in 2020 when facing the extraordinary expenses of mitigating the shock generated by the COVID-19 pandemic.

In addition, Law 1473 of 2011 created the Advisory Committee for the Fiscal Rule, which is an autonomous advisory committee to which the Colombian government must listen on issues related to fiscal stability. However, its pronouncements are not binding. Finally, the law enacting the fiscal rule also established that the existing fiscal institutions (National Development Plan, Medium-Term Fiscal Framework, Medium-Term Expenditure Framework, and the General Budget of the Nation) should be in line with the goals established by the rule (Barreix & Corrales, 2019).

4.3   Ecuador

Ecuador's recent economic history is marked by two watersheds, the first being the discovery of the country's oil reserves at the end of the 1960s, which contributed to the orientation of the Ecuadorian economy towards oil exports and to accelerated growth, and the second being the dollarization implemented in 2000. This decision was made with the objective of ending a period of high inflation that the country had experienced over the previous three decades. Looking beyond these two important events, Ecuador's recent history can be divided into three periods (Cueva & Díaz, 2021). The first period, between 1960 and 1981, was characterized by rapid economic growth, driven by the oil boom, and stable inflation. The second period, between 1982 and 2000 saw an economy with low economic growth and pronounced inflation that averaged 39 percent per year between 1982 and 1999. Finally, the period starting in 2000 was marked by the dollarization of the economy and a new oil boom generated by high international commodity prices, which contributed to the economy's dynamic growth between 2000 and 2017.

Ecuador had an average GDP growth rate of 2.8 percent between 1980 and 1998, during which time the country had periods of minimal and negative growth. However, in 1999 the GDP growth rate declined sharply in the wake of some political instability in the 1990s and some constitutional and central banking reforms. This was one of the worst crises in the country's history (Cueva & Díaz, 2021): output decreased and GDP growth was -4.7 percent.. After this year, the worst outcome growth period in our series is 2020, the pandemic year.

Ecuador's primary balance after 2008 is shown in figure 3b. Cueva and Díaz (2021) explain that even though there was an increase in oil revenues starting in 2007 and revenues from existing taxes increased and new taxes were imposed, expenditures increased which made the primary balance negative for almost all of the 2010s. This is a clear situation where

a priori exogenous positive factors (e.g., commodities super cycle) may not be completely helpful for every country's primary balance dynamic.

In terms of fiscal institutions, in its recent history Ecuador has put in place different fiscal rules in order to improve the performance of its public finances. For example, the country has implemented Spending Rules (since 2010 and revised in 2020), Balance Rules (2003—2009), and Debt Rules (2003—009). In 2018, a reform of the country's fiscal institutionality was introduced. The new fiscal responsibility framework, which is still in the process of implementation, imposes a limit on spending growth and restricts the deficit, while reaffirmingthe objective of keeping the debt burden below 40 percent of GDP in the medium and long term (Arenas de Mesa & Mosqueira, 2021; Barreix & Corrales, 2019).

4.4 Peru

Peruvian GDP growth series show a volatile behavior during the 1970s and 1980s decade, with with consecutive years seeing large swings between positive and negative values. According to Martinelli and Vega (2021), the 1970s were characterized by accelerating inflation and the 1980s by hyperinflation, which came during a period of economic interventionism with the instituting of measures such as price controls and the presence of multiple exchange rates that did not bring the expected positive results. After this difficult period, GDP variation from year to year became less pronounced and there were fewer periods with negative growth from the 1990s forward (Figure 4a). At the beginning of the 1990s, the government implemented a stabilization program intended to achieve a more sustainable evolution of the economy. As explained by Martinelli & Vega (2021), this program put an end to inflationary fiscal financing and put in place policies in accordance with a market-friendly economy (i.e., free exchange market, taxes, tariff dispersion reduction, and privatization of state-owned firms).

It is worth checking out the behavior of the real exchange rate for Peru (Figure 4c) from the early 1990s. It seems to have a change in its variation, becoming less negative and passing the point of positive values, reaching a maximum variation value within the sample in 2003 (5.7 percent). The starting point of this positive values tendency coincides with the free exchange policy initiated as part of the stabilization program.

The available data show that Peru had a less negative primary balance after the stabilization program. Another contributing factor was the Law of Fiscal Prudence and Transparency, enacted in 1999. This law, which was countercyclical in nature, gave Peru the possibility of reserving resources and the opportunity to use them in emergencies in order to have an easier path during economic cycles (Martinelli & Vega, 2021). The following year the primary balance started on a path toward recovery that lasted until 2008. Nonetheless, during the commodities supercycle, the primary balance had a good recovery, though after it finished Peru presented negative numbers again, with the primary deficit exceeding 7 percent of GDP in 2020.

Peru's fiscal institutional framework has undergone significant changes in recent years. Between 2000 and 2013 the country implemented different fiscal rules, notably the Expen-

diture Rule (2000, revised in 2012 and 2013) and the Balance Rule (2000, revised in 2013). In 2013, the Law for the Strengthening of Fiscal Responsibility and Transparency (LFRTF) was enacted and became effective in 2015. This law established new fiscal rules and created an independent fiscal council and the Fiscal Stabilization Fund (sovereign wealth fund). The LFRTF contemplated a cyclically adjusted balance rule, a debt rule, and the continuation of a previously existing revenue rule that applied to the NFPS. In 2016 the Ministry of Economy and Finance (MEF) enacted the Fiscal Responsibility and Transparency Framework for the Non-Financial Public Sector (MRTF) (Arenas de Mesa & Mosqueira, 2021). This new framework included changes such as modifications to the fiscal rules created in 2013; the changing of the Cyclical Adjusted Balance rule to a Fiscal Balance rule and the establishment of limits on debt, fiscal deficit, and expenditure growth, for example; the establishment of a Multiannual Macroeconomic Framework; and the creation of the Public Infrastructure and Public Services Fund (Arenas de Mesa & Mosqueira, 2021; Barreix & Corrales, 2019)

5   Results

5.1   Bolivia

Figure 5 depicts the stochastic volatility of structural shocks for the four variables considered in the TV-SVAR, including the confidence interval computed for the shock's 16th and 84th percentiles of the standard deviation. In the case of Bolivia, the shocks studied do not show high variability in the period considered, a behavior that contrasts with the results obtained for the other countries in the region. This result may reflect the lower development of public economic institutions in Bolivia in comparison with such institutions in the other countries in the Andean region and a lower degree of integration with the global economy. It should be noted that the limited historical availability of information only allows us to estimate the behavior of these shocks from 2005 onward.

Figure 6 shows the impulse response functions of the variables considered in relation to terms-of-trade shocks over time (identified by sign restrictions). The years considered in the study are shown along the x-axis, the magnitude of the shock. along the the y-axis, and the behavior of the impulse response function each year (for the periods 1 to 15) along the z-axis. In line with Bolivia's commodity-exporting position, positive terms-of-trade shocks are associated with an appreciation of the country's exchange rate in the initial period, followed by a depreciation, the effect of which dissolves over time. However, this effect is not pronounced, due to the country's fixed exchange rate system. Regarding the fiscal effect of the terms-of-trade, the results show a procyclical behavior. Thus, positive shocks are associated with a more significant increase in public spending than in income during the period considered. This effect tends to be greater around 2013. Along these lines, output dynamics reflect the positive effect of terms-of-trade shocks on Bolivian economic growth. This effect may be associated with higher public spending and its economic multiplier effect.

5.2   Colombia

In the case of Colombia, the results reveal different volatility levels among the variables' shocks over time. Concerning the terms-of-trade, the results show a sustained increase in the volatility of the shocks during the period 2008—2014. This behavior reflect a greater dependence on oil in the country's export basket and therefore reflect greater sensitivity of the terms-of-trade to the international price of oil and the so-called "commodity super cycle". The results show a peak of uncertainty in 2014 associated with the fall in the international price of oil.

The volatility of shocks related to the RER showed a downward trend during the period considered, although with some volatility peaks before the year 2000. This behavior can be attributed to the country's institutional development in economic terms, which contributed to a gradual reduction in inflation and better monetary policy management. Thus, although the country's exchange rate continued to act as an automatic stabilizer of external shocks, the improved institutional development in monetary terms made it possible to reduce the relative size of these shocks, making for a "stable" fluctuation of the exchange rate.

On the other hand, in terms of the primary balance shocks present relatively constant volatility during the period of analysis. However, there are peaks of uncertainty related to periods of global economic recession (2000, 2008, and 2020).

Finally, in terms of economic growth, the shocks show a slight negative trend, which is in line with the relative macroeconomic stability of the country and the proper functioning of the economic institutions implemented over the last few years.

The impulse response functions reveal important changes in the magnitude of the shocks over time for the different variables. In terms of the behavior of the RER, terms-of-trade shocks had an initial appreciation effect on the rate (-0.6 p.p. to -0.8 p.p.), followed by a slight depreciation effect in subsequent periods. This behavior is present during all the years considered. The impact on initial appreciation became much stronger during the 2008—2015 period, which coincides with one of the periods of the country's most significant oil and coal production and high international prices of these commodities. As a result of the end of the commodity super cycle, the impact of terms-of-trade shocks on the RER returned to levels similar to those evidenced before the mining-energy boom. This high exchange rate sensitivity to external shocks reflects the country's inflation target regime, which allows the exchange rate to act as an automatic stabilizer in the face of external shocks.

Concerning the primary balance, terms-of-trade shocks show a procyclical behavior throughout the years. In this sense, improvements in the terms-of-trade generate a deterioration in the country's fiscal position, which may be associated with a much more pronounced increase in spending than in income generated by better terms-of-trade. However, since 2010 there was a slight reduction in fiscal procyclicality, which coincided with a period of greater exploitation of natural resources in the country. This improvement could be related to the establishment of a fiscal rule and the development of fiscal institutions in the country in order to make good use of resources derived from extractive industries. Despite this progress, there was a tendency toward greater procyclicality in the last few years of the period under study.

Finally, terms-of-trade shocks show a positive effect on economic growth. The most significant impact of these shocks (0.15 p.p) was felt from 2010 to 2015, coinciding with the commodity cycle and the country's mining-energy boom.

5.3 Ecuador

In the case of Ecuador, the behavior related to the volatility of the different shocks remained relatively constant over time (Figure 9). In terms of impulse response functions, except for the RER behavior, the different variables show a changing reaction over time. The dollarization of the economy explains the low effect of terms-of-trade shocks on the RER. However, due to the crisis generated by the pandemic and the local and foreign price differentials, the RER showed a differential impact during 2020. Regarding fiscal variables, positive terms-of-trade shocks generate a higher deficit, in line with a procyclical behavior of fiscal policy. The procyclicality of spending presented a greater magnitude between 2010 and 2013, coinciding with the commodity super cycle. In terms of economic growth, due to the country's commodity-exporting position favorable terms-of-trade shocks are positively related to higher economic growth, especially in the 2000—2003 and 2010—2013 periods.

5.4 Peru

In Peru, the behavior of the volatility of shocks shows high variability, especially in the case of terms-of-trade and RER shocks (Figure 11). The terms-of-trade volatility shows a peak of uncertainty in the period corresponding to the 2008 global financial crisis and the sharp fall in international copper prices that occurred. From this point on, the magnitude of the shocks shows a considerable decrease with some minor peaks in 2014, coinciding with the end of the commodity super cycle. When examining the behavior of the RER, the shocks present high volatility with a significant peak of uncertainty around the year 2003, which coincides with a period of appreciation of the local currency. Finally, the volatility of fiscal and economic growth shocks shows less variation, although with a marked negative trend, mainly in the case of GDP performance. This result may be associated with more economic stability in Peru due to the strengthening of its fiscal institutions in recent years.

When examining the behavior of the terms-of-trade shocks in the variables considered, we observe in all cases a response that varies significantly according to the years considered (Figure 11). In the case of the response in the RER, the exchange rate appreciation effect that occurs after a positive terms-of-trade shock is of a greater magnitude in the period close to the year 2000 (-2.0 p.p.). A trend of greater sensitivity of the RER is also observed as of 2015, although at much lower levels (-1.5 p.p.) than those recorded in 2000. This result may be related to changes in the Peruvian monetary authorities' exchange rate intervention policy to maintain price stability per the established policy objectives.

Regarding the primary balance, terms-of-trade shocks show a differentiated effect depending on the years considered. From 1990 to 1997, trade shocks were associated with a negative response of the primary balance (higher deficit). However, from 2000 on (except for a few

years, notably 2005 and 2010) terms-of-trade shocks resulted in an improvement of the primary balance. This may be because of improved fiscal institutions that allowed the country to implement a countercyclical fiscal policy. In terms of GDP performance, the response of economic growth to terms-of-trade shocks was similar to that of the primary balance. From 2000 on, positive terms-of-trade shocks were associated with positive economic activity.

5.5   Responses to a Terms-of-Trade Shock for Selected Periods

Figures 13 and 14 show the responses of the primary balance and of economic growth, respectively, to terms-of-trade shocks at different points in time. The periods of interest in which the impluse response functions (IRFs) are calculated respond both to periods of global shocks (international financial crisis, commodity super cycle) and periods in which significant changes occurred in the countries' fiscal institutions (fiscal rules going into effect or new institutional arrangements).

Figure 13 reveals that, with the exception of Peru, the countries of the Andean region show an increase in their deficit as a consequence of a positive terms-of-trade shock. In other words, the fiscal balance reflects a procyclical behavior in public finances. Bolivia is the country that shows the greatest sensitivity to the shocks considered, with 2014 being the year with the most pronounced response in its primary balance (-0.66 p.p.) and 2008 the year with the lowest response (-0.6). For this country, no significant heterogeneity is observed between the different time periods considered. In the case of Colombia, the country's primary balance shows a much smaller reaction than that of the other countries considered, with the year 2000 seeing the greatest impact (-0.06 p.p.), which makes sense given the 1999 Colombian economic crisis. The 2014 shock is associated with a lower impact (-0.05). In Ecuador, the response to shocks varies over the period of study. Thus, the largest (negative) impact on the primary balance was felt in 2012, when there was a significant increase in the international price of oil. As mentioned above, this behavior responds to a procyclical policy by the economic authorities. In contrast, in 2018 the impact on the deficit was much lower (-0.27), a result that may be correlated with the development of better fiscal institutions in Ecuador during the last few years of the period under study. Finally, Peru is the country in the region that shows the most variable responses to terms-of-trade shocks. The largest shock for Peru occurred in 2011, with a deficit shock of -3.0 p.p. that also showed high persistence. This strong shock may capture the decline in copper prices as a consequence of the international financial crisis and the other effects of the global economic recession on the local economy. On the other hand, Peru saw an improvement in its primary balance in the face of terms-of-trade shocks in 2000, 2015, and 2018. This result may be associated with the strengthening of Peruvian economic institutions in recent years. In this sense, our results demonstrate significant differences between the effects that terms-of-trade shocks generate in the countries of the region and the heterogeneity of each country's response during the period of study.

In terms of the impact on growth of terms-of-trade shocks, Figure 14 presents behaviors in line with the fiscal policies carried out by the countries. Positive terms-of-trade shocks generate higher economic growth in all countries, with the exception of some years in the case

of Peru (2011, 2012, 2014). This behavior correlates with the higher public spending and infrastructure investment programs undertaken by the countries in the region made possible by mining and energy tax revenues and their multiplier effect on the economy. Ecuador and Peru show the greatest impact on economic growth derived from positive terms-of-trade shocks, while Colombia shows the least sensitivity.

6   Concluding Remarks

This paper studies the impact of terms-of-trade shocks on the fiscal performance and economic growth of the countries of the Andean region, following the methodology proposed by Primiceri (2005), which enables us to study the dynamics of shocks over time. The results show a wide heterogeneity of the effect of these shocks both within and between countries. Within countries, terms-of-trade shocks have a significant time-varying behavior, with a more pronounced effect during the period of the commodity super cycle. In terms of cross-country heterogeneity, the results show in most cases a procyclical behavior of fiscal policy in the region. However, countercyclical behavior is observed during some of the years considered in cases such as Peru. Some of this heterogeneity in the countries' fiscal response could be explained by differences in the development of economic institutions (fiscal rules and councils) for the proper management of revenue generated from extractive industries.

Our results have important policy implications. On the one hand, policymakers should consider the magnitude of shocks and the nonrenewable nature of resources derived from commodity exports when designing an appropriate policy response. Accordingly, the behavior of public spending should be in line with an intertemporal smoothing, avoiding excessive spending in boom periods that subsequently affects fiscal stability. On the other hand, the development of economic institutions that allow for the proper management of revenues from the extractive industries should be one of the priorities of the countries in the Andean region. Although Peru and Colombia have made significant progress in their fiscal institutions, Bolivia and Ecuador lag in this dimension. Finally, developing proper fiscal institutions will be essential in the energy transition process that is advancing in the world. The countries of the region must prepare themselves adequately to avoid significant impacts on their macroeconomic stability and one step that should be taken is the promotion of greater economic diversification.

Figures and Tables

Table 1: Data by country and source

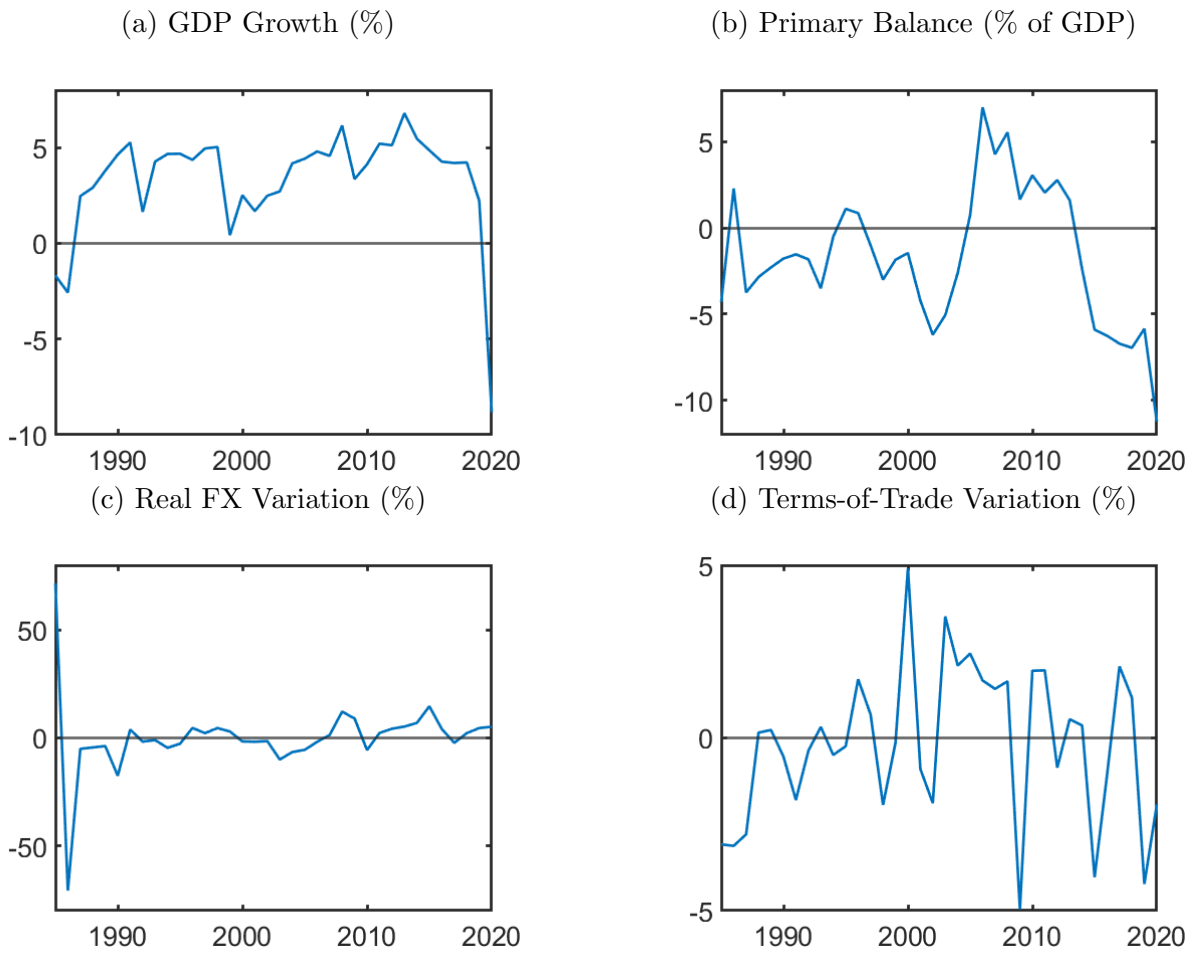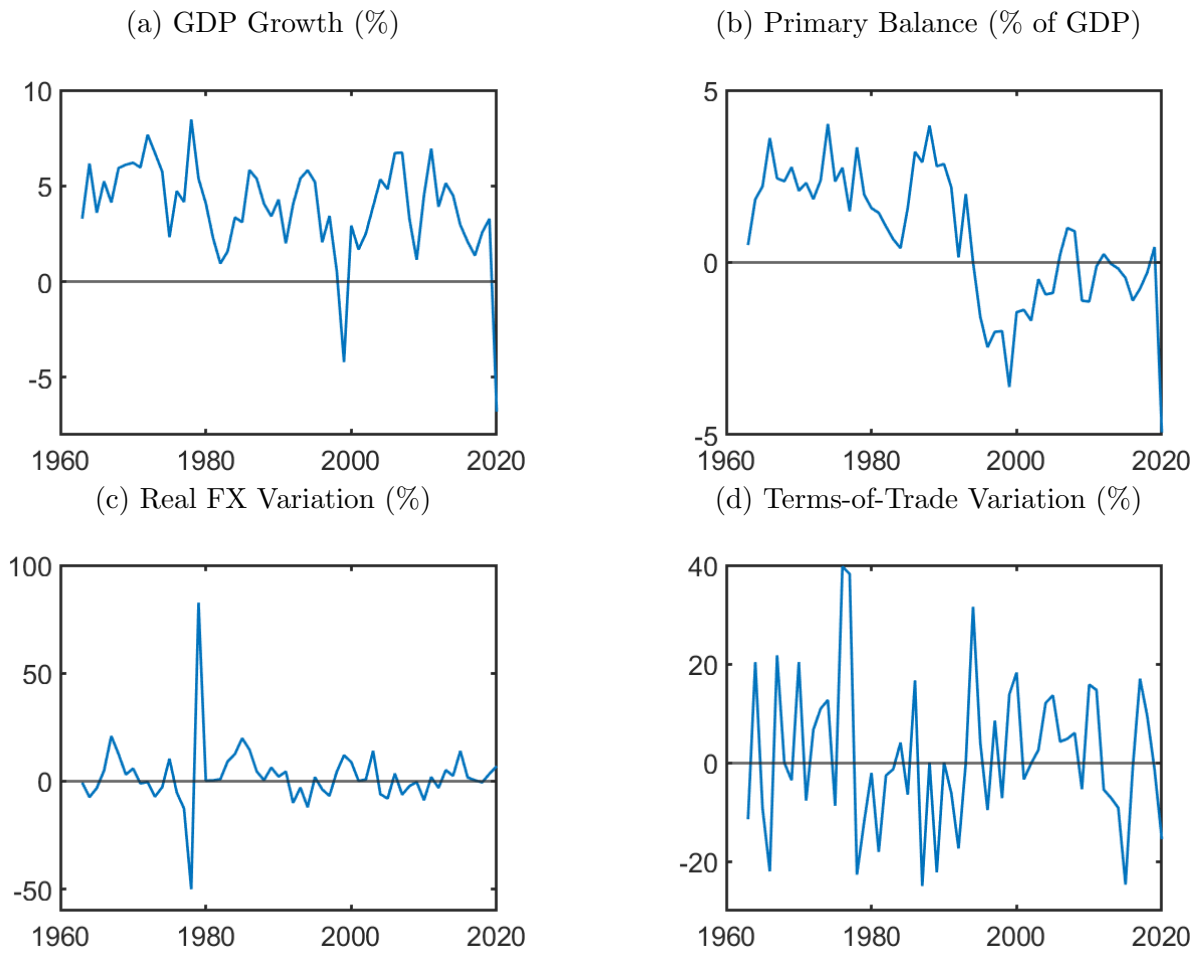| Country	Variable	Source
Bolivia	GDP Growth	World Bank; Central Bank of Bolivia
	Primary Balance	International Monetary Fund
	Real FX Variation	World Bank
	Terms-of-Trade Variation	International Monetary Fund
Colombia	GDP Growth	World Bank; Central Bank of Colombia.
	Primary Balance	International Monetary Fund; Central Bank of Colombia
	Real FX Variation	International Monetary Fund; Central Bank of Colombia
	Terms-of-Trade Variation	International Monetary Fund
Ecuador	GDP Growth	World Bank; Central Bank of Ecuador
	Primary Balance	International Monetary Fund; Central Bank of Ecuador
	Real FX Variation	International Monetary Fund; Central Bank of Ecuador
	Terms-of-Trade Variation	International Monetary Fund
Peru	GDP Growth	World Bank; Central Bank of Peru
	Primary Balance	International Monetary Fund; Central Bank of Peru
	Real FX Variation	International Monetary Fund; Central Bank of Peru
	Terms-of-Trade Variation	International Monetary Fund; Central Bank of Peru
USA	US 10-Yr Yield	Federal Reserve Economic Data

Figure 1: Bolivia Data Series

(a) GDP Growth (%)

(b) Primary Balance (% of GDP)

(c) Real FX Variation (%)

(d) Terms-of-Trade Variation (%)

Figure 2: Colombia Data Series

(a) GDP Growth (%)

(b) Primary Balance (% of GDP)

(c) Real FX Variation (%)

(d) Terms-of-Trade Variation (%)

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Figure 3: Ecuador Data Series

(a) GDP Growth (%)

(b) Primary Balance (% of GDP)

(c) Real FX Variation (%)

(d) Terms-of-Trade Variation (%)
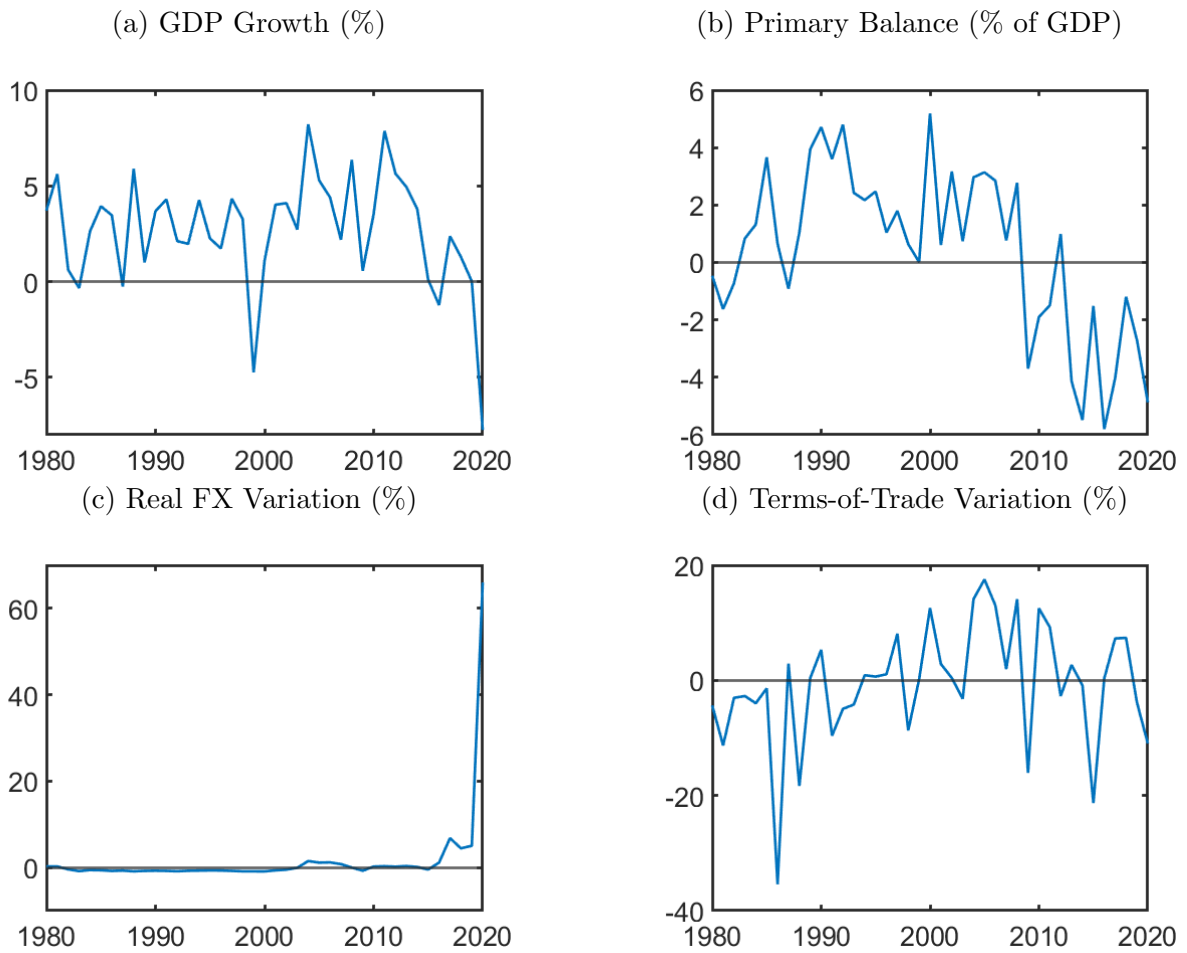
Figure 4: Peru Data Series

(a) GDP Growth (%)

(b) Primary Balance (% of GDP)

(c) Real FX Variation (%)

(d) Terms-of-Trade Variation (%)
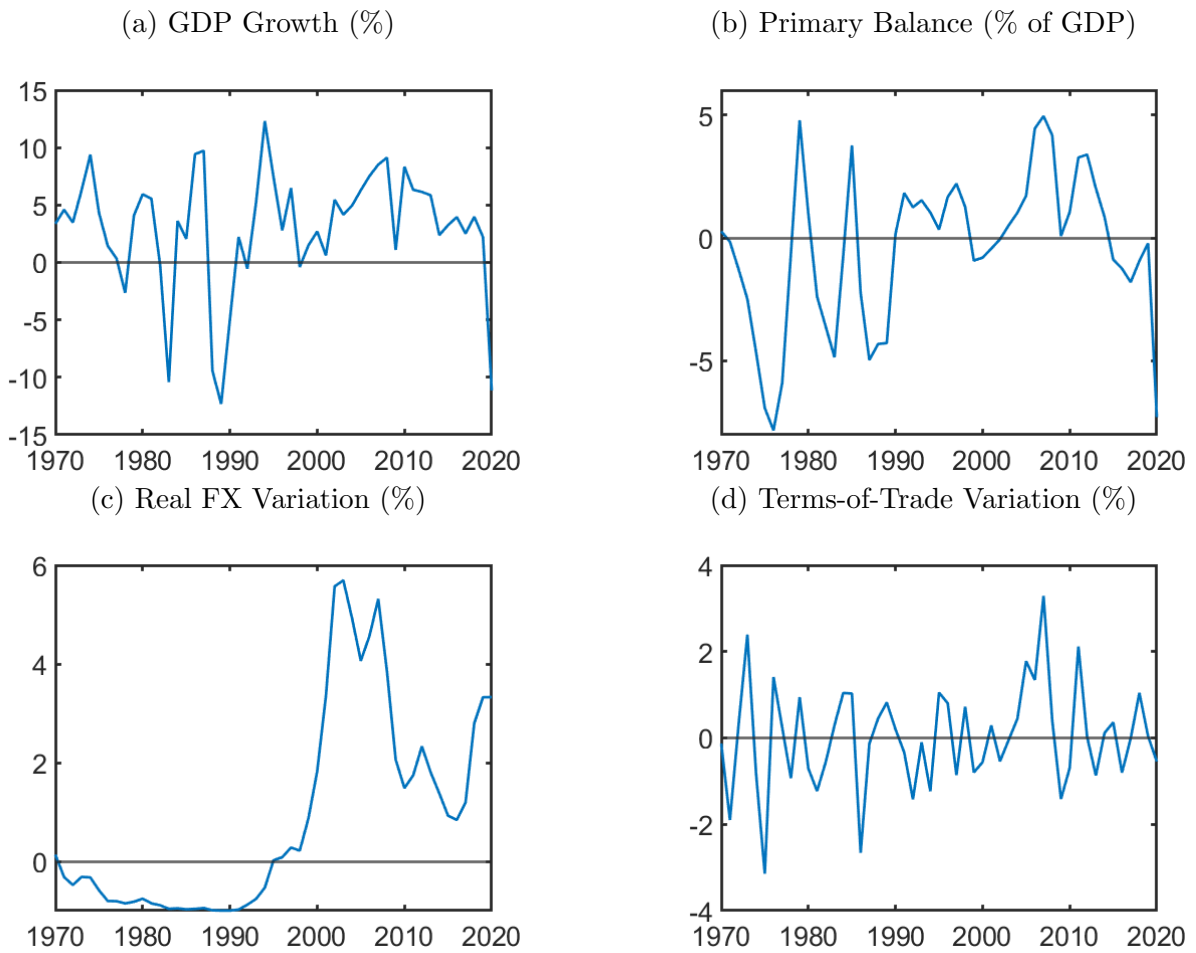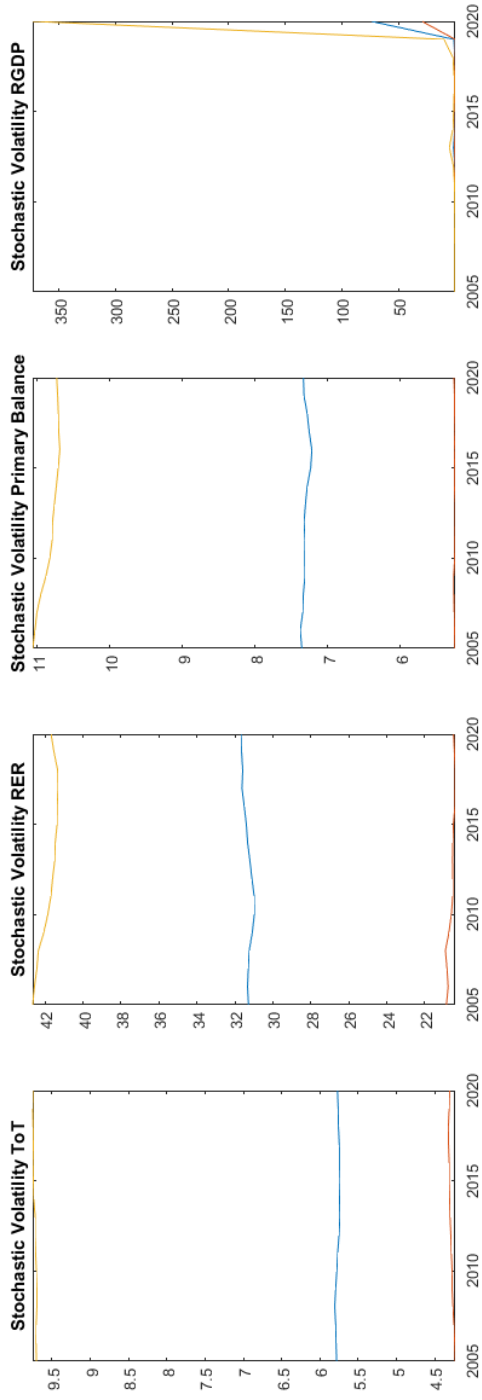
Figure 5: Posterior Estimates for Stochastic Volatility of Structural Shocks, Bolivia
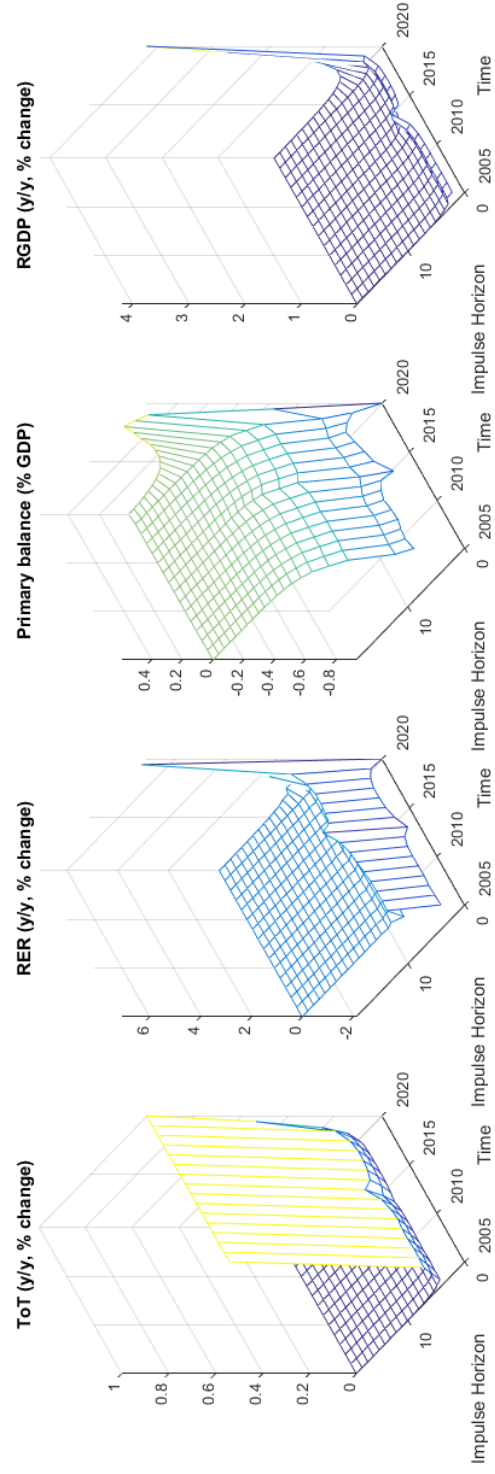
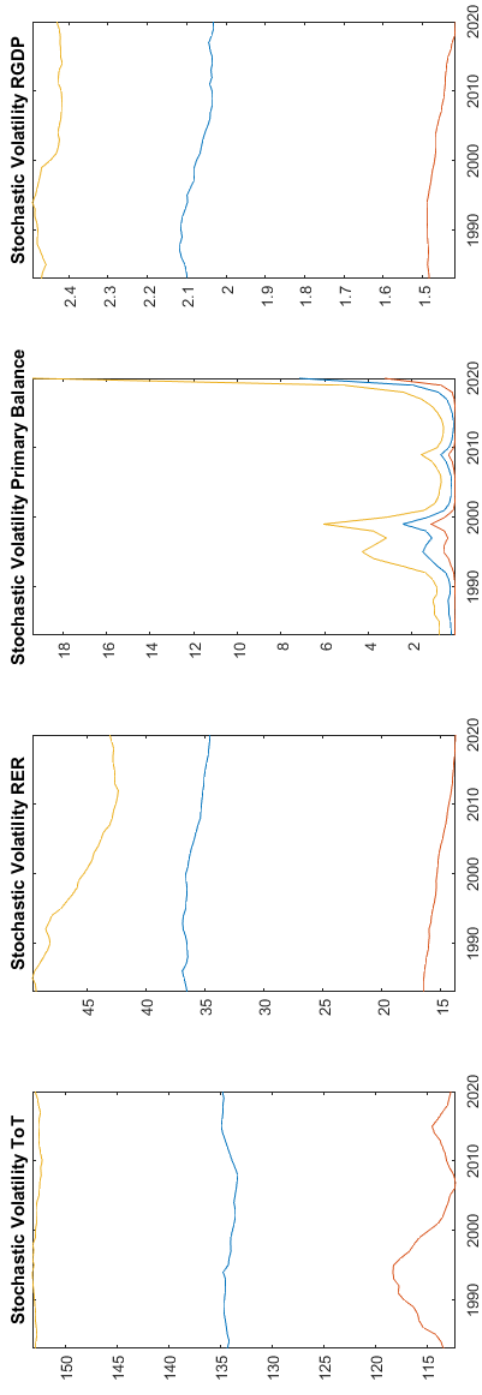Figure 6: Impulse Response Functions, Bolivia

Figure 7: Posterior Estimates for Stochastic Volatility of Structural Shocks, Colombia
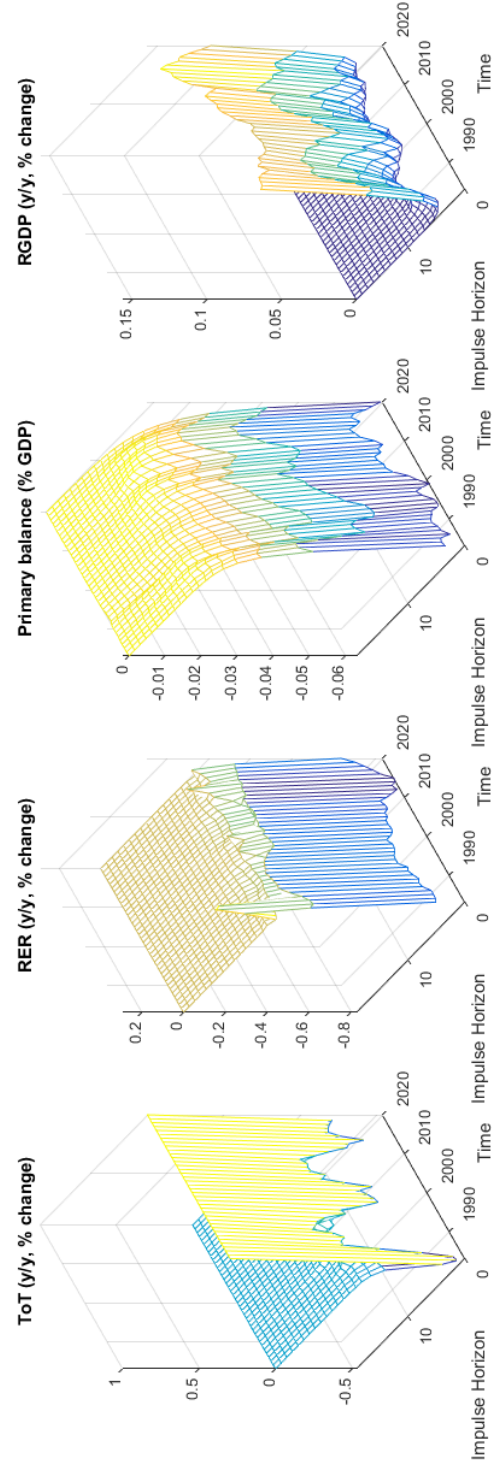

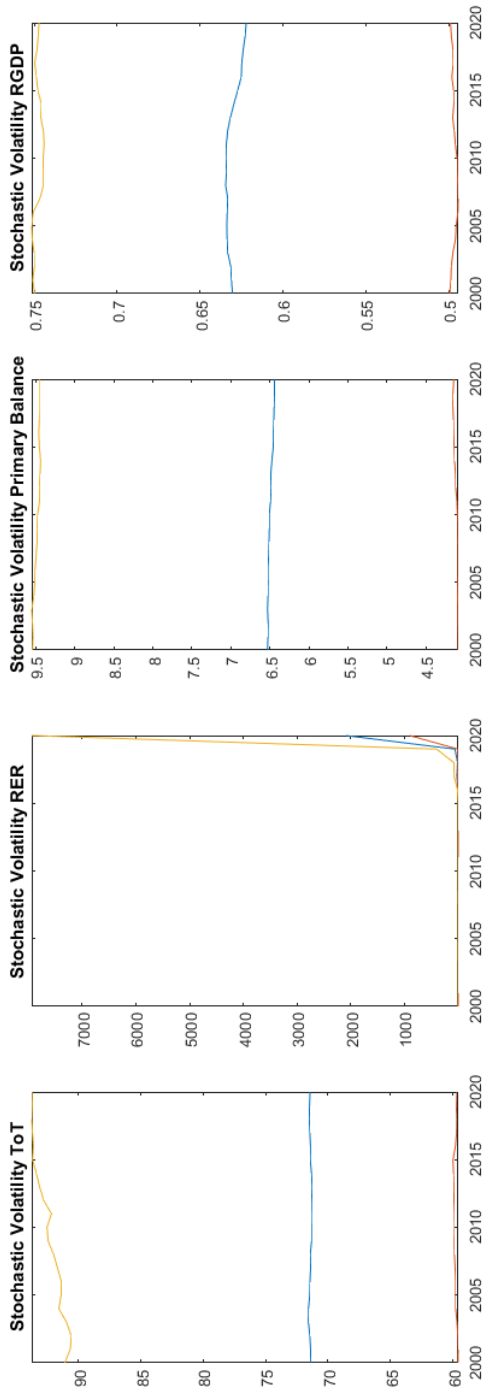Figure 8: Impulse Response Functions, Colombia

Figure 9: Posterior Estimates for Stochastic Volatility of Structural Shocks, Ecuador
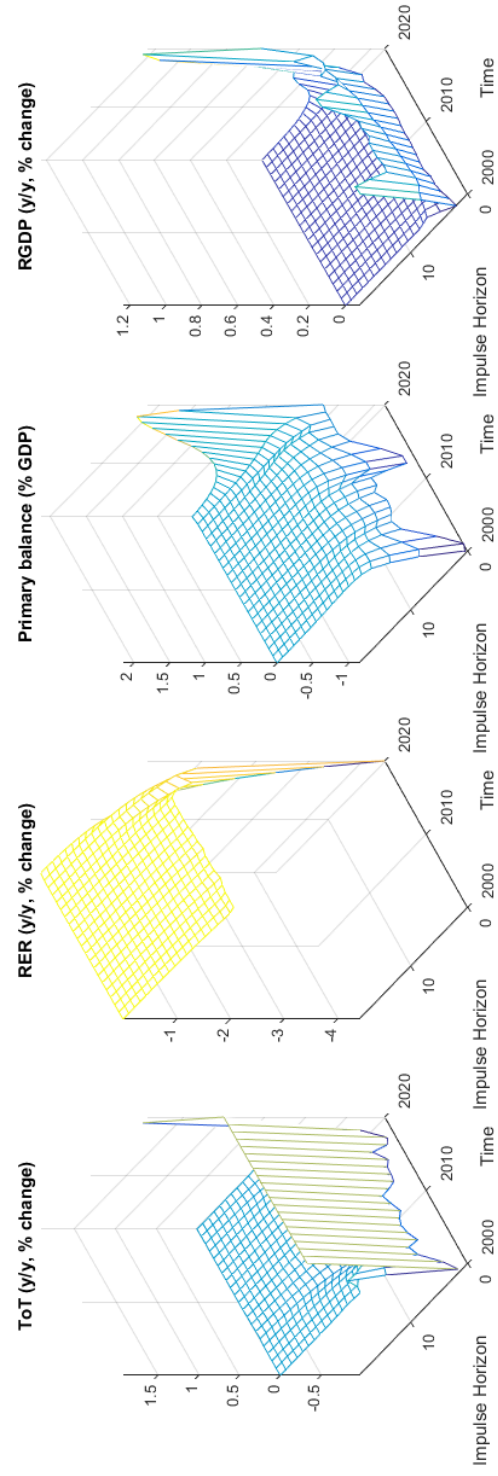

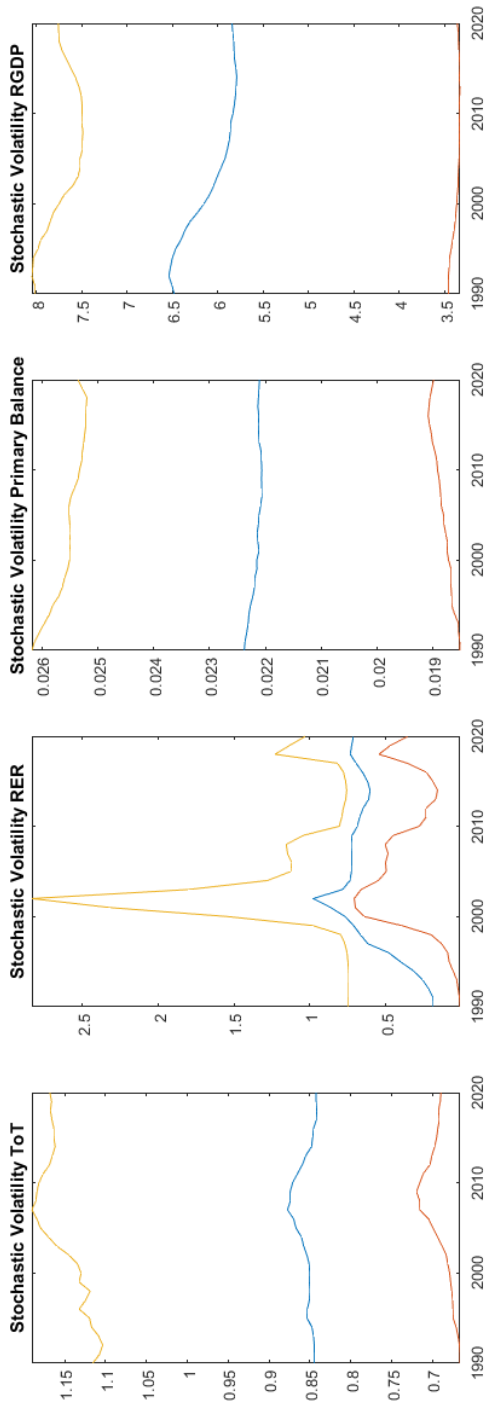Figure 10: Impulse Response Functions, Ecuador

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Figure 11: Posterior Estimates for Stochastic Volatility of Structural Shocks, Peru
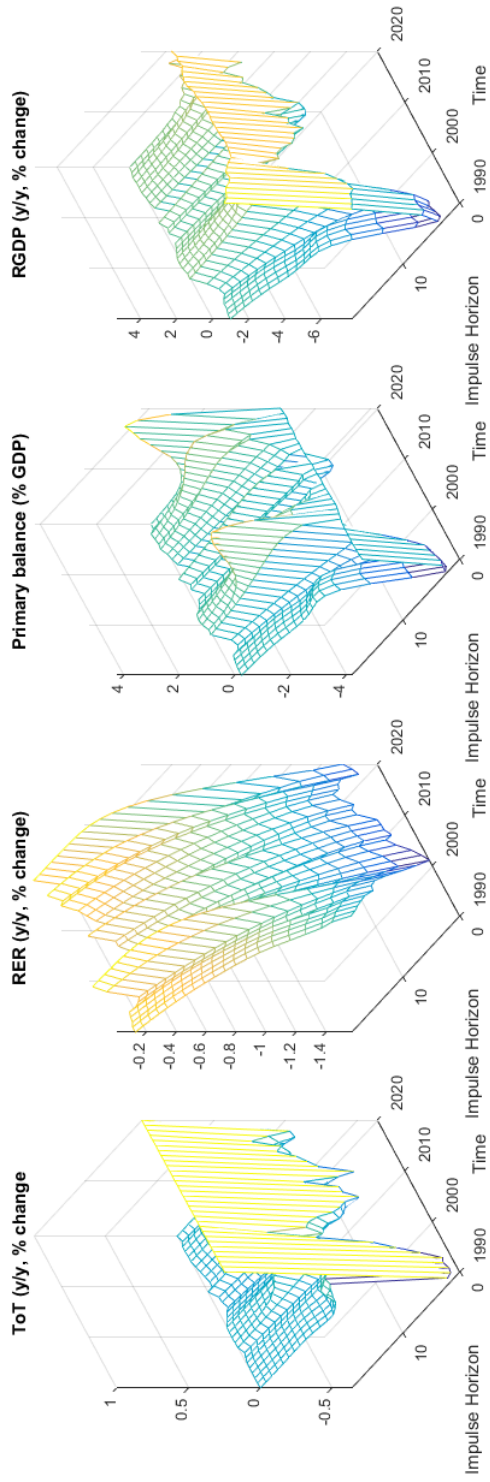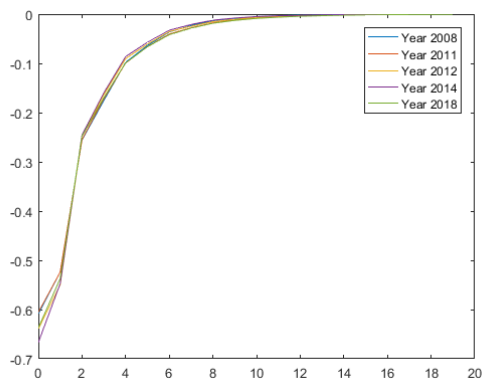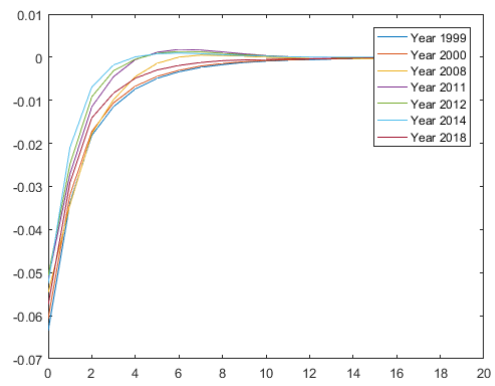

Figure 12: Impulse Response Functions, Peru

Figure 13: Response of the Primary Balance to a Terms-of-Trade Shock
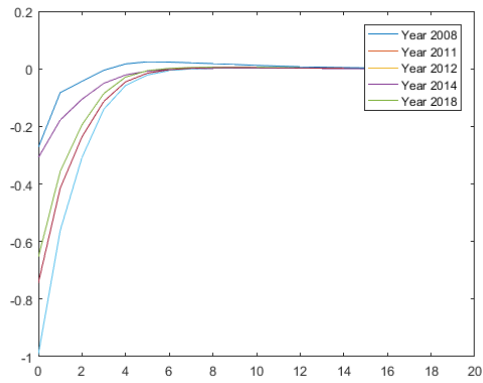
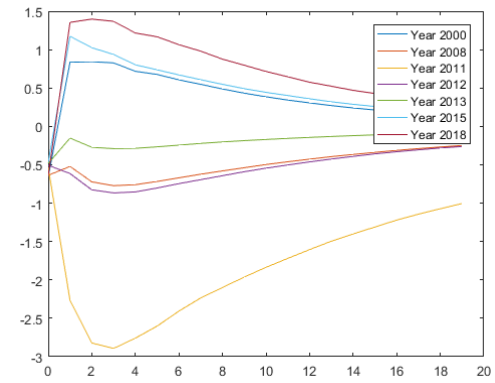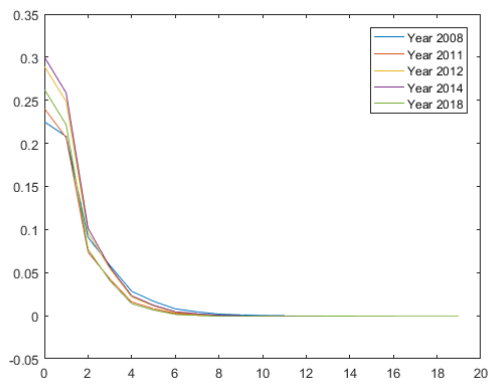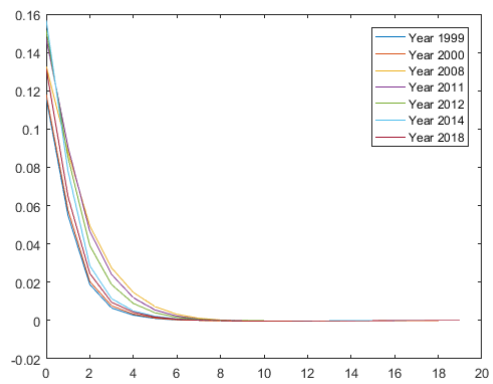(a) Bolivia

(b) Colombia

(c) Ecuador

(d) Peru

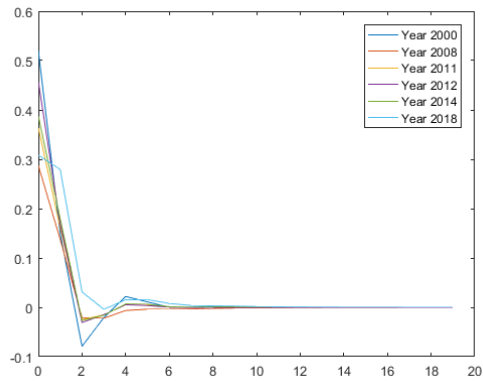Figure 14: Response of the GDP to a Terms-of-Trade Shock
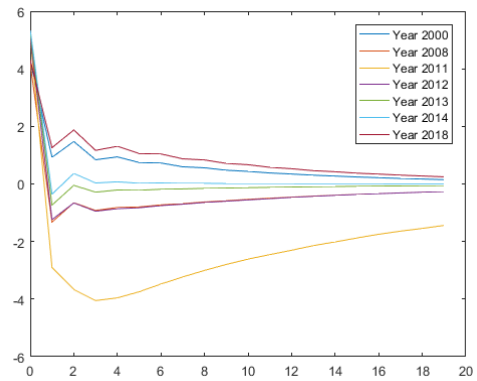
(a) Bolivia

(b) Colombia

(c) Ecuador

(d) Peru

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A   Sampling algorithms for estimation of model parameters

A.1   Drawing coefficient states (B_t's and $a_{ij,t}$'s)

A.1.1   A standard state-space representation

Equations (3) and (4) have, respectively, the state-space form of equations (14), the "observation" equation, and (15), the "transition" equation, below:

$$y_t = H_t \beta_t + \varepsilon_t \tag{14}$$

$$\beta_t = F\beta_{t-1} + u_t, \tag{15}$$

where

$$\begin{bmatrix} \varepsilon_t \\ u_t \end{bmatrix} \sim i.i.d.\ N\left(\begin{bmatrix} 0 \\ 0 \end{bmatrix}, \begin{bmatrix} R_t & 0 \\ 0 & Q \end{bmatrix} \right).$$

In the observation equation the unknown parameters consist of the elements of H not fixed or given as data and the nonzero elements of the covariance matrix R. In the transition equation, the parameters to be estimated are the nonzero and free elements of F, Q, and the state variable, β_t.

A.1.2   Bayesian econometrics: The general procedure

A Bayesian approach to estimating this model is based on three basic steps.

Step 1. The researcher forms a prior belief of the distribution of the parameters to be estimated $P(\beta, F, H, Q, R) \sim d(.)$, where $d(.)$ is a generic *joint* density function of the parameters.

Step 2. The researcher collects data on the observed variables (y) and constructs the likelihood function of the model parameters contained in the data $F(y_t \mid \beta, F, H, Q, R)$.

Step 3. The researcher *updates* the prior belief on the model parameters (step 1) based on the information contained in the data (step 2). Put another away, the researcher combines the prior distribution $P(\beta, F, H, Q, R)$ and the likelihood function $F(y_t \mid \beta, F, H, Q, R)$ to obtain the *joint posterior* distribution $H(\beta, F, H, Q, R \mid y_t)$.

Defined by Bayes' theorem, $H(.)$ is a product of the likelihood $F(.)$ and the joint prior $P(.)$. More formally,

$$H(\beta, F, H, Q, R \mid y_t) \propto (y_t \mid \beta, F, H, Q, R) \times P(\beta, F, H, Q, R)$$

where $\propto$ indicates proportionality.

To proceed further in terms of inference, the researcher has to "isolate" the component of the posterior relevant to the parameter of interest. For example, to conduct inference about β, the researcher has to derive the *marginal* posterior distribution for β from the *joint* posterior distribution of the parameters. Similarly, inference on Q is based on the *marginal* posterior distribution for Q. However, derivation of the *marginal* posterior distribution (from a *joint* posterior distribution) requires analytical integration, a procedure that may prove difficult in complex models such as the one used here. It was the development of simulation methods such as Gibbs sampling that greatly simplified this integration.

A.1.3 Gibbs sampling: A general description

Gibbs sampling is a numerical method that uses draws from *conditional* distributions to approximate *joint* and *marginal* distributions. From Blake and Mumtaz (2017), a general description of Gibbs sampling is the following:

Suppose we have a joint distribution of k variables

$$f(x_1, x_2...x_k).$$

This may, for example, be a joint posterior distribution.

We are interested in obtaining the marginal distributions

$$f(x_i), \ i = 1...k.$$

Assume that the form of the conditional distributions $f(x_i \mid x_j), \ i \neq j$ is known. A Gibbs sampling algorithm with the following steps can be used to approximate the marginal distributions:

(1) Set starting values for $x_1...x_k$
$$x_1^0, ...x_k^0$$
where the superscript 0 denotes the starting values.

(2) Sample x_1^1 from the distribution of x_1 *conditional* on *current* values of $x_2...x_k$
$$f(x_1^1 \mid x_2^0, ...x_k^0)$$

(3) Sample x_2^1 from the distribution of x_2 *conditional* on *current* values of $x_1, x_3, ...x_k$
$$f(x_2^1 \mid x_1^1, x_3^0, ...x_k^0)$$

(4) Sample x_k^1 from the distribution of x_k *conditional* on *current* values of $x_1, x_2, ...X_{k-1}$
$$f(x_k^1 \mid x_1^1, x_2^1, ...x_{k-1}^1)$$

to complete 1 iteration of the Gibbs sampling algorithm.

As the number of Gibbs iterations increases to infinity, the samples or draws from the conditional distributions converge to the joint and marginal distributions of x_i at an exponential rate (for a proof of convergence see **?**). More concretely, by repeating the Gibbs sampling algorithm M times (i.e., a large enough number to achieve convergence) and saving the last H draws of x_i, we obtain H values for $x_1...x_k$. A histogram for $x_1...x_k$ is an approximation for the marginal density of $x_1...x_k$. One crucial thing to note is that the implementation of the Gibbs sampling algorithm requires the researcher to know or assume the form of the conditional distributions $f(x_i \mid x_j)$.

A.1.4   Gibbs sampling in state-space models

Going back to the state-space model of equations (14) and (15) above, a Gibbs sampling algorithm for this problem can be constructed by assuming that the state variable, β_t is *known and observed*. This observation implies the following general Gibbs algorithm:

Step 1. Conditional on β_t, sample H and R from their posterior distributions.

Step 2. Conditional on β_t, sample F and Q from their posterior distributions.

Step 3. Conditional on the parameters of the state-space (F, H, Q, R), sample β_t from its *conditional* posterior distribution.

Step 4. Repeat steps 1 to 3 until convergence is achieved.

As shown by Carter and Kohn (1994), the conditional posterior distribution $H(\beta^T \mid F, H, Q, R, y^T)$, where $\beta^T = [\beta_1, \beta_2, ..., \beta_T]$ is the time series vector of β (the same applies to y^T), can be factored into the following conditional distributions.

$$H\left(\beta^T \mid y^T, F, H, Q, R\right) = H\left(\beta_T \mid y^T, F, H, Q, R\right) \times H\left(\beta_{T-1} \mid \beta_T, y^T, F, H, Q, R\right). \quad (16)$$

Note that the right-hand side splits into the product of (1) the *marginal* distribution of the state-space variable at time T and (2) the distribution of the vector β_{T-1} conditioned on β_T (i.e., its conditional posterior). As shown in **?**, considering the fact that β_T follows a first-order autoregressive or Markov process, equation (16) can rewritten as

$$H\left(\beta^T \mid y^T, F, H, Q, R\right) = H\left(\beta_T \mid y^T, F, H, Q, R\right) \prod_{t=1}^{T-1} H\left(\beta_t \mid \beta_{t+1}, y^t, F, H, Q, R\right),$$

where, given that ε_t and u_t are normally distributed:

$$H(\beta_t \mid \beta_{t+1}, y^t, F, H, Q, R) \sim N(\beta_{t|t+1}, P_{t|t+1}).$$

Then, to draw β's from these distribution we need to calculate its mean, $\beta_{t|t+1}$, and variance, $P_{t|t+1}$, using the Kalman filter:

$$\beta_{t|t+1} = E(\beta_t \mid \beta_{t+1}, y^t, F, H, R, Q),$$
$$P_{t|t+1} = \text{Var}(\beta_t \mid \beta_{t+1}, y^t, F, H, R, Q).$$

Starting from an initial value, that is, a prior of state $\beta_{0|0}$ and variance of the state $P_{0|0}$, a standard Kalman filter consists of the following equations, which are evaluated recursively through time:

$$\beta_{t|t-1} = \mu + F\beta_{t-1|t-1}$$
$$P_{t|t-1} = FP_{t-1|t-1}F' + Q$$
$$\eta_{t|t-1} = y_t - H\beta_{t|t-1}$$
$$f_{t|t-1} = HP_{t|t-1}H' + R$$
$$\beta_{t|t} = \beta_{t|t-1} + K\eta_{t|t-1}$$
$$P_{t|t} = P_{t|t-1} - KHP_{t|t-1},$$

where $K = P_{t|t-1}H'f_{t|t-1}^{-1}$ is the "Kalman gain" (i.e., the weight attached to the prediction error). Running these equations from $t = 1, 2, ...T$ delivers $\beta_{T|T}$ and $P_{T|T}$ at the end of the recursion, which are known as the "updating" equations as they update the initial estimates $\beta_{t|t-1}$ and $P_{t|t-1}$ using information contained in the prediction error, $\eta_{t|t-1}$.

Finally, the draw of β_T and the output of the filter are now used for the first step of the backward recursion (which provides $\beta_{T-1|T}$ and $P_{T-1|T}$) used to make a draw of β_{T-1}. This process can be thought as updating $\beta_{t|t}$ (the Kalman filter estimate of the state variable) for information contained in β_{t+1}. This backward recursion continues until time zero. As shown in Blake and Mumtaz (2017, p. 74), for a given time t, the updating formulas of the backward recursion are

$$\beta_{t|t,\beta_{t+1}} = \beta_{t|t} + P_{t|t}F'(FP_{t|t}F' + Q)^{-1}(\beta_{t+1} - \mu - F\beta_{t|t})$$
$$P_{t|t,\beta_{t+1}} = P_{t|t} - P_{t|t}F'(FP_{t|t}F' + Q)^{-1}FP_{t|t}$$

The above formulas constitute the Carter and Kohn algorithm.

A.1.5   Drawing β_t's

Equations (14) and (15) can be rewritten as the following state-space forms:

$$y_t = X_t' \beta_t + A_t^{-1} \Sigma_t \varepsilon_t$$
$$X_t' = I_n \otimes [1, y_{t-1}', ..., y_{t-k}']$$
(17)

$$\beta_t = \beta_{t-1} + \eta_t.$$
(18)

The observation equation (17) is linear and has Gaussian innovations with known variance. Following the procedure described in Section A.1.4, the conditional posterior distribution $H\left(\beta^T \mid y^T, A^T, \Sigma^T, V\right)$ can be factored as

$$H(\beta^T \mid y^T, A^T, \Sigma^T, V) = H(\beta_T \mid y^T, A^T, \Sigma^T, V) \prod_{t=1}^{T-1} H(\beta_t \mid \beta_{t+1}, y^t, A^T, \Sigma^T, V).$$

where

$$H(\beta_t \mid \beta_{t+1}, y^T, A^T, \Sigma^T, V) \sim N(\beta_{t|t+1}, P_{t|t+1}).$$

Then, to draw β's from this distribution we need to calculate its state mean, $\beta_{t|t+1}$, and state variance, $P_{t|t+1}$:

$$\beta_{t|t+1} = E(\beta_t \mid \beta_{t+1}, y^T, A^T, \Sigma^T, V),$$
$$P_{t|t+1} = \text{Var}(\beta_t \mid \beta_{t+1}, y^T, A^T, \Sigma^T, V).$$

The vector of β_t's can be easily drawn, because it can be computed using the forward (Kalman filter) and the backward (Carter and Kohn algorithm) recursions presented in Section A.1.4.

The starting values of the Kalman filter, that is, the prior of the state $\beta_{0|0}$ and variance of the state $P_{0|0}$ are the following:

Priors for β_t's. The first $T_{0=20}$ observations of the sample (the 20-year period 1962—1981) are used to estimate a standard fixed-coefficient VAR via ordinary least squares (OLS) such that $\beta_{0|o} = \beta_{OLS}$ with a coefficient covariance matrix given by $p_{0|0} = \Sigma_0 \otimes (X_{0t}' X_{0t})^{-1}$, where $X_{0t} = \{Y_{0t-1}, \ldots Y_{0t-p}, 1\}, \Sigma_0 = \frac{(Y_{0t} - X_{0t}\beta_0)'(Y_{0t} - X_{0t}\beta_0)}{T_0}$ and the subscript 0 denotes the fact that this is the training sample.

The last recursion of the Kalman filter provides $\beta_{T|T}$, and $P_{T|T}$ which are the mean variance of the posterior distribution of β_T. After drawing a value from this distribution, the draw is used in the backward recursion to obtain $\beta_{T-1|T}$ and $\beta_{T-1|T}$ from period $t-1$ to 0, that is, to obtain β_t conditional on β_{t+1}.

A.1.6 Drawing the $a_{ij,t}$'s

From equations (8) to (11), state-space formulations for the $a_{ij,t}$'s can be derived. Specifically, the state-space formulation for $a_{12,t}$ is

$$u_{2,t} = -a_{12,t}u_{1,t} + \varepsilon_{2,t}, \text{Var}(\varepsilon_{2,t}) = \sigma_{2,t} \tag{19}$$

$$a_{12,t} = a_{12,t-1} + v_{1t}, \text{Var}(v_{1t}) = S_1 \tag{20}$$

and the state-space formulation for $a_{13,t}$ and $a_{23,t}$ is

$$u_{3,t} = -a_{13,t}u_{1,t} - a_{23,t}u_{2,t} + \varepsilon_{3,t}, \text{Var}(\varepsilon_{3,t}) = \sigma_{3,t} \tag{21}$$

$$\begin{pmatrix} a_{13,t} \\ a_{23,t} \end{pmatrix} = \begin{pmatrix} a_{13,t-1} \\ a_{23,t-1} \end{pmatrix} + \begin{pmatrix} v_{2t} \\ v_{3t} \end{pmatrix}, \text{Var}\left(\begin{pmatrix} v_{2t} \\ v_{3t} \end{pmatrix}\right) = S_2. \tag{22}$$

Because S is block diagonal, the Kalman filter and Carter and Kohn algorithm can be applied equation by equation. The priors of the $a_{ij,t}$ are the following:

Priors for $a_{ij,t}$'s. Let $A = \Sigma_0^{1/2}$ and let A_0 denote the inverse of the matrix A with the diagonal normalized to 1. The initial values for $a_{ij,t}$ (i.e., the initial state $a_{ij,0|0}$) are the nonzero of A_0 with the variance of the initial state set equal to $abs(a_{ij}) \times 10$, following **?**.

Note that without the block diagonality of S it would not be feasible to apply the Kalman filter and Carter and Kohn algorithm equation by equation and draw separately every block of $a_{ij,t}$'s. However, while the block-diagonal assumption of S simplifies the analysis, it is not essential and Primiceri (2005) provides a procedure to deal with an unrestricted S.

A.2 Drawing volatilities

A.2.1 The Metropolis Hastings algorithm

In this section, following Blake and Mumtaz (2017), we briefly describe the Metropolis Hastings (MH) algorithm in a general setting. Suppose that we are interested in drawing samples from the following distribution:

$$\pi(\Phi),$$

where Φ is a $K \times 1$ vector that represents a set of parameters and is referred as the "target density." $\pi(\Phi)$ can be thought of a posterior distribution where the Gibbs sampler is not

operation, because the conditional distributions of different blocks of the Φ parameters are unknown. However, if provided with a value of $\Phi = \Phi^*$ it is possible to evaluate the density of Φ^*, that is, $\pi(\Phi^*)$. If this is the case, the MH algorithm can be used to sample from $\pi(\Phi)$ following the steps described below:

Step 1. Specify a candidate density $q\left(\Phi^{G+1} \mid \Phi^G\right)$ where G indexes the draw of the parameters Φ. One must be able to draw samples from this density.

Step 2. Draw a candidate value of the parameters Φ^{G+1} from the candidate density $q(\Phi^{G+1} \mid \Phi^G)$.

Step 3. Compute the probability of accepting Φ^{G+1} (denoted by α) using the expression

$$\alpha = \min \left(\frac{\pi\left(\Phi^{G+1}\right)/q\left(\Phi^{G+1} \mid \Phi^G\right)}{\pi\left(\Phi^G\right)/q\left(\Phi^G \mid \Phi^{G+1}\right)}, 1 \right).$$

The numerator of this expression is the target density evaluated at the new draw of the parameters $\pi\left(\Phi^{G+1}\right)$ divided by the candidate density evaluated at the new draw of the parameters $q\left(\Phi^{G+1} \mid \Phi^G\right)$. The denominator is the same expression evaluated at the previous draw of the parameters.

Step 4. If the acceptance probability α is large enough, retain the new draw Φ^{G+1}, otherwise retain the old draw Φ^G. We decide if α is large enough by drawing a number u from the standard uniform distribution. If $u < \alpha$, accept Φ^{G+1}; otherwise keep Φ^{G3}.

Step 5. Repeat steps 2 to 4 M times and base inference on the last L draws. Provided convergence exists, the empirical distribution using the last L draws is an approximation to target density.

A.2.2 A simple stochastic volatility model using the MH algorithm

Consider the simple stochastic volatility model

$$y_t = \varepsilon_t \sqrt{\exp\left(\ln h_t\right)}$$
$$\ln h_t = \ln h_{t-1} + v_t$$
$$v_t \sim N(0, g),$$

where h_t is time-varying variance. This is a state-space model where the observation equation is nonlinear in terms of the state variable h_t and therefore the Carter and Kohn algorithm does not apply. **?** suggest applying an independence MH algorithm at each t to sample from

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[3]This essentially means that we accept the draw with probability α if this experiment is repeated many times (e.g., if $\alpha = 0.1$ and we perform 1000 replications we should expect 100 of the 1000 draws to have $u < \alpha$).

the conditional distribution of h_t which is given by $f(h_t \mid h\_t, y_t)$ where the subscript $-t$ denotes all other dates than t.

Because the transition equation is a random walk, the knowledge of h_{t+1} and h_{t-1} contains all the relevant information about h_t. This way, the conditional distribution of h_t can be simplified as

$$f(h_t \mid h_{-t}, y_t) = f(h_t \mid h_{t-1}, h_{t+1}, y_t).$$

To sample from $f(h_t \mid h_{t-1}, h_{t+1}, y_t)$ a date-by-date application of the independence MH algorithm could be performed following steps 2 to 4 as described in Section A.2.1. However, this procedure does not apply for the first and last observations. Therefore, the general algorithm is modified as follows (Blake & Mumtaz, 2017, p. 158):

? show that this density has the following form:

$$f\left(h_t \mid h_{t-1}, h_{t+1}, y_t\right) = h_t^{-0.5} \exp\left(\frac{-y_t^2}{2h_t}\right) \times h_t^{-1} \exp\left(\frac{-\left(\ln h_t - \mu\right)^2}{2\sigma_h}\right),$$

that is, this density is the product of a normal density and a log normal density, with

$$\mu = \frac{\left(\ln h_{t+1} + \ln h_{t-1}\right)}{2}$$

$$\sigma_h = \frac{g}{2}.$$

Then, the Bayesian updating begins according to the following steps:

Step 1. Generate a starting value for $h_t, t = 0...T$ as $\hat{\varepsilon}_t^2$ and set the prior $\bar{\mu}, \bar{\sigma}$ (e.g, $\bar{\mu}$ could be the log of the OLS estimate of the variance of ε_t and $\bar{\sigma}$ could be set to a big number to reflect the uncertainty in this initial guess). Set an inverse-gamma prior for g (i.e., $p(g) \sim \text{IG}\left(\frac{g_0}{2}, \frac{v_0}{2}\right)$). Set a starting value for g.

Step 2. The drawing of h_t proceeds in three steps:

2a for $t=0$, sample the initial value of h_t denoted by h_0 from the log normal density

$$f\left(h_0 \mid h_1\right) = h_0^{-1} \exp\left(\frac{-\left(\ln h_0 - \mu_0\right)^2}{2\sigma_0}\right),$$

where the mean $\mu_0 = \sigma_0 \left(\frac{\bar{\mu}}{\bar{\sigma}} + \frac{\ln h_1}{g}\right)$ and $\sigma_0 = \frac{\bar{\sigma}g}{\bar{\sigma}+g}$.

2b. For each date $t = 1$ to $T - 1$ draw a new value for h_t from the candidate density (call the draw $h_{t,new}$)

$$q\left(\Phi^{G+1}\right) = h_t^{-1} \exp\left(\frac{-(\ln h_t - \mu)^2}{2\sigma_h}\right),$$

where $\mu = \frac{(\ln h_{t+1} + \ln h_{t-1})}{2}$ and $\sigma_h = \frac{g}{2}$. Compute the acceptance probability

$$\alpha = \min\left(\frac{h_{t,\text{ new}}^{-0.5} \exp\left(\frac{-y_t^2}{2h_{t,\text{ new}}}\right)}{h_{t,\text{ old}}^{-0.5} \exp\left(\frac{-y_t^2}{2h_{t,\text{ old}}}\right)}, 1\right).$$

Draw $u \sim U(0,1)$. If $u < \alpha$, set $h_t = h_{t,\text{ new}}$. Otherwise retain the old draw.

2c. For the last time period $t = T$ compute $\mu = \ln h_{t-1}$ and $\sigma_h = g$ and draw $h_{t,new}$ from the candidate density

$$q\left(\Phi^{G+1}\right) = h_t^{-1} \exp\left(\frac{-(\ln h_t - \mu)^2}{2\sigma_h}\right).$$

Compute the acceptance probability

$$\alpha = \min\left(\frac{h_{t,\text{ new}}^{-0.5} \exp\left(\frac{-y_t^2}{2h_{t,\text{ new}}}\right)}{h_{t,\text{ old}}^{-0.5} \exp\left(\frac{-y_t^2}{2h_{t,\text{ old}}}\right)}, 1\right).$$

Draw $u \sim U(0,1)$. If $u < \alpha$, set $h_t = h_{t,\text{ new}}$. Otherwise retain the old draw.

Step 3. Given a draw for h_t, compute the residuals of the transition equation $v_t = \ln h_t - \ln h_{t-1}$. Draw g from the inverse-gamma distribution with scale parameter $\frac{v_t' v_t + g_0}{2}$ and degrees of freedom $\frac{T+v_0}{2}$.

Step 4. Repeat steps 2 and 3 M times. The last L draws of h_t and g provide an approximation to the marginal posterior distributions.

A.2.3   Drawing $(\sigma_{i,t})$'s

Equations (3) and (6) can be rewritten into the following state-space forms:

$$\begin{aligned} A_t(y_t - X_t'\beta_t) &= \Sigma_t \varepsilon_t \\ (y_t - X_t'\beta_t) &= u_t \end{aligned} \tag{23}$$

$$\ln \sigma_{i,t} = \ln \sigma_{i,t-1} + \nu_{i,t}, \ \nu_{i,t} \sim N(0, g_i), \tag{24}$$

where $\ln \sigma_{i,t}$ is time-varying variance. This is a state-space model where the the observation equation is nonlinear in the state variable σ and therefore can be estimated by the MH

algorithm explained above after linearization. Note that because $\Sigma_t \varepsilon_t = A_t u_t$ is contemporaneously uncorrelated, the draws $\ln \sigma_{i,t}$ from the MH algorithm can be done separately for each $i = 1...3$.

Priors for $\sigma_{i,t}$. A starting value for $\sigma_{i,t}$ is obtained from $\hat{u}_{it}^2$. The prior for $\bar{\mu}_i$ can be set equal to the log of the ith diagonal element of Σ_0 and $\bar{\sigma}$ to a large number, 10 in this case, following Blake and Mumtaz (2017).

A.3 Drawing hyperparameters

Conditional on β^T, Σ^T, A^T, and y^T, it is easy to draw posteriors using Gibbs sampling for Q, the block of S, and W, because all of these posteriors are observable. We discuss the priors we used to estimate these conditional distributions below.

A.3.1 Priors

The first $T_0 = 20$ observations (the 20-year period between 1962 and 1981) are used to calibrate the prior distributions.

Prior for Q. The (conjugate) prior for Q is inverse-Wishart[4] $p(Q) \sim IW(Q_0, T_0)$, where Q_0 is the scale matrix and T_0 are the degrees of freedom. This prior is key, because it shapes the amount of time variation allowed for in the VAR model, that is, a large value for the scale matrix Q_0 would imply more fluctuation in β_t.

The scale matrix Q_0 is set equal to $p_{0|0} \times T_0 \times \tau$, where τ is a scaling factor chosen by the researcher. Some studies set $\tau = 3.510^{-4}$, i.e., a small number, to reflect the fact that the training sample is typically short and the resulting estimates of $p_{0|0}$ maybe imprecise. Note that one can control the a priori amount of time variation in the model by varying τ.

Priors for S_1 and S_2. The prior for S_1 is inverse-gamma[5] $p(S_1) \sim IG(S_{10}, T_0)$ and the prior for S_2 is inverse Wishart $p(S_2) \sim IW(S_{20}, T_0)$. We follow **?** who set $S_{10} = 0.001$ and
$$S_{20} = \begin{pmatrix} 0.001 & 0 \\ 0 & 0.001 \end{pmatrix}.$$

Priors for W. An inverse-gamma prior for $W = g_i$, $p(nu_i) \sim IG(g_0, \nu_0)$ is set with scale parameter $g_0 = 0.01^2$ and degrees of freedom $\nu_0 = 1$.

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[4]The appendix in **?** provides a detailed description of the inverse-Wishart distribution.
[5]See **?** for a detailed description of the inverse-gamma and gamma distributions and **?** for a detailed description of algorithms for drawing from these distributions.

A.3.2 Posteriors

Posterior for Q. The hyperparameter Q has an inverse-Wishart conditional posterior distribution $p(Q) \sim IW(\eta\eta' + Q_0, T + T_0)$.

Posteriors for S_1 and S_2. The hyperparameter S_1 has an inverse-gamma conditional posterior distribution $p(S_1) \sim \text{IG}(\frac{v_1 v_1' + S_{10}}{2}, \frac{T + T0}{2})$.

The hyperparameter S_2 has an inverse-Wishart conditional posterior $p(S_1) \sim \text{IW}(v_2 v_2' + S_{20}, T + T0)$.

Posterior for W. The hyperparameters in W have an inverse-gamma conditional posterior distribution $p(\nu_i) \sim IG(\frac{\nu_i \nu_i' + g_0}{2}, \frac{T + \nu_0}{2})$.

A.4 Summary of the sampling chain

Step 1. Set priors and initial values of B^T, $a_{ij,t}$, $\sigma_{i,t}$, Q, the block S, and W:

    1a. Set prior for Q: $p(Q) \sim IW(Q_0, T_0)$ and starting values of the Kalman filter: $\beta_{0|0}$ and $P_{0|0}$.

    1b. Set prior for S_1: $p(S_1) \sim \text{IG}(S_{10}, T_0)$, a prior for S_2: $p(S_2) \sim IW(S_{20}, T_0)$, and starting values of $a_{ij,t}$.

    1c. Set a starting value for for $\sigma_{i,t}$ and a prior for W: ν_i, $p(\nu_i) \sim IG(g_0, \nu_0)$.

Step 2. Conditional on A_t, W_t, and Q, draw β_t using the Carter and Kohn algorithm.

Step 3. Conditional on the draw of β_t, sample Q from $p(Q) \sim IW(\eta\eta' + Q_0, T + T_0)$.

Step 4. Conditional on β_t, W_t, and S_1 and S_2, draw $a_{ij,t}$ using the Carter and Kohn algorithm.

Step 5. Conditional on the draw of $a_{ij,t}$ estimate residuals v_1, v_2, v_2, and sample S_1 from $p(S_1) \sim \text{IG}(\frac{v_1 v_1' + S_{10}}{2}, \frac{T + T0}{2})$ and S_2 from $p(S_2) \sim IW(\frac{\nu \nu_1' + g_0}{2}, \frac{T + \nu_0}{2})$.

Step 6. Using the draw of A_t from Step 4, calculate $\varepsilon_t = A_t u_t$ and draw $\sigma_{i,t}$ using the MH algorithm.

Step 7. Conditional on a draw for $\sigma_{i,t}$, draw W from $p(\nu_i) \sim IG(\frac{\nu_i \nu_i' + g_0}{2}, \frac{T + \nu_0}{2})$.

Step 8. Repeat steps 2 to 7 M times. The last L draws give an approximation to the marginal posterior distribution of the model parameters.

B   Algorithm for sign restrictions

Once past the burn-in stage –the process of dropping $M - L$ Gibbs draws– and marginal posterior distributions of all the model's parameters are obtained, following **?**, the steps presented below are used to calculate the required A_0 matrix with sign restrictions:

Step 1. Draw a $N \times N$ matrix K from the standard normal distribution.

Step 2. Calculate the matrix Q from the QR decomposition of K. Note that Q is orthonormal i.e., $Q'Q = I$.

Step 3. Calculate the Cholesky decomposition of the current draw of $\Sigma = \tilde{A}_0' \tilde{A}_0$.

Step 4. Calculate the candidate A_0 matrix as $A_0 = Q\tilde{A}_0$. Because $Q'Q = I$, this implies that $A_0' A_0$ will still equal Σ, that is, by calculating the product $Q\tilde{A}_0$ we alter the elements of $\tilde{A}_0$ but not the property that $\Sigma = \tilde{A}_0' \tilde{A}_0$. The candidate A_0 matrix has the following form:

$$
A_0 = \begin{pmatrix} 1 & 0 & 0 \\ a_{12} & 1 & 0 \\ a_{13} & a_{23} & 1 \end{pmatrix}.
$$

The second and third rows of this matrix corresponds to the oil price shock impact on the fiscal variable and the nonoil real GDP, respectively. We need to check if $a_{12} < 0$, $a_{13} > 0$, and $a_{23} < 0$. If this is the case, a contemporaneous shock in $oilprice_t$ will lead to a fall in pb_t and an increase in $nonoilrgdp_t$, because the elements $\begin{pmatrix} a_{13} & a_{23} & 1 \end{pmatrix}$ correspond to the current period response of $nonoilrgdp_t$ to an $\varepsilon_{oilprice,t}$ shock. If $a_{12} < 0$, $a_{13} > 0$, and $a_{23} < 0$, the algorithm stops and uses this A_0 matrix to compute a draw of impulse responses. If the restriction is not satisfied, the algorithm goes back to step 1 and tries with a new K matrix.

Step 5. Repeat steps 1 to 4 for every retained Gibbs draw.