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THE INFLUENCE OF GLOBAL UNCERTAINTY AND FINANCIAL SHOCKS, AND SOVEREIGN RISK SHOCK ON THE BRAZILIAN TERM STRUCTURE OF INTEREST RATE

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THE INFLUENCE OF GLOBAL UNCERTAINTY AND FINANCIAL SHOCKS, AND SOVEREIGN RISK SHOCK ON THE BRAZILIAN TERM STRUCTURE OF INTEREST RATE

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SUMÁRIO

ABSTRACT

Global and sovereign risk shocks significantly influence the business cycle in emerging markets. We examine their impact on the nominal and real term structure of interest rates (TIR) and the respective inflation risk premium (*irp*)) using a SVAR model for Brazil that also includes key macroeconomic variables. An adverse global uncertainty shock steepens both nominal and real TIR by reducing short-term yields, while *irp* shows less responsiveness. A positive shock to the US 3-year rate (*us*3*yr*) elevates nominal and real TIR but flattens their slopes due to a lesser increase at longer maturities; meanwhile, *irp* rises and becomes steeper. An adverse sovereign risk shock similarly pushes nominal and real TIR, and *irp*, upward, increasing their slopes. The sign of the covariance of *irp* with economic activity and inflation is shock-dependent, as is the relationship between the covariance of these variables and *irp*. Global uncertainty shocks explain approximately 22% of the forecast error variance (FEV) for 1-year real rate, being less impactful for longer maturities, nominal rates, and *irp*. Shocks to *us*3*yr* account for at least 25% of the FEV for nominal and real rates, and *irp*. Sovereign risk shocks also contribute substantially for FEV of nominal and real yields, and *irp*.

Keywords: Term structure of interest rate, inflation risk premium, sovereign risk, uncertainty, US interest rate, SVAR.

JEL Codes: C32, E43, E44, E47, F41, G15

RESUMO

Choques globais e de risco soberano influenciam significantemente o ciclo de negócios de países emergentes. Examinamos seus impactos na estrutura a termo da taxa de juros (ETTJ) nominal e real e na curva de prêmio de risco inflacionário (*irp*) do Brasil a partir de um modelo SVAR que também contém importantes variáveis macroeconômicas. Um choque adverso na incerteza global eleva a declividade da ETTJ nominal e real por reduzir as taxas de curto prazo, enquanto o risco inflacionário mostra-se invariável ao choque. Um choque positivo na taxa de 3 anos de títulos do governo americano (*us3yr*) desloca a ETTJ nominal e real para cima, deixando-as também menos inclinadas em função do menor aumento das taxas mais longas. A estrutura a termo do risco inflacionário também desloca para cima, mas fica mais inclinada. Um choque adverso de risco soberano move as três curvas para cima e eleva suas declividades. O sinal das covariâncias do *irp* com atividade econômica e com a inflação depende da natureza do choque, o mesmo ocorrendo com a relação entre a covariância das duas últimas variáveis com o *irp*. Choques de incerteza global explicam aproximadamente 22% da variância do erro de previsão (FEV) da taxa real de 1 ano, sendo menos relevante para maturidades mais longas, taxas nominais e *irp*. Choques em *us3yr* explicam aproximadamente 25% da FEV das taxas nominais, reais e do *irp*. Choques no risco soberano também explicam parcela relevante do FEV destas variáveis.

Palavras-chave: Estrutura a termo da taxa de juros, prêmio de risco de inflação, risco soberano, incerteza, taxa de juros dos EUA, SVAR.

1. INTRODUCTION

Sovereign risk shocks and global economic fluctuations play significant roles in shaping the business cycles of emerging economies. Ferreira and Valério (2022) estimate that these factors jointly account for 70%-84% of the forecast error variance (FEV) of GDP in Brazil, Chile, Colombia, and Peru over a 16-quarter horizon, with proportions exceeding 50% for key economic indicators such as the CPI, nominal exchange rate, and monetary policy interest rate. Similar findings are reported by Fernández, González and Rodriguez (2018). Given their substantial impact, it is reasonable to expect these shocks to be important drivers of movements in the term structure of interest rates (TIR).

We study how shocks to sovereign risk, global economic uncertainty, and global finance influence the nominal TIR and two of its components: the real interest rate and the inflation risk premium (*irp*). Instead of relying on factors, our investigation focuses directly on yields for 1, 3, and 4 years, providing insights into the joint effects on the short and intermediate segments of the yield curve at the business cycle frequency. Our focus is on Brazil, one of the largest emerging economies, with highly liquid and well-functioning government bond markets.

Despite considerable interest in understanding the relationship between yield curves and macroeconomic variables, there is a gap in the empirical literature regarding joint analyses of the influence of structural shocks on the macroeconomy, nominal TIR, and its components.[1] We contribute by proposing a VAR model to pursue this unified analysis, which allows us to present results on the following: i) joint responses of key macroeconomic variables and the three term structures (nominal and real yield curve, and the respective inflation risk premium curve) to important shocks affecting an emerging economy; ii) the FEV decomposition of the yields and premiums across maturities; and iii) the sign of conditional covariances, which, despite of playing a central role in general equilibrium asset pricing models, is not much explored for emerging economies, even though they tend to suffer more intensely from flight-to-safety behavior triggered by international and domestic shocks.[2]

In our approach, we focus on a pure econometric analysis that takes yields and inflation premiums as given, incorporating them into a Vector Autoregression (VAR) framework estimated using the Bayesian techniques developed by Sims and Zha (1998) and refined by Zha (1999) for cases involving an exogenous recursive block. This methodology is particularly suited for modeling small open economies such as Brazil. Unlike other macro-finance studies, such as Bekaert, Engstrom and Ermolov (2021) and Moench and Soofi-Siavash (2022), which often rely on factor models to summarize information about term structures and the state of the macroeconomy, we work directly with observable rates and premiums across various maturities. This direct approach allows for a clearer understanding of the dynamics at specific points along the short and medium segments of the yield curve. Our approach is facilitated by the fact that we are not interested in studying shocks targeted to the curves, a common practice aimed at uncovering macroeconomic shocks from innovations in factors, as seen in Moench

[1] Beauregard et al. (2024) develop and estimate an arbitrage-free Gaussian asset pricing model for Mexico, also incorporating nominal and real yields. Their goal, however, is not to evaluate response to shocks, but to study the behavior and determinants of inflation expectation and risk premia after excluding the effects of liquidity in nominal and real markets, which they estimate.

[2] According to Carrière-Swallow and Céspedes (2013), the contractionary effect of a global uncertainty shock is stronger in emerging economies, possibly due to less developed financial sectors.

and Soofi-Siavash (2022). Instead, we focus on shocks that can be easily identified as exogenous to the term structures, simplifying the establishment of precedence in the SVAR.

A notable disadvantage of our approach is the high dimensionality of the VAR model, which reduces the degrees of freedom and makes the inclusion of additional vertices and variables almost prohibitive, particularly given our small sample size, which consists of monthly observations from September 2009 to June 2018. Nonetheless, the probability intervals we estimate still ensure the significance of most impulse response functions, despite occasionally observing very large intervals. Another disadvantage is the necessity to compare the exact response at each vertex to verify the effects on the level, slope, and curvature. However, this limitation provides the advantage of allowing a detailed understanding of each vertex's contribution to movements in these factors. For example, we find that a negative global uncertainty shock steepens the Brazilian nominal and real term structure by more intensely reducing short-term rates. The nominal curve steepens further because the nominal 4-year maturity yield does not react significantly, while the 4-year real yield decreases. A positive innovation to the US Treasury 3-year constant maturity yield (*us*3*yr*) increases the level of the nominal and real TIR but flattens the former due to a more intense rise at the short end, while the real TIR becomes steeper due to a higher increase at longer maturity. Additionally, the level of *irp* rises, and its curvature increases. In the event of an adverse sovereign risk shock, the levels of both nominal and real yield curves rise, with the nominal term structure becoming more convex due to a higher increase in the 3-year yield, while the real term structure steepens. The inflation premium also steepens, with an insignificant move of the 1-year *irp* and a substantial rise at the 4-year vertex, which may reflect uncertainties about future inflation due to fiscal challenges.

Our VAR uses variables regularly employed in canonical pricing models for the term structure of interest rates and in neo-Keynesian DSGE models for small open economies. In particular, our empirical macro-finance model incorporates the following domestic variables: the sovereign spread, nominal exchange rate, CPI, a proxy for monthly GDP, and the monetary policy interest rate, besides the term structure variables. The proposed structure allows comparisons with results from the macro-finance literature, alongside a better comprehension of the general equilibrium channels operating through the economy. The responses to shocks only partially support previous evidence of a negative relation between *irp* and economic activity, as reported by Ang, Bekaert and Wei (2008), Breach, D'Amico and Orphanides (2020), among others. We verify that this negative relation is shock-dependent, being verified in the presence of a sovereign risk shock, mildly observed for 1 and 3-year *irp* when the world economy is exposed to an uncertainty shock, and not present at all when a positive shock to *us*3*yr* strikes, since the Brazilian inflation premiums increase while economic activity does not move significantly. These shock-dependent relationships are consistent with the responses of economic activity and inflation moving, which move in different directions in different directions following shocks to global uncertainty and sovereign risk. This pattern should justify a positive *irp* according to standard pricing models. However, neither economic activity nor the price level is affected by a positive innovation in *us*3*yr*. This is despite a significant increase in *irp* across all maturities studied, suggesting that other economic factors may influence the determination of *irp*.

Our work also communicates with the extensive literature on the influence of global finance on emerging economies. As observed by Akinci (2013), innovations in the global interest rate only modestly influence the business cycle of emerging economies after controlling for global financial risk factors, which also motivates our interest in disentangling the impact of shocks to world economic uncertainty and the world interest rate. We operate this differentiation by incorporating the implied volatility index VIX and the US Treasury 3-year nominal yield in the international exogenous bloc of our global VAR. Along this line, Ferreira and Valério (2022) show the necessity to filter the influence of global variables from the sovereign spread if one intends to understand its impact domestically. This need becomes clear in our results, as the Brazilian sovereign risk significantly responds to global shocks, especially innovations to global uncertainty, responsible for around 35% of the FEV of the Brazilian sovereign risk according to our estimation.

Our investigation also communicates with some empirical works on macro-finance applied to Brazil. Lowenkron and Garcia (2007), Montes and Curi (2017), and Reis (2018) concentrate their analyses on the inflation risk premium, which they find to positively correlate with fiscal uncertainty, a result aligned with our estimated response of *irp* to a sovereign risk shock and its contribution to the FEV of *irp*, especially at longer maturity, around 20%. The first two works also report a positive correlation between *irp* and monetary policy credibility, which we do not evaluate.[3] Exploring other relations, Vicente and Graminho (2015) encounter a significant correlation between the *irp* of 3 and 4-year expiring nominal bonds with the covariance between consumption variation and future inflation. We detect a similar pattern when the system is hit by innovations to world uncertainty and sovereign risk, but not when exposed to shocks to *us*3*yr*.[4]

The subsequent sections of this paper are organized as follows. Section 2 discusses the data. In Section 3, we detail the econometric methodology, including the restrictions to identify the SVAR. Sections 4 and 5 present the analyses of IRFs and Forecast Error Variance, respectively. Finally, Section 6 provides concluding remarks.

2. DATA

We start presenting the variables related to the term structure, moving later to macroecomic variables.

2.1 The term structure of the interest rate

We use the nominal and real term structure of interest rate estimated daily by the Brazilian Association of the Entities in the Capital and Financial Market (ANBIMA)[5] following Svensson (1994). The data used in the estimation correspond to real and nominal future rates on government securities traded in the secondary market from September/21/2009 to June/18/2018[6]. Nominal rates are from two different securities: *Letras do Tesouro Nacional* (LTN), a nominal zero-coupon bond, and *Notas do Tesouro Nacional série F* (NTN-F), which pays coupon every semester but had its price adjusted to

3 Fernandes and Thiele (2015) is another *pure econometric* study applied to Brazil aiming to ascertain the relationship between local and global macroeconomic variables with the term structure of Brazilian break-even inflation (*brei*), while we focus on nominal and real TIR and inflation premium.

4 Our results are only partially comparable, as we do not rely on aggregate consumption but on a measure of economic activity.

5 ANBIMA gently made available the term structure they estimate.

6 The dates are restricted by the information made available to us by ANBIMA when this project started.

eliminate the effect of coupons from prices. The real rates come from the *Notas do Tesouro Nacional - série B* (NTN-B), an inflation protected government security that pays the inflation accrued until the maturity plus a real yield known at the time the bond is acquired[7].

Let $r^N_{t,\tau}$ and $r^R_{t,\tau}$ represent, respectively, the nominal and the real yields of a bond bought at time t and maturing τ years ahead. We consider τ = 1,3,4, or 12, 36 and 48 months ahead, respectively. The associate break-even inflation is defined as $\pi^{brei}_{t,\tau} = r^N_{t,\tau} - r^R_{t,\tau}$ and $\pi^{brei}_{t,\tau} = \pi^{exp}_{t,\tau} + irp_{t,\tau}$ where $\pi^{exp}_{t,\tau}$ is the time t expected inflation accumulated for 12 months up to τ, and $irp_{t,\tau}$ the inflation risk premium evaluated at time t.[8]

The $irp_{t,\tau}$ is obtained after using a survey that the Brazilian Central Bank (BCB) conducts with professional forecasters about expected inflation. In this survey, expected inflation is informed for each of the next 18 months and for the accumulated rate at the end of next calendar years. Our measure of expected inflation evaluated at a particular month t is constructed by averaging daily median expectations reported in month t. [9] Since we do not have forecasts for each month after the 18th, we adopt the following interpolation rule considering that we work with years τ = 1,3,4:

$$\pi^{exp}_{t,\tau} = \begin{cases} 100\left[\prod_{i=0}^{11}(1+\pi^s_{t+i}) - 1\right] & if \quad t \neq 1 \quad and \quad \tau = 1 \\ 100\left[(1+\pi^s_{t,\tau})^{\frac{13-t}{12}}(1+\pi^s_{t,\tau+1})^{\frac{t-1}{12}} - 1\right] & if \quad t \neq 1 \quad and \quad \tau = 3,4 \\ \pi^s_{t,\tau} & if \quad t = 1 \quad and \quad \tau = 1,3,4 \end{cases}$$

where s refers to rates directly observed from the surveys, and $t = 1, 2, \cdots, 12$ correspond to months of the year, with t = 1 being January, t = 2 February, and so long.

Our interest in evaluating the responses in the short and middle part of the yield curve justifies working with 1-, 3-, and 4-years maturity. The absence of surveys about expected inflation for longer time intervals restricts working with longer maturity. The absence of the 2-year maturity securities is due to degrees of freedom concerns, since our VAR already counts with a high number of variables.

Figure 1 displays the evolution of nominal and real rates, *brei*, expected inflation, and *irp*. Longer maturity yields are normally higher, a common pattern that is related to the larger premium charged to hold assets expiring at longer maturities. It is however worth observing that the difference between 3- and 4-year rates is small and even negligible in some months, suggesting that the term premium embedded in these maturities is similar. All boxes show a spike in 2015-2016, years in which

7 The inflation rate that indexes the NTN-B is the IPCA, an official price index that serves as the reference for the inflation target regime and that is computed by the Brazilian Institute of Geography and Statistics.

8 A more rigorous approach would allow for the presence another two terms in the *brei*: a convexity term due to Jensen's inequality and a liquidity premium. However, previous studies (Vicente and Graminho (2015) for Brazil, Joyce, Lildholdt and Sorensen (2010) for UK, Ang, Bekaert and Wei (2008) for USA, among others) verify that these erms are negligible.

9 Caldeira and Furlani (2013), Vicente and Graminho (2015), and Mariani and Laurini (2017) show that the median of the inflation forecasts from the survey of the BCB is a good predictor of future inflation.

Brazil lost the investment grade status according to the three major risk agencies (Fitch, Moody's, and Standard and Poor's), amid an uncontrolled increase in the public debt and high fiscal deficits. During this period, the *brei* increased substantially to values close to 9% (despite of an inflation target of 4.5%) with *irp* going close to 5%.[10]

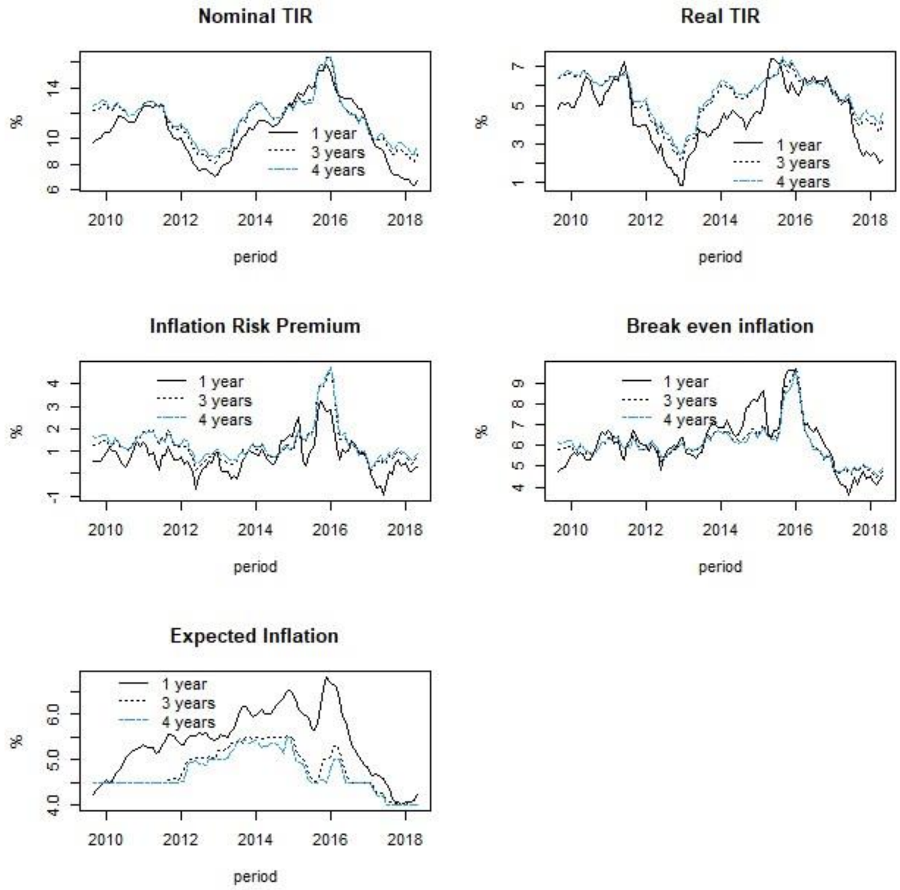


Figure 1 – Nominal interest rate and its components for contracts maturing in 1, 3, and 4 years.

Sources: *Central Bank of Brazil* for the raw data on expected inflation, which was further adapted to obtain values for 3 and 4 years ahead expectation regardless of the month of the year the BCB survey was conducted. *ANBIMA* for nominal and real interest rates, which they obtain after fitting the Svensson (1994) model. The break-even inflation equals the nominal rate minus the real rate, and the inflation risk premium is the break-even inflation minus the expected inflation.

Table 1 – Variance of the nominal yield term structure and its components.

Maturity (τ)	r_τ^N	r_τ^R	π_τ^{brei}	irp_τ	π_τ^{exp}
1	5.592	2.683	1.668	0.581	0.518
3	3.614	1.567	0.943	0.682	0.216
4	3.120	1.333	0.787	0.653	0.175

Source: Prepared by the authors.

10 See Bonomo et al. (2024) for a comprehensive analysis of the events that happened in Brazil during this period.

Table 1 shows the variances of r_{τ}^{N}, $r_{t,\tau}^{R}$, π_{τ}^{brei}, irp_{τ}, and π_{τ}^{exp} for τ = 1,3,4. The variance of *irp* is the only that increases with maturity, which is an expected pattern from a risk measure. The remaining variances fall with maturity, suggesting higher uncertainty on short-run adjustments following shocks compared to more confidence on the fundamentals that determine these variables in the medium run, a result aligned with the findings of Breach, D'Amico and Orphanides (2020) for inflation uncertainty in the US.

2.2 The other variables

We include variables from which we intend to identify shocks or that synthesize the macroeconomic structure of the economy, which correspond in a large extent to fundamentals that determine the nominal TIR and its components. From *macroeconomic structure of the economy* we mean variables that are normally used in standard DSGE models of a small open emerging economy.

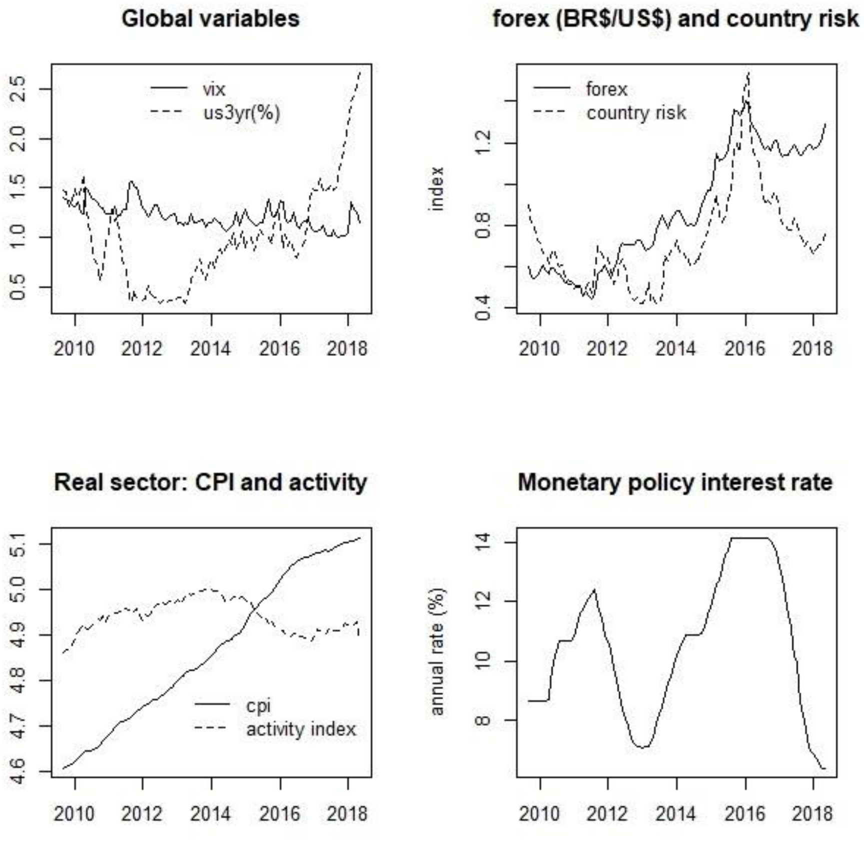


Figure 2 – Global and domestic macroeconomic variables used in the estimation: Sept/2009 - May/2018.

Sources: CBOE for daily *VIX*; FRED - FED Saint Louise for the monthly 3-Year Treasury Constant Maturity Rate (*us*3*yr*); IPEADATA for daily country risk *EMBI −BR*; Brazilian Institute of Geography and Statistics for the *CPI* measure *IPCA*; Brazilian Central Bank for the seasonally adjusted activity index (IBC-BR) and for the daily values of the SELIC, which is the Brazilian monetary authority policy interest rate.

In the domestic macroeconomic bloc, we include the Emerging Markets Bond Index for Brazil - EMBI BR+, computed by the investment bank JP Morgan Chase, as our proxy for sovereign risk[11]. The EMBI BR+ (or *BR.RISK* hereafter) is reported in terms of excess basis points over US dollar denominated sovereign bonds. We work with monthly averages of daily information available at *www.ipeadata.gov.br*. The nominal exchange rate of the Brazilian Real against the US dollar (*FX*) is included because its response to shocks is ultimately passed through to domestic prices. We use the consumer price index IPCA (*P*) computed by the Brazilian Institute of Geography and Statistics (IBGE) because it is the reference index for the inflation target regime. We include the Brazilian Activity Index (IBC-Br) computed by the BCB as a proxy for the monthly GDP (*Y*). This index intends to track the country's GDP after compiling information from monthly surveys of several economic indicators. Finally, we incorporate the Brazilian monetary policy interest rate SELIC (r^{monet} hereafter) after monthly averaging from the daily rates available at *www.ipeadata.gov.br*.

Among the global variables we include the volatility index (*VIX*) to proxy for global economic uncertainty. We use monthly averages from the daily values computed by the CBOE. The 3-Year Treasury Constant Maturity Rate (*us*3*yr*), available from the FRED database, proxies for the level of financial tightness in the world economy. The evolution of all these series can be assessed in figure 2.

3. THE ECONOMETRIC METHODOLOGY

We evaluate the responses to shocks using the following structural vector autoregressive (SVAR) model:

$$AX_t = B_0 + \sum_{h=1}^{p} B_h X_{t-h} + \varepsilon_t$$

where t is a time index; X_t is an $n\times1$ column vector formed by all n endogenous variables; ε_t is another column vector containing n structural shocks; A is an $n\times n$ matrix of instantaneous impact coefficients; B_0 is a $n\times1$ column vector of intercepts; B_h is an $n\times n$ matrix of autoregressive coefficients; and p is the autoregressive order. Structural shocks are assumed to be independent: $\varepsilon_{it} \sim id(\mathbf{0};\sigma_i^2 I_n)$, for $i = 1,...,n$. Pre-multiplying both sides by A^{-1} results in the reduced VAR:

$$X_t = \Phi_0 + \sum_{h=1}^{p} \Phi_h X_{t-h} + e_t$$

11 The relevance of sovereign risk for the business cycle and the pricing of assets is not confined to emerging economies. Itskhoki and Mukhin (2021) find important impact in real and financial variables of developed economies when using an open economy new-keyneisan DSGE model. They show that several exchange rates puzzles (disconnects) are reconcile when sovereign risk shocks are taken into consideration.

where $e_t = A^{-1}\varepsilon_t$, $\Phi_0 = A^{-1}B_0$ and $\Phi_h = A^{-1}B_h$ for $h = 0,1,\cdots p$. Because we estimate the reduced form, it is required to impose restrictions to identify the structural coefficients and shocks. Our identification strategy is presented next.

3.1 An empirical macro-finance model for Brazil

The structure of vector X is $X = [vix, us3yr, br.risk, r_1^R, irp_1, r_1^N, r_3^R, irp_3, r_3^N, r_4^R, irp_4, r_4^N, fx, y, p, r^{cb}]'$, where *vix*, *br.risk*, *fx*, *y* and *p* are natural logarithm of their capital letter counterparts. Since $n = 16$, the dimension of X, Φ_0, and e_t is 16×1, while A, B_h, and Φ_h are 16×16.

Lag Restrictions

We acknowledge the small open economy status of Brazil by imposing zero restrictions on the reduced form VAR to avoid domestic variables from affecting the dynamics of the international variables *vix* and *us*3*yr*. This is accomplished by setting $\Phi_{h,ij} = 0$ for $i = 1,2$, $j \geq 3$, for all h, where i represents a line and j a column. This restriction implies that the international variables may affect each other but are not impacted by the Brazilian variables. This structure justifies our use of the Bayesian VAR procedure developed by Zha (1999), designed to situations where there is bloc exogeneity.[12]

We add an additional restriction on the lag structure of the reduced VAR: $\Phi_{h,3j} = 0$, for $j \geq 4$. This imposes no feedback from the other domestic variables to *br.risk*. Since we do not include the set of domestic variables that ultimately determine debt sustainability, feed-back from other variables to sovereign risk would most likely capture a reverse causality influenced by the high correlation between real and financial variables with *br.risk*.[13]

Impact Restrictions

The structural parameter identification involves imposing zero restrictions on the impact matrix A. We allow ε^{vix} and ε^{us3yr} to contemporaneously affect all domestic variables except *y* and *p*, which, being real sector variables, are less likely to react immediately to these shocks, particularly given the monthly frequency of the data. Similarly, *br.risk* immediately influences all domestic variables except *y* and *p*.

In the international bloc, we follow Bloom (2009) and interpret ε^{vix} as a proxy for global economic uncertainty shock that contemporaneously affects *vix* and *us*3*yr*. This ordering considers that *us*3*yr* is influenced by innovations in all its components (real rate, expected inflation, or *irp*), making it more responsive to various news sources than *vix*, also given the typical response of the Federal Reserve to economic shocks, which may be anticipated by shifts in *us*3*yr*. These arguments lead us to place *us*3*yr* after *vix*.

Beyond the fact that the positive shock ε^{us3yr} represents a tightening to the world financial condition, its exact interpretation is not straightforward since this innovation can potentially be originated on specific shocks to the components of *us*3*yr*. However, results of Ang, Bekaert and Wei

12 Cushman and Zha (1997) use similar econometric methodology to identify the impact of monetary policy shocks in Canada after controlling for the influence of the US policies and variables, which form an exogenous block.

13 Lowenkron and Garcia (2007) also assume strong exogeneity of EMBI-BR that functions as a price of risk in a no-arbitrage model intended to discuss the relation between inflation risk premium and monetary policy. Ferreira and Valério (2022) also impose similar restriction in VARs for Brazil, Chile, Colombia, and Peru.

(2008) showing that 80% of the unconditional variance of 3-month nominal rate of the US government bond and 71% of the 5-year is explained by expected inflation place a high probability that ε^{us3yr} is most likely reflecting news associated to higher expected inflation. The IRFs we estimate suggest a similar interpretation. Important, we also conduct an experiment inverting the order in the international bloc by placing *vix* after *us*3*yr*. We show that the responses produce results hard to reconcile with economic intuition.

Moving back to restrictions on the domestic variables, we allow r^{monet} to react immediately to any shock, as central bankers continuously monitor market conditions, as argued by Cushman and Zha (1997) in the context of Canada. However, this may not be an important restriction, since the BCB monetary policy council, responsible for determining the policy rate, meets in intervals of 45 days, and we are working with monthly data.

Other restrictions are set for estimation purposes, without influencing the IRFs nor the FEVD we report. These restrictions ensure the invertibility of matrix *A*. We adopt a strategy close to a recursive identification from variable r_1^R to r^{monet} in vector *X*. The scheme is not identical to a recursive identification because *y* and *p* do not react contemporaneously to any financial sector shock, forming the real sector bloc with *y* preceding *p*, following the literature.

4. ESTIMATION AND IMPULSE RESPONSE FUNCTIONS

This section reports the IRFs, starting with the impact of global shocks and then moving to sovereign risk shock. Since we work with variables in level, we use the same priors proposed by Sims and Zha (1998) and Zha (1999), which are designed for cases where some non-stationary variables are present in the VAR. Details about these priors are in the Appendix. The VAR is estimated with six lags ($p = 6$), which is sufficient to whiten the residuals of all equations, eliminating concerns regarding non-stationarity.

The shocks under investigation (ε^{vix}, ε^{us3yr}, and $\varepsilon^{br.risk}$) can be understood as disturbances to prices of risk from various sources affecting the term structure, akin to several no-arbitrage affine macro-finance models. However, in our context, *vix*, *us*3*yr*, and *br.risk* are entirely exogenous to the term structure's evolution, while the remaining domestic macroeconomic variables, which may also be viewed as prices of risk, exhibit feedback with the term structure, as in Hördahl, Tristani and Vestin (2006).

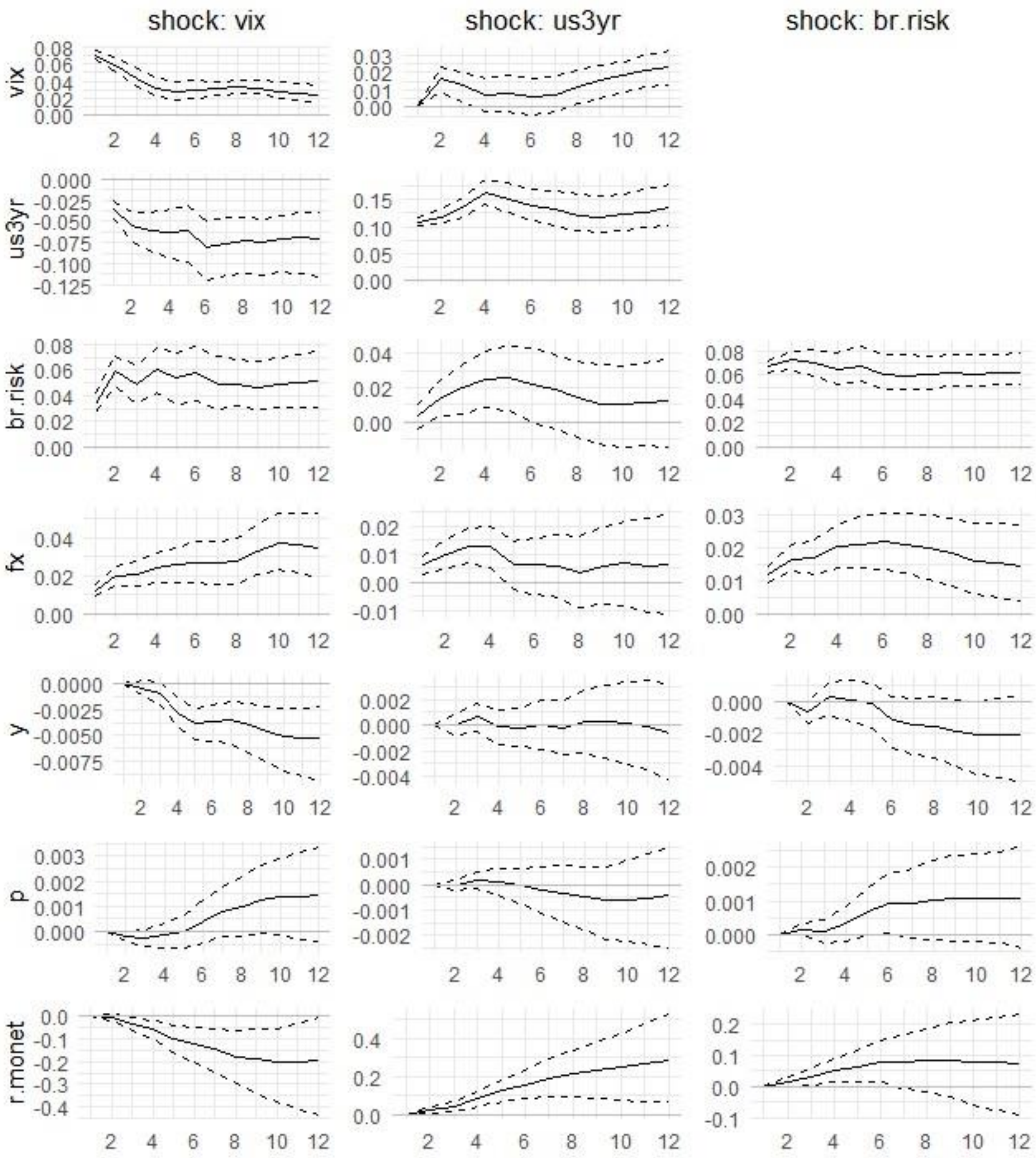
4.1 Global Shocks

We start by discussing the impact of global shocks in the international macro-finance structure of our model, moving later to the influence on macroeconomic variables, before focusing on the term structure.

4.1.1 Response of the International Bloc

The first two columns of Figure 3 show, respectively, responses to ε^{vix} and ε^{us3yr}. The first two lines display the responses of *vix* and *us*3*yr*, in this order.

An unexpected 1 standard deviation rise of ε^{vix} (approximately 8.5%) elevates the market volatility and causes an immediate 3bp drop in *us*3*yr*, as if market participants expected the FED to set nominal rates below pre-shock level in the future. These reactions seem consistent with interpreting a positive ε^{vix} as an adverse uncertainty shock of similar nature as that of the 2008/09 financial crisis, which led the FED to reduce its policy rate to values close to zero amid a recession and freezing of the financial market.[14] A complementary view is that agents fly to safety and increase the purchase of *us*3*yr*, reducing its return.



14 Moench and Soofi-Siavash (2022) show that their "news shock" target to the US nominal yield curve is akin to an uncertainty shock, since they produce similar responses the variables they analyze. Particularly related to our work, they observe that an adverse "news shock" elevates volatility (VOX) and reduces the 2-year nominal yield, which are results very similar to ours.

Figure 3 – Impulse response functions of the international and domestic macroeconomic variables.

Note: Dashed lines represent the 68% probability interval. Source: Prepared by the authors.

An unexpected 1 standard deviation increase in ε^{us3yr} (0.1 or 10 bp) causes $us3yr$ to increase and persistently remains above pre-shock level during the first year following the shock. *vix* raises, likely reflecting higher uncertainty about the future path for monetary policy and, consequently, for economic activity, which are concerns aligned with interpreting ε^{us3yr} as a shock to future expected inflation, matching the findings of Ang, Bekaert and Wei (2008).

We also find some resemblances between our IRFs and results documented by Popescu and Smets (2010) and Caldara et al. (2016) in the context of separating uncertainty and financial shocks, which in their case is related to excess premiums in the financial market.[15] Their adverse financial shock heightens financial tightness and uncertainty, which are the responses we observe following ε^{us3yr}, except that in our case the financial tightness seems to be rooted on news suggesting higher future inflation. Comparing uncertainty shock is less straightforward, as it typically leads to financial conditions tightening in their studies, manifested by increased premiums and spreads, whereas we observe a decrease in $us3yr$ following ε^{vix}. A reconciliation is possible if this reduction in $us3yr$ reflects expectations of smaller future policy rate to offset the tightening in financial conditions reported by them amidst higher uncertainty. This interpretation is also consistent with Ang, Bekaert and Wei (2008), since the decline in $us3yr$ may mirror expectations of low inflation and anemic economic activity that are typically observed during recessions, which coincide with periods of financial strain reported by Popescu and Smets (2010) and Caldara et al. (2016).

We observe limited economic rationale from an alternative identification where $us3yr$ precedes *vix*. Specifically, a sudden increase in ε^{us3yr} causes *vix* to decline, while the positive shock ε^{vix} results in a negligible and insignificant decrease in $us3yr$. In the Appendix, we also show that the responses of Brazilian variables indicate that this alternative ordering elicits reactions that lack economic coherence. We will revisit this topic later.

4.1.2 Response of Domestic Macroeconomic Variables

The responses of the domestic macroeconomic variables start being displayed in the third line of Figure 3. A sudden increase in ε^{vix} leads to higher *br.risk* and a depreciation of *fx*, expected movements following heightened global economic uncertainty. The CPI begins to rise significantly after 5 months, likely a consequence of the currency devaluation. Despite this augment, r^{monet} decreases in the year

15 Popescu and Smets (2010) identify the financial shock using a metric of financial risk premia, while Caldara et al. (2016) rely on the excess bond premium of Gilchrist and Zakrajšek (2012).

following the shock, suggesting a heightened concern about the adverse effects on domestic activity, which falls below its pre-shock trend, a pattern reported by other articles.[16,17]

The positive innovation ε^{us3yr} increases *br.risk* and depreciates *fx*, consistent with the risk-adjusted uncovered interest parity. Unlike the ε^{vix} shock, now both higher *vix* and *us*3*yr* move in the same direction to induce this depreciation. Local economic activity and CPI show no significant response[18], but r^{monet} rises, suggesting again that ε^{us3yr} captures concerns about future inflationary pressures in the USA (and likely globally).

In the appendix, we present responses when *us*3*yr* precedes *vix* in the identification strategy. In this scenario, despite *vix* falling, the positive innovation ε^{us3yr} does not lead to significant reactions in *br.risk* or *fx*, results inconsistent with common knowledge about international financial market dynamics, challenging the economic interpretation for the ε^{us3yr} shock under this alternative identification scheme.

4.1.3 The Domestic Term Structure Response

The responses of nominal and real yields, and *irp* can be assessed in Figure 4. Figures 6 and 7, in the Appendix, allow a visual inspection of the movements in the shapes of all term structures in 1, 3, and 9 months after the shocks.

TIR response to $\boldsymbol{\varepsilon^{vix}}$

Figure 4 illustrates a significant decline in 1, 3, and 4-year real interest rates subsequent to ε^{vix}, signaling a downward shift in the yield curve, despite an insignificant response on impact. The magnitude of these reactions also reveals a steepening of the real TIR curve over time, driven by a more pronounced decrease at shorter maturities. The nominal yield curve also steepens mainly due to the decline in r_1^N, aligning with the decrease in r^{monet} and suggesting an appropriate anticipation of monetary policy. Although r_3^N declines and r_4^N increases, these responses are mostly insignificant[19]. Regarding the inflation risk premium, irp_1 and irp_3 exhibit a modest positive significant response during the first quarter after the shock.

TIR response to $\boldsymbol{\varepsilon^{us3yr}}$

The second column of Figure 4 displays the significant rise in nominal and real rates of all maturities after the shock *us*3*yr*, indicating an upward shift of both curves. These responses align with

16 Ferreira and Valério (2022) also note a decrease in the policy interest rate of Brazil and Colombia following adverse global uncertainty shocks, but an insignificant reaction in Chile and Peru. Bhattarai, Chatterjee and Park (2020) find inconclusive results for a group of Latin American economies. Particularly, using a panel VAR, they observe different short-term policy rate reactions depending on the uncertainty proxy and identification strategy, though they note a clear increase in long-term nominal interest rates.

17 Filho 2014, Barboza and Zilberman (2018), and Ferreira and Valério (2022) report a negative impact of increased global uncertainty on Brazilian economic activity.

18 Our findings align with those of Akinci (2013), who do not detect a significant impact on the economic activity of emerging economies from an increase in a benchmark global interest rate after accounting for risk factors.

19 Our findings contrast with those of Bhattarai, Chatterjee and Park (2020), who, in a panel VAR analysis of 16 emerging countries, observe an increase in long-term nominal interest rates following a sudden rise in VIX.

the increase in r^{monet}, suggesting a correct assessment of the BCB reaction function and supporting the interpretation of $\boldsymbol{\varepsilon}^{us3yr}$ as a global expected inflation shock.

The magnitude of the responses indicates that the upward movement in the nominal TIR is accompanied by a shift in its shape. Initially, a uniform increase across maturities maintains the shape unchanged. However, approximately after 3 months, the curvature begins to alter, as r_3^N experiences a greater increase compared to r_1^N and r_4^N, while r_1^N rises almost twice as much as r_4^N. Subsequently, after 9 months, the response of r_4^N becomes statistically insignificant, whereas the change in r_1^N surpasses its previous impact, even exceeding the effect on r_3^N, which remains significant. These dynamics result in a flatter slope compared to pre-shock levels and immediately after the shock.

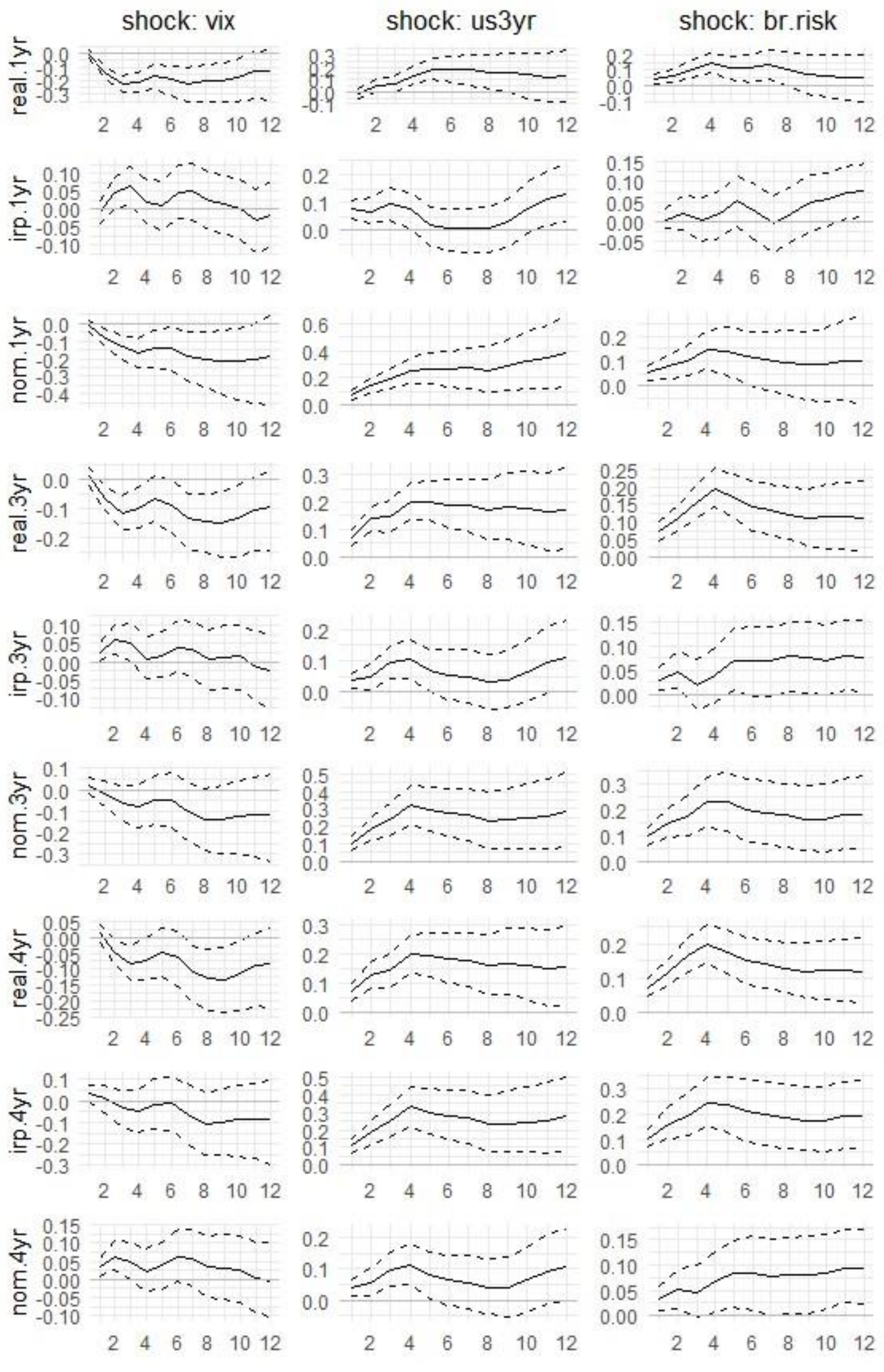


Figure 4 – Impulse response functions of real and nominal yields, and the correspondent inflation risk premiums.

Note: Dashed lines represent the 68% probability interval. Source: Prepared by the authors.

The responses of the real TIR are easier to characterize. It moves upward on impact and in the months that follow, becoming steeper on impact and at least 3 months later due to a more intense rise at longer maturities. After 9 months, however, the real rates of all maturities are almost 16 basis points higher than pre-shock levels, indicating a parallel move upward.

The term structure of *irp* becomes steeper, characterized by a notable increase in irp_3 and irp_4, the latter being approximately 15 bps and 20 bps higher than the former after 3 and 9 months, respectively. Conversely, irp_1 shows no significant reaction. This pronounced elevation at longer maturities is another piece aligned with the interpretation that $\boldsymbol{\varepsilon}^{us3yr}$ reflects an anticipated inflation shock, heightening uncertainty regarding the future trajectory of inflation worldwide and the consequent policy response.

4.2 Sovereign Risk Shock - $\boldsymbol{\varepsilon}^{br.risk}$

The $\boldsymbol{\varepsilon}^{br.risk}$ shock is conventionally interpreted as a sovereign risk shock. We conceptualize it as information influencing perceptions regarding the likelihood of the Treasury defaulting on its financial obligations, encompassing partial default through inflation. Examples include news emerging from election polls, debates in the national parliament concerning subjects impacting public debt dynamics, and announcements of policies with similar ramifications. Signals from any branch of government affecting perspectives on public debt trajectory or macroeconomic policy management, including institutional matters, can also move $\boldsymbol{\varepsilon}^{br.risk}$.

Response of macroeconomic variable

The positive shock $\boldsymbol{\varepsilon}^{br.risk}$ leads to an approximately 7% increase in *br.risk* on impact, persisting above pre-shock levels throughout the first year. *fx* immediately depreciates in 1.2%, peaking at 2.2% six months later. Although *y* declines below its pre-shock trend, significance is only observed after 5 months. Additionally, the CPI increases, likely in response to the currency depreciation, prompting a rise in r^{monet}.

Figure 3 illustrates that the responses to $\boldsymbol{\varepsilon}^{br.risk}$ closely resemble those induced by ε^{vix}, except for the behavior of r^{monet}, which increases in the former case but decreases with ε^{vix}. This disparity may stem from the international uncertainty shock potentially triggering liquidity challenges on a global scale, prompting central banks to implement liquidity injections to prevent financial sector distress. Conversely, in the event of an adverse sovereign risk shock, typically associated with indications of fiscal mismanagement, the central bank may prioritize signaling a firm commitment to price stability to safeguard its reputation, despite short-term GDP costs. Essentially, the central bank may signal the prevalence of monetary dominance.

Response of the term structure

Both real and nominal TIR significant increase immediately following $\boldsymbol{\varepsilon}^{br.risk}$. The IRFs show a continuous upward shift until the first quarter post-shock, followed by a gradual reduction, maintaining term structures slightly lower but still above pre-shock levels after 9 months. Real TIR becomes steeper, driven by a greater increase in longer maturities. Meanwhile, the curvature of nominal TIR shifts, as r_3^N experiences a larger increase compared to r_1^N and r_4^N, with the latter two rising in similar magnitudes.

While irp_1 shows no significant reaction, likely indicating reduced uncertainty about the short-term impact of $\boldsymbol{\varepsilon}^{\boldsymbol{br.risk}}$ on inflation, irp_3 and irp_4 gradually increase over time, contributing to a steeper term structure of *irp* primarily driven by a more pronounced rise in irp_4. This suggests heightened uncertainty regarding public debt sustainability and future obligations, often associated with currency devaluations and inflationary pressures. This interpretation is aligned with the findings of Lowenkron and Garcia (2007), Montes and Curi (2017), and Reis (2018) indicating that the Brazilian inflation premium positively correlates with fiscal uncertainty. Such uncertainty may also explain the augment in longer real yields, while the rise in r_1^R may likely be anticipating a hawkish monetary stance, which we capture by the rise in r^{monet}, to fight the inflation caused by the currency devaluation. The positive response of r_1^N combines these ingredients.

4.3 Inflation risk premium and the macroeconomy

Several studies have documented a negative relationship between economic activity and the *irp* (e.g., Buraschi and Jiltsov (2005), Ang, Bekaert and Wei (2008), Camba-Mendez and Werner (2017), Breach, D'Amico and Orphanides (2020)). Our IRFs only partially support this finding, showing a strong conditional negative relation following $\boldsymbol{\varepsilon}^{\boldsymbol{br.risk}}$ and a milder relation in response to ε^{vix} verified only for irp_1 and irp_3 in the first quarter following the shock. Conditioned on ε^{us3yr}, the premium of all maturities increases, but economic activity remains largely unaffected. These results indicate that the relationship between *irp* and economic activity is both shock-dependent and sensitive to the maturity from which the premium is derived. [20]

The positive relationship between current inflation and *irp* reported by Ang, Bekaert and Wei (2008) is observed in our IRFs conditioned on ε^{vix} and $\boldsymbol{\varepsilon}^{\boldsymbol{br.risk}}$ but not when exposed to ε^{us3yr}. The nature of the shock also affects the relationship between the monetary policy interest rate and *irp*. These variables move in the same direction following ε^{us3yr} and $\boldsymbol{\varepsilon}^{\boldsymbol{br.risk}}$, but in opposite directions following ε^{vix}.[21]

The findings about the relationship between *irp* and economic activity and inflation translate to the relationship between *irp* and the covariance between economic activity and the price level. According to consumer-based asset pricing models applied to nominal TIR, this covariance determines the inflation risk premium[22], with a negative value indicating that agents suffer losses due to inflation in states of low activity, leading them to demand a premium to hold the security. [23] We indeed observe this negative relationship conditioned on ε^{vix} and $\boldsymbol{\varepsilon}^{\boldsymbol{br.risk}}$. However, neither economic activity nor the price

20 Reis (2018) finds a positive but non-significant relationship between *irp* and the output gap using Brazilian data.

21 The conditioned relation between r^{monet} and *irp*, based on responses to specific shocks, cannot be conflated with evidence that more credible monetary policies and central bank independence are associated with smaller inflation premiums, a result verified by several works for Brazil (Lowenkron and Garcia (2007), Joyce, Lildholdt and Sorensen (2010), Andreasen (2012), Doi, Fernandes and Nunes (2017), Montes and Curi (2017)).

22 To be more rigorous, the covariance between the inflation and the variation in consumption, not in economic activity, determines the inflation premium.

23 Ang, Bekaert and Wei (2008) and several other articles find that the inflation premium justifies the average positive sloped nominal yield term structure.

level is impacted by ε^{us3yr}, despite the large significant increase of *irp* in all maturities, suggesting that other economic relations may also contribute to the determination of *irp*.

Consistent with previous research on Brazil showing a positive correlation between *irp* and the sovereign risk (e.g., Lowenkron and Garcia (2007), Montes and Curi (2017), Reis (2018)), these variables respond in the same direction in the presence of all shocks we evaluate. This may be due to concerns about future inflation driven by fiscal pressures, which aligns with findings of a positive correlation between the premium and inflation uncertainty also verified by several scholars (e.g., Durham (2006), Garcia and Werner (2010), Söderlind (2010), D'Amico, Kim and Wei (2018)).

Lastly, we also extend comparison to the findings of Beauregard et al. (2024), even though their analysis is more structural, based on regressing the inflation risk premium of a 5-year Mexican Treasury nominal bond on a set of international variables, while we are interested in responses over the business cycle. They report a positive significant conditional correlation with *vix*, while we also detect a movement in the same direction when exposing to ε^{vix} and ε^{us3yr}. They however find a positive conditional correlation with the 10-year US Treasury nominal interest rate, while we detect a positive co-movement following ε^{us3yr}, but an opposite following ε^{vix}. However, as we show in the next Section, shocks to ε^{us3yr} explain 34.4% of the FEV of our irp_4 for 12 months forecast (which is more comparable to their structural analysis of a 5-year inflation risk premium), while shocks to ε^{vix} explain only 2.4%.

5. RESULTS - FORECAST ERROR VARIANCE DECOMPOSITION (FEVD)

Before delving into the analysis of each shock's importance for the forecast error variance (FEV) of the term structure, it is informative to quantify their impact on *vix*, *us*3*yr*, and *br.risk*. The contribution of ε^{vix} to the FEV of *us*3*yr* fluctuates around 80% for forecasting horizons of 3 to 12 months, whereas ε^{us3yr} accounts for 3.7% to 11.4% of the FEV of *vix*. These proportions underscore the importance of jointly modeling external variables to enhance the measurement and identification of shocks. Furthermore, it is noteworthy that the contribution of ε^{vix} to the FEV of *br.risk* ranges from 32% to 37% for forecasts up to 1 year, with a smaller contribution from ε^{us3yr} (around 4%). Once again, obtaining a more accurate understanding of the impact of ε^{us3yr} necessitates filtering out the influence of global risk factors.

The joint contribution of $\boldsymbol{\varepsilon^{vix}}$, $\boldsymbol{\varepsilon^{us3yr}}$, and $\boldsymbol{\varepsilon^{br.risk}}$

Global shocks contribute significantly to the FEV of nominal TIR. For forecasts ranging from 3 to 12 months, they account for 30%-41% of the FEV of r_1^N, which increases to 37%-46% when augmented by the influence of $\boldsymbol{\varepsilon^{br.risk}}$. Similarly, for r_3^N, global shocks contribute 24%-37%, rising to 39%-54% with $\boldsymbol{\varepsilon^{br.risk}}$. The contribution of global shocks to r_4^N is 17%, expanding to 23%-35% with $\boldsymbol{\varepsilon^{br.risk}}$**.**

In the case of r_1^R, global shocks contribute 28%-37% to forecasts up to 1 year, increasing to 32%-44% with $\boldsymbol{\varepsilon^{br.risk}}$. ε^{vix} and ε^{us3yr} jointly explain 29%-40% of the FEV of r_3^R, rising to 47%-59% after adding $\boldsymbol{\varepsilon^{br.risk}}$. Similar contributions are observed for r_4^R.

For the FEV of irp_1, ε^{vix} and ε^{us3yr} jointly contribute 14%-20% for forecasts from 3 to 12 months, increasing to 17%-26% with the addition of $\boldsymbol{\varepsilon}^{br.risk}$, with similar patterns for irp_3. Notably, larger influences are seen for irp_4: 24%-37% for global shocks, and 41%-57% after incorporating $\boldsymbol{\varepsilon}^{br.risk}$. In summary, the shocks we study explain large proportions of the FEV of r^N, r^R, and *irp*.

FEV from ε^{vix}

ε^{vix} contributes around 10% to the FEV of r_1^N consistently across forecasting horizons, with modest impact on r_3^N and r_4^N. For r_1^R, its contribution ranges from 20.7% to 25.6% over 3 to 12 months, declining to 11.2% for r_3^R at 12 months and 8.5% forr_4^R. Its influence on *irp* is negligible.

FEV from ε^{us3yr}

ε^{us3yr} significantly contributes to the FEV of r_1^N, r_3^N and r_4^N, ranging from 21.4% to 33% across different forecasting periods. Its impact increases with longer forecast periods, reaching approximately 30% for r_3^N and r_4^N at 6 to 12 months. It also explains a notable portion of the FEV for r_1^R, particularly at 6- and 12-months forecasting. For *irp*, *us*3*yr* explains between 11% and 17% across various forecast periods, with larger contributions for irp_4, ranging from 23.9% to 34.4%, which aligns with the interpretation that ε^{us3yr} captures an expected global inflation shock.

FEV from $\boldsymbol{\varepsilon}^{br.risk}$

$\boldsymbol{\varepsilon}^{br.risk}$ significantly drives the FEV of the nominal TIR, particularly for r_3^N and r_4^N, with contributions ranging from 14.3% to 26.7%. Its impact on *irp* is notable, especially for irp_4, where contributions range from 16.8% to 21.1%.

Table 2 – Forecasting error variance decomposition (%)

Shock	forecast horizon	r_1^N	r_3^N	r_4^N	r_1^R	r_3^R	r_4^R	irp_1	irp_3	irp_4
ε^{vix}	3 months	8.5	1.1	4.1	25.6	8.2	4.3	2.8	4.2	0.5
	6 months	9.8	1.4	3.7	23.5	7.1	3.8	2.6	2.8	0.5
	12 months	11.5	4.4	3.5	20.7	11.2	8.5	2.9	2.2	2.4
ε^{us3yr}	3 months	21.4	23.3	12.1	2.4	20.5	21.6	14.7	11.1	23.9
	6 months	29.0	33.1	15.3	11.7	29.2	30.6	12.4	13.9	34.0
	12 months	29.9	32.5	16.0	13.0	28.7	29.5	16.8	14.9	34.4
$\boldsymbol{\varepsilon}^{br.risk}$	3 months	7.1	14.3	6.4	4.9	18.3	22.8	1.6	4.2	16.8
	6 months	8.4	18.6	11.0	7.8	22.9	26.7	3.0	7.5	21.1
	12 months	4.9	16.9	16.0	6.0	17.3	21.3	6.4	11.6	20.0

Note: Forecast error variance for 3, 6, and 12 months ahead forecast. *Source*: Prepared by the authors.

6. FINAL REMARKS

Most works on macro-finance focusing on the term structure of interest rates tend to i) analyze factors instead of actual yields, ii) use endogenously derived components of nominal yields from a pricing model instead of relying on observables, and iii) do not analyze the nominal yields jointly with most of their components. We deviate from these trends and propose a *pure econometrics* macro-finance model for Brazil to study the impact of shocks to domestic sovereign risk, global economic uncertainty and global finance on the nominal TIR of the Brazilian Treasury bonds and two of its components - real interest rate and inflation risk premium at 1-, 3- and 4-years maturity. The VAR model incorporates domestic macroeconomic variables that are typically present in general equilibrium models for small open economies, allowing a better understanding of the channels responsible for driving nominal and real yields at various maturities, alongside the inflation risk premium. This structure also permits comparisons to results and relations of pricing models.

Our FEV decomposition underscores the relevance of the shocks we study for the variability of yields and inflation premiums. World economic uncertainty shocks have a limited impact on the unexpected changes in nominal yields and *irp*, but significantly affect the 1-year real rate (approximately 22% of the FEV for 3-12 months forecasts). Conversely, innovations in the 3-year US nominal interest rate, which seem related to shocks to expected inflation in the US, account for a larger share of variation in the nominal and real TIR, as well as in *irp*. These contributions oscillate between 20% and 34%.

Sovereign risk shocks minimally affect the 1-year nominal rate (up to 8%) but have a more pronounced impact on 3 and 4-year yields (14%-19% and 6%-16%, respectively). Real rates are notably influenced at these maturities (17%-23% and 21%-27%, respectively), the same happening with the *irp* of bonds maturing in 4 years (17%-21%).

The impulse responses show that an adverse global uncertainty shock produces a steeper nominal TIR in Brazil because shorter-term yields fall intensely, while longer maturities do not react. Similar movement is observed for real yields, while the *irp* appears less responsive to this shock. An adverse innovation in global financial condition moves nominal and real TIR upwards and can reduce their slopes due to a smaller rise at longer maturities. The term structure of the *irp* moves upwards and becomes steeper due to of a more intense rise at longer maturities.

An adverse sovereign risk shock drives nominal and real TIR upwards and affects their shapes, with the first becoming more convex and the second steeper. The *irp* curve also becomes steeper due to a larger increase at longer maturities, aligning with uncertainties arising from the fiscal front that ultimately augments the odds of a fiscal dominance scenario.

The IRFs also reveal that the relationship between the inflation risk premium and macroeconomic variables tends to be shock-dependent, so a general sign for the covariances does not fit all cases. For instance, we verify a negative relation between impulses of *irp* and economic activity when Brazil is exposed to shocks to sovereign risk and global economic uncertainty, but not in the presence of global finance shocks.

Shock-dependence is also observed in the relation between *irp* and inflation, being positive in the presence of shocks to sovereign risk and global uncertainty, and nonexistent in the presence of a global finance shock. Finally, we identify that the direction of the relation of *irp* with the covariance between activity and inflation, which forms the inflation risk premium according to consumer-based asset price models, is also shock-dependent, being negative when conditioned to shocks to sovereign risk and global uncertainty, and essentially nonexistent in the presence of a global finance shock that, however, has a huge impact on inflation premiums.

There is an important research agenda to enhance our understanding of the relationship between macroeconomics and the term structure of nominal yields and their components in open economies in general and in emerging countries in particular. On the empirical front, including expected inflation in a framework like ours is an easy and natural extension that would allow the full description of what happens to nominal yields by checking responses in all their components. Following this path requires a longer data range than ours due to degrees of freedom concerns. Along the same line, including extra maturities would provide robustness to the results we have reached. Another natural path is to consider the possibility of state-dependence in the responses, since previous works, like Ang, Bekaert and Wei (2008), have shown how the macro-finance relations are influenced by the state of the economy. This path should complement our findings that several correlations are shock-dependent. Contributions on the theoretical front should attempt to explain the reasons behind the shock-dependent relationships we observe, and the role played by sovereign risk and nominal exchange rate in driving the results.

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APPENDIX

A1. Bayesian priors for the VAR estimation

In order to estimate the Bayesian VAR we use priors suggested by Sims and Zha (1998) and Zha (1999). We combine two unit root priors: the Minnesota prior and the sum-of-coefficients prior: the Minnesota prior and the sum-of-coefficients prior. The Minnesota prior imposes the restriction that coefficients on the first lag has prior mean of 1. This is done by creating variables such that for the *i*th equation a set of $k-1$ dummy observations, indexed by $j = 1, \dots, m$ and $l = 1, \dots, p$, is inserted in the data sample according to

$$y_i(r,j)\,;\; r = 1,\dots,k-1;\; j = 1,\dots,m = \begin{cases} \mu_1\mu_2\sigma_r/l^{\mu_4}, & \text{if } r = j, r \leq m \\ 0, & \text{otherwise} \end{cases}$$

$$x_i(r,s)\,;\; r = 1,\dots,k-1;\; s = 1,\dots,k-1 = \begin{cases} \mu_1\mu_2\sigma_r/l^{\mu_4}, & \text{if } r = s, \\ 0, & \text{otherwhise} \end{cases}$$

where μ_1, μ_2 and μ_4 are hyperparameters. μ_1 controls the overall tightness of $\Phi(0)$; μ_2 controls the relative tightness of $\Phi(L)$ for $L > 0$, and μ_4 controls the tightness on lag decay. We set these hyperparameters at the default values suggested by Sims and Zha (1998) and Zha (1999), which are 1, 0.5 and 1, respectively.

The sum-of-coefficients prior is used in cases where there s a belief that the variables have a unit root, which is captured by setting to 1 the sum of the lagged coefficients of each dependent variable. This is achieved in a system of *m* equations, *l* lags, and *k* coefficients, by introducing m observations, indexed by *i*, of the form:

$$y(i,j)\ ;\ i = 1,\dots,m;\ j = 1,\dots,m = \begin{cases} \mu_5 \bar{y}_{0i}, & \text{if } i = j, \\ 0, & \text{otherwise} \end{cases}$$

$$x(i,s)\ ;\ i = 1,\dots,m;\ s = 1,\dots,k = \begin{cases} \mu_5 \bar{y}_{0i}, & \text{if } i = j, \text{for all } l, \\ 0, & \text{otherwise} \end{cases}$$

where $\bar{y}_{0i}$ is the average of initial values of the variable i and μ_5 is a hyperparameter that controls the weight of the prior. When $\mu_5 \to \infty$, the model can be expressed entirely in terms of differenced data. We follow Sims and Zha (1998) and Zha (1999), and set $\mu_5 = 1$.

A2. Impulse response functions with an alternative ordering

The following figure shows the impulse responses from an alternative ordering in which ε^{vix} does not instantaneously affect *us*3*yr*, but it is influenced by ε^{us3yr}.

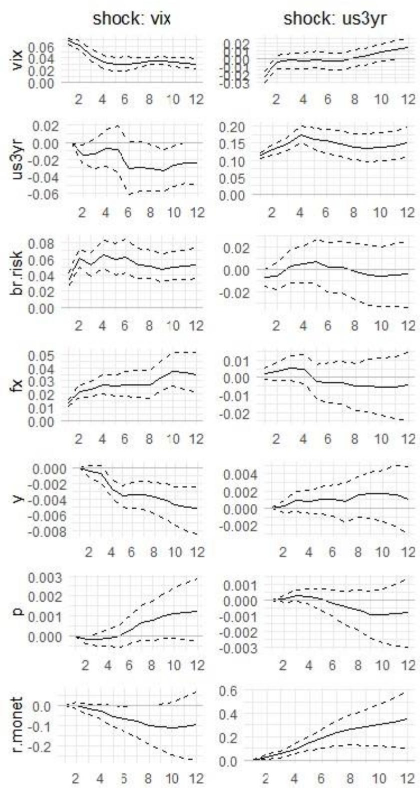

Figure 5 – Impulse response functions of the macroeconomic variables under an alternative identification: *us3yr* preceding *vix*.

Note: Dashed lines represent the 68% probability interval. These impulse responses are based on an alternative identification strategy where *vix* does not affect *us*3*yr* contemporaneously but suffers an immediate influence of a shock to *us*3*yr*. Source: Prepared by the authors.

A3. Term structure shift under different shocks

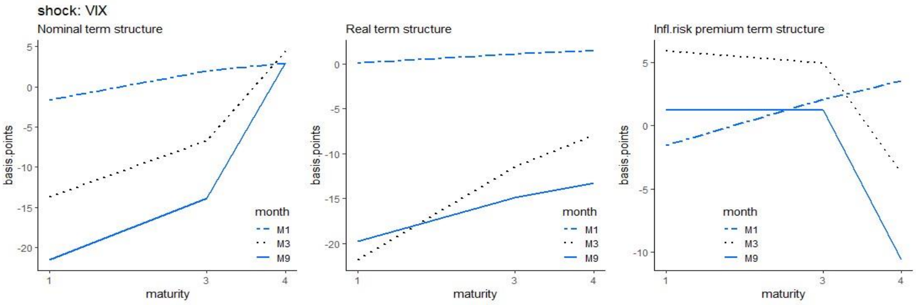

Figure 6 – The impact of a shock ε^{vix} on the term structure of nominal and real rates and of the inflation risk premium.

Note: The numbers indicate basis point reaction from pre-shock levels. *Source*: Prepared by the authors.

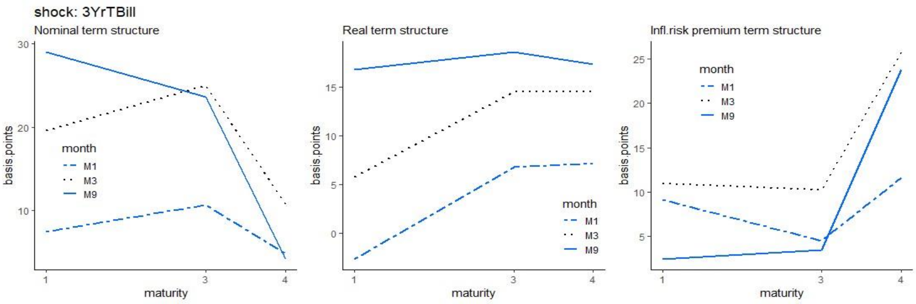

Figure 7 – The impact of a shock ε^{us3yr} on the term structure of nominal and real rates and of the inflation risk premium.

Note: The numbers indicate basis point reaction from pre-shock levels. *Source*: Prepared by the authors.

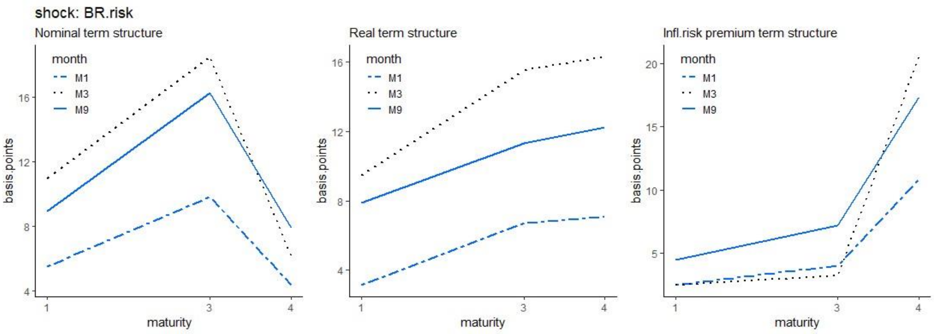

Figure 8 – The impact of a shock $\varepsilon^{br.risk}$ on the term structure of nominal and real rates and of the inflation risk premium.

Note: The numbers indicate basis points reaction from pre-shock levels. *Source*: Prepared by the authors.