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Evaluating The Self-Defeating Fiscal Austerity Hypothesis for a
Dollarized Economy: The Peruvian Case

Tesis para obtener el grado académico de Maestro en Economía que
presenta:

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RESUMEN

Este artículo aborda la hipótesis de la futilidad de las políticas de austeridad fiscal planteada por Attinasi y Metelli (2017) y Cherif y Hasanov (2018), utilizando un enfoque de Vectores Autorregresivos con Parámetros Variables en el Tiempo y Volatilidad Estocástica (TVP-VAR-SV) con datos de Perú entre el segundo trimestre de 2000 y el segundo trimestre de 2024. La idea central es determinar si, tras un choque transitorio de austeridad fiscal, el nivel de deuda como porcentaje del PBI efectivamente se reduce en el largo plazo. En caso de que el nivel de deuda no se reduzca (o incremente) en el largo plazo, diríamos que la política fiscal es fútil durante el periodo analizado. Esta futilidad puede originarse debido al encogimiento de la actividad económica que sigue a una contracción fiscal. Al elegir la economía peruana como objeto de estudio, se puede analizar también el papel que desempeñan las dinámicas del tipo de cambio en un intento de consolidación fiscal, lo cual no ha sido abordado previamente en la literatura. Estimamos una familia de modelos con parámetros variables en el tiempo y volatilidad estocástica para determinar si la introducción de estos elementos es prescindible para un buen ajuste del modelo. En cuanto a la hipótesis de la futilidad de las políticas de austeridad fiscal, encontramos evidencia en contra de esta en los modelos seleccionados, ya que la respuesta acumulada a largo plazo de la razón deuda/PIB se estabiliza 0.6 puntos porcentuales del PIB por debajo de la trayectoria sin choque. Los resultados sugieren la relevancia del canal del tipo de cambio para la efectividad de los choques de austeridad. Sin embargo, sí encontramos que la estabilización fiscal puede resultar inútil cuando se realiza mediante recortes de gasto en lugar de aumentos en los ingresos. Esta ineficacia proviene del perjuicio al crecimiento del PIB causado por los choques de austeridad en el gasto y su corta persistencia, lo que reduce su efecto a largo plazo.



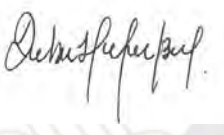
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Evaluating The Self-Defeating Fiscal Austerity Hypothesis for a Dollarized Economy: The Peruvian Case

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Abstract

This article incurr in the self-defeating fiscal austerity hypothesis set forth by Attinasi and Metelli (2017) and Cherif and Hasanov (2018) using a Time Varying Parameters VAR with Stochastic Volatility (TVP-VAR-SV) approach with data for Peru between 2000Q2 and 2024Q2. This hypothesis is best tested using the framework provided by Favero and Giavazzi (2007, 2009), where debt levels are determined in its deterministic dynamics equation. By choosing the peruvian economy as our study subject, we are able to analyse the role played by exchange rate dynamics in a fiscal consolidation attempt, which has not been previously addressed in the literature. We estimate a family of Time Varying Parameters models with Stochastic Volatility to discriminate if the introduction of time-varying parameters or stochastic volatility is expendable for a good fit. In regard to the self-defeating fiscal austerity hypothesis, we find evidence against it in our selected models, as the long run cumulative response of the debt-to-GDP stabilizes 0.6 points of GDP below the no shock path. The results hint at the relevance of the exchange rate channel for the effectivity of austerity shocks. However, we do find that fiscal stabilization might be futile when conducted through spending cuts instead of raising revenues. This futility steems from the detriment of GDP growth caused by spending austerity shocks and the short persistence of them, thus reducing its long-run effect.

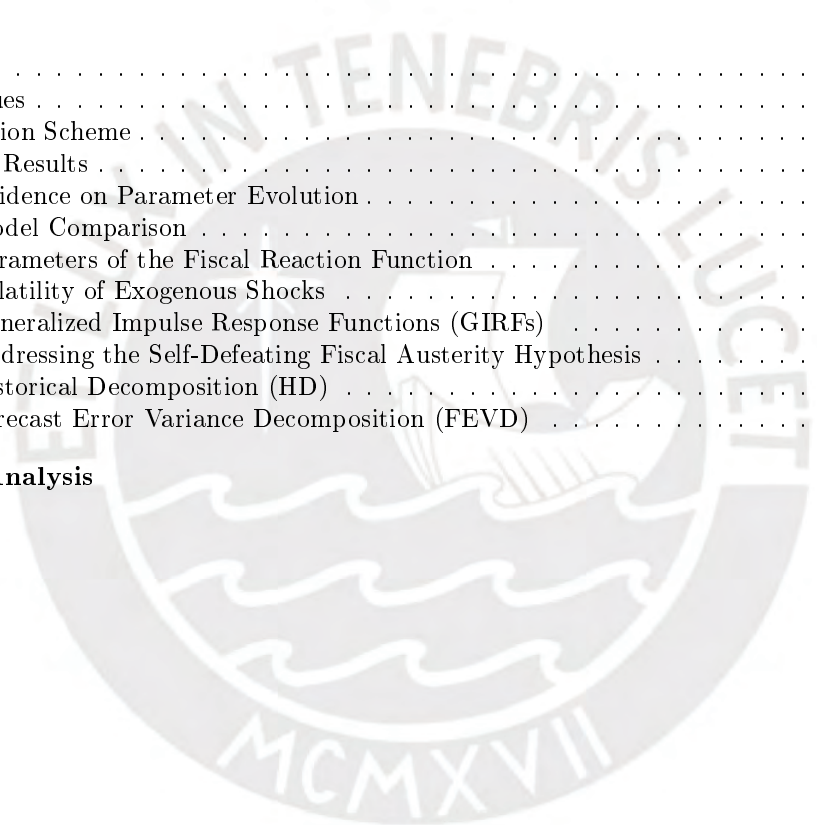
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1 Introduction

Fiscal austerity policies are designed to reduce government debt even at the expense of job losses and output reduction. The implementation and the further effects of austerity policies on macroeconomic variables has been widely studied for advanced economies; see Bianchi et al. (2019), Honohan (2016) or Alesina et al. (2015). However, due to the negative impact of fiscal contractions on economic activity, some authors have examined the possibility that debt as a share of GDP might be unaffected in the long run or, in the worst scenario, increased as a consequence of GDP growth losses; see Cherif and Hasanov (2018), Attinasi and Metelli (2017) or Castro et al. (2015). This analysis has not been yet conducted for an emerging market economy with a partially dollarized such as Peru.

In the core of these analyses lies the fiscal reaction function (henceforth FRF) from which austerity shocks are identified. The FRF has been at the core of the debt sustainability analysis since Bohn (1998). The underlying rationale of FRF estimation is quantifying the government's eagerness, and to a lesser extent ability, to quickly stabilize its financial liabilities either by raising revenues or reducing non-interest expenditure. Therefore, most of the FRF literature has focused in quantifying the response of the primary surplus to debt levels variations and business cycle fluctuations.

However, in a VAR framework the introduction of a FRF accounts for fiscal policy's interactions with other sectors of the economy such as the monetary authority, the business cycle and the external sector. This approach has proved being useful for studying fiscal consolidation processes and the determinants of debt reduction as it allows to simulate the response of debt levels to inflation, growth, monetary and fiscal austerity shocks and to quantify their contribution to debt dynamics; see Hasko (2007), Chung and Leeper (2007), Favero and Giavazzi (2007), Stanova (2015) and Cherif and Hasanov (2018).

In order to inquire on the possibility of futile fiscal austerity policies we estimate and evaluate a family of competing VAR models with time varying parameters and stochastic volatility (TVP-VAR-SV) to find which one fits best the data without incurring in overparametrization; see Chan and Eisenstat (2018). We allow for stochastic volatility (SV) in the process as this approach has proved to be a relevant introduction for the study of macroeconomic dynamics in Peru; see Jiménez et al. (2023), Urbina and Rodríguez (2023), Rodríguez et al. (2023), Rodríguez et al. (2024a) and Rodríguez et al. (2024b). The estimated models, which differ on the subsets of parameters for which time-variability is allowed, are compared using the Marginal Log-Likelihood (ML) and the Deviance Information Criteria (DIC). This part of the analysis allows us to discriminate if the introduction of TVP and/or SV is expendable for a good fit.

This article seeks to contribute to the FRF literature and the austerity shocks discussion in the following dimensions. First, we inquire on the characteristics of the FRF for Peru. The comparison of TVP-VAR-SV models allows us to test for time varying behavior of the peruvian fiscal authority. Quantifying the lagged response of the primary balance to the growth rate and the debt levels is of particular interest as it determines the eagerness of the fiscal authority to stabilize the business cycle and debt levels. Second, we use generalized impulse response functions (henceforth, GIRF) to test for the self-defeating fiscal austerity hypothesis. We exploit the time-varying GIRFs to study the dynamic impact of austerity shocks on debt, growth, the exchange rate and borrowing costs. These responses show the side-effects of austerity policies driven by the fiscal authority. Finally, we estimate the forecast error variance decomposition (FEVD) and the historical decomposition (HD) of the primary balance to inquire on which shocks contributed the most to the fiscal consolidation process in Peru and to asses whether austerity shocks are relevant during the sample period.

In summary, we find evidence both for a stabilizing fiscal behavior and a counter-cyclical reaction of the policy maker in Perú. On the other side, our results provide evidence against the self defeating fiscal austerity hypothesis in Perú, particularly when these stabilization policies are conducted through an increase in revenues. The non-futility of austerity policies are explained because the loss of GDP growth is not big enough to neutralize the other transmission channels of fiscal austerity policies to the reduction of debt as a share of GDP. However, we find that fiscal austerity policies are more prone to be futile when conducted at certain quarter which coincide with the export prices boom and fast GDP growth period in Peru.

This article is divided as follows. Section 2 summarizes the FRF literature in relation to fiscal consolidation processes. Section 3 presents the family of proposed TVP-VAR-SV models, the DIC and log-ML comparison criteria and the computation of GIRFs. Section 4 derives economic insights from the baseline model using the GIRFs, the HDs and FEVDs. Finally, Section 5 disserts about the implications of the results for debt stabilization processes and concludes.

2 Literature Review

The fiscal stabilization literature draws from (i) the FRF literature and (ii) the works on the interaction of fiscal shocks on other macroeconomic variables. The FRF literature tries to quantify the responses of the primary balance to gross debt as a share of GDP and to an economic activity indicator such as the output gap or the growth rate. Bohn (1998) derives the FRF from the tax smoothing model outlined by Barro (1979). Using yearly data for the span between 1916 and 1995, he finds that the US government reacts to a 1% increase in the debt over GDP ratio by raising the primary surplus in 0.054%. Bohn (1998) suggests that a positive fiscal stabilization parameter is a sufficient condition to ensure fiscal sustainability, assessment which has been questioned ever since; see Bohn (2007).

This approach has been constantly revisited for a variety of countries and expanded either by refining the econometric model, allowing non-linearities in the fiscal response to debt levels or the business cycle or allowing for time varying responses. In general terms, most applications find debt-stabilizing reactions and counter-cyclical fiscal policies; see Adedeji and Williams (2007), Ogbeigun and Shoband (2020) for emerging markets and Afonso (2008), Afonso and Jalles (2011), Reicher (2012) for advanced economies. As an exemption, De Mello (2005) finds no counter-cyclical fiscal activism in Brazil and Galvis (2015) finds no evidence of fiscal stabilization policies in Colombia. The issue of non-linear responses to debt levels has been tackled mainly to test for fiscal fatigue. Ghosh et al. (2013) finds evidence of fiscal fatigue for advanced economies, but most applications have rejected this hypothesis; see Checherita-Westphal and Žďárek (2017), Echevarria Icaza (2018) and Di Iorio and Fachin (2022).

In summary, FRF literature points at the relevance of treating heterogeneity across countries (Adedeji and Williams, 2007; Debrund and Kinda, 2016; Ogbeifund and Shobande, 2020; Di Iorio and Fachin, 2022; Stoian et al., 2022) and along time (Afonso, 2008; Burger and Marinkov, 2012; Campos and Cysne, 2019; Berger et al., 2021; Afonso and Coelho, 2022). Both types of heterogeneities can be attributed to changing macroeconomic fundamentals along the business cycle (Aldama and Creel, 2022; Afonso and Coelho, 2022), switching priorities in the policy maker's preferences (Berger et al., 2021; Dubbert, 2022) or perceived financial risk (Afonso and Coelho, 2022). We address this heterogeneities by introducing TVP and SV in our estimation to determine if these are relevant for a proper modelling of the behavior of the fiscal authority.

In regard to the interaction between the fiscal policy and other macroeconomic variables, we deem relevant the links with inflation, output growth, the exchange rate and borrowing costs. The necessity of modelling monetary policy and inflation dynamics in a fiscal framework is justified as interest rates determines the borrowing cost in local currency and inflation helps to erode debt as a share of nominal GDP; see Hasko (2007), Hall and Sargent (2011) and Aizenmann and Marion (2011) on the relevance of inflation in eroding public debt in OECD countries and the US respectively, Saugweme et al. (2023) on the effect of debt levels on inflation and Sims (2011) on the theoretical nexus between fiscal policy and inflation.

Moreover, fiscal position and debt levels might have an impact on prices according to the fiscal theory of the price level (Antonini et al., 2013; Afonso, Alves and Tovar Jalles, 2021) and even an impact on exchange rate dynamics through the expectations mechanism theory which asserts that expansionary fiscal policies under non-Ricardian regimes nudge fiscal authorities into debt monetizing, thus underpinning currency depreciation expectations (Alberola, Cantú, Cavallino and Mirkov, 2022). Therefore, coordination between fiscal and monetary policies or the lack thereof determine how costly demand and supply shocks are mitigated (Ankargren and Shahnazarian, 2019).

The external sector is taken into account because of the importance of commodity prices for economic activity and the peruvian tributary system (Mendoza Bellido and Anastacio Clemente, 2022) and for fiscal fluctuations in emerging economies (Loong Lee, 2020; Urbina and Rodriguez, 2023). On the other hand, the introduction of debt levels in the VAR also allows to study the debt feedback effect on growth dynamics and other macroeconomic fundamentals as pointed by Stanova (2015), Ramzan and Ahmad (2014) and Favero and Giavazzi (2007).

VAR frameworks are suitable for such analysis because they allow us to (i) control for the interaction of debt accumulation with other relevant macroeconomic variables to avoid overestimating the role of fiscal policy and (ii) generate IRFs for our set of variables after an exogenous primary surplus shock. These approaches to the FRF can be divided depending on how debt dynamics are introduced in the model.

The first approach (henceforth linear approach), introduced by Chung and Leeper (2007), model debt dynamics as just another equation inside the VAR model; see Khalid (2007), Nguyen (2013), Tashevska and Tevranski (2017), Lau and Lee Syn-Yee (2018), Yusuf and Segun (2019), Loong Lee (2020), Ramadan (2020), Kadir and Tsen (2021), Afonso, Alves and Tovar Jalles (2021) and Dubbert (2023). Under this approach, both the debt feedback effect on growth, inflation and borrowing costs and the response of debt to these

variables are assumed to be linear.

On general terms, this literature has corroborated the relevance of growth and inflation shocks on debt reduction processes; Kadir and Tsen (2021) find that inflation and growth shocks are relevant for explaining debt level reduction in Malaysia and Afonso et al. (2021) find that shocks to the output gap persistently reduce debt levels in 28 European Union economies.

In regard to the self-defeating fiscal austerity hypothesis, evidence using this approach is ambiguous. Tashevska and Tevronski (2017) find that debt decreases significantly after an austerity shock in Macedonia but Ramadan (2020) find that debt increases up to 6 quarters after an austerity shock in Egypt, thus finding evidence for self-defeating fiscal austerity policies hypothesis.

Evidence of fiscal stabilization policies in emerging countries is also mixed. Nguyen (2013) India's fiscal policy reacts not only to debt and output, but also to interest rate levels. However, Tashevska and Tevronski (2017) do not find a significant response of the fiscal surplus to debt levels in Macedonia and Lau and Lee Syn-Yee (2018) find evidence of fiscal stabilization policies in Malaysia but not in Thailand and the Philippines.

A second approach (henceforth, non-linear approach), introduced by Favero and Giavazzi (2007), models public debt dynamics as an equation outside the VAR. To account for public debt dynamics this modelling relies on the deterministic debt accumulation equation, which is a non-linear function of growth, interest rate, inflation and primary surpluses. In this framework, debt as a share of GDP is introduced as an exogenous variable in the VAR to account for the debt-feedback effect on the macroeconomic system; see Favero and Giavazzi (2009), Afonso and Sousa (2011a, 2011b, 2012), Stanova (2015), Attinasi and Metelli (2017), Cherif and Hasanov (2018) and Moreno (2019).

Favero and Giavazzi (2007) assert that the introduction of the debt accumulation equation, which is derived from the intertemporal budget constraint of the government, prevent the computation of IRF along unstable debt paths as the government might adjust the primary balance to keep the debt-to-GDP ratio stable. Using this framework for US data, Favero and Giavazzi (2009) find that the non-linearity that arises from the debt accumulation equation is more relevant as the feedback effect of debt on other variables becomes stronger. By using data for a sample of EU member countries, Stanova (2015) further develops this argument by pointing that initial debt levels matter when using this approach because the debt feedback effect on growth might dampen the effect of expansionary fiscal policies on output growth. Afonso and Sousa (2011a, 2011b, 2012) point that the introduction of debt dynamics is relevant as the models that explicitly include the deterministic debt growth equation generate more persistent responses of GDP to fiscal shocks.

Moreno (2019) compares both approaches for the peruvian economy and does not find a significant difference in the IRFs of debt levels to any of the structural shocks considered; namely real economic growth, exchange rate, growth inflation, export prices and interest rates shocks. However, the author does not compute confidence intervals for the IRFs nor delves deep into the discussion of the impact of austerity shocks on debt dynamics, thus limiting the insights inferable from his model.

In regard to the self-defeating fiscal austerity hypothesis, which is the main interest of this document, Attinasi and Metelli (2017) find that austerity shocks effectively reduce the debt-to-GDP ratio in the short run. However, debt levels return to the no-shock after some periods, implying the long run futility of austere fiscal policy. Nevertheless, they find that the effect of a revenues driven austerity shock on debt is more persistent than that of an austerity shock driven by spending cuts. Cherif and Hasanov (2018) test this hypothesis for the US and also find that debt levels return to the no-shock path after some periods, although these authors do not decompose the analysis between spending and revenues shocks.

The causes that make fiscal austerity policies futile have been explored in the literature. Calcagno (2012) asserted that fiscal adjustments with regressive distribution effects might reduce the disposable income of population groups with higher consumption propensities, which would imply a weakening of the aggregate demand. Bianchi et al. (2019) suggest that taxation driven fiscal austerity might introduce distortions that deter investments in technological progress, thus leading to slower GDP growth and rising debt levels. The concept of fiscal multipliers play a vital role in determining the effectiveness of austere fiscal policies or lack thereof, as these mediate between the austerity shock and the nominal growth channel of debt levels reduction. Alesina (2015) point that fiscal adjustments in Europe have been very costly in terms of output loss, although measures driven by taxation have stronger negative effects in these countries. For the peruvian case, Jiménez et al. (2023) find that spending multipliers are consistently bigger than taxation multipliers between 1994 and 2018, results which align with those of Rossini et al. (2012) and Vyturina and Leal (2016). Therefore, we would expect the nominal growth channel to be milder for taxation driven austerity shocks.

Although it is not explicitly noted in the aforementioned literature, the debate about the fiscal austerity

hypothesis is intermingled in its core with the discussion about the time series properties of debt levels and whether or not this variable has a unit root. Antonini et al. (2013) point that a unit root might be present in debt levels when the increase of government revenues does not match the spending shock to which it responds. Marcat and Scott (2009) attribute this persistent behavior to the incompleteness of stock markets, under which a state uses debt as a buffer stock to smooth fluctuations along the business cycle and repays it slowly, thus inducing high persistence in debt levels. As it is highly unlikely that debt levels could grow without limits, we should deny the presence of a unit root. However, we pair with Antonini et al. (2013) in the assertion that the debt-to-GDP ratio is likely to display a unit root when it does not exceed extreme values, which is the case of Peruvian debt levels in the chosen sample.

We incur in the fiscal austerity hypothesis literature and expand it by also taking into account the role played by the depreciation rate channel which is not addressed by the aforementioned authors as foreign exchange risk is not relevant for countries that issue most of its debt in local currency. In this sense, the Peruvian economy is an ideal case as its public debt is heavily dollarized. Moreover, we exploit the advantages of the TVP-VAR-SV models to address the treatment of time variability of fiscal behavior pointed out in the uniequational literature. However, we expand the TVP-VAR-SV framework by introducing the deterministic debt dynamics equation to test the self-defeating fiscal austerity hypothesis while we properly measure the collateral effects of austerity shocks on GDP growth, inflation and the depreciation rate.

3 Methodology

3.1 The Fiscal Reaction Function

A FRF is written as follows:

$$pb_t = \beta_{\Delta d}^{FRF} \Delta d_{t-1} + \beta_g^{FRF} g_{t-1} + \mathbf{Z}'_t \boldsymbol{\theta} + \epsilon_t^{FRF}$$

where pb_t represents the fiscal primary balance-to-GDP ratio, g_t refers to an economic activity indicator such as the output gap or the real growth rate of the GDP, Δd_t refers to a gross debt variable, \mathbf{Z}_t is a vector with other variables considered and ϵ_t^{FRF} is the austerity shock. Under our notation, $\beta_{\Delta d}^{FRF}$ is the fiscal stabilization parameter. A positive value for this parameter indicates that the fiscal authority reacts to an increase in debt as a share of GDP by generating surpluses in the primary balance. Parameter β_g is referred to as the fiscal activism parameter, for which positive value indicates counter-cyclical fiscal policies. This equation is estimated as a part of the SVAR model. Therefore the variables in \mathbf{Z}_t are the other variables included in the VAR. An outline of the estimated VAR models is conducted in 3.2.

In regard to other determinants in \mathbf{Z}_t , the literature has considered (i) variables related to borrowing costs such as the interest rate or the interest burden (Debrun and Kinda, 2016) and (ii) external variables such as terms of trade, global economic activity and commodities prices (Loong Lee, 2020; Moreno, 2020; Campos and Cysne, 2019). For our specification, we consider the implicit local currency interest rate (i_t^{lc}) and the implicit foreign currency rate (i_t^{fc}) to control for borrowing costs. The inflation of GDP's deflator (π_t) is included because of its role in determining the real costs of borrowing and its eroding effect on debt as a share of nominal GDP. Moreover, we add the depreciation rate (e_t) as public debt in Peru is partially dollarized during the chosen sample. Finally, the VAR includes the export price index growth (epi_t) to account for the dynamics of the external sector.

3.2 Econometric Model

We estimate a family of TVP-VAR-SV models. In this Section, we discuss the general model, that allows for time variability in intercepts, lagged variables coefficients, contemporaneous impact coefficients and volatilities. All the estimated models are particular cases of this one, varying on the restrictions imposed on them.

Using the notation of Chan and Eisenstat (2018), we denote the $n \times 1$ vector of observations as \mathbf{y}_t . The model in structural form can be written as:

$$\mathbf{B}_0 \mathbf{y}_t = \boldsymbol{\mu}_t + \mathbf{B}_1 \mathbf{y}_{t-1} + \dots + \mathbf{B}_p \mathbf{y}_{t-p} + \boldsymbol{\epsilon}_t, \quad \boldsymbol{\epsilon}_t \sim \mathcal{N}(\mathbf{0}, \boldsymbol{\Sigma}_t) \quad (1)$$

$$\mathbf{h}_t = \mathbf{h}_{t-1} + \boldsymbol{\xi}_t, \quad \boldsymbol{\xi}_t \sim \mathcal{N}(\mathbf{0}, \boldsymbol{\Sigma}_h) \quad (2)$$

where \mathbf{B}_{0t} denotes the contemporaneous impact coefficients $n \times n$ matrix, \mathbf{B}_{jt} is the j -th lag VAR coefficient matrix and $\boldsymbol{\Sigma}_t = \mathbf{I}_n \mathbf{h}_t$ is a diagonal matrix with the log-volatilities (\mathbf{h}_t) which follow a random walk process where \mathbf{h}_0 is estimated.

A reduced form for (1) can be written in the following manner:

$$\mathbf{y}_t = \tilde{\mathbf{X}}_t \boldsymbol{\beta}_t + \mathbf{W}_t \boldsymbol{\gamma}_t + \boldsymbol{\epsilon}_t, \quad \boldsymbol{\epsilon}_t \sim \mathcal{N}(\mathbf{0}, \boldsymbol{\Sigma}_t) \quad (3)$$

where $\boldsymbol{\beta}_t = \text{vec}((\boldsymbol{\mu}_t, \mathbf{B}_{1t}, \dots, \mathbf{B}_{pt})')$, $\tilde{\mathbf{X}}_t = \mathbf{I}_n \otimes (1, \mathbf{y}'_{t-1}, \dots, \mathbf{y}'_{t-p})$ and $\boldsymbol{\gamma}_t$ is a $k_\gamma \times 1$ vector with the free elements of \mathbf{B}_{0t} stacked in a row. As a consequence, \mathbf{W}_t is a $n \times k_\gamma$ matrix with contemporaneous variables such that:

$$\mathbf{W}_t = \begin{bmatrix} 0 & 0 & 0 & \dots & 0 & 0 & \dots & 0 \\ -y_{1t} & 0 & 0 & \dots & 0 & 0 & \dots & 0 \\ 0 & -y_{1t} & -y_{2t} & \dots & 0 & 0 & \dots & 0 \\ \vdots & \vdots & \vdots & \ddots & \vdots & \vdots & \ddots & \vdots \\ 0 & 0 & 0 & \dots & -y_{1t} & -y_{2t} & \dots & -y_{(n-1)t} \end{bmatrix}$$

where y_{it} is the i -th variable in \mathbf{y}_t . Given this, we can write this reduced form VAR model as a space-state model:

$$\mathbf{y}_t = \mathbf{X}_t \boldsymbol{\theta}_t + \boldsymbol{\epsilon}_t \quad \boldsymbol{\epsilon}_t \sim \mathcal{N}(\mathbf{0}, \boldsymbol{\Sigma}_t) \quad (4)$$

$$\boldsymbol{\theta}_t = \boldsymbol{\theta}_{t-1} + \boldsymbol{\eta}_t, \quad \boldsymbol{\eta}_t \sim \mathcal{N}(\mathbf{0}, \boldsymbol{\Sigma}_\theta) \quad (5)$$

where matrix \mathbf{X}_t is the concatenation of $\tilde{\mathbf{X}}_t$ and \mathbf{W}_t and $\boldsymbol{\theta}_t = (\boldsymbol{\gamma}'_t, \boldsymbol{\beta}'_t)'$. The time-varying parameters $\boldsymbol{\theta}_t$ follow a random walk process.

3.3 Competing Models

We obtain a family of models by imposing different sets of restrictions to (1). Since (1) allows for time variability in intercepts, lagged variables coefficients, contemporaneous impact coefficients and volatilities, we refer to it as the unconstrained model from which the rest are particular cases. We consider 6 restricted versions of the TVP-VAR-SV model: (i) TVP-VAR which assumes constant variances ($\mathbf{h}_t = \mathbf{h}_0$), (ii) TVP-VAR-R1-SV which assumes that intercepts and lagged variables coefficients are constant ($\boldsymbol{\beta}_t = \boldsymbol{\beta}_0$), (iii) TVP-VAR-R2-SV which assumes that the contemporaneous impact coefficients are constant ($\boldsymbol{\gamma}_t = \boldsymbol{\gamma}_0$), (iv) TVP-VAR-R3-SV which allows only for time variability in intercepts and volatilities, (v) CVAR-SV allows only for stochastic volatilities and (vi) CVAR which is a standard bayesian VAR.

3.3.1 Prior Distributions

We assume that the initial values of the time-varying parameters and the logarithm of the standard deviations follow pairwise-independent Normal distributions: i.e. $\boldsymbol{\theta}_0 \sim \mathcal{N}(\mathbf{a}_\theta, \mathbf{V}_\theta)$, $\mathbf{h}_0 \sim \mathcal{N}(\mathbf{a}_h, \mathbf{V}_h)$. The state variables follow independent multivariate random walk processes with variance-covariance matrixes $\boldsymbol{\Sigma}_\theta = \text{diag}(\sigma_{\theta_1}^2, \dots, \sigma_{\theta_n}^2)$, $\boldsymbol{\Sigma}_h = \text{diag}(\sigma_{h_1}^2, \dots, \sigma_{h_n}^2)$. The diagonal elements in $\boldsymbol{\Sigma}_\theta$ and $\boldsymbol{\Sigma}_h$ follow independent Inverse-Gamma distributions: i.e. $\sigma_{\theta_i}^2 \sim \mathcal{IG}(v_{\theta_i}, S_{\theta_i})$, $\sigma_{h_j}^2 \sim \mathcal{IG}(v_{h_j}, S_{h_j})$ for $i = 1, \dots, k_\theta$, $j = 1, \dots, k_h$.

3.4 Estimation Algorithm

We use a Gibbs Sampling algorithm to estimate the aforementioned models with the precision sampling method developed by Chan and Jeliakzov (2009). The algorithm proposed by Chan and Einsestat (2018) proceeds as follows: (i) draw samples from $(\boldsymbol{\theta} | \mathbf{y}, \mathbf{h}, \boldsymbol{\Sigma}_\theta, \boldsymbol{\Sigma}_h, \boldsymbol{\theta}_0, \mathbf{h}_0) \sim \mathcal{N}(\hat{\boldsymbol{\theta}}, \mathbf{K}_\theta^{-1})$, where $\hat{\boldsymbol{\theta}} = \mathbf{K}_\theta^{-1} (\mathbf{X}' \boldsymbol{\Sigma}^{-1} \mathbf{y} + \mathbf{H}'_\theta \mathbf{S}_\theta^{-1} \mathbf{H}_\theta \boldsymbol{\alpha}_\theta)$, $\mathbf{K}_\theta = \mathbf{X}' \boldsymbol{\Sigma}^{-1} \mathbf{X} + \mathbf{H}'_\theta \mathbf{S}_\theta^{-1} \mathbf{H}_\theta$ with $\boldsymbol{\alpha}_\theta = \mathbf{H}_\theta^{-1} \tilde{\boldsymbol{\alpha}}_\theta$ and matrices \mathbf{H}_θ , \mathbf{X} , \mathbf{S}_θ , $\boldsymbol{\Sigma}$ and $\tilde{\boldsymbol{\alpha}}_\theta$ are defined in Appendix A of Chan and Eisenstat (2018a), (ii) we get draws of \mathbf{h} from $(\mathbf{h} | \mathbf{y}, \boldsymbol{\theta}, \boldsymbol{\Sigma}_\theta, \boldsymbol{\Sigma}_h, \boldsymbol{\theta}_0, \mathbf{h}_0)$ using the auxiliary mixture sampler of Kim et al. (1998), (iii) using the conditional distributions of the diagonal elements in $\boldsymbol{\Sigma}_\theta$, draw samples from $(\sigma_{\theta_i}^2 | \mathbf{y}, \boldsymbol{\theta}, \mathbf{h}, \boldsymbol{\theta}_0, \mathbf{h}_0) \sim \mathcal{IG}(v_{\theta_i} + \frac{T}{2}, S_{\theta_i} + \frac{1}{2} \sum_{i=1}^T (\theta_{it} - \theta_{i,t-1})^2)$, (iv) using the conditional distributions of the diagonal elements in $\boldsymbol{\Sigma}_h$, draw samples from $(\sigma_{h_j}^2 | \mathbf{y}, \mathbf{h}, \boldsymbol{\theta}, \boldsymbol{\theta}_0, \mathbf{h}_0) \sim \mathcal{IG}(v_{h_j} + \frac{T}{2}, S_{h_j} + \frac{1}{2} \sum_{i=1}^T (h_{it} - h_{i,t-1})^2)$, (v) draw samples from $(\boldsymbol{\theta}_0 | \mathbf{y}, \mathbf{h}, \boldsymbol{\Sigma}_\theta, \boldsymbol{\Sigma}_h) \sim \mathcal{N}(\hat{\boldsymbol{\theta}}_0, \mathbf{K}_{\theta_0}^{-1})$, where $\hat{\boldsymbol{\theta}}_0 = \mathbf{K}_{\theta_0}^{-1} (\boldsymbol{\Sigma}_\theta^{-1} \boldsymbol{\theta}_1 + \mathbf{V}_\theta^{-1} \boldsymbol{\alpha}_\theta)$ and $\mathbf{K}_{\theta_0} = \boldsymbol{\Sigma}_\theta^{-1} + \mathbf{V}_\theta^{-1}$, (vi) draw samples from $(\mathbf{h}_0 | \mathbf{y}, \mathbf{h}, \boldsymbol{\Sigma}_\theta, \boldsymbol{\Sigma}_h) \sim \mathcal{N}(\hat{\mathbf{h}}_0, \mathbf{K}_{h_0}^{-1})$, where $\hat{\mathbf{h}}_0 = \mathbf{K}_{h_0}^{-1} (\boldsymbol{\Sigma}_h^{-1} \mathbf{h}_1 + \mathbf{V}_h^{-1} \boldsymbol{\alpha}_h)$ and $\mathbf{K}_{h_0} = \boldsymbol{\Sigma}_h^{-1} + \mathbf{V}_h^{-1}$, and finally (vii) repeat steps (i)-(vi) N times. Prior values for \mathbf{a}_θ , \mathbf{V}_θ , \mathbf{a}_h , \mathbf{V}_h , v_{θ_i} , v_{h_j} , S_{θ_i} and S_{h_j} are specified in Section 4.2.

3.5 Bayesian Model Comparison Criteria

Following Chan and Einsestat (2018) we consider the Bayes factor based on the log-marginal likelihood and the deviance information criterion (henceforth DIC) for model comparison purposes.

3.5.1 Marginal Likelihood

We define $BF_{ij} = p(\mathbf{y}|M_i)/p(\mathbf{y}|M_j)$ as the Bayes factor, where $p(\mathbf{y}|M_k) = \int p(\mathbf{y}|\boldsymbol{\theta}_k, M_k)p(\boldsymbol{\theta}_k, M_k)d\boldsymbol{\theta}_k$ is the marginal likelihood for the model M_k . Hence, if $BF_{ij} > 1$, we say that model M_i is preferred BF_{ij} -times by the data over model M_j . Chan and Eisenstat (2015) propose a marginal likelihood estimation based on a cross-entropy method which uses an importance sampling from the following:

$$\hat{p}_{IS}(\mathbf{y}) = \frac{1}{N} \sum_{n=1}^N \frac{p(\mathbf{y}|\boldsymbol{\theta}_n)p(\boldsymbol{\theta}_n)}{g(\boldsymbol{\theta}_n)}. \quad (6)$$

If we use $g(\boldsymbol{\theta}_n) = g^* = p(\mathbf{y}|\boldsymbol{\theta}_n)p(\boldsymbol{\theta}_n)/p(\mathbf{y})$, we would obtain an estimator with variance equal to zero, such that:

$$\hat{p}_{IS}(\mathbf{y}) = \frac{1}{N} \sum_{n=1}^N \frac{p(\mathbf{y}|\boldsymbol{\theta}_n)p(\boldsymbol{\theta}_n)}{p(\mathbf{y}|\boldsymbol{\theta}_n)p(\boldsymbol{\theta}_n)p(\mathbf{y})} = p(\mathbf{y}). \quad (7)$$

Therefore, we choose g close enough to g^* to have an efficient estimator. For this purpose, we consider a parametric family $\{f(\boldsymbol{\theta}; v)\}$ indexed by the parameter vector v within which we find the importance density $f(\boldsymbol{\theta}; v^*)$ that is closest to g^* . This can be found by solving:

$$\hat{v}_{CE}^* = \arg \max \frac{1}{L} \sum_{l=1}^L \log f(\boldsymbol{\theta}_l; v). \quad (8)$$

where $\{\boldsymbol{\theta}_l\}_{l=1}^L$ are draws obtained from the posterior density. Once \hat{v}_{CE}^* is computed, we generate a random sampling $\{\boldsymbol{\theta}_l\}_{l=1}^N$ from the density $f(\boldsymbol{\theta}; \hat{v}_{CE}^*)$ and estimate the marginal likelihood by using the importance sampling $p(\mathbf{y})$. For further details on the marginal likelihood estimation see Section 4.2 of Chan and Eisenstat (2018a).

3.5.2 Deviance Information Criterion

The DIC proposed by Spiegelhalter et al. (2002) has the advantage of being relatively insensitive to prior distributions. This criterion is based on the deviance which, following Chan and Grant (2016), is defined as $D(\boldsymbol{\theta}) = -2\log p(\mathbf{y}|\boldsymbol{\theta}) + 2\log h(\mathbf{y})$, where $p(\mathbf{y}|\boldsymbol{\theta})$ is the likelihood and $h(\mathbf{y})$ is a function of the data.

Given that the DIC can be conceptualized as a trade-off between model fit and model complexity, we can calculate it as the sum of the posterior mean deviance and the effective number of parameters, so that $DIC = D^*(\boldsymbol{\theta}) + p_D$, where $D^* = -2E_{\boldsymbol{\theta}}(\log p(\mathbf{y}|\boldsymbol{\theta})|\mathbf{y}) + 2\log h(\mathbf{y})$ and $p_D = D^*(\boldsymbol{\theta}) - D(\hat{\boldsymbol{\theta}})$. In this way, we can write $DIC = 2D^*(\boldsymbol{\theta}) - D(\hat{\boldsymbol{\theta}})$ by assuming that $h(\mathbf{y}) = 1$. Consequently, $DIC = -4E_{\boldsymbol{\theta}}(\log p(\mathbf{y}|\boldsymbol{\theta})|\mathbf{y}) + 2\log p(\mathbf{y}|\hat{\boldsymbol{\theta}})$. So, given a set of competing models, we choose the one with the lowest DIC. For further detail on DIC estimation see Section 4.1 of Chan and Eisenstat (2018a).

3.6 Generalized Impulse Response Functions

For the purpose of precisely quantifying the impact of austerity shocks we set up a hybrid model where the macroeconomic fundamentals are determined in a VAR with debt feedback (9) and the first difference of the debt to GDP ratio is determined in the debt equation identity (10); see Cherif and Hasanov (2018), Stanova (2015), Afonso and Sousa (2012) and Favero and Giavazzi (2007)¹. We define $\mathbf{y}_t = [\mathbf{x}'_t|\Delta d_t]'$ to isolate debt dynamics from the other variables in the VAR in the following manner:

$$\hat{\mathbf{A}}_{0t}\mathbf{x}_t = \hat{\mathbf{C}}_t + \hat{\mathbf{A}}_{1t}\mathbf{x}_{t-1} + \hat{\boldsymbol{\Omega}}_t\Delta d_{t-1} + \mathbf{v}_t \quad (9)$$

$$\Delta d_t = \frac{(1-\delta)i_t^{lc} + \delta i_t^{fc} + \delta(1+i_t^{fc})e_t - \pi_t(1+g_t) - g_t}{(1+g_t)(1+\pi_t)}d_{t-1} - pb_t \quad (10)$$

¹Although this authors use the debt to GDP ratio without differencing.

where the coefficients matrixes $\hat{\mathbf{A}}_{0t}$, $\hat{\mathbf{A}}_{1t}$ and $\hat{\mathbf{\Omega}}_t$ are the corresponding submatrices of the estimated matrices $\hat{\mathbf{B}}_{0t}$, $\hat{\mathbf{B}}_{1t}$ and $\hat{\boldsymbol{\mu}}_t$ respectively. The correspondance between these matrices are given in Appendix 1.

This framework allows us to adress the nonlinearity of the data generating process of debt. However in such a nonlinear framework the computation of IRFs are not straightforward because they depend on the initial values of \mathbf{x}_t and d_t to which we will refer as \mathbf{x}_0 and d_0 . We recurr to the computation of non-linear impulse response functions (henceforth GIRF) outlined by Koop (1996). These are defined as the difference between the conditional expected values path of a given variable X_n under the materialization of a shock ($v = 1$) and the conditional expected values parth for the same variable without the shock ($v = 0$) to which we refer as the no-shock path:

$$GIRF(X_n|\mathbf{x}_0, d_t) = E[X|\mathbf{x}_0, d_0, v = 1] - E[X|\mathbf{x}_0, d_0, v = 0] \text{ for } n = 0, 1, 2... \quad (11)$$

The algorithm for computing the GIRF works as follows: (i) assume values for x_0 and for d_0 ; (ii) compute an expected no shock path conditional on d_0 and \mathbf{x}_0 ; (iii) simulate a path conditional on the materialization of a shock in any of the structural shocks of the model and (iv) compute confidence intervals by repeating this process for all simulations of the model.

Adyacent to the computation of GIRF, we decompose the GIRF of debt to an austerity shock as follows: (i) the interest rates channel, (ii) the depreciation rate channel, (iii) the nominal growth channel and (iv) the primary balance channel. This decomposition is done using a rearranged version of (10). The contribution of each channel to the IRF is the difference between its contribution in the shock (s) and the no-shock (ns) path.

$$\Delta d_t^s - \Delta d_t^{ns} = (\hat{i}_t^s - \hat{i}_t^{ns}) + (\hat{e}_t^s - \hat{e}_t^{ns}) + (\hat{y}_t^s - \hat{y}_t^{ns}) + (\hat{pb}_t^s - \hat{pb}_t^{ns}) \quad (12)$$

where \hat{i}_t stands for the interest rates channel, \hat{e}_t for the depreciation rate channel, \hat{y}_t for the nominal growth channel and \hat{pb}_t for the primary balance channel. Further details on the decomposition of equation (12) can be found in Appendix 2.

4 Results

4.1 Data

The variables for the model are chosen taking into account the main determinants of the fiscal position for resource rich countries such as export prices (Mendoza Bellido and Anastacio Clemente, 2022), debt levels and economic activity (Bohn, 1998). Short term debt drivers are also included as endogenous to the model to inquire on how austerity shocks might affect debt dynamics through diverse mechanisms. These drivers include inflation (Hall and Sargent, 2011; Aizenmann and Marion, 2011), monetary policy (Hasko, 2007), GDP growth (Cherif and Hasanov, 2017) and the exchange rate growth (Consejo Fiscal del Perú, 2023).

The proposed models are estimated using quarterly data for 2000Q2 to 2023Q2. This period covers the process of debt reduction from 48.5% to 18.0% of GDP between 2000 and 2014 and the subsequent rise up to 36.2% in 2021 due to the financial requirements for the fiscal stimulus injected during the COVID crisis (Consejo Fiscal del Perú, 2023) as shown in Figure 1. The sample also includes the commodities prices boom between 2004 and 2014 (Mendoza Bellido and Anastacio Clemente, 2022). This dynamic is depicted in Figure 1, from which a correlation of 0.290 between the EPI growth and the primary balance can be observed.

The baseline model is composed of seven variables: the EPI growth (epi_t), the real GDP growth rate (g_t), nonfinancial public sector's primary surplus as a share of quarterly GDP (pb_t), local currency depreciation rate (e_t), GDP's deflator inflation rate (π_t), the implicit interest rate of the debt issued in local currency (i_t^{lc}), the implicit interest rate of the debt issued in foreign currency (i_t^{fc}) and the first difference of the gross debt as a share of yearly GDP which will be referred as debt levels growth (Δd_t). We chose to model debt levels as a first difference as the debt-to-GDP ratio displays an unit root behavior when it does not excede threshold levels; see Antonini et al. (2013). These series are obtained from the peruvian's central bank webpage.

4.2 Prior Values

We use the following set of initial values for the mean of the initial value of the time varying coefficients as well for the mean of the initial log-standard deviation we set $\mathbf{a}_\theta = \mathbf{0}$ and $\mathbf{a}_h = \mathbf{0}$. The variance-covariance

matrices for these coefficients are given by $\mathbf{V}_\theta = 10 \times \mathbf{I}_{k_\theta}$ and $\mathbf{V}_h = 10 \times \mathbf{I}_{k_n}$. The degrees of freedom of the Inverse Gamma distributions assumed are relatively small: $\nu_{\theta_i} = \nu_{h_j} = 5, \forall i, j$ and $S_{\theta_i} = 0.01^2$ for the lagged variable coefficients and $S_{\theta_i} = 0.01^2$ for intercepts, contemporaneous impacts parameters and log-standard deviations.

4.3 Identification Scheme

We impose contemporaneous effect restrictions under a recursive scheme to identify structural shocks. We consider a triangular factorization that imposes a unit contemporaneous response of variables to their own structural shocks. We consider the criteria outlined in the literature to order the variables. The primary balance is allowed to respond contemporaneously to export price and growth shocks as these directly increase tax revenues (Mendoza Bellido and Anastacio Clemente, 2022). The exchange rate can react contemporaneously to the fiscal position as in mentioned by Alberola et al. (2022). Inflation can react contemporaneously to growth and exchange rate shocks as in Winkelried (2013). Finally, debt levels react contemporaneously to all the variables in the model as in IMF (2013).

4.4 Empirical Results

4.4.1 Evidence on Parameter Evolution

Table 1 presents a set of tests for time variability of the parameters of the TVP-VAR-SV model. Changes in the parameter and volatilities distribution are evaluated using the trace test proposed by Cogley and Sargent (2005), the Kolmogorov-Smirnov test (henceforth KS-test) and a t -test. The trace test assesses whether the prior trace of Σ_h differs statistically from its median posterior trace. The KS-test determines whether, at two different point in time, the underlying distribution of each parameter is the same. The t -test follows the same rationale as the KS-test and differs only by assuming that the underlying distribution is normal.

We consider the following quarters for comparison purposes: an observation after the introduction to fiscal rules (2000Q3), a post-financial crisis observation (2009Q1), a post-commodities prices boom pre-COVID observation (2016Q2) and a post-COVID observation (2024Q2) which is the last period of the sample. Table 1 displays the results of these three tests for $\mathbf{B}_{0,t}$, $\mathbf{B}_{i,t}$ and Σ_t in the TVP-VAR-SV model.

Notably, all volatilities vary over time. The trace test shows that the trace of the posterior median of Σ_h is statistically significant. Specifically, at 84% confidence, this trace is between 0.38 and 0.74. Both the KS-test and the t -test coincide that all volatilities vary between the selected sample periods.

On the other hand, the KS-test suggests that the distributions of most elements of $\mathbf{B}_{0,t}$ are time-varying, which suggests the adequacy of the TVP-VAR-SV or the TVP-VAR-R1-SV models. Finally, most elements of $\mathbf{B}_{i,t}$ change between both the beginning and end of the sample period. However, we find that almost a third of these parameters have not changed prior and post-COVID. The t -test delivers a similar insight, although more conservative.

In sum, we conclude that a TVP-VAR-SV approach is suitable for this sample. As most $\mathbf{B}_{0,t}$ and $\mathbf{B}_{i,t}$ parameters change over time, we argue that there could be an evolution in the transmission mechanism between the economic sectors explicitated in this model. Moreover, a change in volatilities seems likely as practically all elements of Σ_h changed between every period analyzed.

4.4.2 Model Comparison

For the estimation of each model we use 10 parallel markov chains performing 11,000 simulations from which the first 1000 are discarded as burn-in simulations. Then, we take the first of each 10 draws so there are 10,000 posterior draws for calculating the Log-ML and the DIC. We choose a 1 lag model according to the BIC obtained from estimating the CVAR model.

Table 2 shows both the mean and the standard deviation of the Log-ML and the DIC for all models detailed in Sections 3.1 and 3.2. According to the Log-ML, we should select the TVP-VAR-SV and CVAR-SV models. However, the TVP-VAR-SV performs worse than the CVAR-SV according to this criteria as the BF in favor of the CVAR-SV is 4.635 as shown in Table 3. We note that the models with constant variances performed the worst according to these criteria. The BF in favor of the worst performing model with SV (the TVP-VAR-R3-SV) against the TVP and the CVAR is 2.224×10^{44} and 6.195×10^{42} respectively. Accordingly, we argue in favor of the necessity of modelling time varying volatilities for this sample. In regard to the introduction of TVP, evidence is mixed. Allowing all parameters to be time varying worses the model independently of the introduction of SV as the CVAR-SV is preferred to the CVAR and the TVP-VAR-SV is

preferred to the TVP-VAR. However, allowing all parameters to vary is preferable to allowing only a subset of the parameters to vary as the TVP-VAR-SV is preferred over the TVP-VAR-R1-SV, the TVP-VAR-R2-SV and the TVP-VAR-R3-SV.

If we take the DIC into account, the CVAR-SV and TVP-VAR-R3-SV model should be selected. In this case the TVP-VAR-R3-SV performs notably better than the TVP-VAR-R1-SV and TVP-VAR-R2-SV, which suggests that allowing for changing μ_t is more efficient than allowing for varying γ_t or β_t . However, allowing for all parameters to be time-varying is not supported by the DIC as the relatively constrained versions of the TVP-VAR-SV perform better and have a higher p_D .

In brief, we consider the CVAR-SV, which allows for stochastic volatility as our baseline model. As a robustness check, we compare the results of the CVAR-SV against those of the relatively unconstrained TVP-VAR-SV and TVP-VAR-R3-SV models and the CVAR.

4.4.3 Parameters of the Fiscal Reaction Function

In regard to the FRF, from which we derive insight into the fiscal policymaker's behavior, we observe in Figure 2 that the value of the fiscal stabilization parameter (β_d^{FRF}) barely varies. Between 2000 and 2024 the parameter fluctuates around a mean of 0.160 and has been significantly different from zero during the sample period. According to the interpretation of Bohn (1998), this implies that the fiscal policy maker takes into account fiscal sustainability during this period because debt level increases are addressed with the generation of fiscal surpluses. Specifically, a 1% increase in debt as a share of GDP generates a 0.16% rise in the primary balance as a share of GDP. This response is higher than the estimations made by Moreno (2019) for a sample between 2000 and 2018, which calculate a fiscal stabilization response between 0.08% and 0.13%.

In respect to the fiscal activism parameter (β_g^{FRF}), we find that it fluctuates around 0.12 between 2000 and 2012. This implies that a 1% fall in the yearly real GDP growth is met with an expansionary fiscal policy that generates a 0.12% of GDP in additional deficit the next quarter; which is smaller than the reaction of 0.15% found by Moreno (2021). After 2012, this parameter followed a decreasing path until stabilizing around 0.07 in 2021. This would have implied a weaker fiscal activism if the contemporaneous impact coefficient of the GDP growth rate on the primary balance (γ_g^{FRF}) had not increased parallelly. Parameter γ_g^{FRF} first reaches statistically significant values in 2012Q1 and steadily increases until reaching a maximum value at 0.087 in 2020Q2, after which it stabilizes at 0.065. This points out that fiscal policy has become more countercyclical after 2010 and that the response to shocks stemming from the goods market are addressed immediately, which is a sign of an efficient, although not necessarily sustainable, counter-cyclical fiscal policy.

Adjacent to the fiscal activism parameter, it is noteworthy that the fiscal policy stance does not seem to take into account export prices directly. Although the γ_{epi}^{FRF} parameter has evolved over time, the 84% confidence interval has only excluded zero between 2006Q1 and 2009Q4, when epi_t^{FRF} reached its highest values. From this we infer that the fiscal policy maker only took into account the export prices during the quarters in which this grew at the fastest rate in the sample in order to generate fiscal surpluses. The loss of significance of this parameter coincides with the increase of γ_g , which implies that at this point the priority was switched from the generation of surpluses to the execution of counter-cyclical policies.

On the other hand, unlike Nguyen (2013) we do not find consistent evidence of other components in the fiscal reaction function as the interest rate, exchange rate and inflation lag-coefficients (β_{imn}^{FRF} , β_{ime}^{FRF} , β_e^{FRF} and β_π^{FRF} , respectively) have not been statistically different from zero during the sample period except for a period between 2005 and 2010. The overall lack of significance for parameters β_{imn}^{FRF} , β_{ime}^{FRF} and β_e^{FRF} imply that the fiscal policy maker does not take into account borrowing costs as neither a local monetary tightening nor a relatively higher cost of borrowing in foreign currency nudges into the generation of fiscal surpluses. Additionally, an statistically non-negative parameter β_π^{FRF} points out that the fiscal policy maker does not seem to take advantage of inflationary periods which would reduce the real cost of expansionary fiscal policies, of which debt levels reduction in 2021 and 2022 are an striking example.

To conclude the analysis of the FRF, it is noteworthy that its most notable changes along the sample period have addressed its counter-cyclicity; as it became stronger and more efficient during the post-boom years when it had to address the natural disasters of 2017 or the COVID-pandemic in 2020. In terms of its eagerness to stabilize debt levels, we can not point at any change, which hints at the irrelevance of the periodic changes of fiscal rules which took place during the sample period. Moreover, we could hypothesize that the periodical change of fiscal rules were designed to avoid significant changes in the fiscal behavior.

4.4.4 Volatility of Exogenous Shocks

Figure 3 displays the median of the standard deviations of the domestic variables structural shocks for our selected models. In general terms, the CVAR-SV and the TVP-VAR-R3-SV coincide in calculating the volatility, while the TVP-VAR-SV consistently undermeasures it. On the other hand, as the CVAR estimates an average volatility for the whole sample, it abstracts the dynamics of varying uncertainty along time. In the equations which model the behavior of a policy maker the volatility measures the extent of discretionarity in the policy making, i.e. the share of variation of a policy variable that is not attributable to a policy rule.

The volatility of austerity shocks, which are identified as shocks in the FRF, reduced slowly from a maximum of 1.080% of quarterly GDP (roughly 0.27 % of yearly GDP) in 2000Q2 until 1.001% of quarterly GDP (0.25% of yearly GDP) in 2004Q4, after it rises again up to 1.063 % of quarterly GDP (0.26% of yearly GDP). As parameter γ_g^{FRF} rises after 2010Q1, a reduction in the volatility of the austerity shocks is observed until it hits an all-time low at 0.893 of quarterly GDP (0.23% of yearly GDP) towards the end of the commodity price boom. However, from 2015Q1 onwards, it has steadily increased ever since even surpassing pre-fiscal rules levels after the COVID pandemic where it reached 1.276% (0.32% of yearly GDP). We suspect that the increase in fiscal policy's volatility is derived from the fiscal rules suspension in 2017 and 2021, as the 1% upper limit for the fiscal deficit and the 30% limit for debt levels were lifted to allow for stronger expansionary fiscal policies; see Mendoza Bellido and Anastacio Clemente (2022). These expansionary fiscal policies are driven to meet negative supply shocks stemming from the effects of the coastal El Niño. This aligns with the increase in the volatility of inflation and GDP growth, which began in 2017Q1. In conclusion, the fiscal authority has not adopted a predictable behavior as during most of the sample period the discretionary component of fiscal policy has accounted for more than 1.00% of quarterly GDP in each period.

On the other hand, the volatility of the debt shock, which is interpreted as the materialization of contingent debt, the effect of debt management operations or the use of government's resources to pay for its liabilities, has steadily decreased since 2000 but has risen strongly after 2020. The first periods of the sample show a decrease of debt levels volatility, which were at its highest point in 1.885% of yearly GDP at the beginning of the sample period. After remaining relatively constant around 1.300% of GDP between 2004Q1 and 2008Q2, this volatility reduced sharply along the export prices boom and hit its lowest point at 0.482% of yearly GDP in 2013Q1. However, it has steadily increased ever since, first during the slowing of the economic cycle between 2013Q1 and 2016Q2, later during the negative supply shocks era between 2017Q2 and 2019Q4 and finally during the COVID-19 crisis, where it hit a peak at 1.86% of GDP. In recent years, this volatility is stabilizing around 1.068 % of yearly GDP, barely below its pre-COVID-19 levels. The volatility of the shocks of the rest of debt-drivers, such as the GDP growth the inflation rate and the implicit interest rates, follow a similar path; the only exception being the depreciation rate, which hit an all-time high at 3.495% in 2008Q4 but has steadily reduced ever since, only having increased slightly after 2020Q1.

4.4.5 Generalized Impulse Response Functions (GIRFs)

Figure 4 displays the median GIRFs of pb_t to all exogenous shocks, thus giving us insights on the behavior of the fiscal policy. According to our selected models, the fiscal position barely reacts to export prices shocks in the TVP-VAR-SV model. However, it does react positively to them in the TVP-VAR-R3-SV and the CVAR-SV models. This suggests that the correlation observed in Figure 1 can be partially attributed to an eagerness to save the revenues generated during the export price booms but also to other transmission mechanisms. These mechanisms are most likely through the GDP growth. As export prices dynamize the local economic activity and accelerates GDP growth, the fiscal authority reacts counter-cyclically by generating surpluses. In this sense, the FRF has a strong fiscal activism component which is supported by all selected models. On average, the primary surplus rises in 0.139% during 8 quarters after a 1% unexpected increase of GDP growth. In regard to the fiscal stabilization component, we find a statistically significant stabilizing response. According to our selected models, after an elevation of debt levels in 1% of GDP the fiscal authority rises the surplus in about 0.120% in the following quarter. This suggests a stabilizing, although brief, response of the fiscal authority to debt accumulation which alligns with the fiscal stabilization parameter between 0.08 and 0.13 found by Moreno (2019).

The GIRFs also show how the FRF takes indirectly into account borrowing costs. According to our selected models, after a 1% increase in the implicit interest rate for debt in foreign currency the fiscal authority reacts by raising the balance in about 0.226% of GDP. The length of this response depend on the selected model as it fades immediately according to the TVP-VAR-R3-SV model but it lasts up to 5 quarters in the CVAR-SV model. This is likely a direct response of the fiscal policy to accumulation of debt in foreign currency.

The primary balance also reacts to exchange rate innovations. It decreases around 0.04% of GDP after a 1% depreciation shock, implying that periods of currency depreciation, which derives in higher inflation, contractionary monetary policy and lower growth rates, are followed by fiscal expansionary fiscal policies.

On the other hand, Figure 5 shows how debt levels growth react to the structural shocks in the selected models. As expected, debt levels growth rise after an increase of the implicit interest rates, although this increase fades in the third quarter. The effect of depreciation shocks on debt last longer, as debt rises up to four quarters afterwards. This can be attributed to the partial dollarization of the peruvian debt, which exposes the fiscal position to exchange rate hazard. On the other hand, GDP growth and inflation help to erode debt levels as noted by Aizenman and Marion (2011). Finally, both export prices and austerity shocks accelerate the debt reduction persistently. However, the persistence of export price shocks on debt levels is stronger as it also dynamizes economic activity and rises GDP growth, whereas austerity shocks reduce them.

Notably, in the linear IRFs displayed in Figure 5 we also observe how the linear modelling of debt dynamics produces a worse identification of shock dynamics on debt. First, according to the standard IRFs the impact of i_t^c shocks on Δd_t is negative, which is counter-intuitive. Second, the effect of epi_t shocks on debt are strongly overestimated. While epi_t shocks reduce debt through the nominal growth channel, the increase in government spending driven by this debt reduction dampens the overall effect of epi_t shocks on Δd_t . Finally, g_t shocks have a more persistent effect in the non-linear framework as these bolster the generation of primary surpluses through the counter-cyclical fiscal activism channel apart from eroding debt as a share of GDP.

In regard to the reaction of the model variables to austerity shocks, in Figure 6 we find a fiscal austerity puzzle in the first quarters after a contraction, after which the positive reaction of g_t in the CVAR-SV model is statistically significant. As this shock reduces the aggregated demand, π_t falls accordingly. The depreciation rate experiences a short term reduction after an austerity shock. This effect has been found both for advanced and emerging economies; see Kim and Roubini (2008), Nguyen (2013) and Alberola et al. (2022). According to Alberola et al. (2022), fiscal prudence increases trust in the local currency as it signals that the fiscal policy maker will not monetize its debt, thus reducing the risk premium and allowing for the appreciation of the local currency in the short term.

In Figure 7 and Figure 8 are displayed the GIRFs for an alternative model where pb_t is decomposed between spending (s_t) and revenues (r_t). Figure 7 shows the effect of spending-driven austerity shocks on the model's variables while Figure 8 shows the analogous for taxation driven austerity shocks. As spending austerity shocks reduce aggregate demand, g_t and π_t fall as expected. The puzzle observed in Figure 6 is also observed for the impact of taxation driven austerity shocks on GDP growth, as g_t rises up to 6 quarters after an r_t shock. This puzzle is partly explained by the increase in spending that follows an r_t shock. However, it is most likely that the reduction in π_t and e_t , which are directly linked to the price of goods, allows for for a more dynamic economic activity. The impact of austerity shocks on the implicit interest rates is not so clear. As austerity shocks reduce debt, which is the denominator in the construction of the i_t^{lc} and i_t^{fc} variables, these variables increase contemporaneously by an accounting effect. However, in the case of s_t shocks these reduce i_t^{lc} afterwards in the CVAR-SV model, which would imply that the government has emitted debt at lower rates. A similar dynamic is observed for the the impact of s_t on i_t^{fc} in the TVP-VAR-SV model only.

Figure 9 displays the GIRF of the model's variables to debt shocks. Unlike Sims (2008) and Malukeke and Odhiambo (2023), we do not find that a rise in debt levels growth accelerate inflation. However, we do find that GDP growth contracts after a Δd_t shock. This nexus implies that growing debt levels might dampen the possibility of using nominal GDP growth for fiscal stabilization. The GIRFs also suggest that Δd_t shocks have no significant effect on e_t , i_t^{lc} nor i_t^{fc} . In summary, there is evidence for a debt-feedback effect, implying the relevance for modelling such channel as proposed by Stanova (2015).

Figure 10 shows quarter-specific IRFs of pb_t to the rest of the shocks. We note a growing persistence of the effect of epi_t and g_t shocks on pb_t between 2000Q3 and 2016Q2 according to the TVP-VAR-SV model, which implies a better management of revenues and a more consistent counter-cyclical fiscal policy, respectively. However, this response has become weaker after the COVID-19 pandemic, as pb_t barely reacts to an epi_t shock in 2024Q2. We also observe a strengthening of the response of pb_t to g_t shocks between 2000Q2 and 2024Q2, which alligns with the increase of γ_g^{FRF} discussed in earlier sections. With respect to the response of pb_t to Δd_t shocks, it barely varies as the $\beta_{\Delta d}^{FRF}$ does.

Figure 11 displays the quarter-specific IRFs of debt levels. According to the TVP-VAR-SV, the response of Δd_t to g_t became more persistent between 2000Q3 and 2016Q2, implying a better management of the fiscal surpluses generated during the export price boom. However, in line with the reaction of pb_t to epi_t shocks, this reaction has become negligible in recent years. We also observe that austerity shocks effects on debt have increased its persistence between 2006Q3 and 2024Q2, which imply that debt consolidation processes

are more sustained in recent periods.

Apart from the aforementioned evolution of the effect of austerity shocks on debt levels, Figure 12 shows that the effect of pb_t shocks on e_t has sharply increased after the COVID-19 pandemic. This suggests a strengthening of the risk premium channel of fiscal policy set forth by Alberola et al. (2022), which imply a loss of trust of foreign investors to the proper management of Peruvian fiscal finances. Additionally, the impact of pb_t shocks on π_t has also reduced along the sample period, implying a weakening of the nominal growth channel of austerity shocks. In respect to the debt-feedback channel, Figure 13 shows no evidence of a consistent evolution of the impact of debt shocks on π_t , i_t^{lc} or i_t^{fc} . However, it shows that the debt-feedback channel on g_t increased its persistence from 2000Q3 and 2009Q1 and decreased it sharply towards 2016Q2. Afterwards, the persistence of this effect has stabilized around the 2006Q3 levels. The response of pb_t to Δd_t shocks changed slightly during the sample period. In 2000Q3 the peak response of pb_t was around 0.17, after which it stabilized around 0.14. Moreover, the persistence of this response has become stronger after 2009Q1.

4.4.6 Addressing the Self-Defeating Fiscal Austerity Hypothesis

The self-defeating fiscal austerity hypothesis set forth by Attinasi and Metelli (2017) and Cherif and Hasanov (2018) suggests that, after an austerity shock, the debt-to-GDP ratio returns to the no-shock path after some quarters. Panel A of Figure 14 shows the cumulative response of the first difference of the debt-to-GDP ratio to austerity shocks, which in sum is the response of the debt-to-GDP ratio. For the computation of these responses, initial values of 30% for the debt-to-GDP ratio (d_0) and the sample means for the rest of the variables in \mathbf{y}_t are assumed. We do not find evidence for this hypothesis as the debt-to-GDP ratio does not return to the no shock path in any of the selected models.

After an austerity shock of 1% of GDP in the primary balance, the debt-to-GDP ratio diverges from the no shock path until reaching its lowest point between the 6th and 15th quarter afterwards depending on the model. The magnitude of the lowest point varies considerably between models, as the TVP-VAR-SV identifies it at -0.568 after 6 quarters and the CVAR-SV model identifies it at roughly -1.000 after 9 quarters. Afterwards, the debt-to-GDP ratio leans towards the no-shock path but does not return to it, nor the 68% confidence interval does. Finally, the debt-to-GDP ratio stabilizes at -0.403 in the TVP-VAR-SV model and in -0.645 in the CVAR-SV model.

Panel B of Figure 14 shows the median decompositions of the cumulative GIRF of debt to an austerity shock in the selected models. As we observe, a 1% of GDP fiscal effort does not imply an equal reduction of the debt-to-GDP ratio due to the nominal growth channel in the CVAR-SV model. After 20 quarters, the primary balance channel, which is interpreted as the fiscal effort, accounts for a reduction of the debt-to-GDP ratio of 0.93%. The appreciation of the local currency which follows a fiscal consolidation process account for a subsequent reduction of 0.13%. We also observe a positive although negligible contribution of the interest rates channel of 0.04% of GDP. We attribute this increase to the reduction in inflation, which imply higher real implicit rates. Finally, the negative effect of the fiscal austerity on GDP growth through the aggregate demand dampens this effect in 0.16%. Thus after 20 quarters the 1% austerity shock implies a reduction in the debt-to-GDP ratio of at least 0.79%. However, after 40 quarters the negative contribution of the primary balance has reduced to 0.87% and the positive contribution of the nominal growth channel has intensified up to 0.17%; whereas the exchange rate and interest rates channel have remained constant. Therefore, after 40 periods the overall effect of an austerity shock on the debt-to-GDP ratio has reduced to -0.64% .

On the other hand, Panel A of Figure 15 shows the quarter-specific responses of the debt-to-GDP ratio to an austerity shock. We find that fiscal austerity is futile during the peak of the commodity prices boom (2006Q3 and 2009Q1), as the response of the debt-to-GDP ratio strongly leans to the no shock path after 20 quarters. However, this response diverges from the no shock path in the years that followed, as economic conditions in Peru worsened due to the coastal El Niño and the COVID pandemic. Interestingly, the long run effect of austerity shocks on the debt-to-GDP ratio found in the TVP-VAR-SV model for 2016Q2 and 2024Q4 align with that found by the CVAR-SV model. The TVP-VAR-SV model estimates a long run reduction of the debt-to-GDP ratio of 0.64% for 2016Q2 and 0.66% for 2024Q2.

From the Panel B of Figure 15 we can deduce which are the mechanisms through which an austerity shock is self-defeating in 2006Q3. We note that, apart from the weakening of the primary balance and the exchange rate channels to -0.25% and 0.02% of GDP respectively, the nominal growth channel is stronger at this quarter (0.26% of GDP). In this quarter, the effect of an austerity shock on g_t is negative (see Figure 10) and its effect on inflation is stronger. This negative impact on GDP growth drives the exchange rate channel upwards in spite of a reduction of the depreciation rate through the risk premium channel set forth

by Alberola et al. (2022). A similar pattern is observed for 2009Q1. In 2024Q2 the austerity shock was not self-defeating as the interest rates and nominal growth channels are milder and the exchange rate and primary balance channels are stronger. This suggests the relevance of controlling the interest rates and exchange rates channel for the effectiveness of austere fiscal policy. This can be achieved through the emission of bonds at fixed interest rates and an increase the share of debt emitted in foreign currency, although this last point can be discussed due to other negative effects of dollarization on debt dynamics uncertainty.

Figure 16 displays sensitivity analysis for the cumulative GIRF of pb_t shocks on debt levels. Panel A shows a sensitivity analysis to different values of the debt-dollarization ratio (δ), as higher values of this parameter strengthen the exchange rate channel. A higher dollarization ratio implies a faster and stronger reduction following an austerity shock. If the debt is fully dollarized ($\delta = 1$), the debt-to-GDP ratio would fall 1.1% of GDP after 10 quarters and 0.76% of GDP after 40 quarters following a 1% austerity shock, whereas in the fully de-dollarized case ($\delta = 0$) this impact is damped to 0.83% and 0.53% respectively. From another viewpoint, this implies that a higher dollarization ratio is desirable when austere fiscal-stabilization policies are conducted whereas a lower one is preferred when expansionary fiscal policies are needed.

Panel B of Figure 16 conducts a similar sensitivity analysis of this GIRF to different initial values of the debt-level (d_0). Higher debt levels are more difficult to handle because of its negative impact on g_t , which are added to the aforementioned negative effect of austere fiscal policies on the aggregate demand. Up to 10 quarters after the shock it seems that austere fiscal policies perform slightly better when debt levels are higher. However, afterwards we observe that a higher level of debt implies a less impactful austere fiscal policy, thus undermining the possibility of conducting fiscal consolidation processes when debt levels are very high. Under an initial debt level of 100% of GDP the long run impact of an austerity shock is a reduction of barely 0.30% of GDP, whereas under an initial debt-to-GDP ratio of 25% this impact widens up to a reduction of 0.67%. We also note that this effect is not linear, as the greater the debt level the greater loss in efficiency of the austerity shock.

As we have seen in previous figures, the futility (or lack thereof) of fiscal austerity depends on the relative size of the nominal growth channel, the magnitude of which is dependent on the values of fiscal multipliers. The austerity shock we identify in the FRF equation assumes that this multiplier is the same for both government spending and revenues shocks. This assumption is rather unlikely as government spending has a higher multiplier effect on GDP than taxation in Perú; see Jiménez et al. (2024) and Rossini et al. (2012). Therefore we estimate an alternative model where pb_t is disaggregated in government spending (g_t) and revenues (r_t), thus allowing us to discern if fiscal austerity is futile depending on whether it is conducted using government revenues or spending.

Figure 17 displays the cumulative GIRF of debt levels to spending and taxation austerity shocks for the selected models. We observe that spending austerity shocks are futile after 18 quarters in the TVP-VAR-SV model, after 15 quarters in the TVP-VAR-R3-SV model and after 12 quarters in the CVAR-SV model, as the 68% confidence interval includes zero after these horizons. Spending-austerity might have a relatively mild effect in the short term, as the debt-to-GDP ratio decreases between 0.32% and 0.56% of GDP after the shock depending on the chosen model. However, in all selected models, the distribution of the GIRFs return to the no-shock path afterwards. On the other hand, the impact of taxation-austerity policies are not futile, as the cumulative GIRFs confidence intervals do not return to the no-shock path at any quarter. Moreover, the CVAR-SV model estimates a stronger long-run effect of taxation fiscal austerity on the debt-to-GDP ratio than the one found in the standard model. After 40 quarters, the taxation driven austerity shock reduces the debt-to-GDP ratio in roughly 0.99%, whereas in the aggregated model this long run effect was 0.64% of GDP. These results hint at the desirability of using revenues driven austerity for fiscal consolidation processes.

In Figure 18 we observe the median decomposition of these shocks. We identify 3 reasons for the difference between the effect of government spending and revenues driven austerity shocks. First, the nominal growth channel for taxation driven austerity shocks is negligible, while it accounts for an increase of 0.33% of GDP for spending driven austerity shocks. As we have seen in Figures 7 and 8, this difference stems from the stronger negative impact of s_t shocks on g_t . These results align with those of Jiménez et al. (2023), as spending multipliers are vastly greater than the taxation ones. Second, the exchange rate channel is 6 times greater for taxation austerity shocks, implying that the risk premium channel set forth by Alberola et al. (2022) is stronger when fiscal consolidation processes are revenues driven. Finally, taxation driven shocks improve primary balances more persistently. While spending austerity shocks are reverted after 10 quarters and reduce its contribution to 0.4% of GDP in the long run, taxation driven austerity shocks are more persistent adding up to 1.2% of GDP over the same horizon.

In Figure 19 we observe that the evolution of these responses in the TVP-VAR-SV model have followed the

same pattern as in the baseline model, where the effectiveness of austerity shocks reduced sharply in 2006Q3 and 2009Q1, implying a null effect after 14 quarters in the case of government spending driven austerity shocks and a mild long run reduction of 0.1% of GDP in the case of taxation driven austerity shocks, and a subsequent increase in 2016Q2 and 2024Q, although in this case the effect does not match the one found in the CVAR-SV model. Figure 20 shows the decomposition of these responses for spending driven austerity shocks in the TVP-VAR-SV model. The banishing of the hump in the primary balance channel is the main driver of the greater efficacy of spending austerity shocks in 2016Q2 and 2024Q2. Such a hump implies that, some quarters after an austerity spending shock, the fiscal authority incurs in higher spending. In respect to the decomposition of the responses to taxation driven austerity shocks, Figure 20 shows that its main driver is the widening of the exchange rate channel towards 2024Q2. This implies the relevance of generating trust in investors by conducting recaudatory reforms.

In summary, we do find evidence against the fiscal austerity futility hypothesis when it is driven by temporary increases in revenues. Fiscal austerity works efficiently when the surpluses are sustained over a long period of time, the negative effect on real growth and inflation is milder and the exchange rate channel is taken advantage of. Spending driven austerity does not seem an appropriate instrument as its persistence is shorter, its effect on nominal growth is superior and does not generate an appreciation of the local currency through the risk premium channel.

4.4.7 Historical Decomposition (HD)

Panel A of Figure 21 displays the historical decomposition of the primary balance for the selected models. We observe a brief period of expansionary fiscal policy in the first years of the sample. After 2005Q1, where we observed that parameter γ_{epi}^{FRF} became statistically significant for some quarters, there is a clear tendency of the fiscal activism to react counter-cyclically to epi_t shocks. We note that the biggest generation of surpluses are driven by epi_t shocks during the commodities price boom between 2006 and 2014 according to all models. This result aligns with those of Urbina and Rodríguez (2022), who find a strong dependency of taxes to commodity prices. During this period, the contribution of epi_t shocks accounted for up to 2.832 % of quarterly GDP (roughly 0.708% of yearly GDP) according to CVAR-SV model. Nevertheless, there are efforts to strengthen the fiscal position far beyond what was explained by the FRF as we observe a series of fiscal austerity shocks in the range from 0.413% to 0.744 % of quarterly GDP (roughly 0.103 to 0.186 % of yearly GDP) during the period prior to the 2009 financial crisis. The magnitude of this effort varies between models, as the TVP-VAR-SV estimates an effort of up to 0.432% which is considerably lower than the 0.744 % estimated by the CVAR-SV. The CVAR-SV and the TVP-VAR-R3-SV suggest that some efforts to accumulate fiscal savings were made after the 2009 financial crisis, as there are austerity shocks as big as 0.767 % of quarterly GDP (roughly 0.192 % of yearly GDP) in the 2011Q1-2012Q4 period.

After 2014Q4, as γ_g was rising and the EPI growth pace was slowing, the PB shows a fiscal expansionary stance. According to all selected models, this expansionary policy was mainly driven by epi_t shocks, while the TVP-VAR-R3-SV attributes a higher participation to expansionary fiscal shocks in some quarters. From 2017Q1 until 2019Q4, most models coincide in that fiscal policy was discretionally determined as the expansionary fiscal shocks explained most of its variations. In this period we observe the -2.507% of quarterly GDP expansionary shock during the coastal El Niño, which was aimed to stabilize the economy. Nevertheless, from 2020Q1 onwards GDP growth shocks have gained importance over pb_t and epi_t innovations. One one hand, the magnitude of g_t shocks nudged the counter-cyclical nature of the FRF to incur in deficits to stabilize the economy during the COVID-19 crisis in all selected models. On the other hand, the contribution of epi_t shocks to pb_t have become negligible in the TVP-VAR-SV model, which is the only selected model that allows β_t coefficients to vary. This implies a weakening of the capacity of the mining industry to generate revenues and to dynamize the economy, which is coherent with the fact that record high copper prices during 2021 and 2022 were not enough to prevent g_t from slowing down.

In Panel B of Figure 21 we observe the HD of e_t during the sample period. As expected, this variable dynamics are driven mainly by its own shocks and by epi_t shocks. This is particularly clear between 2004Q1 and 2013Q4, where epi_t shocks account for appreciations of the local currency of up to 6,045% in the TVP-VAR-SV model. The role of fiscal policy is negligible during most of the sample period in the TVP-VAR-SV model but not in the other selected models, as the austerity shocks during the export price boom accounted for local currency appreciations of up to 1.721% in the TVP-VAR-R3-SV model and 1.570% in the CVAR-SV model. All selected models coincide in the assertion that, from 2020Q2 until 2021Q4, pb_t shocks contributed to increase e_t . The TVP-VAR-SV model estimates a contribution of 1.272% in 2020Q3 but the CVAR-SV model places it in 1.629% in the same quarter. Afterwards, pb_t shocks become irrelevant again in the

TVP-VAR-R3-SV and the TVP-VAR-SV models, but contribute 1.314% in 2024Q2 in the CVAR-SV model.

Figure 22 displays the HD of g_t and π_t in Panel A and Panel B respectively, which are the variables that make the nominal growth channel for debt level's reduction. We observe that g_t is mainly driven by epi_t shocks in the pre-COVID sample, after which g_t shocks predominate. The contribution of pb_t shocks is negligible during the whole sample, being the strongest contribution of 0.418% in 2022Q4 in the TVP-VAR-SV model and of 0.402% in 2022Q3 in the CVAR-SV model. On the other hand, pb_t shocks are more relevant in the HD of π_t . All models coincide that pb_t austerity shocks contributed to reduce inflation levels during the export price boom. The TVP-VAR-SV estimates a contribution of up to -0.628% in 2007Q3 while the CVAR-SV model estimates a contribution of -1.039% in the same quarter. From 2016Q1 onwards, as the policy maker incurred in an expansionary fiscal policy, all models coincide that pb_t deficit shocks accounted for increments in π_t . For instance, during the COVID crisis in 2020Q4, pb_t shocks accounted for 0.931% inflationary pressure in the CVAR-SV model and 0.812% in the TVP-VAR-SV model.

4.4.8 Forecast Error Variance Decomposition (FEVD)

Panel A of Figure 23 illustrates the FEVD of pb_t over twenty ($h = 20$) quarters. This decomposition allows us to identify the contribution of structural shocks to long-term macroeconomic fluctuations. We observe in all models epi_t shocks predominate over domestic variable shocks during most of the sample period. The relevance of these shocks reached its peak at 92.7% in 2009Q4 according to the CVAR-SV model. The contribution of pb_t shocks in the FEVD of the primary balance has been minor in comparison, as this fluctuated between 3.0% and 10.0% between 2000Q2 and 2019Q4. Nevertheless, the magnitude of the g_t shocks during the COVID-19 crisis increased the share of g_t in this FEVD up to 76.0% in 2021Q2. After 2020Q1, the shares of contributions seem to return to pre-COVID patterns with epi_t shocks around 78.2% and pb_t shocks around 10.5%. On the other hand, the rest of the shocks combined have never accounted for more than 10.0% of the FEVD of pb_t .

In Panel B of Figure 23 we observe the FEVD of e_t . All selected model coincide in attributing a minor relevance to pb_t shocks before the COVID years. However, afterwards the relevance of pb_t shocks rises up to 6, 5% of long run fluctuations in the TVP-VAR-SV model, being the most relevant domestic variable in the decomposition. This implies that as fiscal policy has become more volatile after the COVID crisis, it has become a source of uncertainty in the currency market as theorized by Alberola et al. (2022).

Figure 24 displays the FEVD of g_t and π_t in Panel A and Panel B respectively. A similar pattern is observed, as the share of uncertainty attributed to pb_t shocks became negligible after the export price boom but rose afterwards, reaching its highest values after the COVID crisis. In the decomposition of π_t , the share of uncertainty attributed to austerity shocks reaches 5.9% in 2024Q2 according to the TVP-VAR-SV model, at least twice as large as the share it had before the COVID years.

In summary, although epi_t shocks remain the main driver of uncertainty for the variables in our model, the relevance of pb_t shocks has risen sharply in recent years. This implies a volatile fiscal policy that might have collateral effects on inflation and exchange rate dynamics.

5 Robustness Analysis

To validate our baseline results, we present the following robustness exercises: (i) alternative ordering ($\mathbf{y}_t = (epi_t, e_t, \pi_t, i_t^c, i_t^f, g_t, pb_t, \Delta d_t)'$), (ii) change in lag structure ($p = 2$), (iii) only pre-COVID crisis sample (from 2000Q2 until 2019Q4) and (iv) change in the external variable (real growth rate of the GDP of the United States, henceforth g_t^{usa}). As we observe in Table 3, in all cases the model chosen by the Log-ML and the DIC is the CVAR-SV model; except for the $p = 2$ model, where the Log-ML criteria chooses the TVP-VAR-SV.

As the main focus of these analysis is the cumulative impact of pb_t shocks on d_t using GIRFs, this is the point of comparison. In Figure 25 are displayed the proposed robustness checks and in Figure 26, the decomposition of these using equation 12. Our baseline model predicts that, conditional on an initial debt-to-GDP ratio and initial values for \mathbf{y}_t in its sample mean, an austerity shock will decrease d_t up to 1.000% of GDP in the 9th quarter. Afterwards, the debt-to-GDP ratio rises slowly until stabilizing around -0.600% of GDP below the no-shock path in the long run, creating a hump-shaped response.

Under an alternative ordering the cumulative GIRF is not hump-shaped although it stabilizes around -0.600% of GDP as in the baseline model, see Panel A of Figure 25. In Panel A of Figure 26, we observe that there are a couple of reasons for this divergence. First, pb_t shocks don't seem to reduce nominal GDP

growth in the first 25 quarters. Second, a pb_t shock does not generate primary surpluses in the first 5 quarters after the shock according to this model.

Our second robustness check analyses the exclusion of the post-COVID period from the sample. As we observe in Panel B of Figure 26, the hump shaped observed in the baseline model is partially lost. However, both models stabilize the long run cumulative responses of d_t to pb_t shocks around the same level. Panel B of Figure 26 offers insights for this difference. As we observe, the behavior of the nominal growth, the primary balance and the interest rates channel are roughly similar; the only notorious difference being that of the exchange rate channel. In the pre-COVID model there is no role for the exchange rate channel of fiscal stabilization, which fluctuates nearly zero. On the other hand, the exchange rate channel is taken advantage of in the full model in the first 8 quarters after the pb_t shock. This suggests that the exchange rate channel gains relevance when modelling a sample with periods of high macroeconomic uncertainty, as is the post-COVID sample. During these periods the exchange rate might react more sharply to the fiscal policy stance.

Changing the lag structure to $p = 2$ does not modify the hump-shaped response observed in the baseline model. However, the long run effect of pb_t shocks on d_t increases up to 0.600% below the no shock path as we see in Panel C of Figure 25. In Panel C of Figure 26 we observe a couple of difference in how this effect is decomposed. First, although the $p = 2$ model also attributes a role for the exchange rate channel, it does consider it weaker. Second, the primary balance channel is stronger in the $p = 2$ model, as it attributes it a contribution of 1.000% of GDP of the long run reduction, whereas in the baseline model this contribution is not larger than 0.85% of GDP.

Finally, Panel D of Figure 25 displays the robustness analysis for a change in the selected external variable. In this case the impact of austerity shocks is notoriously stronger, reaching a reduction of d_t of 1.600% in the long run. In terms of the decomposition, the main difference with the baseline model is a stronger pb_t channel. The underlying reason is that this model attributes a stronger persistence of austerity shocks on pb_t , which might be overestimated because epi_t dynamics are omitted.

6 Conclusions

This paper inquires into the properties of the peruvian FRF between 2000Q2 and 2024Q4 and the effectiveness of austerity shocks on the reduction of the debt-to-GDP ratio using a TVP-VAR-SV framework. Our results shed light on the time-varying behavior of the peruvian fiscal authority. First, we find an stabilizing response of the fiscal authority to debt levels growth in line with Moreno (2019). We also find a counter cyclical behavior of the fiscal policy during the whole sample period. Moreover, this counter cyclical response of the primary balance has become faster after 2010, implying a more efficient fiscal activism. However, there is no consistent evidence for the presence of borrowing costs in the FRF. In terms of the volatility of austerity shocks, we find a reduction during the years of the export prices boom and a subsequent increase to pre-fiscal rules levels after the coastal El Niño of 2017 and the COVID-19 pandemic crisis of 2020. In recent years fiscal policy remains a source of uncertainty for the economy as the standard deviation of austerity shocks reaches 1.3% of quarterly GDP (roughly 0.33% of yearly GDP).

In regard to the fiscal austerity futility hypothesis, we find evidence against it in all selected models. After an austerity shock, debt remains around 0.5% of GDP below the no-shock path in the TVP-VAR-SV and the TVP-VAR-R3-SV models. In the CVAR-SV this effect is even stronger, estimating a long run reduction of debt levels of 0.6% of GDP. The effectiveness of fiscal austerity policies rely on minding the nominal growth reduction channel, the persistence of austerity shocks and the utilization of the exchange rate channel. Moreover, we find that high dollarization rates amplify the effectiveness of fiscal austerity policies when the exchange rate channel is taken advantage of. The long-run effectiveness of fiscal austerity is also enhanced when the initial debt level is lower, which alligns with the debt feedback hypothesis of Stanova (2015). However, we do find that fiscal austerity shocks are self defeating if these are conducted through spending cuts, as the cumulative GIRF of debt to spending driven austerity shocks strongly lean to the no-shock path, whereas taxation driven austerity shocks reduce the debt-to-GDP ratio in 1.0% in the long run.

Finally, our HD and FEVD analysis point at the relevance of epi_t shocks between 2004 and 2014, period during which was conducted the fiscal stabilization process which reduced the debt-to-GDP ratio from 50% to below 20%. This implies that the role of austerity shocks for the debt consolidation process in Perú are negligible. In recent years, the contribution of EPI growth shocks to the primary balance has reduced and the relative importance of GDP growth shocks has increased. This leaves the possibility for a local management of debt through fiscal policy.

These results hint at the following policy recommendations. First, fiscal consolidation should be conducted by raising taxation revenues instead of reducing spending, as the multiplier effects of taxation on GDP are vastly inferior to those of spending shocks. Because of this high multiplier effect of spending shocks, fiscal austerity conducted through spending is self-defeating in all chosen models. Additionally, taxation shocks have a greater persistence, thus amplifying the long run effect of austerity shocks; whereas spending austerity shocks are short lived and tend to reverse to expansionary policies in the medium run, thus minding the effectiveness of spending driven austerity in the long run. Second, a higher debt dollarization rate is desirable when fiscal consolidation policies are conducted in order to take advantage of the exchange rate channel. Although we advocate for the utilization of revenues as the policy instrument for fiscal consolidation, we note that at some quarters during the export price boom its effect was vastly mindered. We note that during this period the exchange rate channel greatly decreased, probably because of the relatively great influx of foreign currency into the local economy as a result of commodities trade.

Future research on this topic should incorporate different identification schemes that allow for a two-way simultaneous impact between the primary surplus and the GDP growth rate as done by Blanchard and Perotti (2002). Another improvement on this model could be done by introducing more external variables such as the FED interest rate and the chinese GDP growth rate, which would enrich the analysis of the impact of monetary and demand shocks on domestic debt dynamics. Finally, this framework should be evaluated for samples with relatively high debt-to-GDP ratios under unsustainable debt paths, as in this case the time series properties of debt levels might change and the utilization of the deterministic debt levels equation might not hold.



Appendix 1

We use the estimates of the coefficients of equation (1) for the construction of the matrices in equation (9). Let n be the number of variables in our model, then equation (1) in an SVAR of n variables. As we exclude the n -th equation, corresponding to Δd_t from the model, equation (9) becomes a SVAR of $n - 1$ variables with a pseudo-exogenous variable Δd_t . This implies subtracting the n -th column and the n -th row from matrices \hat{B}_{0t} and \hat{B}_{jt} . Matrix $\hat{\Omega}_t$ is extracted from \hat{B}_{jt} for the feedback coefficients of Δd_t in equation (9).

$$\hat{B}_{0t} = \left[\begin{array}{c|c} \hat{A}_{0t} & \hat{B}_{0t}^{(1:n-1,n)} \\ \hline \hat{B}_{0t}^{(n,1:n-1)} & \hat{B}_{0t}^{(n,n)} \\ \hline 1 \times n-1 & 1 \times 1 \end{array} \right]$$

$$\hat{B}_{1t} = \left[\begin{array}{c|c} \hat{A}_{1t} & \hat{\Omega}_t \\ \hline \hat{B}_{1t}^{(n,1:n-1)} & \hat{B}_{1t}^{(n,n)} \\ \hline 1 \times n-1 & 1 \times 1 \end{array} \right]$$

Matrix $\hat{B}_{0t}^{(1:n-1,n)}$ is read as the submatrix built by extracting the intersection between first ' $n - 1$ 'th rows and the n th column of \hat{B}_{0t} . Another way of expressing the previous equalities is:

$$\hat{A}_{0t} = \hat{B}_{0t}^{(1:n-1,1:n-1)}$$

$$\hat{A}_{1t} = \hat{B}_{1t}^{(1:n-1,1:n-1)}$$

$$\hat{\Omega}_t = \hat{B}_{1t}^{(1:n-1,n)}$$

Appendix 2

The International Monetary Fund (2013) proposes a set of decompositions for Δd_t for analytical purposes. We use a modified version of one of the decompositions:

$$\Delta d_t = \underbrace{\frac{(1-\delta)i_t^{lc} + \delta i_t^{fc}}{(1+\pi_t)(1+g_t)} d_{t-1}}_{\hat{i}_t} - \underbrace{\frac{g_t + \pi_t(1+g_t)}{(1+\pi_t)(1+g_t)} d_{t-1}}_{\hat{y}_t} + \underbrace{\frac{\delta e_t(1+i_t^{fc})}{(1+\pi_t)(1+g_t)} d_{t-1}}_{\hat{e}_t} + \underbrace{(-pb_t)}_{\hat{p}b_t} \quad (13)$$

As pb_t shocks can affect debt through any of these channels, we deem relevant to decompose the cumulative GIRF of Δd_t to austerity shocks as follows:

$$\Delta d_t^s - \Delta d_t^{ns} = \underbrace{(\hat{i}_t^s - \hat{i}_t^{ns})}_{\text{Interest rates}} + \underbrace{(\hat{y}_t^s - \hat{y}_t^{ns})}_{\text{Nominal growth}} + \underbrace{(\hat{e}_t^s - \hat{e}_t^{ns})}_{\text{Depreciation rate}} + \underbrace{(\hat{p}b_t^s - \hat{p}b_t^{ns})}_{\text{Primary balance}} \quad (14)$$

where Δd_t^s is the evolution of Δd_t in the shocked path ($k = s$) and Δd_t^{ns} is the analogous in the no-shock path ($k = ns$).

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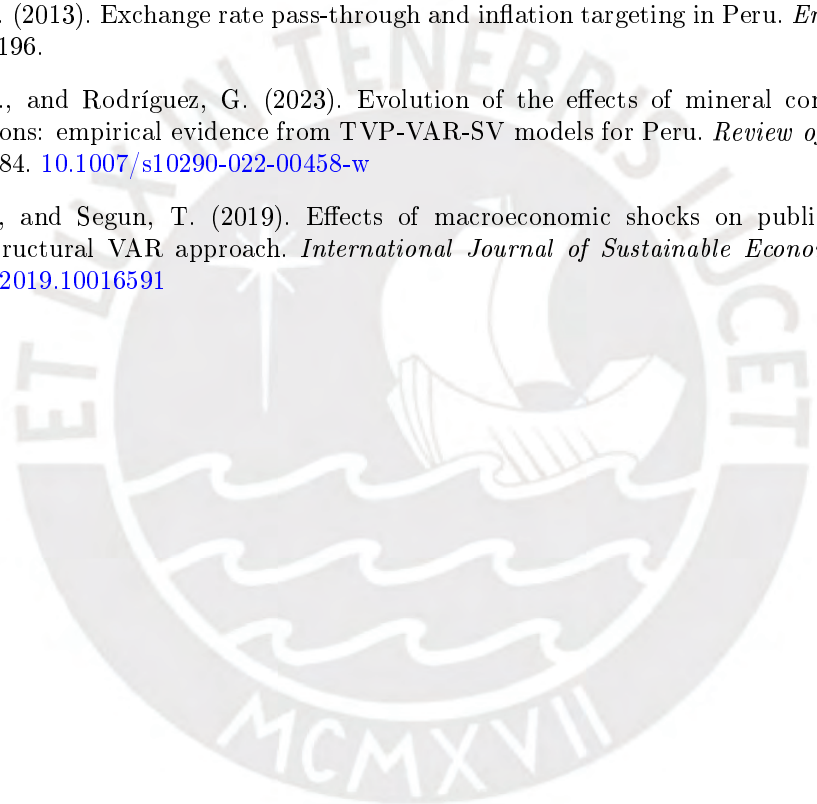
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Tables and Figures

Table 1. Tests for Time Variation in Coefficients and Volatility

Test	Coefficients	2000Q3-2009Q1	2009Q1-2016Q2	2016Q2-2024Q2	Full Sample
Kolmogorov-Smirnov	$B_{0,t}$	26/28	27/28	28/28	28/28
	$B_{i,t}$	61/72	58/72	60/72	66/72
	Σ_t	8/8	8/8	8/8	8/8
t -test	$B_{0,t}$	24/28	27/28	24/28	28/28
	$B_{i,t}$	51/72	61/72	50/72	65/72
	Σ_t	8/8	8/8	8/8	8/8
Trace test	Trace	16% perc	50% perc	84% perc	
	0.32	0.38	0.53	0.74	

$B_{0,t}$ represent contemporaneous impact coefficients, $B_{j,t}$ represent intercept and lagged variables coefficients and h_t represent variance of the structural innovations.



Table 2. Model Selection Criteria

Panel A. Log-ML and DIC

Model	Log-ML	SD	Rank	DIC	SD	p_D	SD	Rank
TVP-VAR-SV	-1923.237	0.254	2	3301.799	1.020	109.164	0.545	5
TVP-VAR	-2032.033	0.214	7	3498.185	1.370	98.2154	0.787	7
TVP-VAR-R1-SV	-1923.934	0.367	3	3250.396	0.909	117.832	0.441	3
TVP-VAR-R2-SV	-1925.299	0.356	4	3272.902	1.272	112.380	0.522	4
TVP-VAR-R3-SV	-1929.921	1.142	5	3216.100	0.699	122.588	0.348	2
CVAR-SV	-1921.702	0.078	1	3204.831	0.354	122.356	0.238	1
CVAR	-2028.453	0.024	6	3394.497	0.328	103.958	0.159	6

Each Log-ML estimate is based on 10,000 evaluations of the integrated likelihood using 10 parallel chains, where the importance sampling density is constructed using 10,000 posterior draws after a burn-in period of 1,000. Each DIC estimate (and the corresponding numerical standard error) is computed using 10 parallel chains; each consists of 10,000 posterior draws after a burn-in period of 1,000. The integrated likelihood and the DIC are evaluated every 10th post burn-in draw—a total of 10,000 evaluations.

Panel B. Bayes Ratio

	TVP-VAR-SV	TVP-VAR	TVP-VAR-R1-SV	TVP-VAR-R2-SV	TVP-VAR-R3-SV	CVAR-SV	CVAR
TVP-VAR-SV	1	1.778×10^{47}	2.007	7.861	799.091	0.215	4.955×10^{45}
TVP-VAR	5.627×10^{-48}	1	1.132×10^{-47}	4.421×10^{-47}	4.503×10^{-45}	1.216×10^{-48}	0.030
TVP-VAR-R1-SV	0.490	8.848×10^{46}	1	3.918	398.135	0.101	2.465×10^{45}
TVP-VAR-R2-SV	0.124	2.252×10^{46}	0.254	1	101.526	0.025	6.294×10^{44}
TVP-VAR-R3-SV	0.008	2.224×10^{44}	0.004	0.004	1	0.007	6.195×10^{42}
CVAR-SV	4.635	8.244×10^{47}	9.318	36.428	3.710×10^3	1	2.291×10^{46}
CVAR	2.016×10^{-46}	35.852	4.059×10^{-46}	1.588×10^{-45}	1.615×10^{-43}	4.354×10^{-47}	1

Coefficients must be interpreted as bayes factor between the model in the column over the model in the row.

Table 3. Model Selection Criteria for Robustness Exercises

	Log-ML	SD	Rank	DIC	SD	p_D	SD	Rank
Alternative Ordering								
TVP-VAR-SV	-1922.088	0.116	2	3304.369	1.225	112.252	0.655	5
TVP-VAR	-2022.546	0.279	7	3483.429	2.531	98.6463	0.713	7
TVP-VAR-R1-SV	-1922.098	0.652	3	3247.263	0.580	118.088	0.264	3
TVP-VAR-R2-SV	-1926.987	0.494	4	3286.392	0.852	111.446	0.466	4
TVP-VAR-R3-SV	-1930.182	1.512	5	3224.043	0.822	121.997	0.647	2
CVAR-SV	-1921.386	0.057	1	3212.054	0.370	122.703	0.160	1
CVAR	-2021.220	0.039	6	3393.764	0.256	103.825	0.130	6
2-Lag Model								
TVP-VAR-SV	-1730.056	0.341	1	2919.983	2.148	114.632	0.854	5
TVP-VAR	-1808.915	0.290	6	3103.736	1.340	101.832	0.759	7
TVP-VAR-R1-SV	-1738.897	0.575	5	2873.671	1.198	119.974	0.531	3
TVP-VAR-R2-SV	-1734.454	0.647	3	2896.398	2.094	116.675	0.535	4
TVP-VAR-R3-SV	-1736.328	0.920	4	2853.873	0.971	124.708	0.390	2
CVAR-SV	-1731.826	0.140	2	2844.026	0.563	126.748	0.223	1
CVAR	-1812.531	0.033	7	3023.641	0.246	104.708	0.135	6
Pre-COVID Sample								
TVP-VAR-SV	-1509.921	0.306	2	2524.944	1.419	102.158	0.682	6
TVP-VAR	-1533.558	0.368	6	2539.458	1.235	97.4670	0.608	7
TVP-VAR-R1-SV	-1510.769	0.553	3	2483.577	0.635	110.260	0.279	4
TVP-VAR-R2-SV	-1513.370	0.450	4	2506.354	1.038	103.998	0.443	5
TVP-VAR-R3-SV	-1516.295	0.996	5	2459.282	0.811	114.521	0.246	3
CVAR-SV	-1505.515	0.051	1	2442.330	0.416	116.982	0.183	1
CVAR	-1533.580	0.024	7	2457.867	0.242	103.178	0.105	2
Change in external variable								
TVP-VAR-SV	-2076.658	0.426	4	3319.945	1.393	165.075	0.596	5
TVP-VAR	-2173.827	0.304	7	3501.858	2.416	150.847	0.913	7
TVP-VAR-R1-SV	-2065.914	0.495	2	3209.995	0.642	187.262	0.439	3
TVP-VAR-R2-SV	-2079.731	0.422	5	3288.154	1.431	168.587	0.819	4
TVP-VAR-R3-SV	-2076.578	1.582	3	3164.742	0.672	193.935	0.296	2
CVAR-SV	-2060.684	0.074	1	3163.901	0.479	195.378	0.251	1
CVAR	-2156.376	0.095	6	3327.953	0.346	165.336	0.170	6

Each Log-ML estimate is based on 10,000 evaluations of the integrated likelihood using 10 parallel chains, where the importance sampling density is constructed using 10,000 posterior draws after a burn-in period of 1,000. Each DIC estimate (and the corresponding numerical standard error) is computed using 10 parallel chains; each consists of 10,000 posterior draws after a burn-in period of 1,000. The integrated likelihood and the DIC are evaluated every 10th post burn-in draw—a total of 10,000 evaluations.

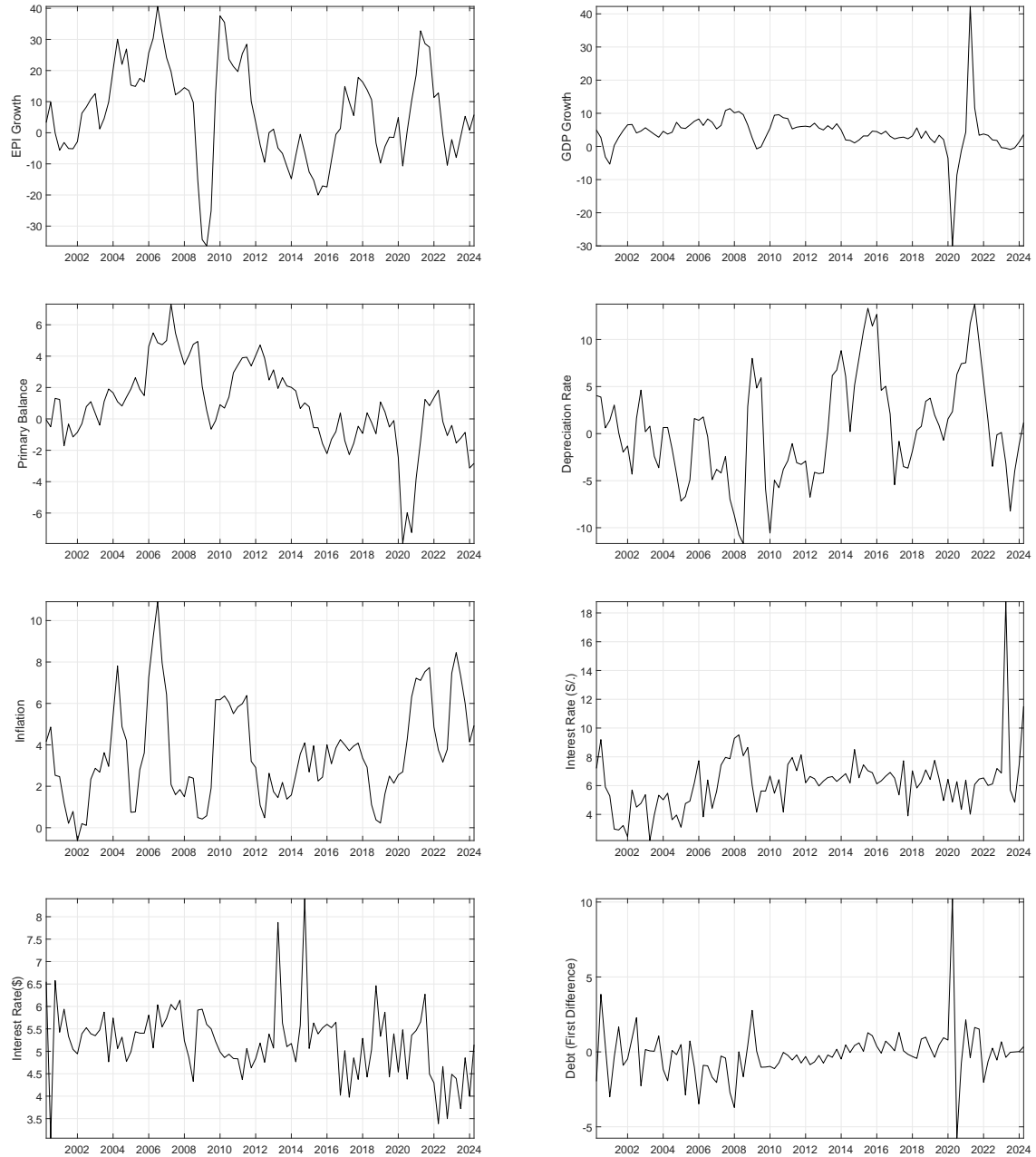


Figure 1. Model Variables Series

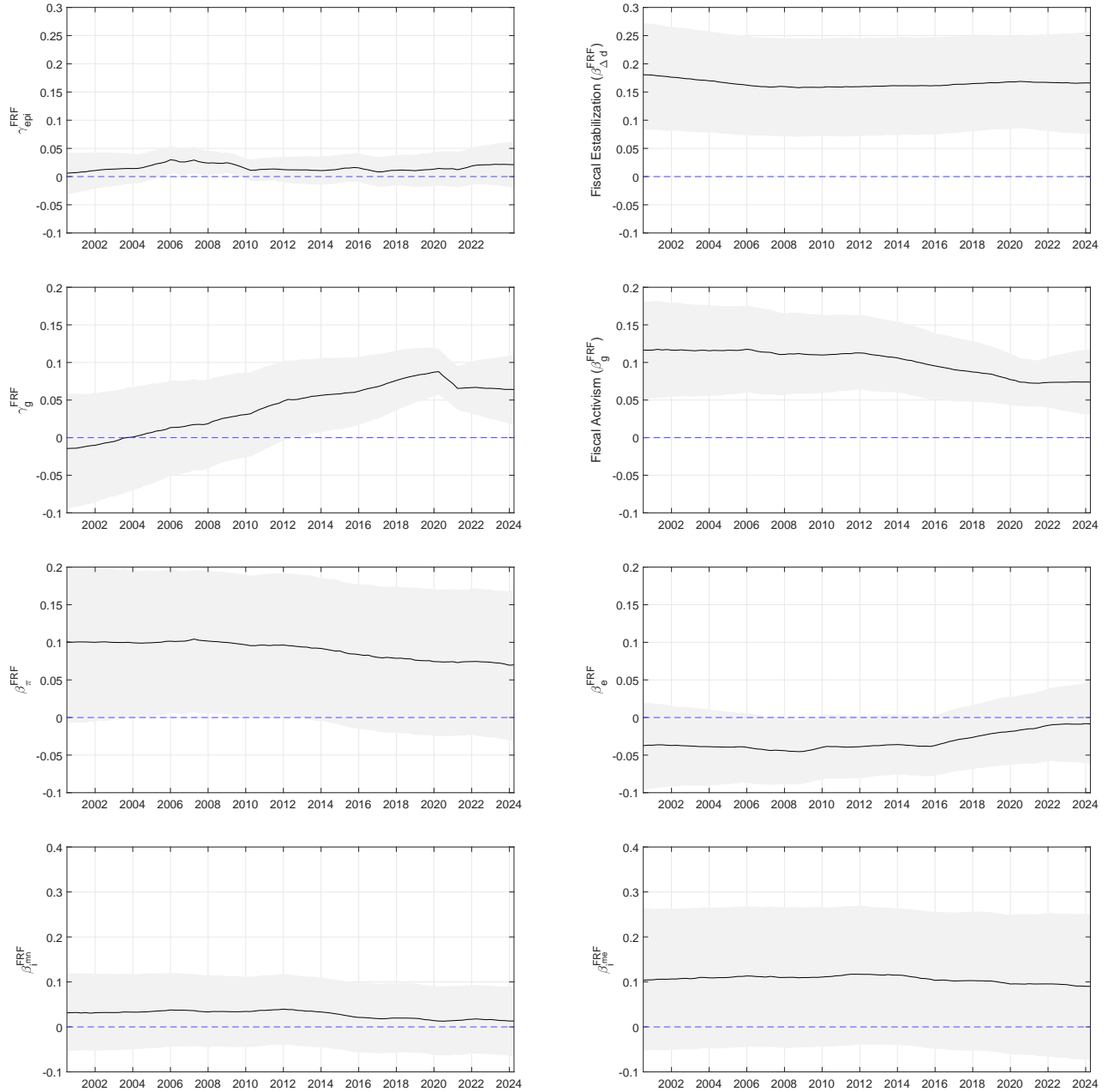


Figure 2. FRF Parameters Evolution in the TVP-VAR-SV Model

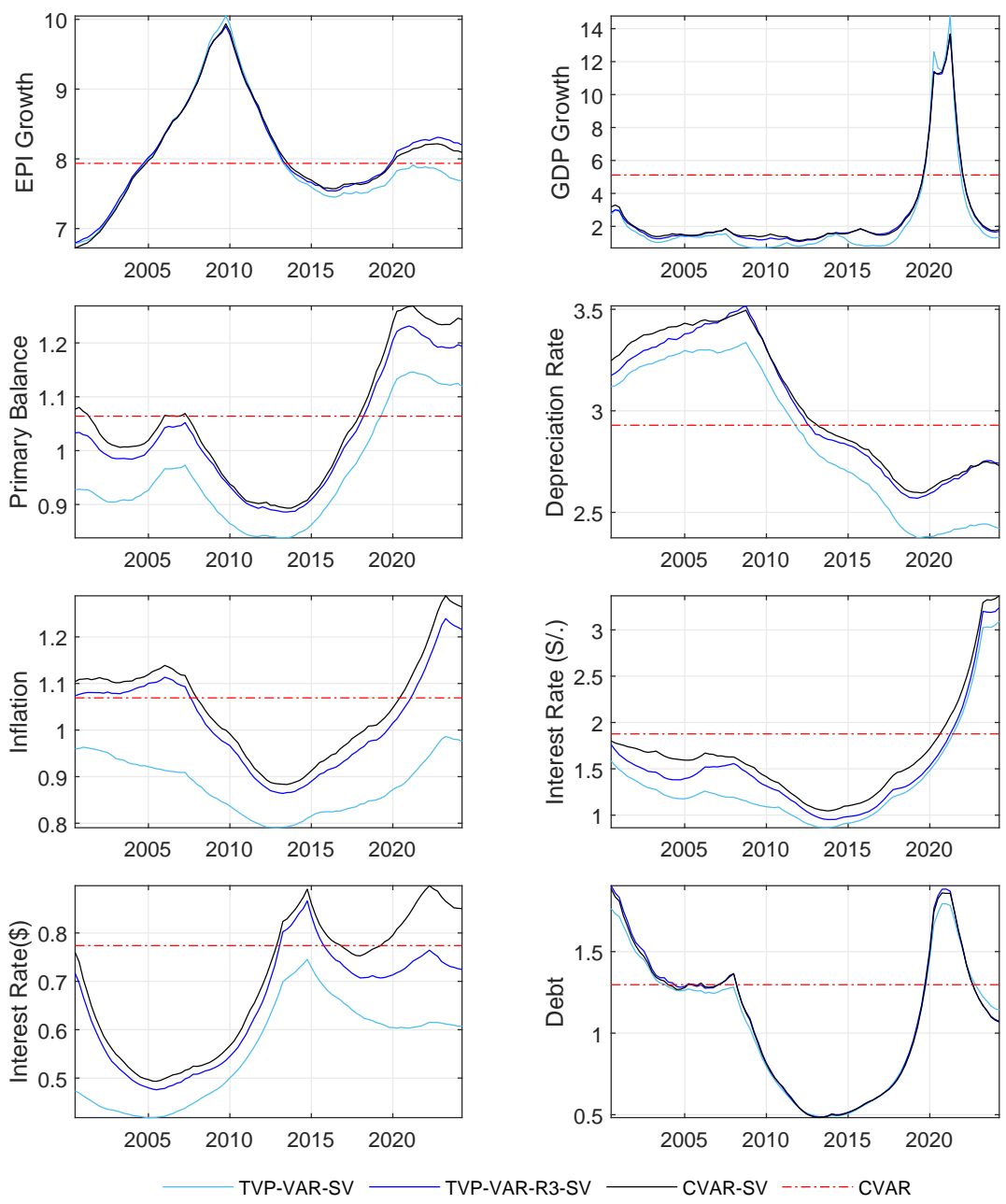


Figure 3. Time-Varying Median Standard Deviation of Structural Shocks

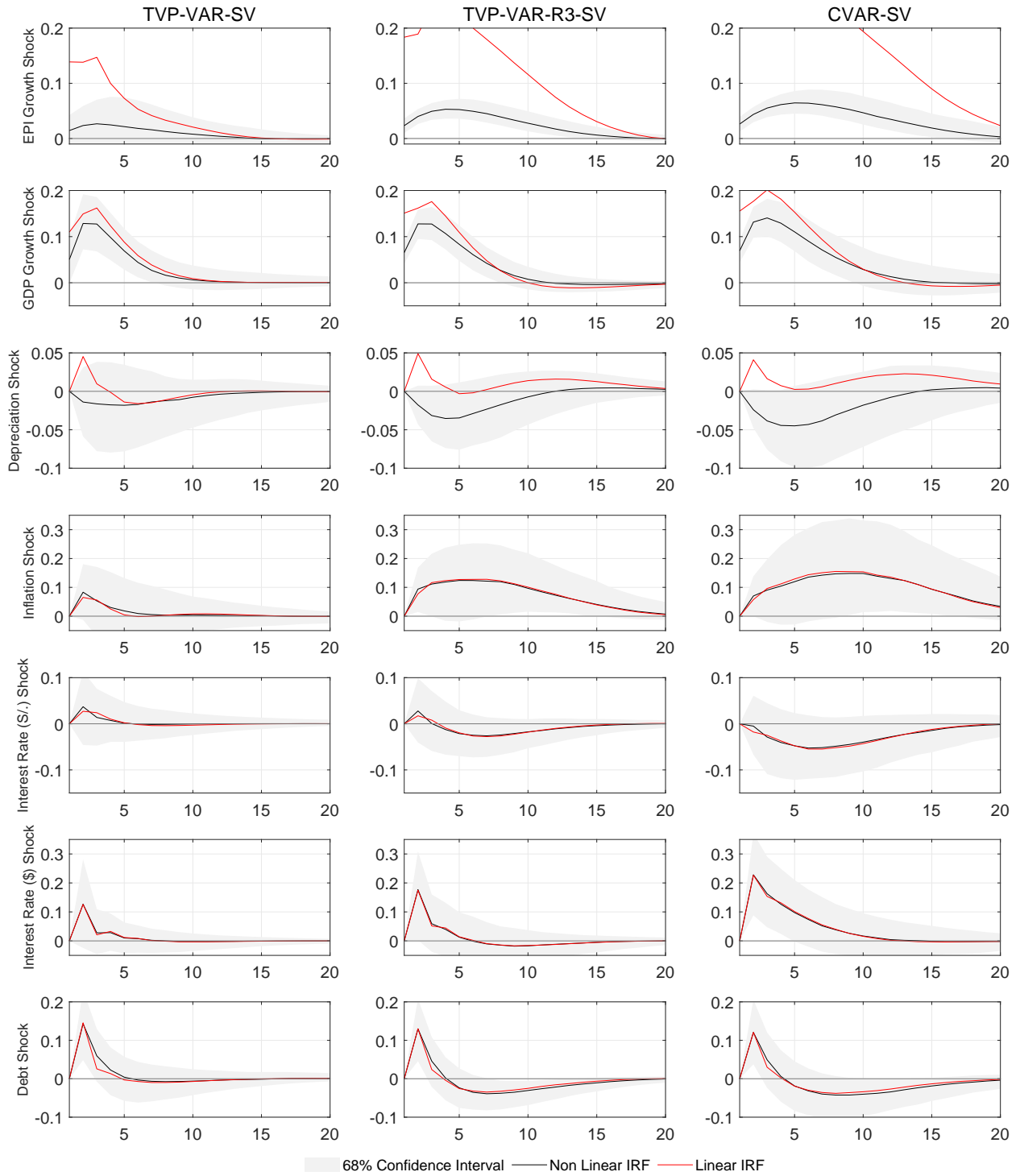


Figure 4. Median GIRF of Primary Balance (% of GDP) to Structural Shocks

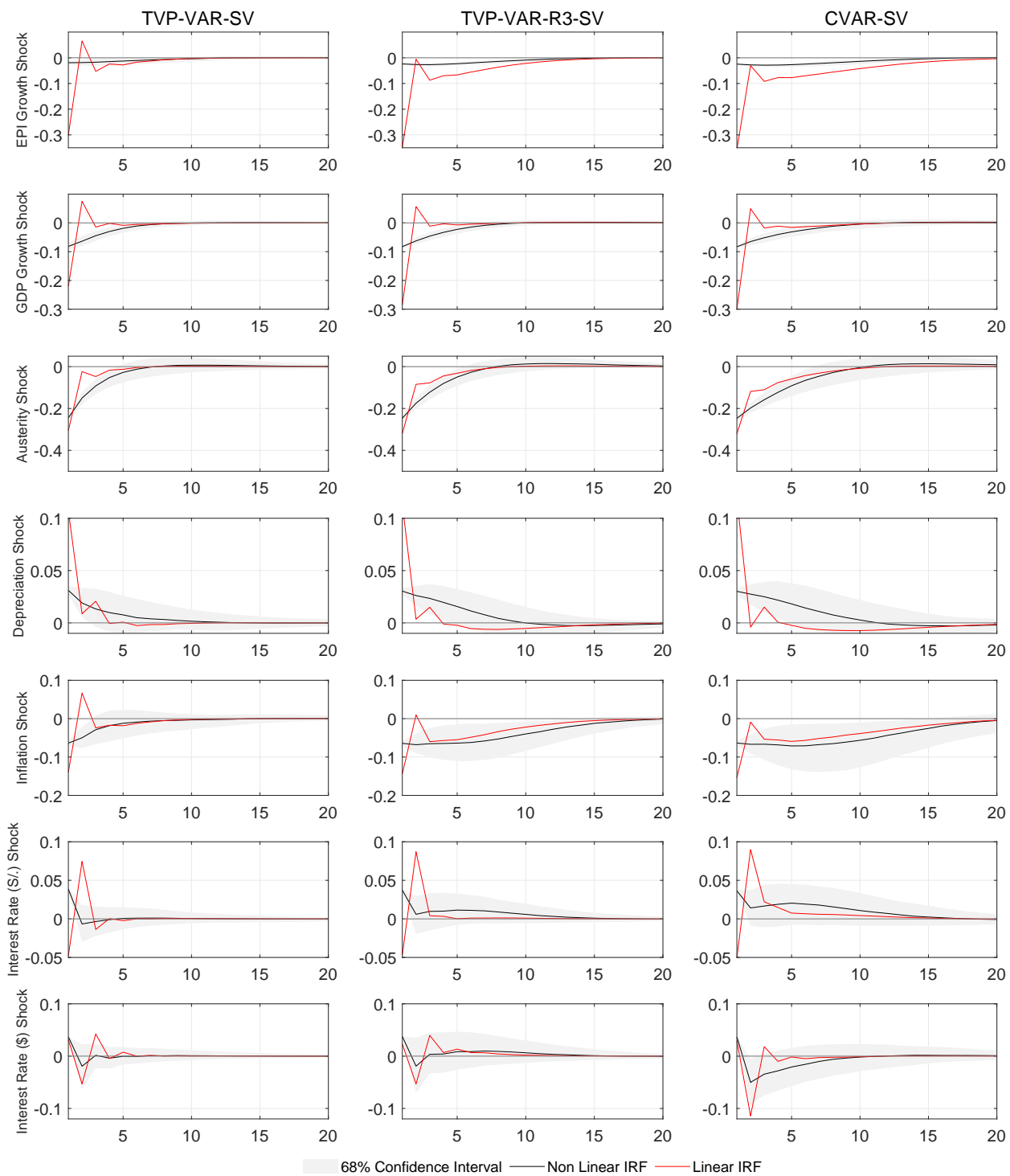


Figure 5. Median GIRF of Gross Debt (% of GDP) to Structural Shocks

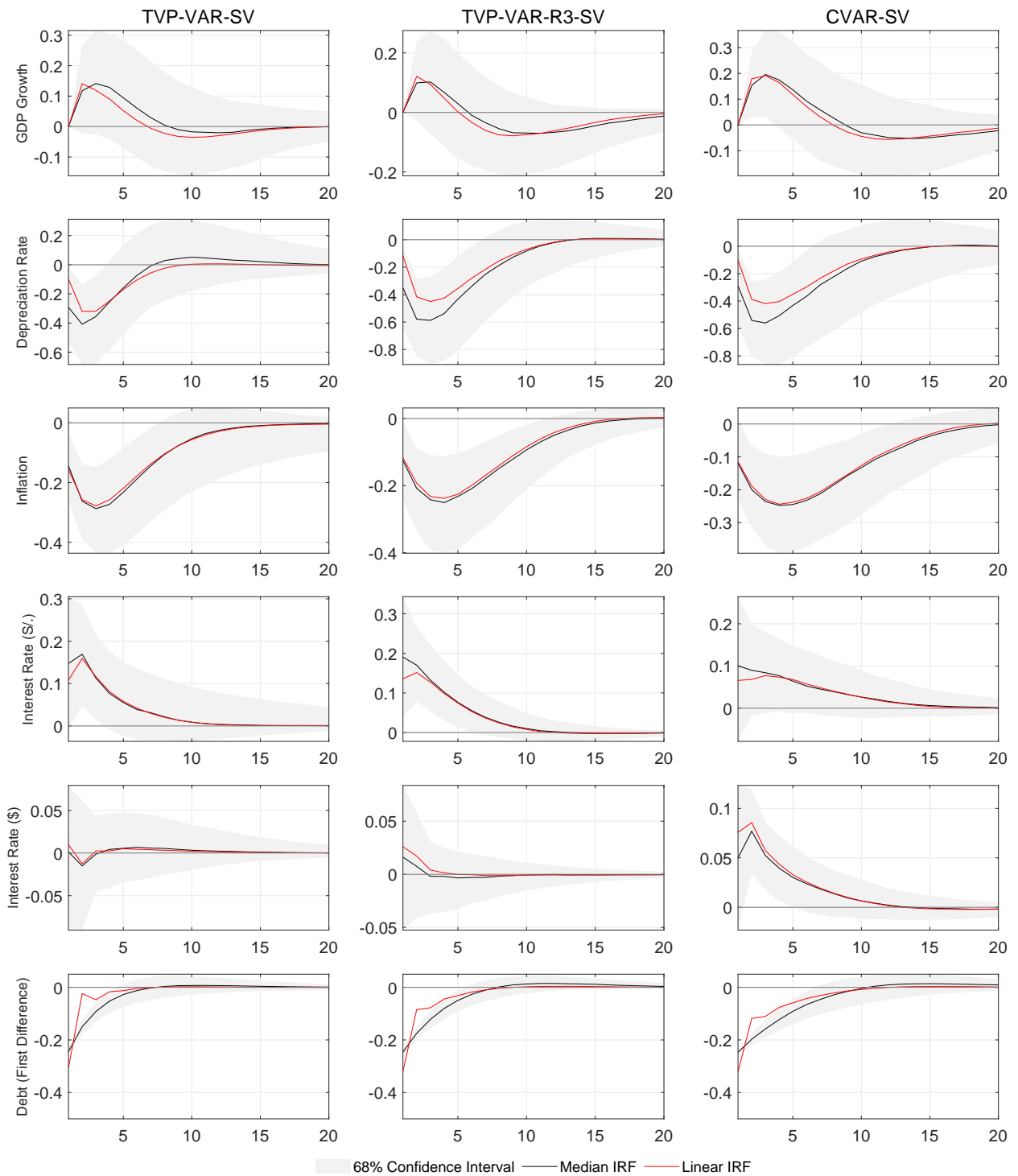


Figure 6. Median GIRF of All Variables to Austerity Shocks

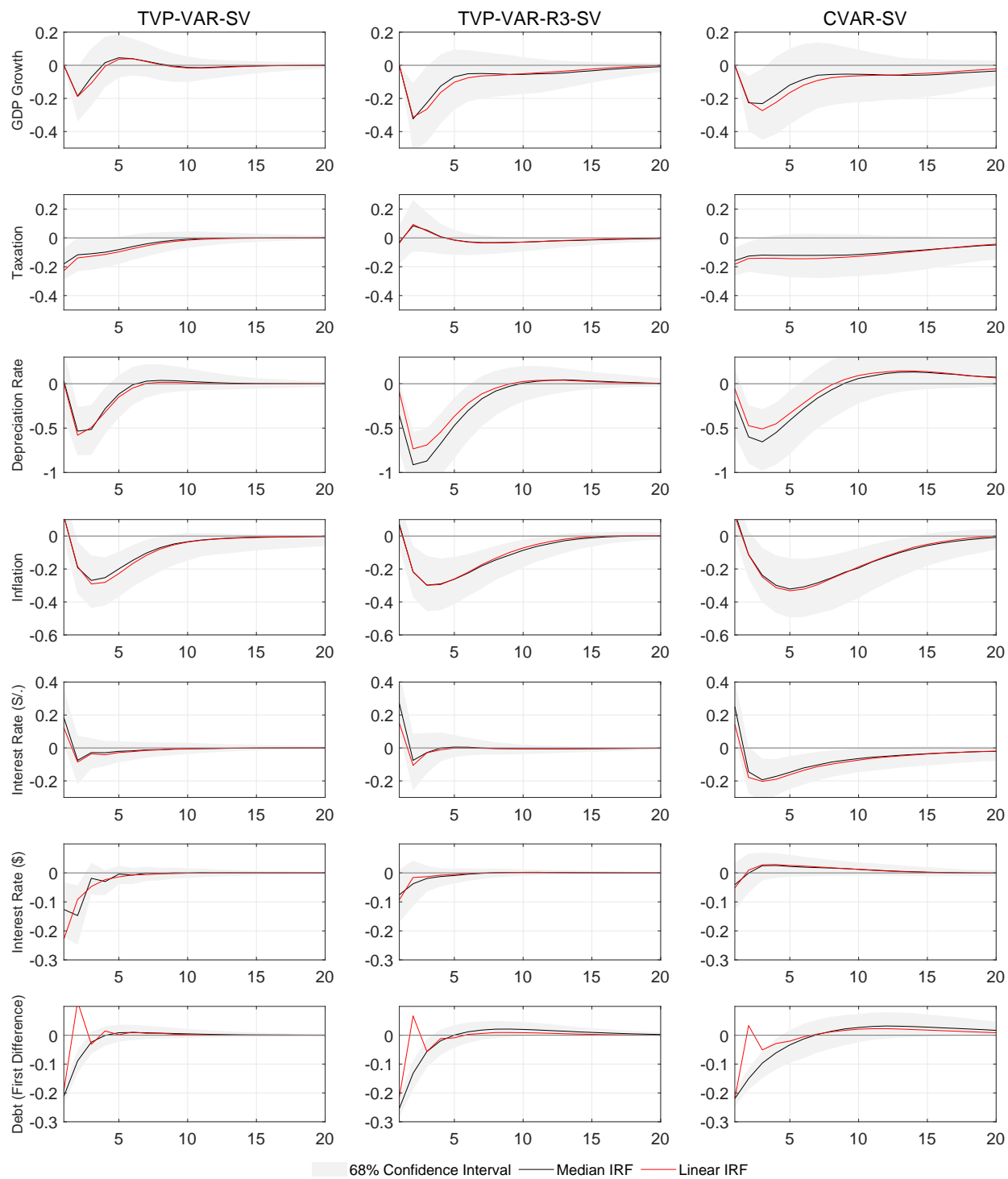


Figure 7. Median GIRF of All Variables to Spending Austerity Shocks

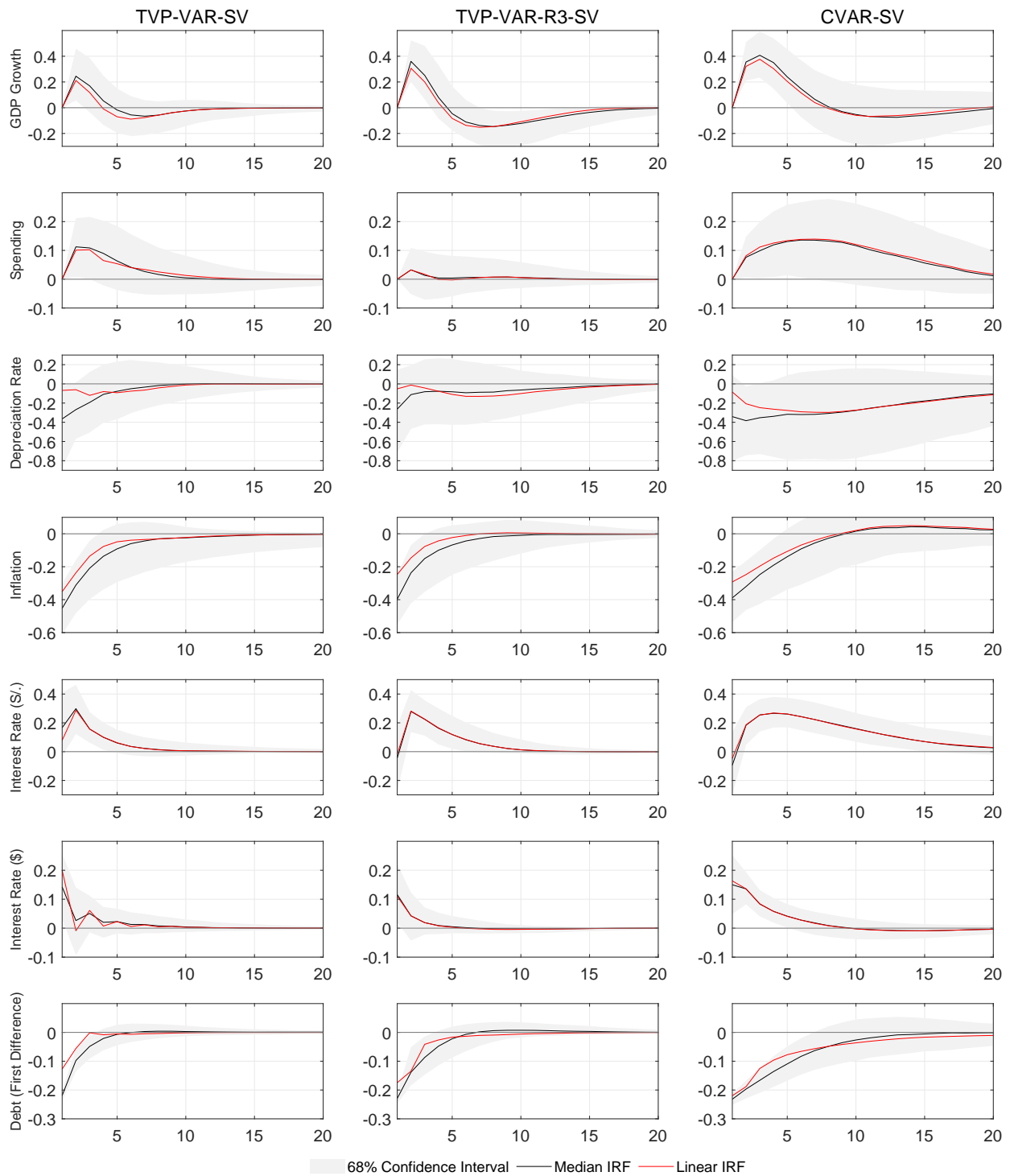


Figure 8. Median GIRF of All Variables to Taxation Austerity Shocks

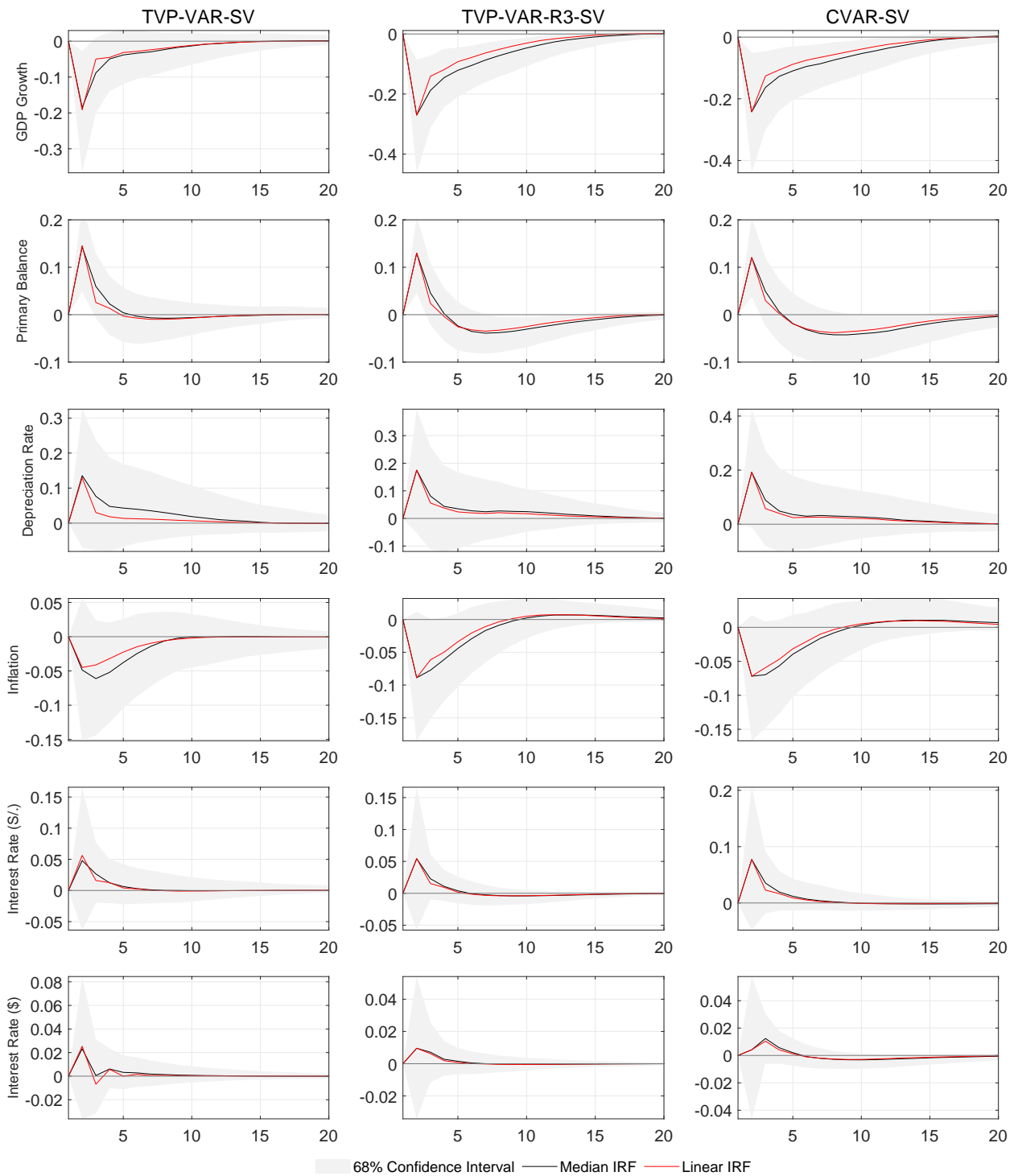


Figure 9. Median GIRF of All Variables to Debt Shocks

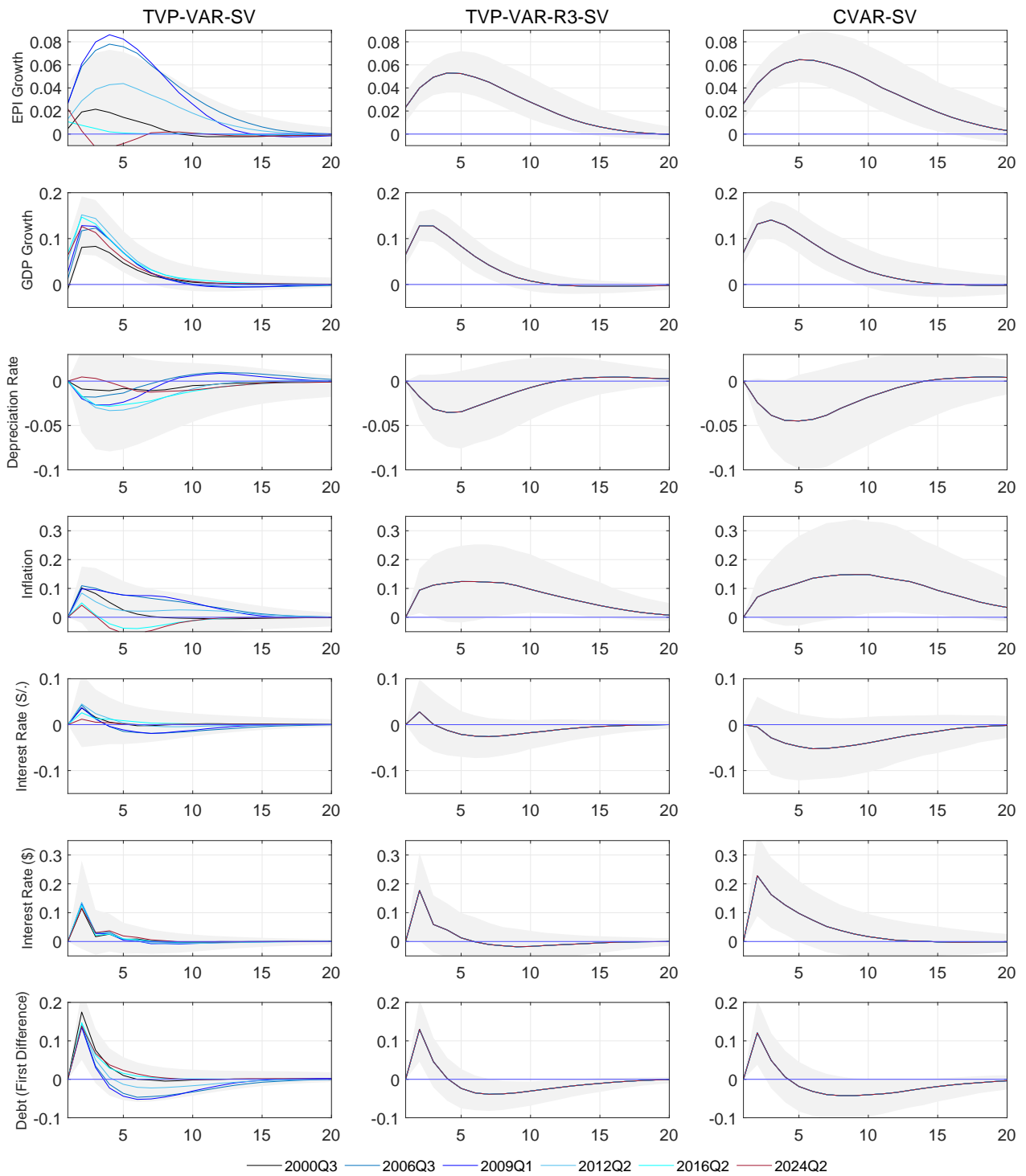


Figure 10. Quarter-Specific Median GIRFs of Primary Balance (% of GDP) to Structural Shocks

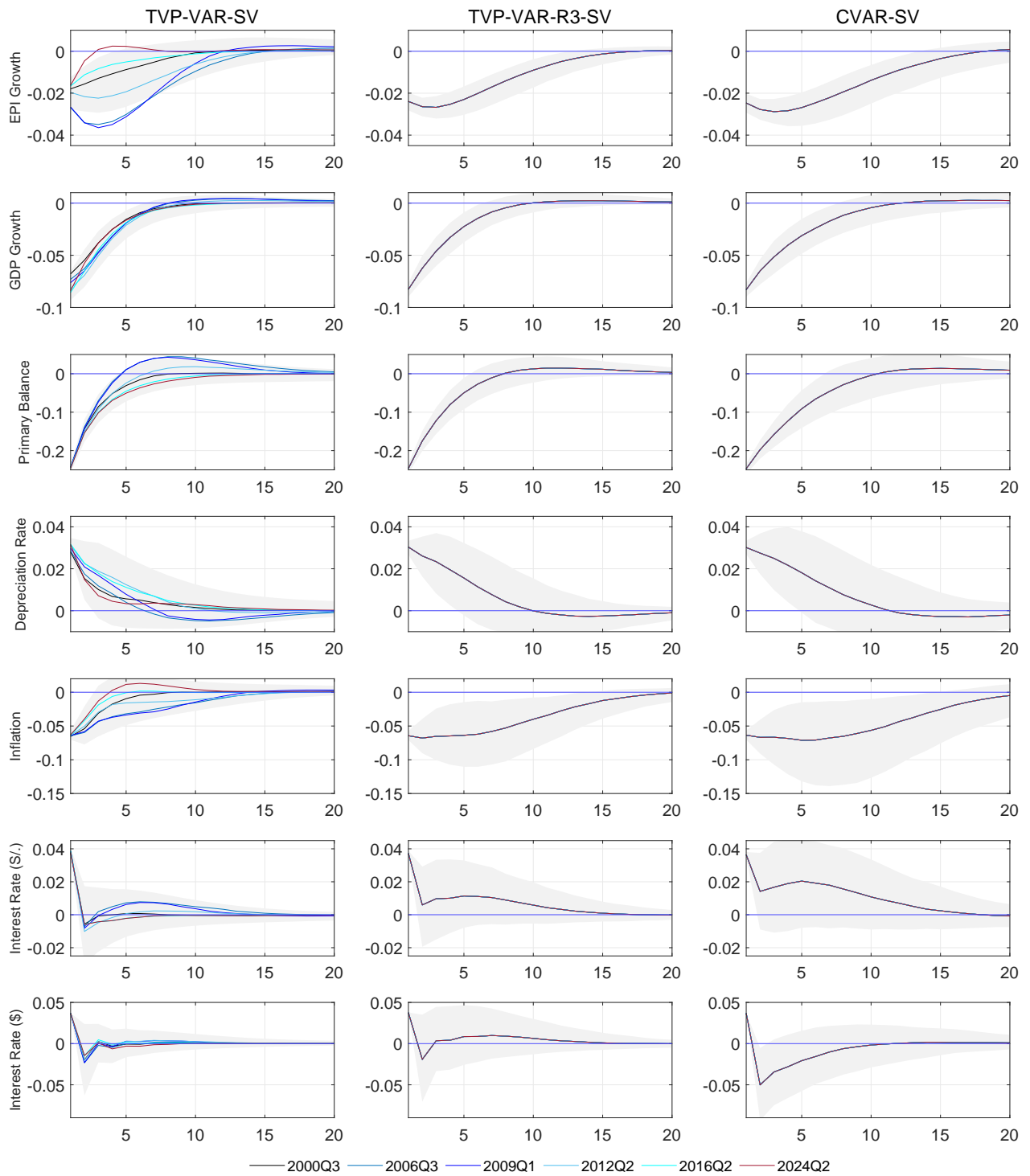


Figure 11. Quarter-Specific Median GIRFs of Gross Debt (% of GDP) to Structural Shocks

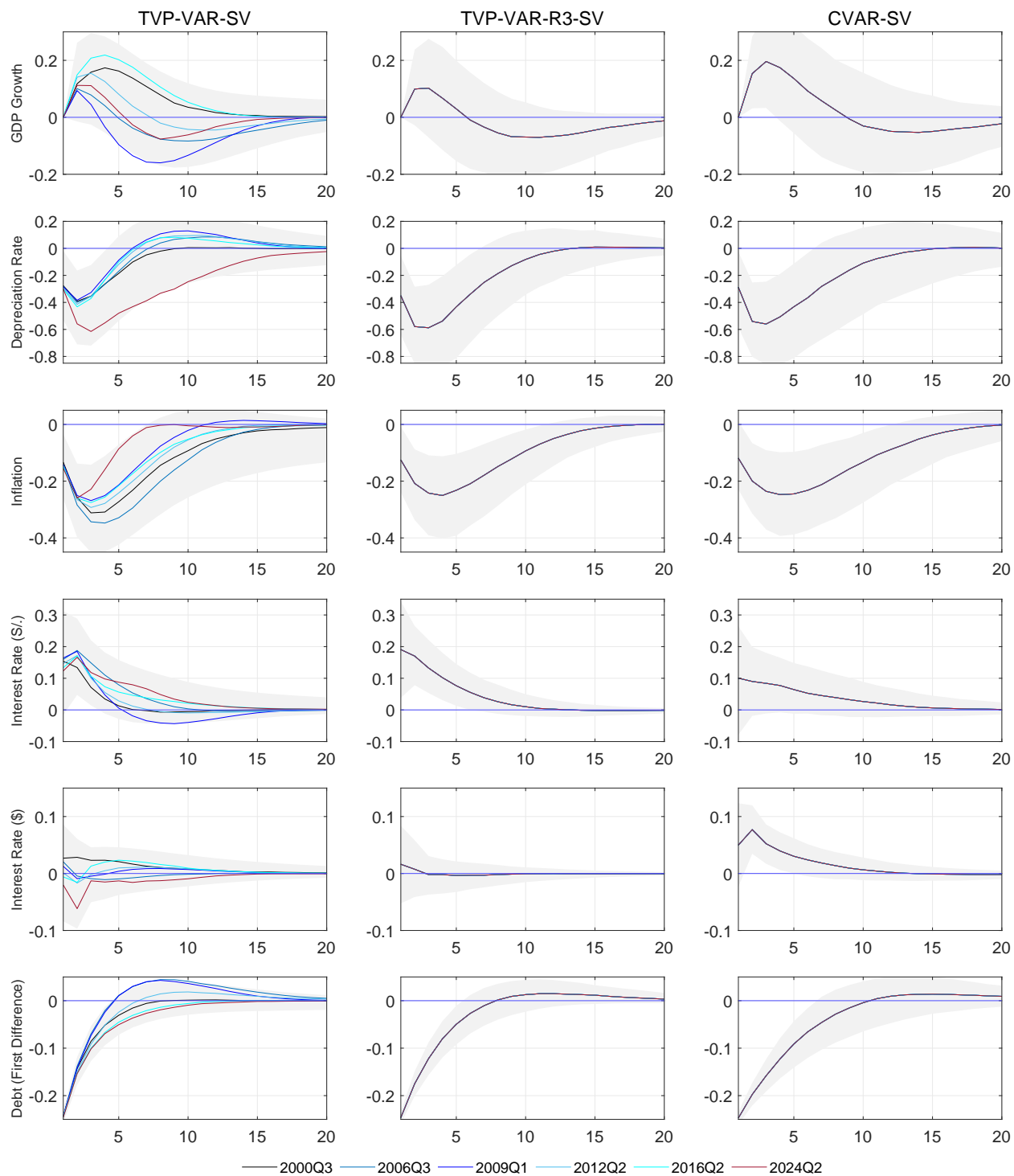


Figure 12. Quarter-Specific Median GIRFs of Model Variables to Austerity Shocks

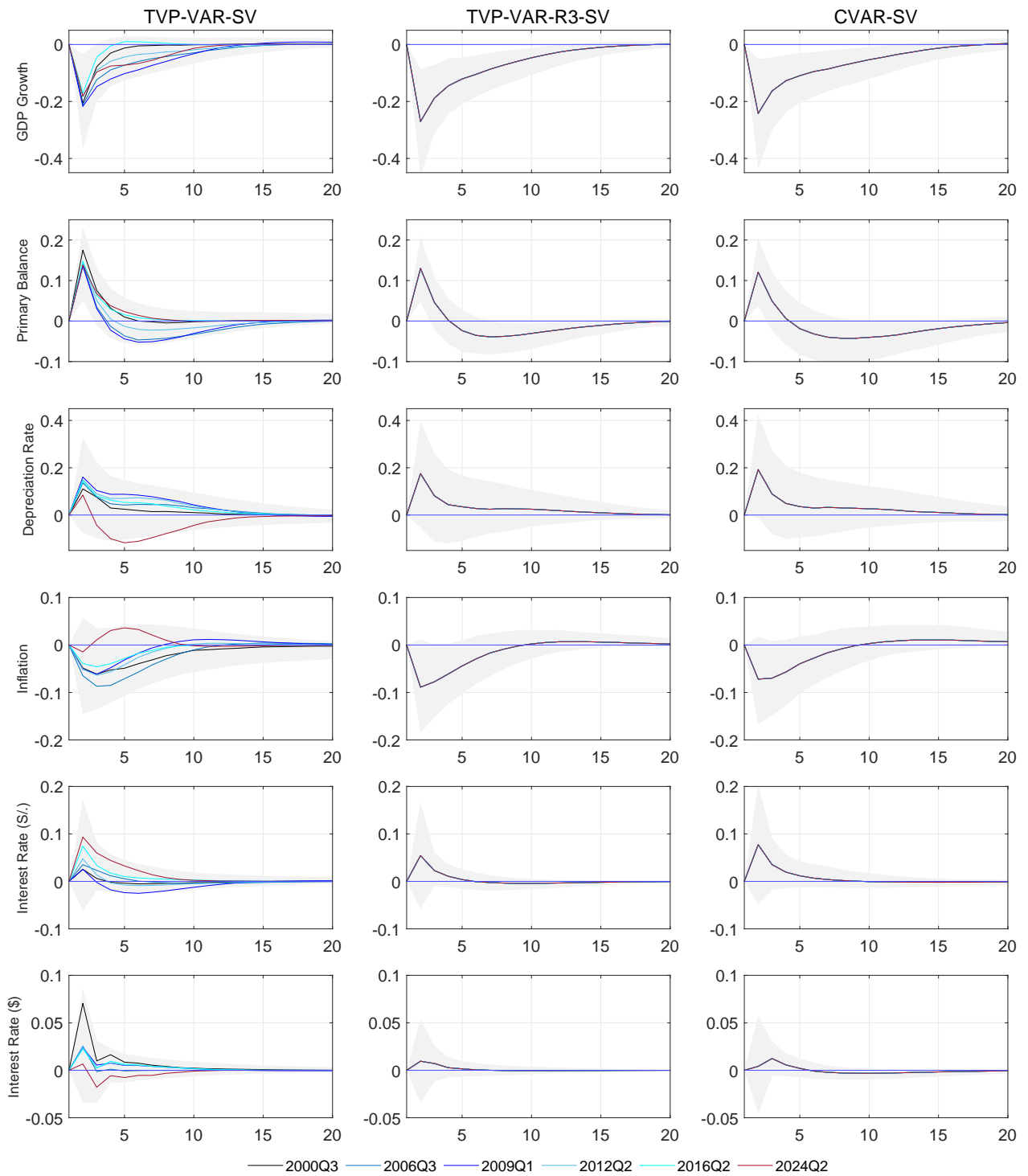


Figure 13. Quarter-Specific Median GIRFs of Model Variables to the First Difference of Debt Shocks

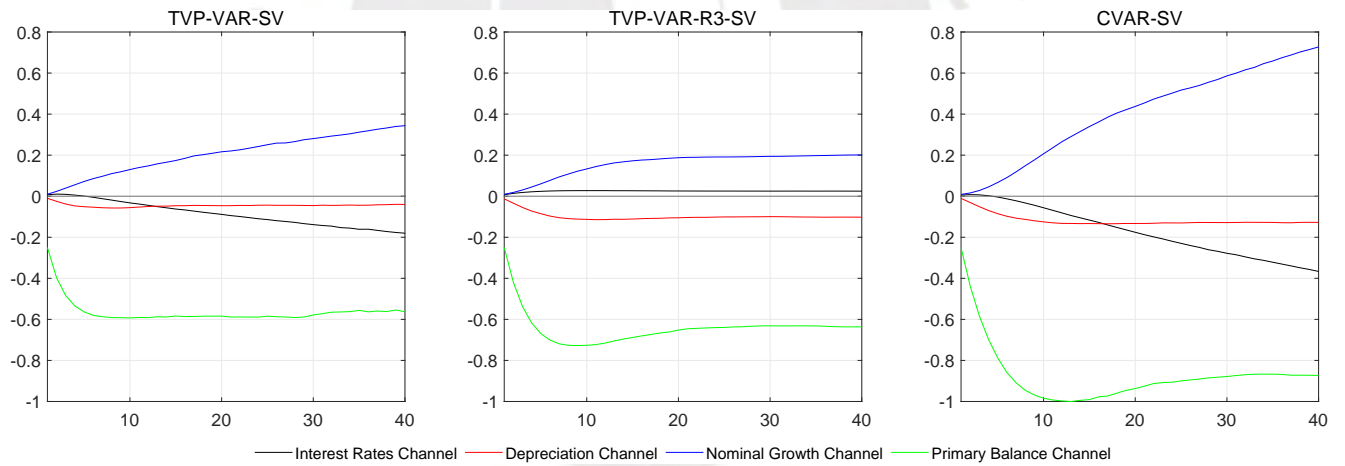
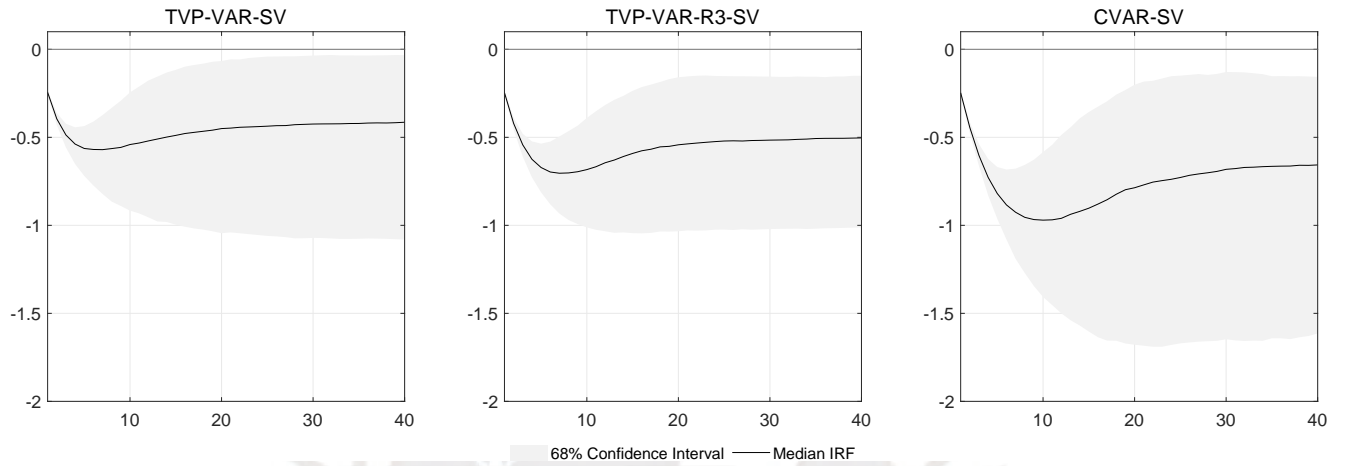
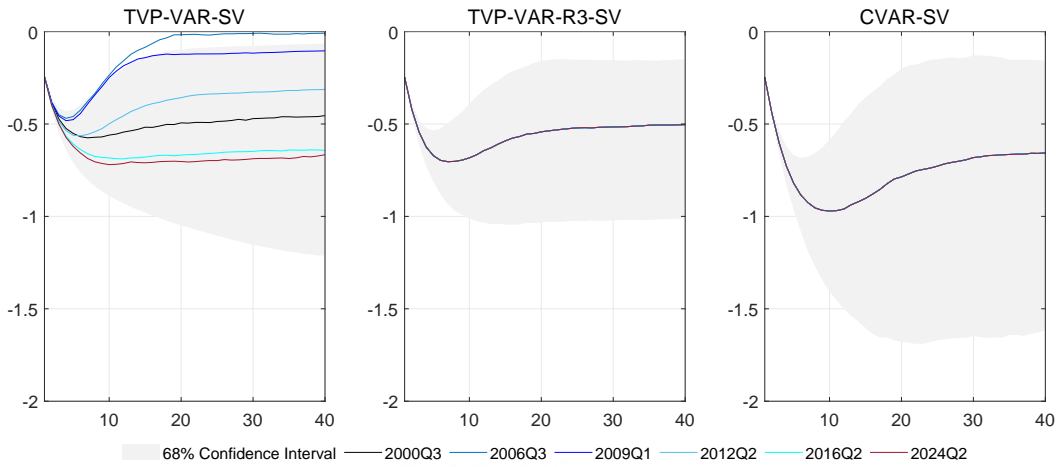
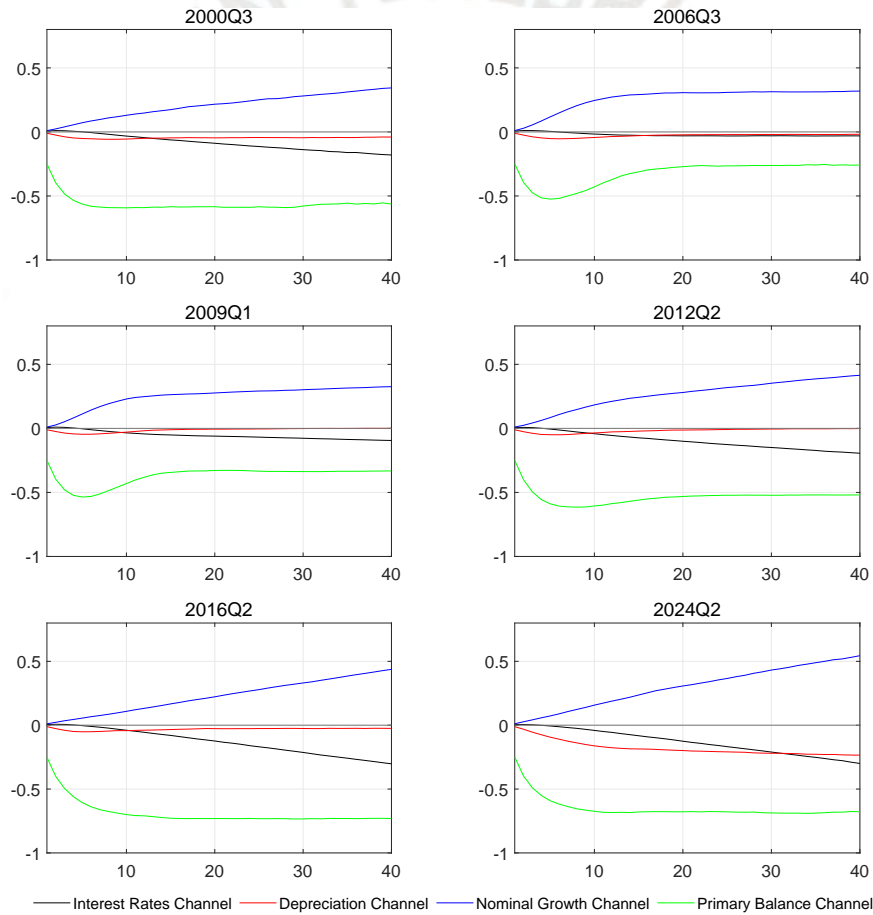


Figure 14. Effect of Austerity Shocks on the Debt-to-GDP Ratio

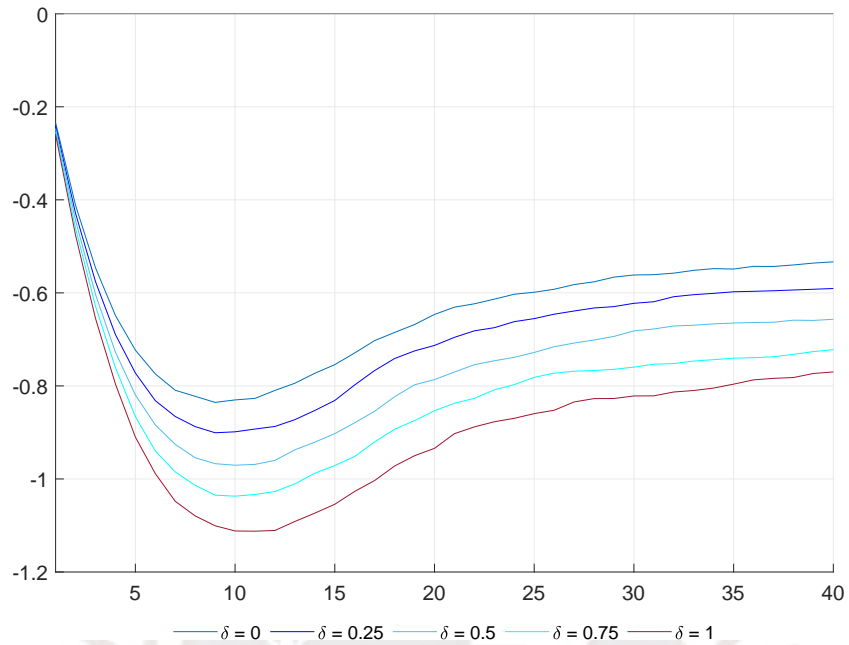


Panel A. Quarter-Specific Median Cumulative Response of the Debt-to-GDP Ratio to Austerity Shocks

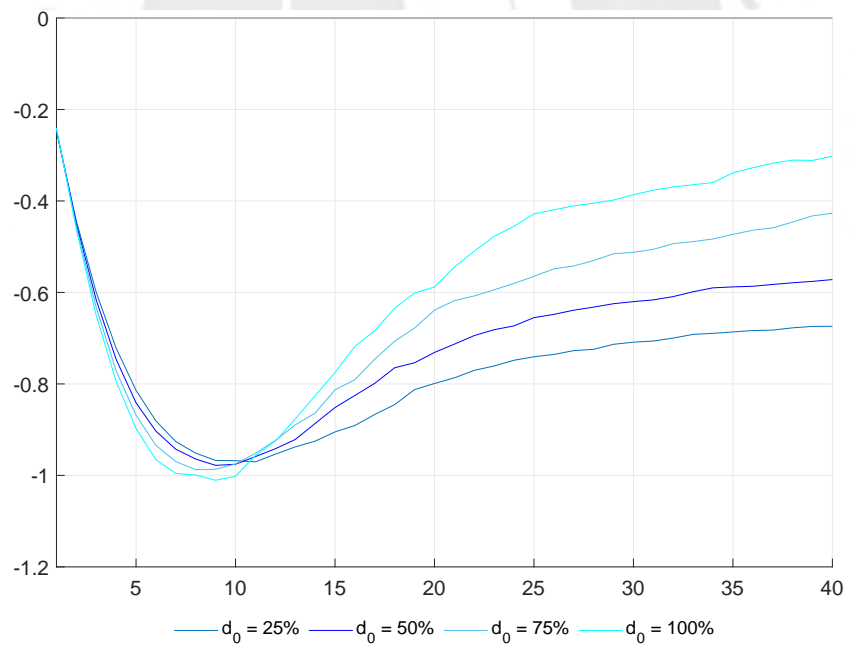


Panel B. Decomposition of the Cumulative Median GIRF of the Debt-to-GDP Ratio to an Austerity Shock in the TVP-VAR-SV Model

Figure 15. Time Varying Effects of Austerity Shocks on the Debt-to-GDP Ratio

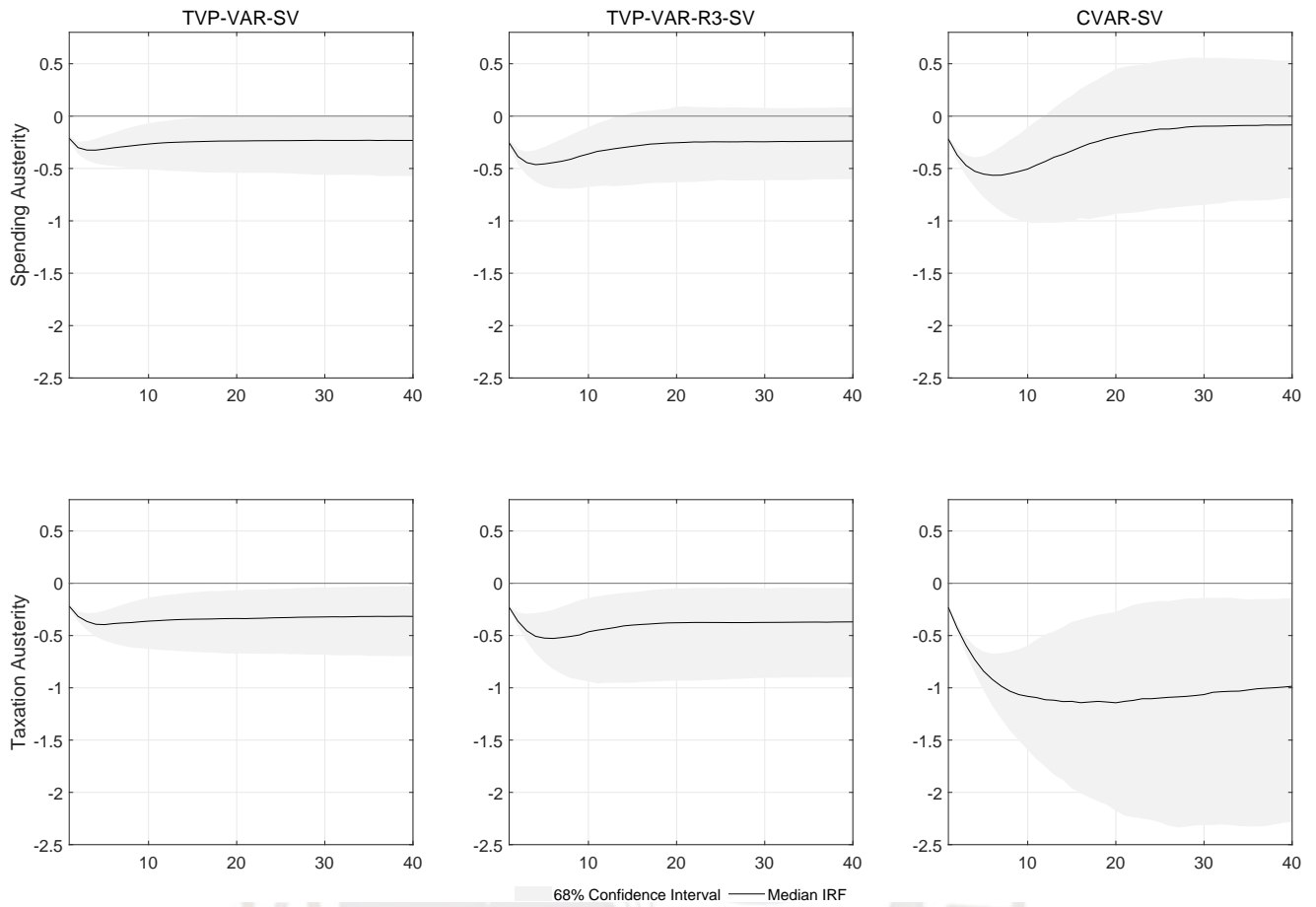


Panel A. Dollarization Level Sensitivity Analysis in the CVAR-SV Model

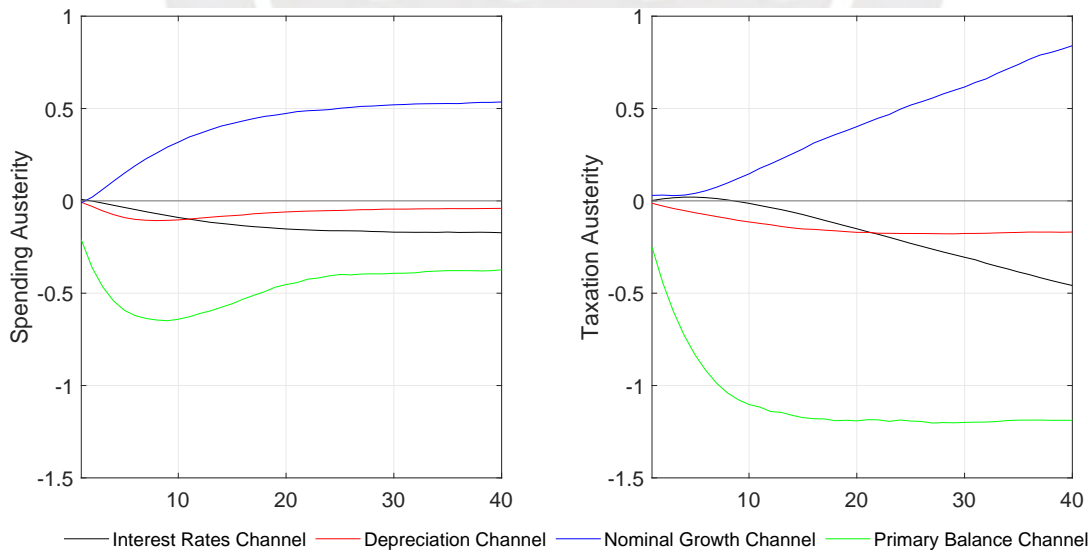


Panel B. Initial Debt Level Sensitivity Analysis in the CVAR-SV Model

Figure 16. Sensitivity Analysis for the Effect of Austerity Shocks on the Debt-to-GDP Ratio



Panel A. Cumulative Response of the Debt-to-GDP Ratio to Spending and Taxation Austerity Shocks



Panel B. Decomposition of the Cumulative GIRF of the Debt-to-GDP Ratio to Spending and Taxation Austerity Shocks in the CVAR-SV Model

Figure 17. Effect of Austerity Shocks on the Debt-to-GDP Ratio

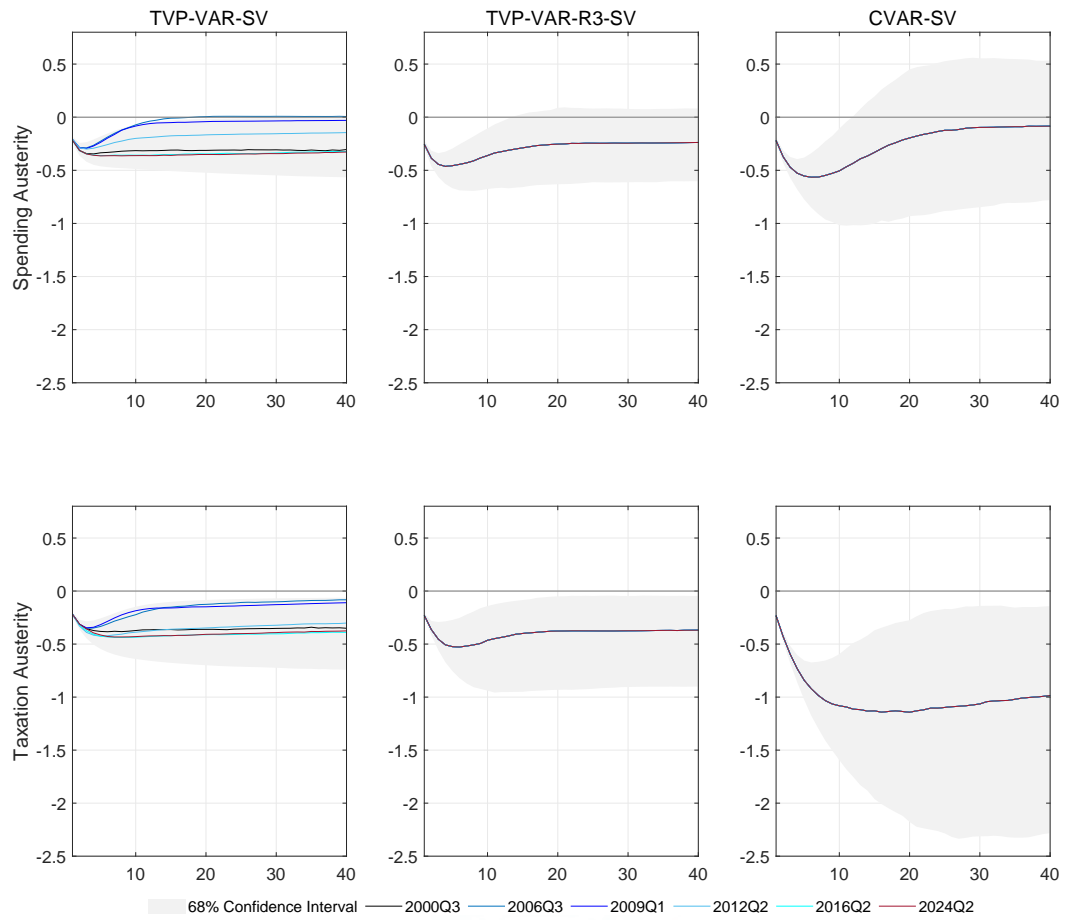


Figure 18. Cumulative Response of the Debt-to-GDP Ratio to Spending and Taxation Austerity Shocks

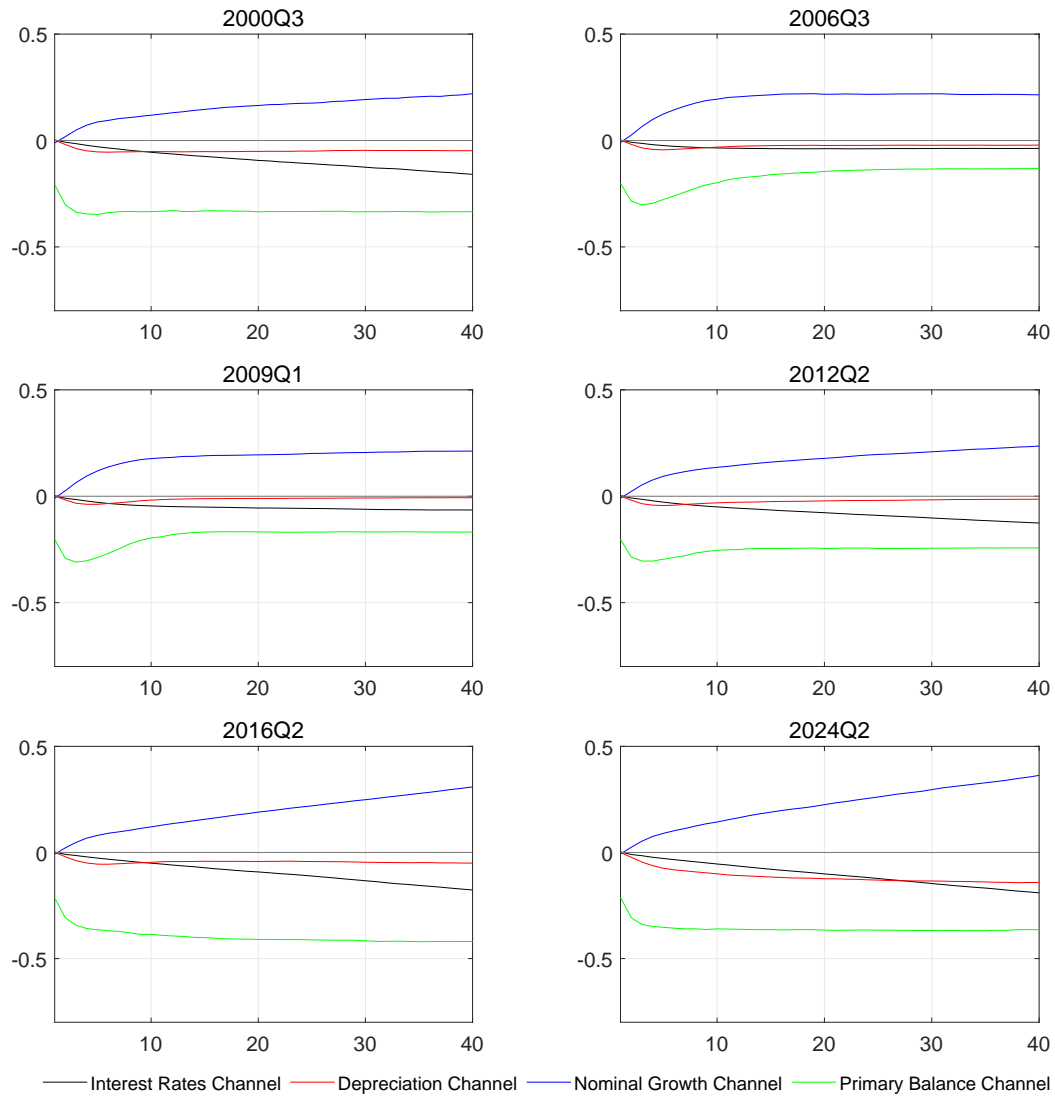


Figure 19. Decomposition of the Cumulative GIRFQ of Debt to Spending Austerity Shocks in the CVAR-SV Model

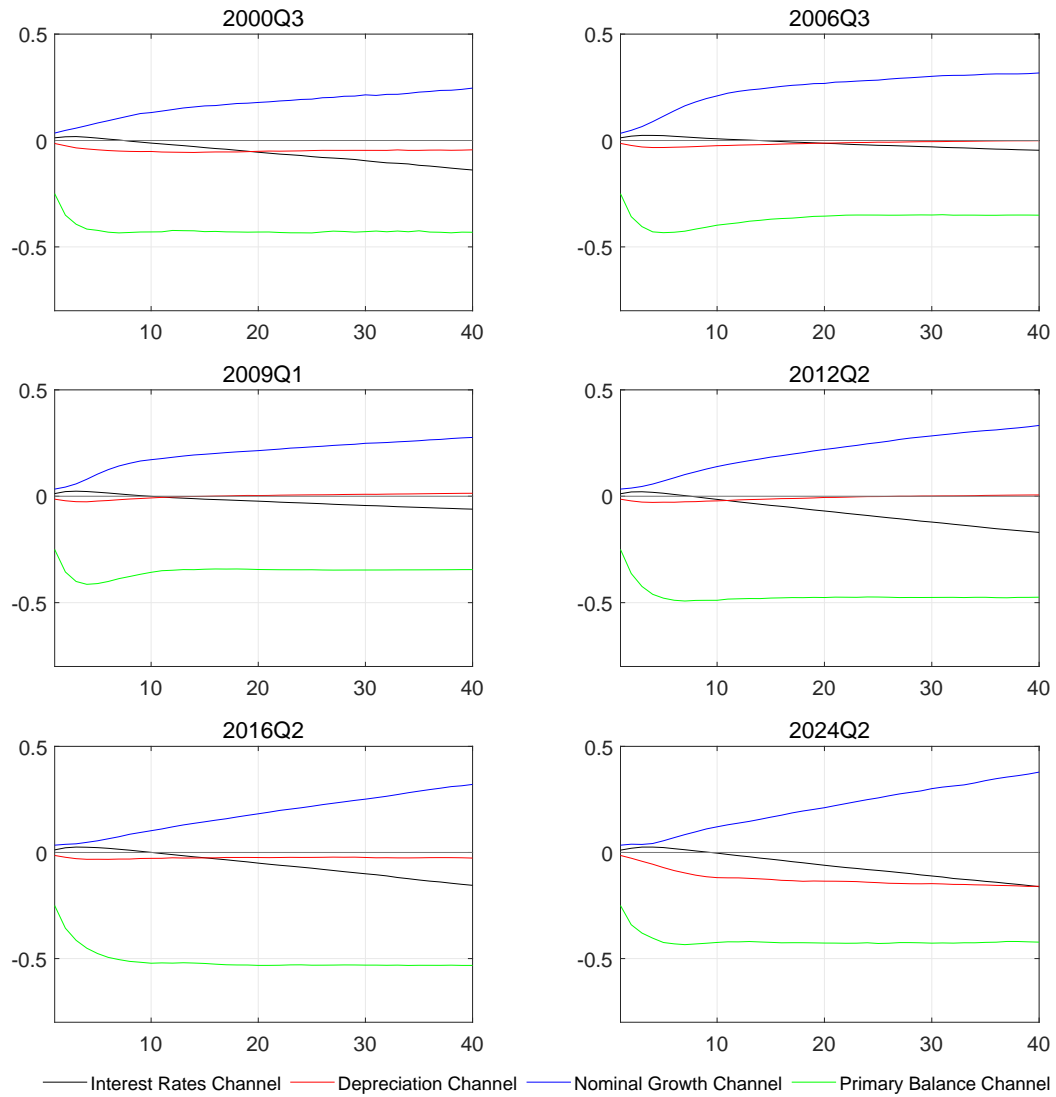
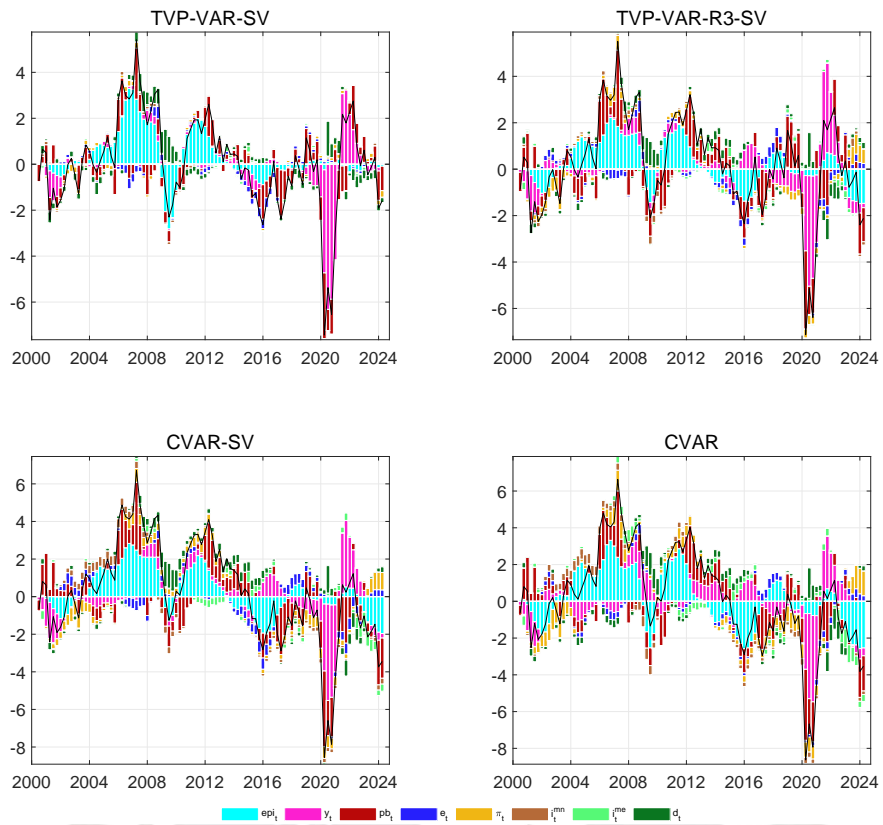
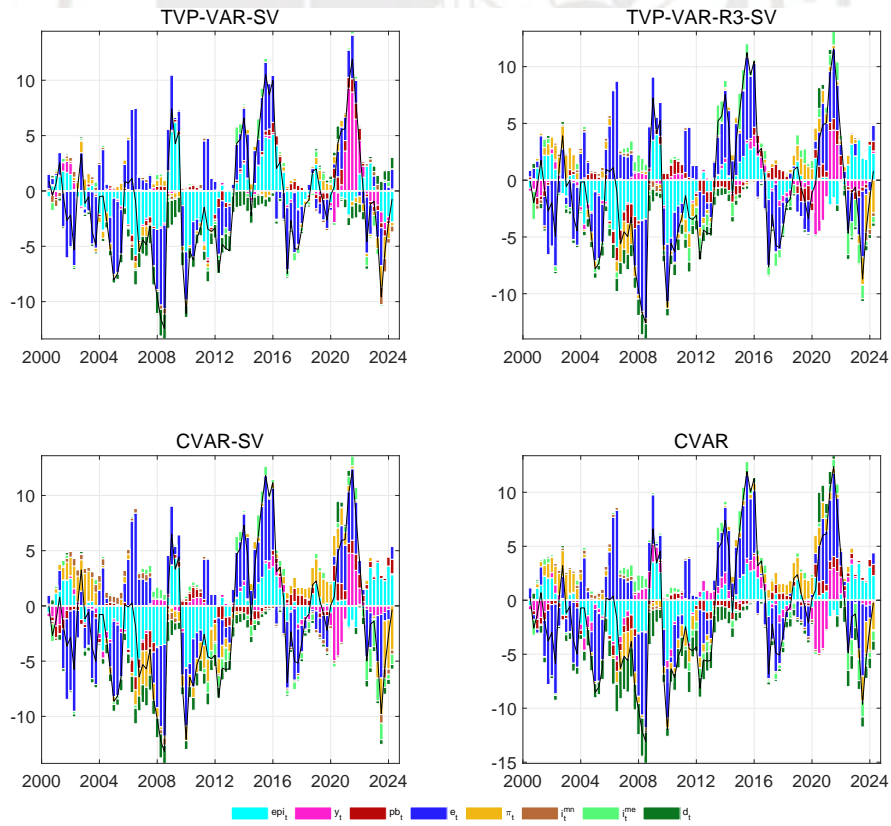


Figure 20. Decomposition of the Cumulative GIRF of Debt to Taxation Austerity Shocks in the CVAR-SV Model



Panel A. Primary Balance (% of GDP)



Panel B. Depreciation Rate

Figure 21. Historical Decomposition of Primary Balance and the Depreciation Rate

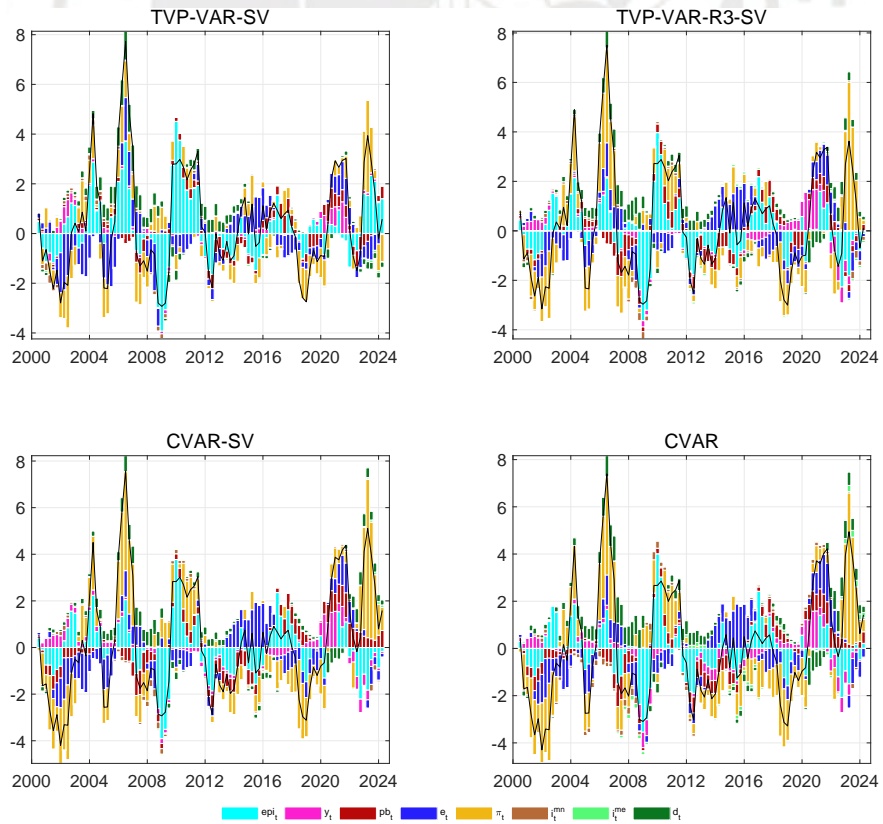
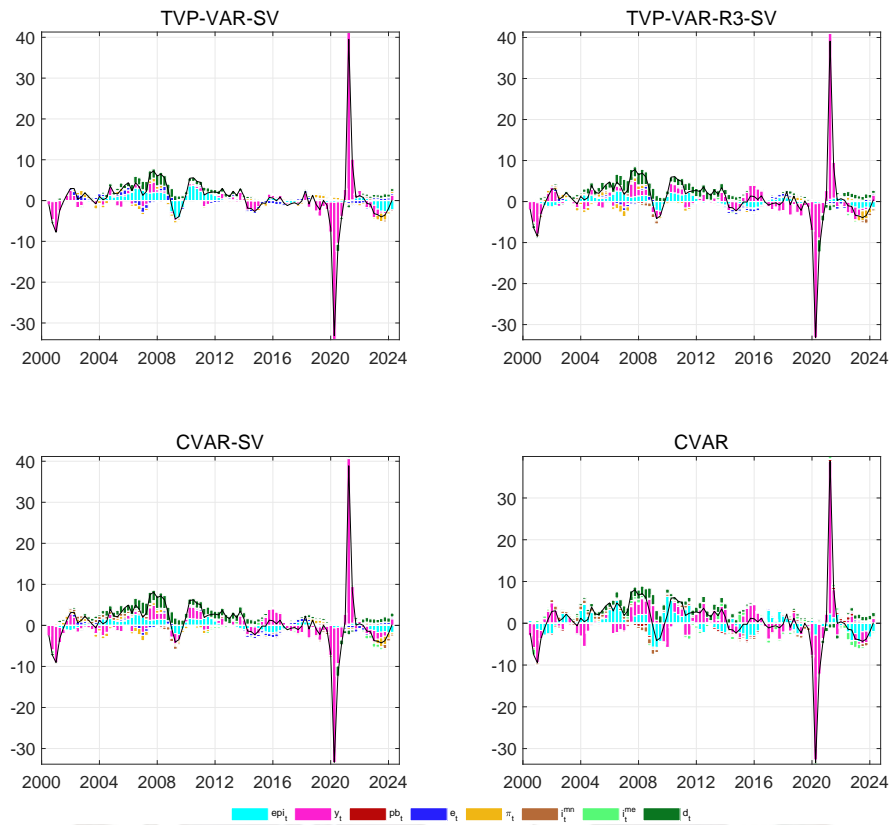
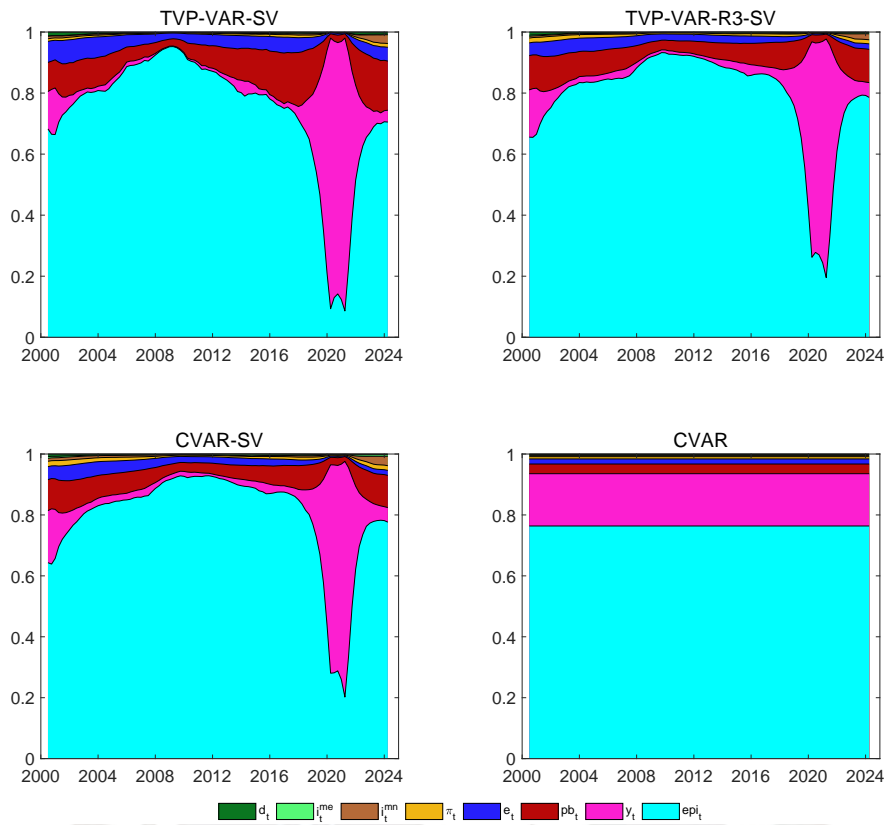
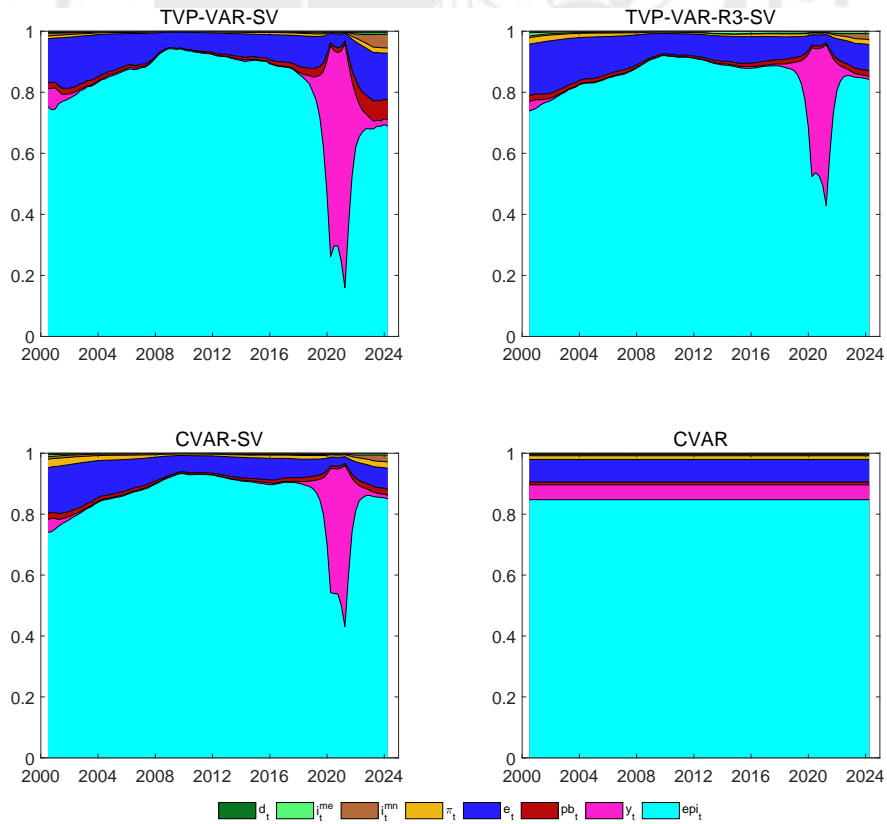


Figure 22. Historical Decomposition of GDP Growth and the Inflation Rate

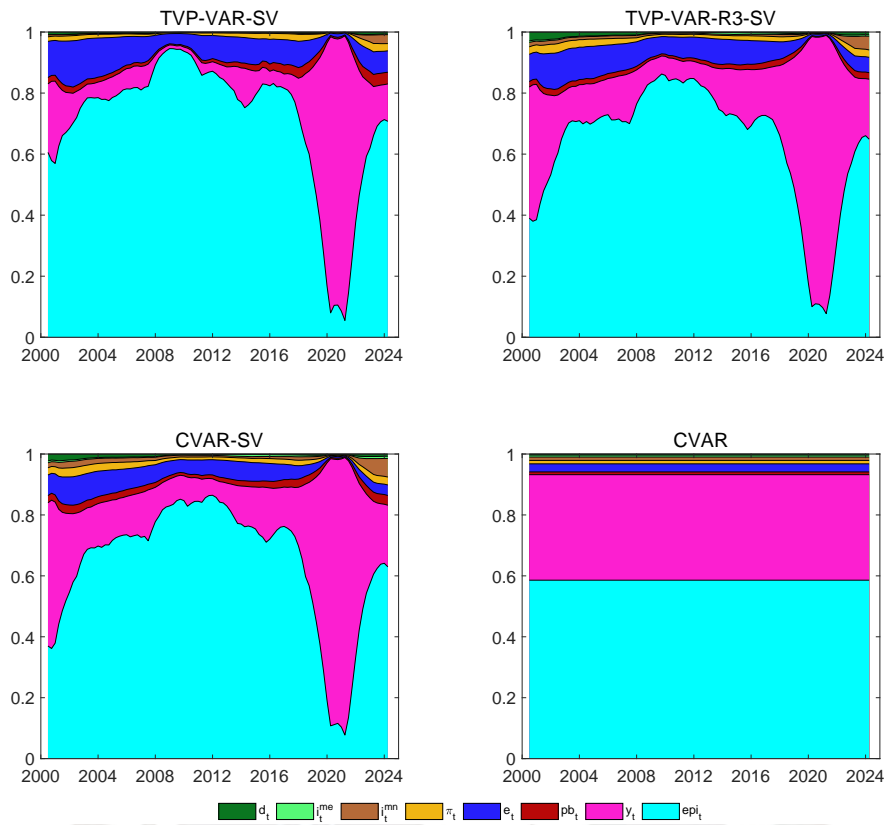


Panel A. Primary Balance (% of GDP)

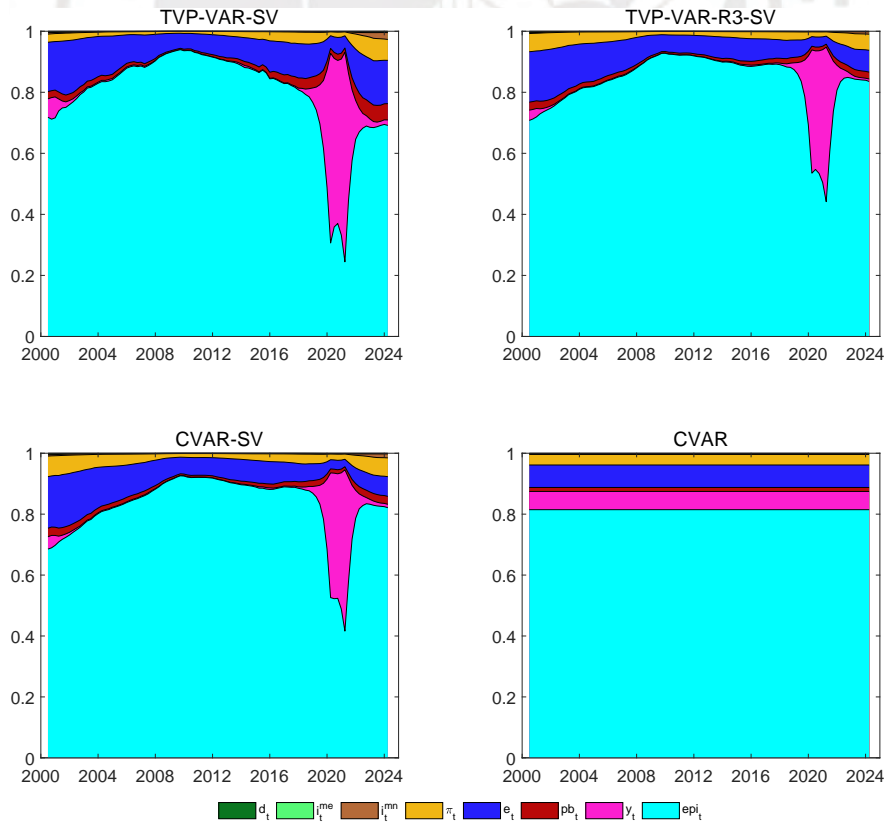


Panel B. Depreciation Rate

Figure 23. FEVD of Primary Balance and the Depreciation Rate



Panel A. Real GDP Growth



Panel B. Inflation Rate of the GDP Deflator

Figure 24. FEVD of GDP Growth and the Inflation Rate

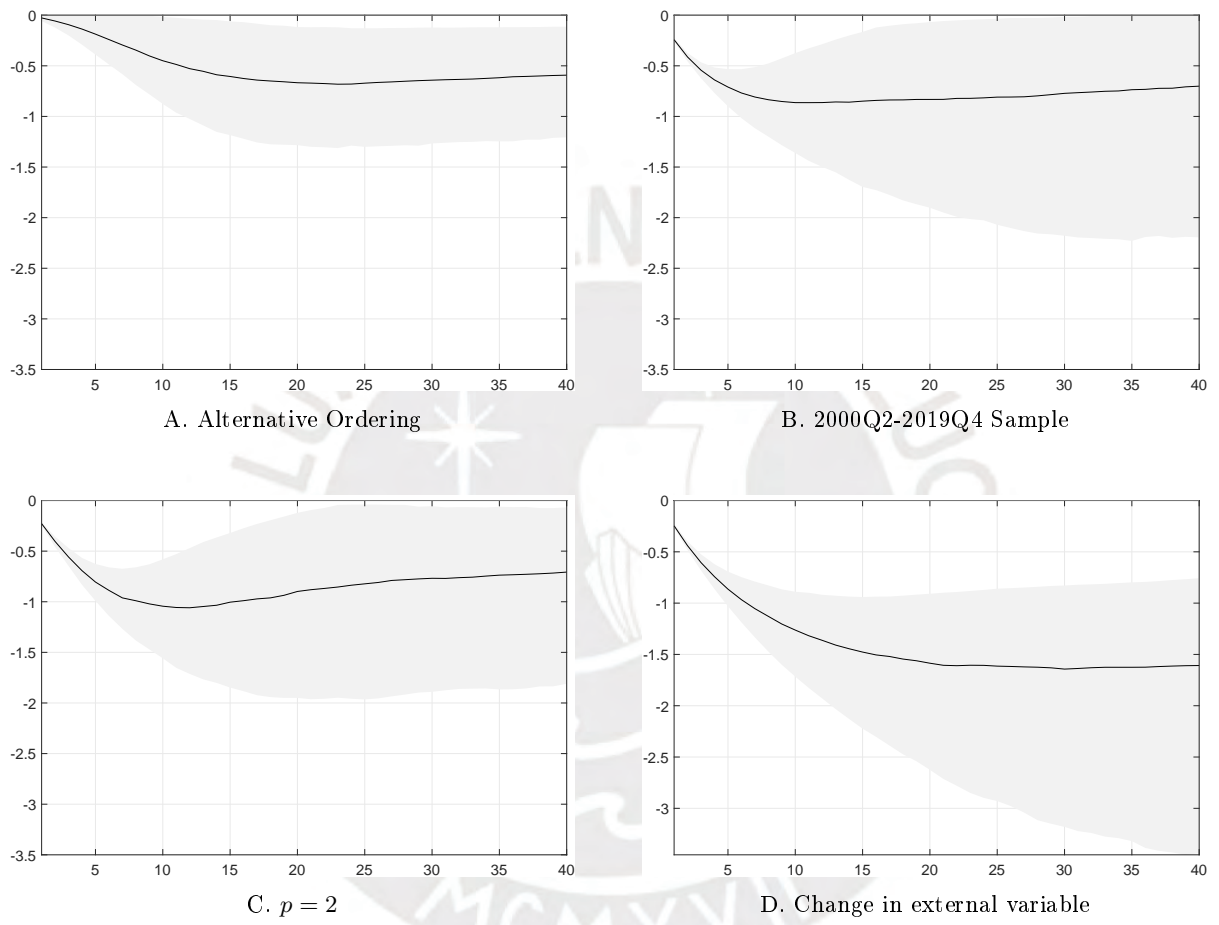
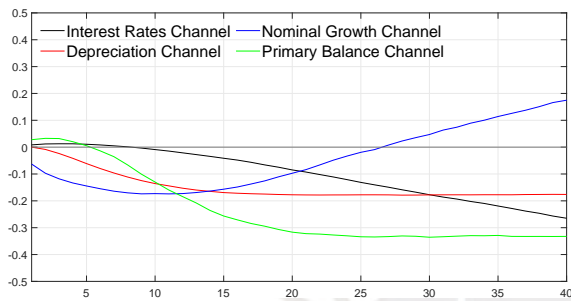
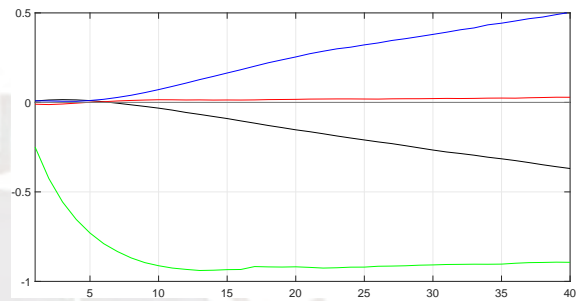


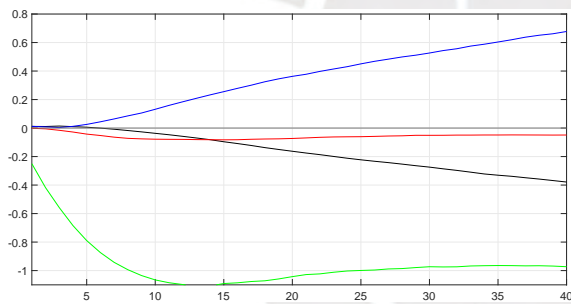
Figure 25. Sensitivity Analysis for the Cumulative Median GIRF of the Debt-to-GDP Ratio to Austerity Shocks in the CVAR-SV Model



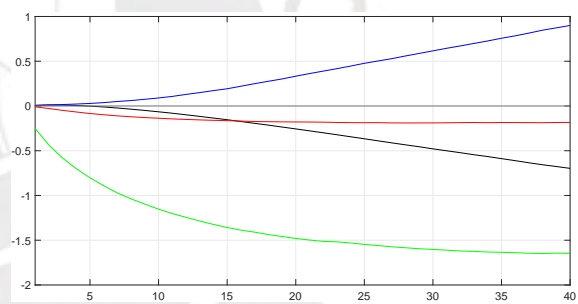
A. Alternative Ordering



B. 2000Q2-2019Q4 Sample



C. $p = 2$



D. Change in External Variable

Figure 25. Sensitivity Analysis for the Decomposition of the Cumulative Median GIRF of the Debt-to-GDP Ratio to Austerity Shocks in the CVAR-SV Model