Alba, Carlos; Carrillo, Julio A.; Ibarra-Ramírez, Raúl

Working Paper

Information effects of US monetary policy announcements on emerging economies: Evidence from Mexico

Working Papers, No. 2024-14

Provided in Cooperation with:
Bank of Mexico, Mexico City

Suggested Citation: Alba, Carlos; Carrillo, Julio A.; Ibarra-Ramírez, Raúl (2024) : Information effects of US monetary policy announcements on emerging economies: Evidence from Mexico, Working Papers, No. 2024-14, Banco de México, Ciudad de México

This Version is available at:
https://hdl.handle.net/10419/305402

Banco de México

Working Papers

N° 2024-14

Information Effects of US Monetary Policy Announcements on Emerging Economies: Evidence from Mexico

Carlos Alba
Banco de México

Julio A. Carrillo
Banco de México

Raúl Ibarra
Banco de México

September 2024

Documento de Investigación
2024-14

Working Paper
2024-14

Information Effects of US Monetary Policy Announcements on Emerging Economies: Evidence from Mexico*

Carlos Alba†
Banco de México

Julio A. Carrillo‡
Banco de México

Raúl Ibarra§
Banco de México

Abstract: This paper analyzes, using a VAR model, the effects of US central bank monetary policy announcements, and information shocks from this authority regarding its economic outlook on Mexican financial and macroeconomic variables. Shocks are identified by combining a high-frequency strategy with sign restrictions, which exploits the co-movement between the policy rate and the stock market in the US around FOMC announcements. A restrictive monetary policy shock in the US is identified by an increase in the interest rate and a drop in stock prices, while a positive information shock is identified when both variables rise simultaneously. The results show that positive information shocks from the US central bank improve financial conditions in Mexico, appreciate the peso/dollar exchange rate, lower the sovereign risk premium and forex volatility, and increase stock prices, real activity and prices in Mexico. In contrast, restrictive US monetary policy shocks tighten financial conditions, and reduce real activity and prices in Mexico.

Keywords: Monetary policy, international policy transmission, high-frequency identification, central bank information, VAR model

JEL Classification: E43, E52, E58, F42

Resumen: Este artículo analiza, mediante un modelo VAR, los efectos de los anuncios de política monetaria del banco central de EUA y los choques de información de esta autoridad acerca de sus perspectivas económicas sobre variables financieras y macroeconómicas de México. Los choques se identifican combinando una estrategia de alta frecuencia con restricciones de signo, que aprovecha el comovimiento entre la tasa de política y el mercado accionario en EUA alrededor de los anuncios del FOMC. Un choque restrictivo de política monetaria en EUA se identifica por un aumento en la tasa de interés y una caída en los precios de las acciones, mientras que un choque de información positivo se identifica cuando ambas variables aumentan simultáneamente. Los resultados muestran que choques positivos de información por parte del banco central de EUA mejoran las condiciones financieras en México, aprecian el tipo de cambio peso/dólar, reducen la prima de riesgo soberano y la volatilidad cambiaria y aumentan los precios de las acciones, la actividad real y los precios en México. En contraste, los choques restrictivos de política monetaria en EUA aprietan las condiciones financieras y reducen la actividad real y los precios en México.

Palabras Clave: Política monetaria, transmisión internacional de política, identificación de alta frecuencia, información del banco central, modelo VAR

---

*We thank the anonymous referees, the participants of the 2024 Canadian Economics Association Conference, and Alejandrina Salcedo for their constructive comments. The views expressed in this paper correspond to the authors and do not necessarily reflect those of Banco de México.

† Dirección General de Investigación Económica. Email: calbaf@banxico.org.mx
‡ Dirección General de Investigación Económica. Email: jcarrillo@banxico.org.mx
§ Dirección General de Investigación Económica. Email: ribarra@banxico.org.mx

1 Introduction

Monetary policy decision announcements not only reveal information about the policy interest rate, but also about the central bank's assessment of the economic outlook. Previous literature has shown that an announcement of an increase in the policy rate typically raises both short and long-term interest rates and, in general, may lead to a tightening in financial conditions (Gürkaynak et al., 2005; Gertler and Karadi, 2015). In addition, higher interest rates could also be a response to an improved economic outlook, as assessed by the central bank. In this scenario, the latter tightens monetary policy to prevent demand pressures on inflation and thereby mitigate inflationary risk premia. Consequently, higher interest rates may even result in improved financial conditions. In general, central bank announcements include information about the economic outlook, which can influence private expectations about the economy and the future path of interest rates. As it has become widely recognized, this information may bias the estimated effects of monetary policy. In this regard, a growing part of the literature has focused on disentangling these two components of central bank announcements, conventional monetary policy shocks and information shocks, in order to assess their effects on the economy (Jarociński and Karadi, 2020). As the global economy becomes increasingly interconnected, a central question for researchers and policymakers in emerging market economies (EMEs) is how central bank information shocks, which stem from surprises associated with the Federal Reserve's assessment of the US economy, can affect the transmission of monetary policy across international borders.[1]

This paper analyzes the spillover effects of US monetary policy and central bank information shocks on the Mexican economy, a large EME that has strong economic links with the US.[2] For this purpose, we estimate a Bayesian vector autoregression (VAR) model. This

[1] Throughout the paper, we refer to financial asset price changes surrounding monetary policy announcements as "surprises".

[2] Mexico is a small open economy that heavily depends on the US as a trading partner. About 80% of Mexican exports are sold to the US and almost half of the foreign direct investment that Mexico receives originates from the US (Carrillo et al., 2020). In fact, for a group of large EMEs including Brazil, Chile, Malaysia, and South Africa, among others, Mexico is characterized by the strongest trade relations with the US (Fink and Schüler, 2015). Additionally, in 2023, Mexico became the principal trading partner of the US, surpassing China and Canada. According to our calculations based on data from the United States Census Bureau, Mexico–US trade in 2023 represented 15.6% of all goods exported and imported by the US; the Canada–US share followed at

model allows us to complement monthly economic variables with high-frequency variables reflecting financial market surprises associated with US monetary policy announcements. In particular, we resort to a two-country VAR model with block-exogeneity, which assumes that the US economy is block-exogenous with respect to the Mexican economy. This assumption has been widely used in the literature to analyze the propagation of external shocks into EMEs [see, e.g., Canova (2005), Fink and Schüler (2015) and Carrillo et al. (2020)]. The model combines two approaches to shock identification: sign restrictions and high frequency identification (HFI).[3] As in Jarociński and Karadi (2020), we distinguish between standard US monetary policy shocks and contemporaneous central bank information shocks by imposing sign restrictions on high-frequency data of short-term interest rates and stock prices in a narrow window around the US Federal Open Market Committee (FOMC) announcements.[4] We then obtain the dynamic responses of financial and macroeconomic Mexican variables to these shocks, including sovereign risk, interest rates, stock prices, exchange rates, economic activity, and prices.

Regarding conventional monetary policy shocks, a broad range of models predicts that when monetary authorities unexpectedly increase the policy rate, the expected value of future dividends decreases, and the discount rate with which such dividends are discounted also increases. This, in turn, leads to a decrease in stock prices. Thus, we identify a monetary policy shock through a negative co-movement between interest rates and stock price changes in

15.2%, and the China–US share was 11.3%. Furthermore, significant financial integration exists between the Mexican economy and the US. For instance, measures of the cost of external finance and the term premium show a correlation of approximately 75% and 67%, respectively. Additionally, short-term nominal interest rates in Mexico and the US display a strong co-movement, exhibiting an unconditional correlation of 68% (Carrillo et al., 2020).

[3] The literature on HFI is already large and expanding. Some of the early empirical studies assessing the impact of high-frequency financial market surprises around monetary policy announcements on asset prices include Kuttner (2001), Gürkaynak et al. (2005) and Bernanke and Kuttner (2005), among others. Subsequent studies have focused on evaluating the effects of these surprises also on the macroeconomy [see, e.g., Campbell et al. (2012), Gertler and Karadi (2015), Paul (2020), Nakamura and Steinsson (2018), and Miranda-Agrippino and Ricco (2023)].

[4] See, e.g., Kerssenfischer (2022), Breitenlechner et al. (2021), and Andrade and Ferroni (2021), who also employ a similar approach to identify information and monetary shocks. The methodology is closely related to proxy VARs that employ external instruments to identify the shocks of interest [see Stock and Watson (2012), Mertens and Ravn (2013) and Gertler and Karadi (2015), among others]. This strategy is also in line with recent studies that highlight a signaling channel of monetary policy (Cesa-Bianchi and Sokol, 2022; Melosi, 2017).

a narrow window around FOMC announcements. In particular, we use the current-month fed funds futures to measure changes in expectations about short-term interest rates, and the S&P 500 index to measure changes in stock valuation within a half-hour window around policy announcements. Conversely, if interest rates and stock prices move in the same direction after the release of a monetary policy decision, then it is likely that this decision was accompanied by information regarding the economic outlook. This reaction corresponds to a central bank information shock. If the co-movement between the policy rate and stock prices is positive, this might signal that the central bank has tightened monetary policy to partly counteract the effect of the positive news and prevent demand pressures on inflation (Jarociński and Karadi, 2020). As emphasized by Kerssenfischer (2022), in this case, the negative discount rate effect on stock prices is outweighed by the positive cash flow effect.

Numerous studies have analyzed the propagation of conventional US monetary policy shocks to EMEs. There is significant evidence indicating that an unexpected increase in US interest rates, i.e. a contractionary monetary policy shock, can result in a tightening of financial conditions in these economies [see Uribe and Yue (2006), Mackowiak (2007), Vicondoa (2019), and Carrillo et al. (2020), among others].[5] According to Dahlhaus et al. (2018), in the presence of global financial market integration, changes in US asset prices and yields stemming from monetary policy decisions can be mirrored in corresponding domestic financial market variables. This, in turn, can affect consumption, investment and other variables through the monetary policy transmission mechanism. The importance of this financial channel for the propagation of monetary policy shocks to Mexico is highlighted by studies such as Canova (2005) and Carrillo et al. (2020).[6] In particular, the latter find that a US restrictive

[5] Uribe and Yue (2006) and Vicondoa (2019), for instance, find that US interest-rate shocks induce an increase in the risk of sovereign default as well as a contraction of output and investment in EMEs. According to Uribe and Yue (2006), increases in US interest rates affect domestic variables in EMEs mostly through their effects on the cost of borrowing that these countries face in international financial markets. Canova (2005) and Ilzetzki and Jin (2013) find that a contractionary monetary policy shock in the US induces a depreciation of the exchange rate in EMEs. In turn, Mackowiak (2007) finds that US monetary policy shocks affect interest rates and the exchange rate in EMEs quickly and strongly.

[6] Other channels that may play a role in the international transmission of monetary policy are the exchange rate and trade channels. Canova (2005), Blanchard et al. (2010), and more recently Iacoviello and Navarro (2019) present a detailed discussion of these channels. In particular, the exchange rate channel implies that a monetary contraction in the US leads to an appreciation of the dollar, via the uncovered interest parity condition, that generates a demand substitution between domestic and foreign-produced goods. Thus, real output in EMEs

monetary policy shock leads to an increase in the sovereign risk premium, a rise in the term premium, and a depreciation of the Mexican peso/US dollar exchange rate. Additionally, Mexican real activity might experience a persistent slowdown after an unexpected monetary tightening in the US (Carrillo et al., 2020).

Some studies have also analyzed how central bank information shocks regarding the economic outlook can impact macroeconomic outcomes [see Kerssenfischer (2022), Jarociński and Karadi (2020) and the references therein]. These studies generally suggest that central bank information shocks have effects that are opposite to those of monetary policy shocks on the economy. Therefore, these information shocks may attenuate the estimated responses of various variables to monetary policy shocks.[7] In fact, monetary policy announcements are typically accompanied by central bank communication regarding the economic outlook.[8] Consequently, if this information triggers changes in private sector expectations regarding interest rates and the macroeconomy, it can potentially lead to a bias in the estimated effects of monetary policy, as highlighted by Jarociński and Karadi (2020).

The main contribution of this paper is to analyze how central bank information shocks from US monetary policy announcements affect the Mexican economy. To the best of our

might rise, boosted by cheaper exports. The trade channel, on the other hand, implies that a contractionary monetary policy reduces incomes and expenditures in the US, thus leading to a lower demand for imported goods, and reducing activity abroad. Overall, the strength of this channel should depend on the trade exposure with the United States.

[7] In particular, Kerssenfischer (2022) finds that monetary policy shocks that neglect any potential information channel have puzzlingly insignificant small effects on stock prices and economic expectations in the euro area. According to this author, this is the result of two forces offsetting each other. While contractionary monetary policy lowers stock prices and expected inflation, an improved economic outlook raises them. In this line, Jarociński and Karadi (2020) find that disregarding the information content of the central bank announcements leads to a puzzle in which financial conditions in the euro area improve significantly after a monetary policy tightening, contradicting standard theory. In addition, these authors find that, in the case of the US, the responses of some macroeconomic and financial variables are muted because the central bank information shocks attenuate the estimated responses of these variables to a monetary policy shock. According to Jarociński and Karadi (2020), an additional bias is that the interest rate responses to a monetary shock are larger and more persistent due to the presence of information shocks, which have higher and more persistent effects on interest rates.

[8] This paper complements the literature that aims to quantify the impact of central bank information revelation on the economy [see, e.g., Hansen and McMahon (2016), Campbell et al. (2017), and Lakdawala and Schaffer (2019), among others]. However, instead of using private information proxies created from analyzing the language of announcements or obtained from the differences between the FRB staff and private sector forecasts, the approach used in this paper relies on the information-processing power of the markets and identifies information shocks by analyzing the high-frequency co-movement of interest rate and stock market surprises around FOMC announcements (Jarociński and Karadi, 2020).

knowledge, this is one of the first papers that carries out this analysis for an emerging economy by identifying monetary and information shocks using sign restrictions and HFI in a Bayesian VAR model. As in Jarociński and Karadi (2020), Breitenlechner et al. (2021), and Andrade and Ferroni (2021), among others, this empirical strategy relies on the information inherent in the high-frequency co-movement of interest rates and stock prices around FOMC announcements.[9] The fact that the empirical evidence has underlined the importance of these shocks for affecting the US economy provides motivation for analyzing their transmission into its southern neighbor, a representative EME with a strong degree of financial and commercial integration with this country. In addition, the empirical evidence suggests that ignoring central bank information shocks may lead to biased measurements of the estimated effects of monetary policy shocks. In this regard, we also obtain impulse responses to monetary policy shocks that are purged from the effects of the central bank information shocks. By analyzing the effects of these shocks on the Mexican economy, we also contribute to the discussion about the international propagation of US monetary policy shocks to EMEs.

The main results indicate that positive central bank information shocks in the US improve financial conditions in Mexico, lower the sovereign risk premium, appreciate the Mexican peso/US dollar exchange rate, lower forex volatility, and boost the stock price index. Additionally, the 3-month and the 10-year government bond yields tend to increase a few months after this shock. These information shocks also positively affect economic activity and the price level in Mexico, which help to explain the behavior of interest rates. While our analysis does not delve into individual channels of the transmission mechanism, we conjecture that both trade and financial channels likely play crucial roles in transmitting these shocks to the Mexican economy. Specifically, in response to the unexpected good news in the US economic outlook, we hypothesize that private sector decisions in both countries may positively influence aggregate demand in Mexico through different channels, such as the current

[9] Carrillo et al. (2020) also analyze the effects of US central bank information shocks on the Mexican economy. However, differently from us, these authors do not use high-frequency financial data to separately identify information and monetary policy shocks. In addition, they focus on the co-movement of the US term premium and interest rates to distinguish between the aforementioned shocks. These authors also apply a slightly different set of sign restrictions about the responses of some variables to the monetary and information shocks. Further details about our methodology to identify the shocks are given in Section 2.

account, foreign direct investment, domestic consumption, and investment. The responses of domestic interest rates in turn seem to be consistent with the scenario in which the domestic central bank tightens monetary policy to counteract the spillover effects stemming from a stronger US economy. These effects may lead to demand pressures that could outweigh the impact of the exchange rate appreciation on inflation.

In turn, we find that a restrictive monetary policy shock that is purged from information effects is followed by a tightening of financial conditions, as well as a contraction in real activity and a lower inflation in Mexico. In line with Mackowiak (2007), Iacoviello and Navarro (2019), and Carrillo et al. (2020), among others, the response of economic activity in Mexico looks considerably similar to that in the US after a purged monetary policy shock, particularly in terms of the direction, persistence, and shape of the output's impulse response function. In addition, we also find that the 10-year government bond yield increases after a purged monetary policy shock. In turn, the 3-month interest rate features a somewhat muted response. Thus, our results indicate that distinguishing between monetary policy and central bank information shocks may be important for a better understanding of the international transmission of US monetary policy. In fact, the effects of both shocks are considerably different. Thus, not accounting for the presence of central bank information shocks may lead to biased measurements of the international transmission of monetary policy. In particular, a key difference from the standard HFI of monetary policy shocks, which does not control for the information content of the announcements, is that the purged monetary policy shock leads to a significant contraction in real activity and a lower price level, as well as a tightening in financial conditions in Mexico. The latter is characterized by higher sovereign risk premium, a more depreciated Mexican peso/US dollar exchange rate, higher forex volatility, and lower stock price index.

The remainder of this paper is organized as follows. Section 2 describes the VAR model for a small open economy and the data used in the estimation. The estimation results are reported and discussed in Section 3. The last section concludes and discusses topics for future research.

2 Methodology

To analyze the impact of FOMC announcement surprises on macroeconomic and financial conditions in Mexico, we estimate a VAR model. Specifically, we employ a two-country VAR model with block-exogeneity, which has been widely used in the literature to analyze the propagation of external shocks into EMEs [see Canova (2005), Fink and Schüler (2015), and Carrillo et al. (2020), among others]. The model is estimated using Bayesian techniques, assuming that the US economy is block-exogenous with respect to the Mexican economy. This assumption is supported by the data, as Mexican variables do not appear to Granger cause US variables (Carrillo et al., 2020).[10] The model allows us to combine two approaches to shock identification: HFI and sign restrictions. In particular, we disentangle monetary policy shocks from contemporaneous central bank information shocks using high-frequency changes of interest rates and stock market surprises in a narrow window around the policy announcement in the US. We then obtain the dynamic responses of Mexican variables to these shocks. The sample period covers a single monetary-policy regime in Mexico, namely inflation targeting, which has allowed to reach lower and more stable values of inflation and other nominal variables in this country. The VAR model is estimated using monthly data from December 2001 to June 2019.

2.1 A VAR Model for a Small Open Economy

The reduced form representation of the VAR model is as follows:

$$\boldsymbol{Y}_t = \boldsymbol{c} + \sum_{p=1}^{P} \boldsymbol{B}^p \boldsymbol{Y}_{t-p} + \boldsymbol{u}_t \tag{1}$$

[10] Canova (2005) also verifies this assumption by running individual VAR models for eight EMEs, including Mexico, and examining the exogeneity of the US block with respect to the block of EMEs variables. The results of this exercise show that the null hypothesis that current and lagged values of EMEs variables have zero coefficients in the US block is not rejected for any of the eight countries.

or

$$\begin{pmatrix} \boldsymbol{m}_t \\ \boldsymbol{y}_t^* \\ \boldsymbol{y}_t \end{pmatrix} = \begin{pmatrix} 0 \\ \boldsymbol{c}_{y^*} \\ \boldsymbol{c}_y \end{pmatrix} + \sum_{p=1}^{P} \begin{pmatrix} 0 & 0 & 0 \\ B^p_{y^*m} & B^p_{y^*y^*} & 0 \\ B^p_{ym} & B^p_{yy^*} & B^p_{yy} \end{pmatrix} \begin{pmatrix} \boldsymbol{m}_{t-p} \\ \boldsymbol{y}^*_{t-p} \\ \boldsymbol{y}_{t-p} \end{pmatrix} + \begin{pmatrix} \boldsymbol{u}_t^m \\ \boldsymbol{u}_t^{y^*} \\ \boldsymbol{u}_t^y \end{pmatrix} \tag{2}$$

where $\boldsymbol{Y}_t$ represents a vector of macroeconomic and financial variables, $\boldsymbol{c}$ is a vector of constants, $\boldsymbol{B}^p$ is the coefficient matrix at lag p, and $\boldsymbol{u}_t \sim N(0, \boldsymbol{\Sigma})$ is a vector of white noise residuals normally distributed. The vector $\boldsymbol{Y}_t$ consists of both high-frequency surprises, denoted by $\boldsymbol{m}_t$, and monthly variables $\boldsymbol{y}_t^*$ and $\boldsymbol{y}_t$. In particular, $\boldsymbol{m}_t$ is a vector of N_m surprises in US financial instruments, $\boldsymbol{y}_t^*$ is a vector of N_{y^*} US variables, and $\boldsymbol{y}_t$ is a vector of N_y domestic (Mexican) variables. As in Jarociński and Karadi (2020) and Breitenlechner et al. (2021), $\boldsymbol{m}_t$ is constructed by adding up the intraday surprises occurring in month t on days with FOMC announcements (assuming zero for months with no announcements). In addition, we assume that the high-frequency surprises are exogenous with respect to the monthly variables in Eq. (2). Thus, we set the corresponding coefficients of $\boldsymbol{m}_t$ in the coefficient matrices and in the vector of constants to zero as in Jarociński and Karadi (2020) and Breitenlechner et al. (2021). This implies that $\boldsymbol{m}_t = \boldsymbol{u}_t^m$.[11] Likewise, following Fink and Schüler (2015) and Carrillo et al. (2020), we assume that $\boldsymbol{y}_t^*$ is block exogenous with respect to the domestic variables, i.e. the dynamics of $\boldsymbol{y}_t$ does not affect any of the values in $\boldsymbol{y}_t^*$.[12] We include one lag in our baseline specification, as is common in studies analyzing the international transmission of US monetary policy.[13]

We estimate the model using a Bayesian approach, as it is standard in the sign restrictions literature. In particular, we follow the literature on Bayesian VAR models and use the Minnesota prior of Litterman (1979). The details are provided in Appendix A. We use Gibbs sampling techniques, which allow us to approximate the marginal posterior distribu-

[11] As long as the financial market surprises are unpredictable, the zero restrictions are plausible as specified above (Jarociński and Karadi, 2020).

[12] This assumption is common in the structural VAR literature on the effects of US monetary policy shocks in EMEs (Mackowiak, 2007).

[13] See, for instance, Canova (2005), Georgiadis (2016), and Dahlhaus et al. (2018), among others. As a robustness check, we also estimate the model using two and three lags. The results reported in section 3 are, in general, consistent with our baseline specification.

tions of the parameters of the model by sampling from their conditional distributions. As in Jarociński and Karadi (2020) and Breitenlechner et al. (2021), we report the results based on 2,000 draws from the Gibbs sampler. In particular, the estimation is implemented with 10,000 iterations of the Gibbs sampling procedure of which the first 2000 are dropped as burn-in and every fourth iteration is saved afterwards.[14]

2.2 Identification

As in Jarociński and Karadi (2020), Breitenlechner et al. (2021), Andrade and Ferroni (2021), and Kerssenfischer (2022), we combine HFI and sign restrictions to identify monetary policy shocks and central bank information shocks. Table 1 provides a summary of our identifying restrictions. Our approach is based on two key assumptions about the announcement surprises $\boldsymbol{m}_t$. First, we assume that $\boldsymbol{m}_t$ are affected only by the two types of announcement shocks, i.e., monetary policy and central bank information shocks, and not by any other shock. This assumption is justified given that $\boldsymbol{m}_t$ are measured in a narrow time window around FOMC announcements.[15] Thus, it is unlikely that shocks unrelated to central bank announcements systematically occur at the same time (Jarociński and Karadi, 2020).

Second, we impose sign restrictions on the impact responses of the surprise measures to identify monetary policy shocks and information shocks. Specifically, we assume that a conventional tightening monetary policy shock is associated with an increase in the short-run interest rate and a decline in stock prices. This relationship is consistent with most economic models, which suggest that a monetary policy tightening results in a contraction that reduces the expected future value of dividends. Moreover, higher interest rates increase the discount rate at which future dividends are discounted. Consequently, the stock price, which is the present discounted value of future dividends in standard asset pricing theory, decreases. Conversely, a central bank information shock is associated with an increase in both the short-run interest rate and stock prices. This positive co-movement likely reflects information con-

[14] We diagnose convergence of the Gibbs sampler by inspecting the sequence of retained draws. We find that the use of a chain ten times longer gave essentially the same results.

[15] This is a standard assumption in the literature on high frequency identification [see e.g. Kuttner (2001) and Gürkaynak et al. (2005)].

Variable	Shock		
	Monetary Policy	CB Information	Other
Intraday Frequency			
Interest rate	+	+	0
Stock price index	-	+	0
Low Frequency			
$\boldsymbol{y}_t^*$			
$\boldsymbol{y}_t$			

Table 1: Identifying Restrictions in the Baseline VAR Model

Notes: Restrictions on the contemporaneous responses of variables to shocks. +, –, and 0 denote the respective sign restrictions and zero restrictions. Blank spaces mean that no restriction is imposed on the variable.

tained in the central bank's announcement regarding the economic outlook (Jarociński and Karadi, 2020). Finally, note that we do not impose restrictions on any macroeconomic and financial variables contained in vectors $\boldsymbol{y}_t^*$ and $\boldsymbol{y}_t$, as outlined in Table 1.

We compute the posterior draws of the shocks and the associated impulse responses assuming a uniform prior on the space of rotations conditionally on satisfying the sign restrictions as in Rubio-Ramirez et al. (2010). In this case, the implemented restrictions only provide set identification, that is, conditionally on each draw of the model parameters there are multiple values of shocks and impulse responses that are consistent with the restrictions. Thus, when computing uncertainty bounds, we take all these values into account weighting them according to a uniform prior on rotations.[16]

Finally, to analyze how the presence of central bank information shocks can affect the standard HFI of monetary policy shocks, we also estimate the responses of the variables in the VAR model to a monetary policy shock without explicitly identifying the information shock. To do so, we use all interest rate surprises as proxies for monetary policy shocks and disregard the accompanying stock price movements in the VAR identified with the Cholesky decomposition. Specifically, we order the interest rate surprise first in Eq. (2). In this case, the identifying restrictions are:

[16] As Jarociński and Karadi (2020) emphasize, a uniform prior on rotations is less restrictive than imposing sign restrictions by means of a penalty function approach as in Uhlig (2005).

$$COV(m_t, \varepsilon_t^{MP}) > 0 \tag{3}$$

$$COV(m_t, \varepsilon_t^{i}) = 0 \text{ for all } \varepsilon_t^{i} \text{ other than } \varepsilon_t^{MP} \tag{4}$$

where m_t denotes the interest rate surprise and ε_t^{MP} the monetary policy shock. Identifying restrictions (3) and (4) have been used by Barakchian and Crowe (2013), Gertler and Karadi (2015), and Jarociński and Karadi (2020), among others.

2.3 Data Description

In this subsection, we provide a description of the variables included in the VAR model. This model is estimated at monthly frequency for the period December 2001 to June 2019. Asset-price changes around FOMC announcements are obtained from an updated version of the data used in Gürkaynak et al. (2005) and available from Gürkaynak et al. (2021), including 147 announcements within this period.[17] The standard monetary policy surprise measure, known as target surprise, is defined as the difference between the announced target fed funds rate and expectations derived from fed funds futures contracts (Kuttner, 2001). In particular, the target surprise is computed using the change in the current month's fed funds futures rate during a 30-minute window around the FOMC announcement, from 10 minutes before to 20 minutes after the announcement. This measure has been widely utilized in the literature and enables us to capture exogenous innovations to the current policy interest rates (Kuttner, 2001; Gurkaynak et al., 2005; Hausman and Wongswan, 2011; Gurkaynak et al., 2021). As it is common in the HFI literature, we assume that within this narrow window, only monetary and information shocks systematically influence financial market surprises. As the payout for fed funds futures contracts is based on the average effective fed funds rate prevailing over the month specified in the contract, the change in the fed funds futures rate is adjusted by a

[17] In February 1994, the FOMC began regularly issuing a press release after every meeting. These releases provide detailed information on the committee's policy decisions regarding changes to the target range for the federal funds rate, including the rationale behind these decisions. Additionally, the press releases contain the FOMC's assessment of the state of the financial markets and the economy. The announcement dates and times correspond to those of the respective press releases.

factor that depends on the number of days in the month affected by the change in the target fed funds rate. In particular, for an FOMC announcement on day d of a month with D days, the target surprise is given by $TS_t = (D/(D-d))(ff_{t+20} - ff_{t-10})$, where ff_{t+20} and ff_{t-10} are the fed funds futures rates 20 minutes after and 10 minutes before the announcement, respectively.[18]

We follow the methodology of studies such as Jarociński and Karadi (2020) and Breitenlechner et al. (2021) and use the change in the S&P 500 index, which comprises 500 large companies, as our baseline measure of the stock price surprise. In particular, the change in this index is computed between 10 minutes before and 20 minutes after the FOMC announcement. As Jarociński and Karadi (2020) emphasize, by using this narrow window, we ensure that the "pre-FOMC announcement drift" in the S&P 500 documented by Lucca and Moench (2015) does not affect our measurement.[19]

The rest of variables included in the VAR model, together with their respective sources and transformations, are presented in Table 2. Considering that output and price measurements reflect developments over the entire month rather than purely at a point in time, all low frequency financial variables described in this table are included as monthly average figures.[20] The variables used in our analysis belong to different categories, including external financial conditions, foreign economic activity and prices, country risk, money market, debt, stocks, foreign exchange, domestic economy activity and prices. We selected these variables following previous studies on the transmission of monetary policy and information shocks, such as Gertler and Karadi (2015), Jarociński and Karadi (2020), Breitenlechner et al. (2021), and Andrade and Ferroni (2021), and also studies on the international transmis-

[18] For robustness, we replace the one-month fed funds futures with the three-month fed funds futures. Moreover, following studies such as Gurkaynak et al. (2005), we also replace our baseline measure with the first principal component of surprises in the current month and three-month fed funds futures, and the two-, three-, and four-quarters ahead three-month eurodollar futures. The results reported in Section 3 are consistent with our baseline specification.

[19] As Lucca and Moench (2015) show for the period 1994 to 2011, this index tended to increase by 49 basis points on average in the 24 hours prior to scheduled FOMC announcements. However, this drift is uncorrelated with the responses of either the fed funds futures or the S&P 500 to the announcements within the half-hour windows (Lucca and Moench, 2015). We confirm that our sample contains no discernible drift, with an average 30-minute S&P 500 return of about 0.6 basis points and a standard deviation of 43 basis points.

[20] The exception is the shadow interest rate considering that data are just available as end-of-month figures.

sions of US monetary policy including Mackowiak (2007), Vicondoa (2019), Iacoviello and Navarro (2019), and Carrillo et al. (2020), among others. For the US, we use the shadow interest rate by Wu and Xia (2016), i_t^*, which allows us to account for the unconventional monetary policies implemented by the Federal Reserve after the global financial crisis.[21] In particular, the shadow interest rate mirrors the federal funds rate in times of conventional monetary policy, that is, when the latter is above its effective lower bound. However, the shadow interest rate is not bounded below by 0 percent.[22] In addition, we use the monthly average of the S&P 500 as the stock price index. The measures of real output and the price level in the US are the Industrial Production Index and the Consumer Price Index, Y_t^* and P_t^*, respectively. Finally, as an indicator of financial conditions in this country, we include the excess bond premium EBP_t introduced by Gilchrist and Zakrajšek (2012). This premium is the average corporate bond spread that is purged from the impact of default compensation and, as argued by Gilchrist and Zakrajšek (2012), provides an effective measure of investor sentiment or risk appetite in the corporate bond market (Favara et al., 2016).[23]

Regarding the macroeconomic and financial variables of Mexico, we include the 10-year government bond yield (i_t^{10y}) and the 3-month interest rate on Mexican Treasury bills, CETES, i_t^{3m}.[24] In addition, we include the 5-year credit default swap (CDS) as an indicator

[21] In particular, the Federal Open Market Committee targeted the federal funds rate between 0 to 0.25 percent from December 16, 2008, to December 15, 2015. In this "zero lower bound" environment, some studies have used shadow rate models to characterize the term structure of interest rates or quantify the stance of monetary policy (Kim and Singleton, 2012; Bauer and Rudebusch, 2016; Krippner, 2013 and Wu and Xia, 2016).

[22] It is worth mentioning that the inclusion of the shadow rate in the VAR model is not intended to capture the Federal Reserve's forward guidance strategy. The main reason for incorporating this rate is to avoid introducing non-linearity into the model, which would arise if we used the federal funds rate during periods constrained by the effective lower bound. Although the shadow rate is a useful measure of the overall stance of monetary policy during unconventional times, it does not capture the effects of forward guidance during conventional times, as it closely tracks the federal funds rate in those periods. A more suitable variable for capturing forward guidance effects across the entire sample might be the one-year rate, as employed by Jarociński and Karadi (2020), or the two-year rate, based on the arguments of Gertler and Karadi (2015), Swanson and Williams (2014) and Hanson and Stein (2015), among others, who suggest that the Federal Reserve's forward guidance strategy seems to operate with a roughly two-year horizon. However, since the focus of our study is not on forward guidance, we have chosen not to include these longer-term rates in our model.

[23] This variable also aggregates high-quality forward-looking information about the economy (Gilchrist and Zakrajšek, 2012; Favara et al., 2016). Thus, it improves the reliability and forecasting performance of small-scale VARs (Caldara and Herbst, 2019).

[24] CETES are debt issued by the Federal Government through the Ministry of Finance and the Central Bank of Mexico. We use zero coupon interest rates for both maturities in order to obtain comparable interest rates, as each bond pays different coupons for each maturity.

of country/default risk.[25] Because CDSs provide insurance on a bond default, we can infer the market's estimate of the likelihood of default directly from the price of this variable. The Mexican Stock Exchange Index is included to capture the Mexican stock market behavior. As indicators of the exchange rate market, we include the nominal peso-dollar exchange rate and the volatility implied in one-month options of the Mexican peso. We use the Overall Indicator of Economic Activity (IGAE by its Spanish acronym), Y_t, as an indicator for domestic output.[26] Domestic price level is measured by the consumer price index P_t. The variables Y_t and Y_t^* are seasonally adjusted by their respective statistical offices. For the case of domestic prices, P_t, this variable is seasonally adjusted with the X13-ARIMA method. Finally, the sample period starts in 2001M12 as data on bond yields for some maturities are available from this date onwards. Hence, the beginning of our sample coincides with the adoption of an inflation targeting regime by Banco de México from 2001.[27] In turn, the sample finishes in 2019M06 as data on monetary policy and stock price surprises from Gürkaynak et al. (2021) are available until this date. This implies that the COVID-19 period is excluded from the analysis. Thus, we avoid the influence of the extreme observations of the COVID-19 period on our estimates.

[25] A CDS is a financial derivative that provides insurance against sovereign default. Its price is comparable to the payment of an insurance premium against such an event. An increase in its price reflects a rise in the assessment made by financial market participants that the risk of default materializes. For robustness, we also estimate the model using an alternative country-risk indicator, particularly the Mexico's EMBI plus spread. The results reported in Section 3 are consistent with our baseline specification.

[26] This indicator employs the methodology and the conceptual framework of the national accounts, in particular, GDP. The correlation between the Overall Indicator of Economic Activity and GDP for Mexico, both variables measured in quarterly percentage changes, is 0.99. IGAE is subject to revisions. That is, the data actually available to the central bank at a particular month may differ from the final revised values released by the statistical office. Although it would be of interest to conduct the empirical exercise with real time data in order to provide further information, this type of data for Mexican output is unavailable. Thus, we use revised data in our estimations.

[27] According to Chiquiar et al. (2010), inflation in Mexico went from being a non-stationary process to being a stationary process around the end of 2000 or the beginning of 2001. In addition, Gaytán and González García (2007) find that monetary policy transmission mechanism seems to have presented a structural change after this period. Low and stable inflation has provided certainty to financial contracts, reducing the risk premium of interest rates, and allowing for longer term contracts.

Category	Series	Definition	Transformation	Source
		US surprises contained in vector $\boldsymbol{m}_t$		
Money market	m_t^1	Monetary policy surprise	No transformation	Gürkaynak et al. (2021)
Stock market	m_t^2	Stock price surprise	No transformation	Gürkaynak et al. (2021)
		US variables contained in vector $\boldsymbol{y}_t^*$		
Money market	i_t^*	Shadow interest rate	No transformation	Wu and Xia (2016)
Stock market	$SP500_t$	S&P 500	$\log SP500_t$	Bloomberg
Financial conditions	EBP_t^*	US excess bond premium	No transformation	Favara et al. (2016)
Economic activity	Y_t^*	US Industrial Production Index, monthly	$\log Y_t^*$	FRED
Prices	P_t^*	Consumer Price Index, monthly	$\log P_t^*$	FRED
		Mexican variables contained in vector $\boldsymbol{y}_t$		
Debt market	i_t^{3m}	3-month Treasury Certificates interest rate	No transformation	Banco de México
	i_t^{10y}	10-year government bond yield	No transformation	Banco de México
Country risk	CDS_t	5-year credit default swap (CDS)	No transformation	Bloomberg
Stock market	IPC_t	Mexican Stock Market Index	$\log IPC_t$	Grupo BMV
FX market	e_t	Exchange rate	$\log e_t$	Banco de México
	σ_t	Volatility implied in one-month options of the Mexican peso	No transformation	Bloomberg
Economic activity	Y_t	Global Economic Activity Index, monthly	$\log Y_t$	INEGI
Prices	P_t	Consumer Price Index, monthly	$\log P_t$	INEGI

Table 2: Data Description

3 The International Transmission of US Monetary Policy and Central Bank Information Shocks

3.1 Impulse Response Functions

In this section, we present the impulse responses of the variables in the VAR model to both monetary policy and central bank information shocks, identified using sign restrictions on the high-frequency co-movement of interest rates and stock prices. Additionally, we report the estimated responses to a monetary policy shock identified using the standard Cholesky identification scheme, without explicitly identifying the central bank information shock. In

all cases, the size of the shock is one standard deviation. Due to coefficient restrictions in the VAR system, the announcement surprises in m_t are independent and identically distributed [see Eq. (2)]. That is, they only respond to shocks on impact, and their impulse response function is zero in all other periods. Table 3 reports their impact responses. By construction, the impact responses satisfy the sign restrictions. Specifically, the pure monetary policy shock is associated with a 1.5 to 5.0 basis points (bp) increase of the current-month fed funds futures and a 13.0 to 40.9 bp drop in the S&P 500 index in the 30-minutes window. In turn, the central bank information shock is associated with a 1.3 to 4.9 bp increase in the current-month fed funds futures and a 10.1 to 40.3 bp increase in the S&P 500 index.[28,29] With regard to the monetary policy shock identified by the standard HFI, Panel B of Table 3 shows that this shock is associated with an increase of approximately 5.2 bp in the fed funds futures, which is the only high-frequency surprise included in m_t.[30] As mentioned in Section 2, in this case, we use all the surprises in the fed funds futures as proxies for monetary policy shocks and disregard the accompanying stock price movements.

We begin our analysis by presenting the impulse response functions of US variables to

[28] Figure B.1 in Appendix B presents the monetary policy and central bank information shocks over time, showing that these shocks are distributed throughout our sample period rather than being clustered in specific intervals. This distribution suggests that both types of shocks may play an important role in shaping economic outcomes. Two notable episodes are a sequence of negative central bank information shocks occurred in the aftermath of the dot-com bubble (2001-2002) and the 2008 financial crisis. In both instances, the Federal Reserve significantly reduced the fed funds rate in response to worsening economic conditions and heightened uncertainty. These actions, coupled with the Fed's pessimistic economic outlook, prompted market participants to lower their growth expectations, resulting in positive co-movements of interest rates and stock market changes. For the common sample period, our estimated shocks exhibit a strong correlation with those identified by Jarociński and Karadi (2020), who utilize three-month fed funds futures instead of current-month fed funds futures. Specifically, the correlation is approximately 0.93 and 0.82 for monetary policy shocks and central bank information shocks, respectively. As explained before, our rationale for using one-month fed funds futures stems from our primary focus, which is not on studying forward guidance effects, in contrast to Jarociński and Karadi (2020). Rather, our interest lies primarily in conventional monetary policy shocks associated with exogenous innovations in the current policy rate level.

[29] Although the variance of the monetary policy surprise experiences a significant reduction during the zero lower bound period, surprises related to the S&P 500 contribute to explain an important part of the variation in the shocks during this period. In addition, as mentioned in Section 2.3, the VAR model includes the shadow interest rate by Wu and Xia (2016). This rate allows us to account for the unconventional monetary policies implemented by the Federal Reserve in the aftermath of the global financial crisis.

[30] As is common in the HFI literature, the policy rate surprises are small relative to the raw changes, indicating that most of the Federal Reserve's policy rate decisions are anticipated by market participants (Nakamura and Steinsson, 2018). According to Gurkaynak et al. (2005), unanticipated changes in the fed funds rate capture only a small fraction of the monetary policy news associated with FOMC announcements.

High Frequency	Panel A. Sign Restrictions				Panel B. Standard HFI	
	Monetary Policy Shock		CB Information Shock		Monetary Policy Shock	
Variable	Mean	$(16^{pct}, 84^{pct})$	Mean	$(16^{pct}, 84^{pct})$	Mean	$(16^{pct}, 84^{pct})$
One-month fed funds futures	3.8	(1.5, 5.0)	3.6	(1.3, 4.9)	5.2	(4.9, 5.4)
S&P 500	-31.0	(-40.9, -13.0)	29.2	(10.1, 40.3)		

Table 3: Impact Responses of Announcement Surprises to Shocks

Notes: Posterior means and posterior percentiles 16 and 84. The shocks are measured in basis points.

monetary policy and central bank information shocks using a sign restriction identification approach. Panel A of Figure 1 shows the results for a 24-month horizon along with 68% highest posterior density intervals (HPDIs) associated with the 16th and 84th percentiles of the distribution.[31] Our findings are consistent with recent studies such as Jarociński and Karadi (2020) and Breitenlechner et al. (2021), which identify two distinct shocks with different effects on the US economy. Our results show that a monetary policy shock leads to an immediate and significant increase of about 4.3 bp in the policy rate. Moreover, stock prices experience a persistent drop, while the EBP increases by about 1.0 and 0.1 percent, respectively, suggesting a tightening of financial conditions. Additionally, real output and the price level experience persistent declines of about 0.2 and 0.1 percent, respectively. These results are in line with the standard transmission channels of monetary policy and are consistent with previous empirical studies using HFI by Gertler and Karadi (2015), Ramey (2016), Jarociński and Karadi (2020), and Miranda-Agrippino and Rey (2020), among others. In accordance with Jarociński and Karadi (2020), the effects of these shocks on stock prices, real output, and the price level exhibit persistence, although they return to their baseline levels after a few years.

Regarding the central bank information shock, the results from the second column of Figure 1 indicate that this shock has a greater and more persistent impact on the policy rate than the monetary policy shock. Moreover, financial conditions appear to improve follow-

[31] For results at a 72-month horizon, please refer to Figures B.8 and B.9 in Appendix B. In the main text, we have chosen to present the results for a 24-month horizon to better illustrate the short-term responses of the analyzed variables.

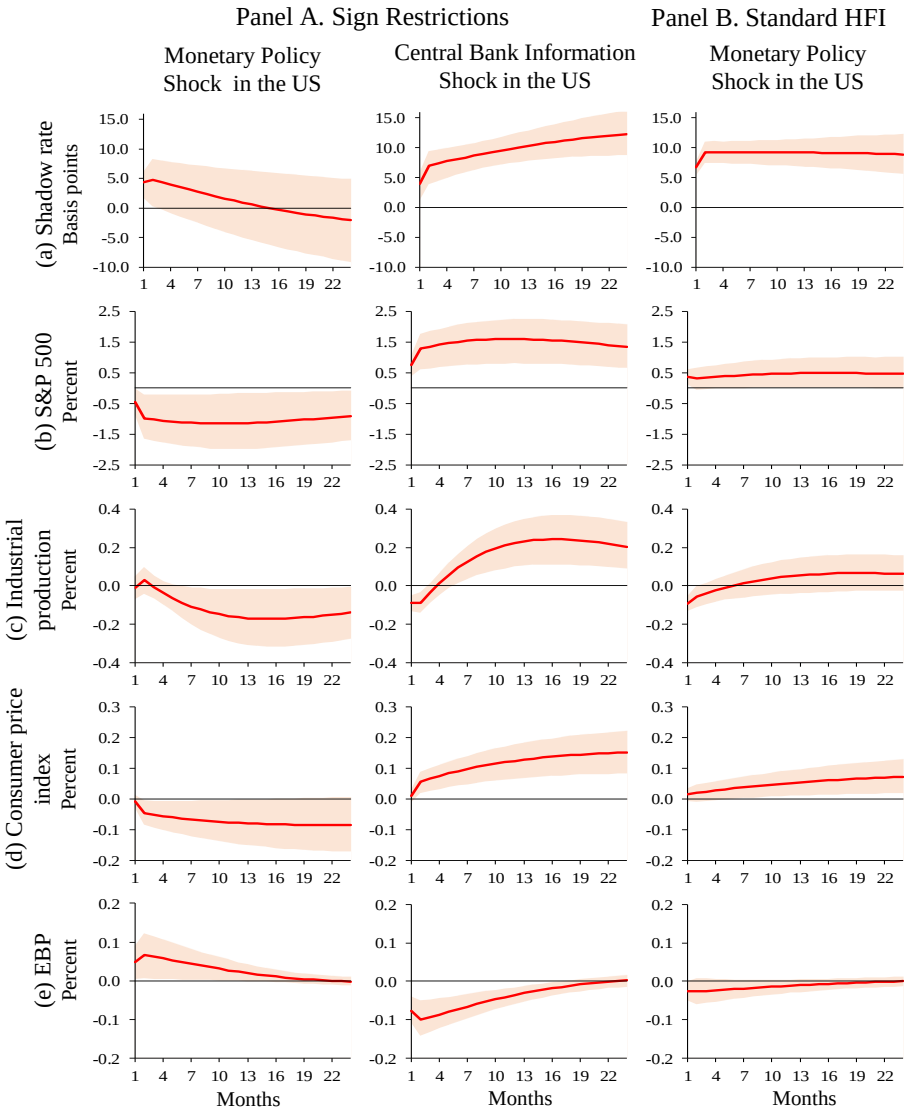

Figure 1: Impulse Response Functions of US Variables to the Monetary Policy and Central Bank Information Shocks

Notes: Responses are presented for a 24-month horizon with the associated 68% highest posterior density intervals (HPDIs).

ing the information shock. Specifically, the shock has a positive persistent effect on stock prices of about 1.5 percent and a negative effect on the excess bond premium of about 0.1 percent two months after the impact. Additionally, the central bank information shock leads to a significant increase in output and the price level. These responses are consistent with the scenario in which the central bank communicates positive news about the economy and

tightens monetary policy to partly offset the effects of the news and prevent demand pressures on inflation, as noted by Jarociński and Karadi (2020). The positive news may lead private decisions to positively influence aggregate demand, resulting in an increase in real output and the price level. The improved economic conditions may increase the expected value of future dividends and the risk appetite in the corporate bond market, leading to an increase in stock prices and a decrease in the EBP.

Moving to Panel B of Figure 1, we present the responses of the US variables to a monetary policy shock without explicitly identifying the central bank information shock. As can be seen, there are notable differences between the responses presented in Panel B of Figure 1 and the "pure" responses to a conventional monetary policy shock reported in the first column of this figure, which are purged from the impact of central bank information shocks. The results reveal some puzzling findings. In particular, the consumer price index appears to increase after the monetary policy shock. Furthermore, the monetary policy shock appears to result in an increase in the S&P 500, contradicting standard theory, and reduce the excess bond premium, thus leading to an improvement in financial conditions. Despite the increase in stock prices and the improvement in financial conditions, real output decreases in response to this shock, reaching its lowest point immediately. An additional bias in the standard HFI approach is that the policy rate response is higher and more persistent. In line with Jarociński and Karadi (2020) and Breitenlechner et al. (2021), this puzzling results could be due to the presence of the central bank information shocks, which seem to have greater and longer-lasting effects on interest rates.

We now turn to the Mexican variables, the primary focus of this study. Panel A of Figure 2 shows the responses of these variables to both monetary policy and information shocks in the US. As can be seen in the first column of this figure, a conventional monetary policy shock in the US is followed by a persistent decrease in real output in Mexico, with domestic prices decreasing by approximately 0.1 percent 15 months after the shock. In line with Mackowiak (2007), Iacoviello and Navarro (2019), and Carrillo et al. (2020), the dynamics of economic activity in Mexico after a monetary policy shock is similar to that observed in the US (see Panel A of Figure 1). Specifically, real output falls as much as industrial production in the US,

with a median peak effect of approximately 0.2 percent, which is reached 17 months after the shock. This decrease shows persistence, as real output continues to decline for approximately two years. On the other hand, regarding the effects of the central bank information shock, we find that a positive co-movement of interest rates and stock prices in the US leads to an increase in output and the price level in Mexico. In particular, domestic prices increase by around 0.1 percent 12 months after the shock. While our analysis does not delve into individual transmission channels, we conjecture that in response to unexpected good news in the US, private decisions in both countries could positively influence aggregate demand in Mexico through various channels, including the current account, foreign direct investment, domestic consumption and investment, among others.

We now analyze the responses of the exchange rate and its volatility to the monetary and information shocks. Our results are in line with those of Mackowiak (2007) and Carrillo et al. (2020), who find that the Mexican peso depreciates after a monetary policy shock in the US. In particular, we find that the exchange rate and its volatility increase by approximately 0.5 percent and 0.5 percentage points, respectively, two months after the shock. This finding may be explained by the behavior of international investors who actively seek higher yields. In response to an increase in US interest rates, these investors tend to allocate their capital towards US assets, which in turn leads to an appreciation of the US dollar. In contrast, following a central bank information shock, we find that the exchange rate and its volatility decrease by approximately 0.5 percent and 1.0 percentage point, respectively, immediately after the shock. These results may suggest that favorable news about the US economy could stimulate investment in the local currency market by economic agents. This is consistent with Baek (2006), De Vita and Kyaw (2008) and Forbes and Warnock (2012), who suggest that a better foreign economic outlook that also benefits the domestic economy may trigger portfolio flows to EMEs, thereby leading to an appreciation of their local currency.

As Panel A of Figure 2 illustrates, a conventional monetary policy shock leads to increases in both the 10-year and 3-month government bond yields, by approximately 9 and 4 basis points respectively, one month after the shock. Thus, the response of the 3-month yield is somewhat muted. Given the important degree of financial integration between both countries,

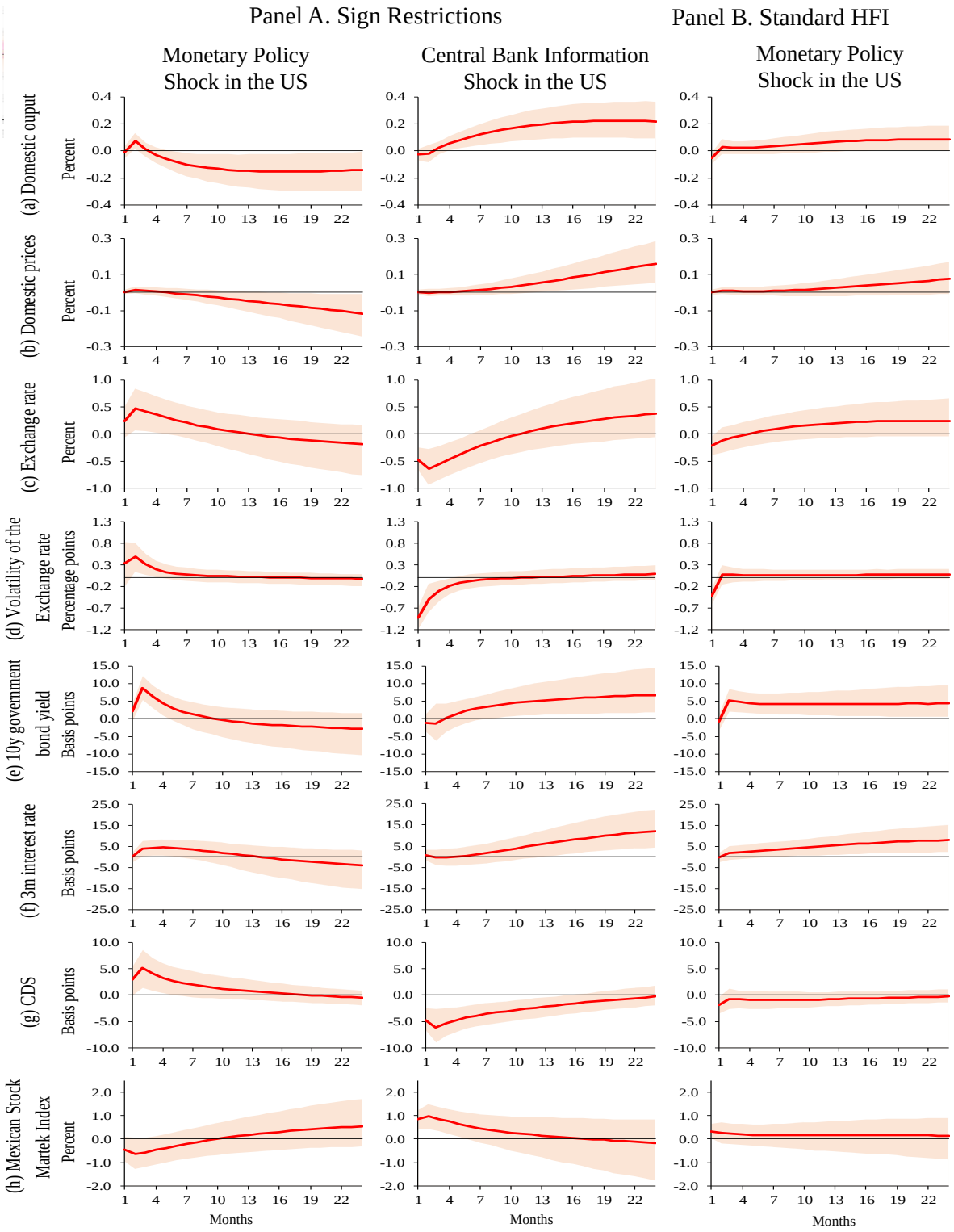

Figure 2: Impulse Response Functions of Mexican Variables to the Monetary Policy and Central Bank Information Shocks in the US

Notes: Responses are presented for a 24-month horizon with the associated 68% highest posterior density intervals (HPDIs).

domestic interest rates are expected to be influenced by interest rates in the US (Banco de México, 2022). This result may be explained by a process of financial arbitrage between domestic and foreign bonds, in particular for those of medium and long maturities. Thus, the differences in the magnitudes of the responses between the two rates could be explained by various factors that affect long-term interest rates, such as inflation expectations, the expected path for short-term interest rates, and the term premium.[32] If the monetary authority raises short-term interest rates and economic agents expect further increases in the future, long-term interest rates may adjust upwards. Furthermore, if the term premium increases in this scenario, long-term interest rates may increase even more than short-term interest rates. In particular, the response of the term premium to monetary policy announcements may depend on the US term premium's reaction to such announcements.[33]

According to Carrillo et al. (2020), the response of the US term premium seems to depend on whether investors perceive the announcements' information as negative or positive about the economic outlook. The term premium tends to rise if investors perceive that the economy is heading towards a downturn, as in the case of a conventional monetary policy shock. Conversely, the term premium tends to decrease if investors perceive that the news accompanying the policy announcement indicate a favorable economic outlook, as in the case of an information shock. Thus, if both premiums in the US and Mexico move in the same direction, long-term interest rates are likely to increase even more than short-term interest rates following a monetary policy shock. Conversely, after a central bank information shock, we would expect the opposite to occur. In fact, our findings indicate that the yields on the 10-year bond and the 3-month CETES tend to increase with a lag following a central bank information shock. In this case, however, the median response of the short-term interest rate is larger than that of the long-term rate 24 months after the shock. The long-term rate increases by approximately 7 bp and the effects are very persistent. In contrast, the short-term

[32] The term premium is the additional compensation that investors require for holding longer-term financial instruments instead of short-term ones.

[33] As discussed by Caceres et al. (2016) and Carrillo et al. (2020), changes in US interest rates may transmit to long-term interest rates in small open economies primarily through unexpected changes in the US term premium.

interest rate rises by about 12 bp and reverts to the baseline at a slower pace. These results are consistent with the scenario in which, because of unexpected good news about the US economy, the term premium falls and the central bank tightens monetary policy to prevent demand pressures on inflation.

In line with previous studies such as Uribe and Yue (2006), Vicondoa (2019), and Carrillo et al. (2020), our findings indicate that a monetary policy shock in the US results in an increase in the sovereign risk premium in Mexico, as measured by the price of the 5-year CDS (see Panel A of Figure 2). Specifically, we observe a 6 bp increase in the CDS price two months after the shock. This result may be linked to the reduction in investors' risk appetite due to the rise in US interest rates (Kamin and Von Kleist, 1999), which may lead them to demand higher premiums for holding higher-risk assets in their portfolios. Conversely, we find that the sovereign risk premium decreases by approximately 6 bp after the central bank information shock. It is worth noting that the impact of this shock on the CDS price seems to be more long-lasting than that of the monetary shock, suggesting that US central bank communication can have an important influence on the market's assessment of the likelihood of sovereign default in the Mexican economy. Specifically, positive news about the US economy may lower the sovereign risk premium in Mexico by generating expectations of improved financial and macroeconomic conditions in the country.

As can be seen in Panel A of Figure 2, a conventional monetary policy shock leads to a decline in Mexican stock prices as indicated by the median response. This result could be associated with the effects of this shock on both expectations on future economic activity and short-term interest rates in Mexico. Specifically, the expected decline in economic activity associated with the monetary policy shock may result in a reduction of the expected value of future dividends, while the increase in interest rates may elevate the discount rate applied to these dividends. Consequently, the stock prices decrease. In contrast, an interest rate increase accompanied by a stock price increase in the US, indicating positive news about the US economy, results in higher stock prices in Mexico. In such a case, the high expected value of future dividends, resulting from improved expectations about the Mexican economy following the information shock, may offset the increase in the discount rate with which

these dividends are discounted. This result reaffirms the importance of US central bank communication to affect the expectations of economic activity in Mexico and subsequently influencing stock prices.

The results presented previously highlight the relevance of central bank communication, suggesting that neglecting this channel in the estimation could lead to biased estimates of monetary policy shocks and obscure our understanding of the transmission of US monetary policy to an EME like Mexico. To shed further light on this matter, Panel B of Figure 2 shows the responses of Mexican variables in the VAR model to a monetary policy shock without explicitly identifying the central bank information shock. In this case, the results do not seem to be consistent with a branch of the literature [Canova (2005), Uribe and Yue (2006), Ilzetzki and Jin (2013) and Vicondoa (2019)], which finds that US monetary policy shocks induce contractionary effects on real activity in EMEs, as well as a tightening of financial conditions in such economies. As can be seen in Panel B of Figure 2, the monetary policy shock produces a delayed positive effect on economic activity in Mexico. In turn, we observe decreases in the sovereign risk premium, the exchange rate, and its volatility, as well as an increase in stock prices following a contractionary monetary policy shock. An additional difference in the standard HFI is that the interest rate responses are more persistent. These results may be explained by the presence of the central bank information shocks, which seem to have greater and more persistent effects on all analyzed variables. This, in turn, could affect the long-term dynamics of interest rates.

In sum, our results highlight the importance of central bank communication in the US to influence macroeconomic outcomes into the Mexican economy. Strong relationships in trade, investment, and financial sectors might help explain the co-movement between economic activity, prices, and interest rates in both countries. In turn, our results indicate that distinguishing between monetary policy and central bank information shocks is also important for understanding the international transmission of US monetary policy. In fact, the effects of both shocks are considerably different and not accounting for the presence of central bank information may lead to biased measurements of international monetary shocks. A key difference from the standard HFI of monetary policy shocks that does not control for

the information content of the announcements is that the purged tightening monetary policy shock lead to a significant contraction in real activity and a lower price level, as well as a tightening in financial conditions in both countries.

3.2 Central Bank Information about Aggregate Supply

Up to this point, we have identified monetary shocks and information shocks by using sign restrictions on high- frequency surprises associated with the co-movement of interest rates and stock prices around FOMC announcements. Our analysis suggests that information shocks behave like aggregate demand shocks, as both output and price levels move in the same direction after the shock. However, as discussed in Jarocinski and Karadi (2020) and Andrade and Ferroni (2021), central bank communication encompasses not only factors that influence demand but also those affecting supply, such as the level of technology and potential output. A key characteristic of these shocks is that output and prices move in the opposite direction.

In this exercise, we distinguish between central bank information shocks related to demand and supply. To achieve this, we add a new high-frequency inflation surprise measure to the vector m_t and impose a set of additional restrictions. As in Jarocinski and Karadi (2020) and Breitenlechner et al. (2021), we use the daily change in the two-years-ahead break-even inflation rate on the day of the FOMC announcement. This variable reflects changes in market-based inflation expectations around policy announcements and is constructed by taking the difference between the two-year constant-maturity yields of nominal and real (inflation-protected) Treasuries (Gürkaynak et al., 2010).[34] Table 4 presents our new set of identifying restrictions. By examining the co-movement of stock prices, which are closely related with the economic outlook, and inflation expectations, which measure perceived price pressures, we are able to differentiate between central bank information shocks related to aggregate demand and those related to aggregate supply. If stock prices and inflation expectations co-move positively, we categorize it as a demand shock. Conversely, if they co-move negatively, we categorize it as a supply shock. Note that after a monetary policy

[34] Data are available since 2004 and are obtained from Gürkaynak et al. (2010).

Variable	Shock			
	Monetary Policy	CB Information about Demand	CB Information about Supply	Other Other
High Frequency				
Interest rate	+	+		0
Stock price index	-	+	+	0
Break-even inflation	-	+	-	0
Low Frequency				
$\boldsymbol{y}_t^*$				
$\boldsymbol{y}_t$				

Table 4: Identifying Restrictions in the VAR Model with Central Bank Information about Supply

Notes: Restrictions on the contemporaneous responses of variables to shocks. +, –, and 0 denote the respective sign restrictions and zero restrictions. Blank spaces mean that no restriction is imposed on the variable.

tightening, inflation expectations are expected to fall. As in Jarocinski and Karadi (2020) and Breitenlechner et al. (2021), we leave the fed funds futures surprise unrestricted after a central bank information shock related to supply.

Figure 3 reports the responses of Mexican variables to the three shocks we now identify. As can be seen, the responses to the monetary policy and central bank demand information shocks are robust to adding a new high-frequency observable and a third shock. In general, the new central bank supply information shock we added does not account for much of the variability of the macroeconomic and financial variables, as witnessed by the near-zero impulse responses. Specifically, central bank information shocks related to supply lead to a significant decrease in the sovereign risk premium and the volatility of the exchange rate. Additionally, they result in a significant increase in the stock price index, also implying an improvement in financial conditions. In line with Jarocinski and Karadi (2020) and Breitenlechner et al. (2021), our results suggest that interest rate and stock market surprises seem to be sufficiently informative to identify monetary policy and central bank information shocks, which seem to relate more to the demand side. In this vein, high-frequency surprises in break-even inflation rates add marginal independent information. Overall, our previous conclusions remain robust also under this more refined identification.

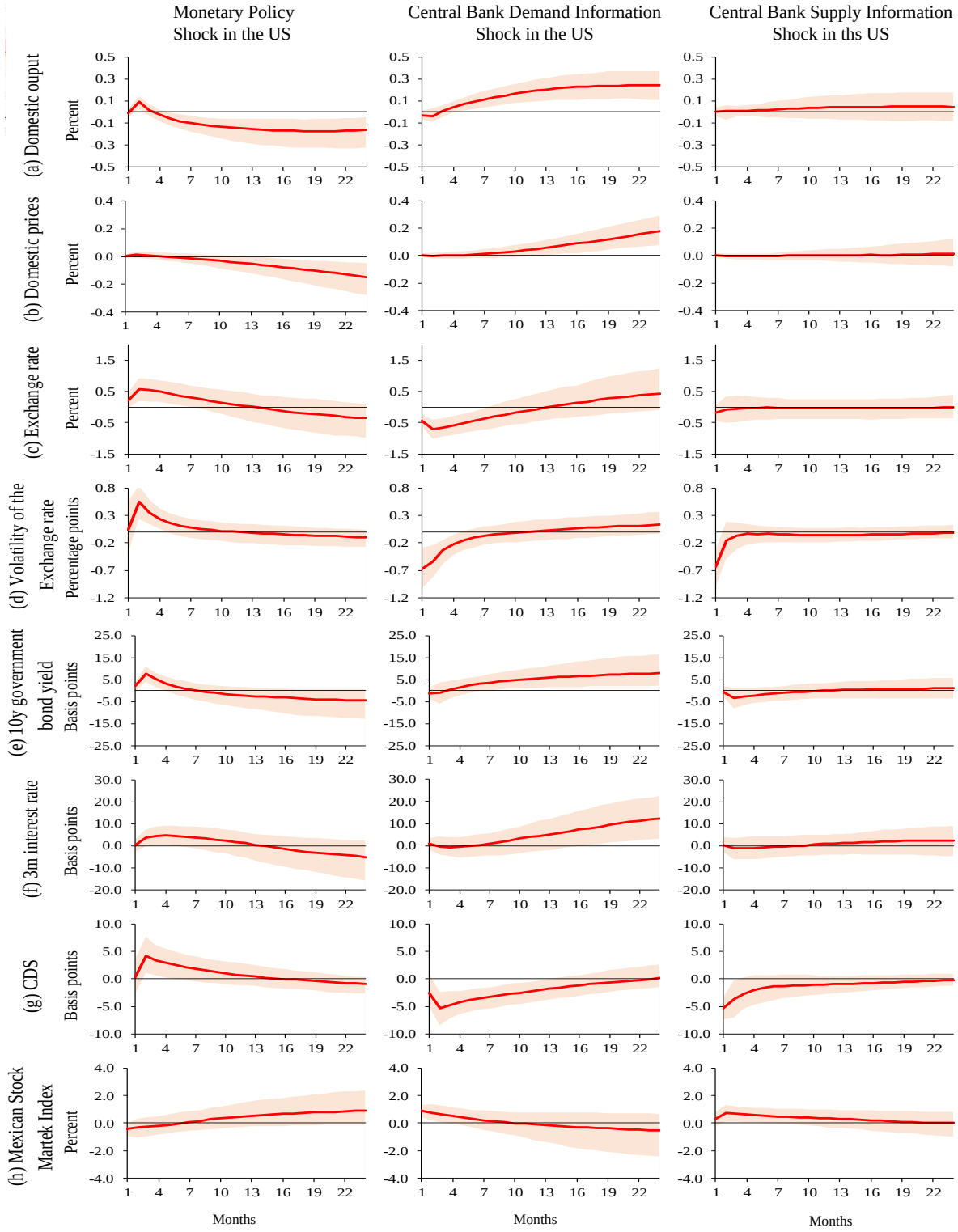

Figure 3: Impulse Response Functions of Mexican Variables to a Monetary Policy Shock and Central Bank Information Shocks related to Demand and Supply in the US

Notes: Responses are presented for a 24-month horizon with the associated 68% highest posterior density intervals (HPDIs).

3.3 Robustness Exercises

In this subsection, we present several robustness checks. In particular, we analyze the responses of additional variables included in the baseline model. We also investigate the sensitivity of our results by considering additional lags and alternative measures of monetary policy surprises. Finally, we estimate the model using an alternative country-risk indicator. The impulse response functions of the Mexican variables are reported in this section, while those corresponding to the US variables are available from the authors upon request.

3.3.1 Additional Variables

Figure B.2 of the Appendix B reports the responses of additional variables that we add, one by one, to the baseline model. In particular, we include the inflation expectations for the next 12 months and four years in Mexico.[35] We find that the two shocks that we identify by sign restrictions also have opposite effects on these variables. In particular, a monetary policy shock generates an instantaneous decrease in the expected rate of inflation for the next 12 months and a delayed negative effect on the inflation expectations for the next 4 years. In turn, a central bank information shock is followed by an increase in both rates. The effects of this shock on the expected rate of inflation for the next 12 months, however, seem to be more persistent than those on the inflation expectations for the next 4 years. Finally, the responses of these variables to a monetary policy shock without identifying the central bank information shock indicate that both rates significantly increase after this shock. In line with our benchmark results, this is because of the presence central bank information shocks, which seem to have higher effects on both variables.

3.3.2 Alternative Monetary Policy Surprises

In these exercises, we replace our baseline measure of the interest rate surprise with alternative measures used in the literature. In particular, we employ the change in the three-month

[35] Data are obtained from the Survey of Professional Forecasters published by Banco de México. Capistran et al. (2010) find that forecasts of inflation taken from Banco de México's survey of professional forecasters outperform forecasts from traditional benchmarks such as univariate and multivariate time series models.

fed funds future as in Jarocinski and Karadi (2020). This measure combines surprises about actual rate setting and near-term forward guidance, so it constitutes a broad measure of the overall monetary policy stance (Jarocinski and Karadi, 2020). In addition, we also use the target factor of Gurkaynak et al. (2005). Namely, this is the first principal component of the surprises in fed funds futures and eurodollar futures with one year or less to expiration in a 30-minute window around an FOMC announcement.[36] The advantage of this measure is that it is a broader indicator of forward guidance (Jarocinski and Karadi, 2020).

The results from these exercises are respectively reported in Figures B.3 and B.4 of the Appendix B. In general, using alternative monetary policy surprises changes very little the impulse response functions to the purged monetary policy shock and the central bank information shock. In addition, we find that the standard HFI mixes the monetary policy shocks with central bank information shocks as in our baseline specification. By using the change in the three-month fed funds future and the target factor of Gurkaynak et al. (2005), we find that the responses of output, price level, exchange rate, country risk and stock prices are muted because the central bank information shocks, which have the opposite effects of monetary policy shocks, attenuate the estimated responses of these variables to a monetary policy shock. Consistent with our previous findings, an additional bias in the standard HFI is that the interest rate responses are more persistent.

3.3.3 Alternative Specification with Additional Lags

For robustness checks, we also estimate the model using two and three lags. The results are reported in Figures B.5 and B.6 of the Appendix B. Although the estimated impulse responses exhibit some loss of significance due to the higher number of parameters being estimated, the qualitative results generally align with our baseline specification. Specifically, they indicate that positive information shocks appear to improve financial conditions. Additionally, 3-month and 10-year government bond yields tend to rise after this shock. Information shocks also positively affect economic activity and the price level in Mexico. In contrast, restrictive

[36] In particular, the target factor is constructed using the following five indicators: the current-month fed funds future, the 3-month fed funds future, and the eurodollar futures at the horizons of two, three and four quarters.

monetary policy shocks in the US tighten financial conditions, contract real activity, and lower prices in Mexico. Finally, not accounting for the presence of central bank information shocks may lead to biased measurements of the international transmission of monetary policy.

3.3.4 Alternative Country-Risk Indicator

Finally, we estimate the model using an alternative country-risk indicator instead of the CDS. In particular, we used the Mexico's EMBI plus spread obtained from Bloomberg, which reflects the difference between the yields on sovereign bonds issued by the local government and bonds issued by governments of the industrialized world with identical currency denomination and maturity. In particular, the EMBI+ index includes US dollar and other external currency denominated Brady bonds, Eurobonds, and traded loans issued by sovereign entities. The results from this exercise, reported in Figure B.7 of the Appendix B, are consistent with our benchmark results.

4 Conclusion

Central bank announcements simultaneously reveal information not just about monetary policy but also about the central bank's assessment of the economic outlook. In this article, we studied the effects of FOMC announcement surprises on macroeconomic and financial conditions in Mexico. The analysis is carried out by estimating impulse-response functions using a Bayesian VAR model. This model allows us to combine two approaches to shock identification, namely HFI and sign restrictions. In particular, we separate conventional monetary policy shocks from concurrent central bank information shocks based on the high-frequency co-movement of interest rate and stock market surprises. We then obtain the dynamic responses of financial and macroeconomic variables to these shocks.

The main results indicate that US central bank information shocks lead to an improvement in financial conditions and higher interest rates in Mexico after some months. Specifically, these shocks are followed by a decrease in the sovereign risk premium, an appreciation of the Mexican peso/US dollar exchange rate and fall of its volatility, as well as an increase

in the stock index. Moreover, we find that the 3-month and the 10-year government bond yields tend to increase a few months after this shock. The central bank information shocks also appear to have a persistent positive effect on activity. These effects may lead to demand pressures that could outweigh the impact of the exchange rate appreciation on prices. In turn, these demand pressures may help to explain the behavior of interest rates. While our analysis does not delve into individual transmission channels of the transmission mechanism, we conjecture that both trade and financial channels likely play crucial roles in transmitting these shocks to the Mexican economy. Specifically, in response to the unexpected good news in the US economic outlook, we hypothesize that private sector decisions in both countries may positively influence aggregate demand in Mexico through different channels, such as the current account, foreign direct investment, domestic consumption, and investment. The responses of domestic interest rates in turn seem to be consistent with the scenario in which the domestic central bank tightens monetary policy to counteract the spillover effects stemming from a stronger US economy, which could potentially lead to demand pressures on inflation.

In turn, our results indicate that distinguishing between conventional monetary policy and central bank information shocks may be important for a better understanding of the international transmission of US monetary policy. In fact, the effects of both shocks are considerably different and not accounting for the presence of central bank information may lead to biased measurements of the international transmission of monetary policy shocks. In particular, an important difference from the standard HFI of monetary policy shocks, which does not control for the information content of the announcements, is that the purged monetary policy shock leads to a significant contraction in real activity and a lower price level, as well as a tightening in financial conditions in Mexico, the latter characterized by higher sovereign risk premium, a depreciation of Mexican peso/US dollar exchange rate, higher forex volatility, and lower stock price index. In contrast, the standard HFI of monetary policy shocks leads to opposite results. In line with Mackowiak (2007), Iacoviello and Navarro (2019), and Carrillo et al. (2020), among others, the economic activity dynamics looks considerably similar to that in the US after a pure monetary policy shock. Finally, we also find that the 10-year and the 3-month government bond yields tend to increase after a pure monetary policy shock,

although in the case of the latter the response is smaller and somewhat muted. Given the important degree of financial integration between both countries, domestic interest rates are expected to be influenced to some extent by interest rates in the US.

This study has significant implications from an economic policy perspective. Our findings highlight the relevance of US central bank communication in influencing macroeconomic outcomes in Mexico. Therefore, the lessons learned from the evidence for this economy could be valuable for the design of economic policy in other countries, as their domestic variables are influenced by the US monetary policy due to the size and importance of its economy. Our results suggest that policymakers should consider closely monitoring US interest rates and stock markets fluctuations around FOMC announcements to disentangle the informational content of such announcements and respond appropriately to external disturbances.

Further research could explore forward guidance effects arising from US announcements on the Mexican economy. Another area of research would be to examine a larger set of financial variables and macroeconomic variables, including the current account, foreign direct investment, domestic consumption, and investment.

References

[1] Andrade, P., and Ferroni, F. (2021). Delphic and odyssean monetary policy shocks: Evidence from the euro area. *Journal of Monetary Economics* **117**, 816-832.

[2] Baek, I. M. (2006). Portfolio investment flows to Asia and Latin America: pull, push or market sentiment? *Journal of Asian Economics* **17**(2), 363-373.

[3] Banco de México (2022). Quarterly inflation report. April-June.

[4] Bauer, M. D., and Rudebusch, G. D. (2016). Monetary policy expectations at the zero lower bound. *Journal of Money, Credit and Banking*, **48**(7), 1439-1465.

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

[6] Bernanke, B. S., and Kuttner, K. N. (2005). What explains the stock market's reaction to Federal Reserve policy? *The Journal of Finance* **60**(3), 1221-1257.

[7] Blanchard, O. J., Faruqee, H., Das, M., Forbes, K. J., and Tesar, L. L. (2010). The initial impact of the crisis on emerging market countries. *Brookings Papers on Economic Activity* (1), 263-323.

[8] Breitenlechner, M., Gründler, D., and Scharler, J. (2021). Unconventional monetary policy announcements and information shocks in the US. *Journal of Macroeconomics* **67**, 1-22.

[9] Caceres, C., Carriere-Swallow, M. Y., Demir, I., and Gruss, B. (2016). *Unconventional monetary policy announcements and information shocks in the US*. International Monetary Fund.

[10] Caldara, D., and Herbst, E. (2019). Monetary policy, real activity, and credit spreads: Evidence from Bayesian proxy SVARs. *American Economic Journal: Macroeconomics* **11**(1), 157-92.

[11] Campbell, J. R., Evans, C. L., Fisher, J. D., Justiniano, A., Calomiris, C. W., and Woodford, M. (2012). Macroeconomic effects of federal reserve forward guidance. *Brookings Papers on Economic Activity* **42**(1), 1-80.

[12] Campbell, J. R., Fisher, J. D., Justiniano, A., and Melosi, L. (2017). Forward guidance and macroeconomic outcomes since the financial crisis. *NBER Macroeconomics Annual* **31**(1), 283-357.

[13] Canova, F. (2005). The transmission of US shocks to Latin America. *Applied Econometrics* **20**(2), 229-251.

[14] Carrillo, J. A., Elizondo, R., and Hernández-Román, L. G. (2020). Inquiry on the transmission of US aggregate shocks to Mexico: a SVAR approach. *Journal of International Money and Finance* **104**, 1-24.

[15] Cesa-Bianchi, A., and Sokol, A. (2022). Financial shocks, credit spreads, and the international credit channel. *Journal of International Economics* **135**, 1-18.

[16] Chiquiar, D., Noriega, A. E., and Ramos-Francia, M. (2010). A time-series approach to test a change in inflation persistence: the Mexican experience. *Applied Economics* **42**(24), 3067-3075.

[17] De Vita, G., and Kyaw, K.S. (2008). Determinants of capital flows to developing countries: a structural VAR analysis. *Journal of Economic Studies* **35**(4), 304-322.

[18] Dahlhaus, T., Hess, K., and Reza, A. (2018). International transmission channels of US quantitative easing: Evidence from Canada. *Journal of Money, Credit and Banking* **50**(2-3), 545-563.

[19] Dedola, L., Rivolta, G., and Stracca, L. (2017). If the Fed sneezes, who catches a cold? *Journal of International Economics* **108**, S23-S41.

[20] Favara, G., Gilchrist, S., Lewis, K. F., and Zakrajsek, E. (2016). Recession risk and the excess bond premium. (No. 2016-04-08). Board of Governors of the Federal Reserve System (US).

[21] Fink, F., and Schüler, Y. S. (2015). The transmission of US systemic financial stress: Evidence for emerging market economies. *Journal of International Money and Finance* **55**, 6-26.

[22] Forbes, K.J., and Warnock, F.E. (2012). Capital flow waves: surges, stops, flight, and retrenchment. *Journal of International Economics* **88**(2), 235–251.

[23] Gaytán, A., and González García, J. R. (2007). Cambios estructurales en el mecanismo de transmisión de la política monetaria en México: un enfoque VAR no lineal. *Monetaria* **30**(4), 367-404.

[24] Gertler, M., and Karadi, P. (2015). Monetary policy surprises, credit costs, and economic activity. *American Economic Journal: Macroeconomics* **7**(1), 44-76.

[25] Gilchrist, S., and Zakrajšek, E. (2012). Credit spreads and business cycle fluctuations. *American Economic Review* **102**(4), 1692-1720.

[26] Georgiadis, G. (2016). Determinants of global spillovers from US monetary policy. *Journal of International Money and Finance* **67**, 41-61.

[27] Gürkaynak, R. S., Kara, A. H., Kısacıkoğlu, B., and Lee, S. S. (2021). Monetary policy surprises and exchange rate behavior. *Journal of International Economics* **130**, 1-24.

[28] Gürkaynak, R. S., Sack, B., and Swanson, E. T. (2005). Do actions speak louder than words? The response of asset prices to monetary policy actions and statements. *International Journal of Central Banking* **1**(1), 55-93.

[29] Gürkaynak, R. S., Sack, B., and Wright, J. H. (2010). The TIPS yield curve and inflation compensation. *American Economic Journal: Macroeconomics* **2**(1), 70-92.

[30] Hansen, S., and McMahon, M. (2016). Shocking language: Understanding the macroeconomic effects of central bank communication. *Journal of International Economics* **99**, S114-S133.

[31] Hanson, S. G., and Stein, J. C. (2015). Monetary policy and long-term real rates. *Journal of Financial Economics* **115**(3), 429-448.

[32] Hausman, J., and Wongswan, J. (2011). Global asset prices and FOMC announcements. *Journal of International Money and Finance* **30**(3), 547-571.

[33] Iacoviello, M., and Navarro, G. (2019). Foreign effects of higher US interest rates. *Journal of International Money and Finance* **95**, 232-250.

[34] Ilzetzki, E., and Jin, K. (2013). The puzzling change in the transmission of U.S. macroeconomic policy shocks. lSE Working Papers.

[35] Jarociński, M., and Karadi, P. (2020). Deconstructing monetary policy surprises—the role of information shocks. *American Economic Journal: Macroeconomics* **12**(2), 1-43.

[36] Kadiyala, K. R., and Karlsson, S. (1997). Numerical methods for estimation and inference in Bayesian VAR-models. *Journal of Applied Econometrics* **12**(2), 99-132.

[37] Kamin, S. B., and Von Kleist, K. (1999). The evolution and determinants of emerging markets credit spreads in the 1990s. BIS Papers, **68**.

[38] Kerssenfischer, M. (2022). Information effects of euro area monetary policy. *Economics Letters* **216**, 1-5.

[39] Kim, D. H., and Singleton, K. J. (2012). Term structure models and the zero bound: an empirical investigation of Japanese yields. *Journal of Econometrics*, **170**(1), 32-49.

[40] Krippner, L. (2013). Measuring the stance of monetary policy in zero lower bound environments. *Economics Letters*, **118**(1), 135-138.

[41] Kuttner, K. N. (2001). Monetary policy surprises and interest rates: Evidence from the Fed funds futures market. *Journal of Monetary Economics* **47**(3), 523-544.

[42] Lakdawala, A., and Schaffer, M. (2019). Federal reserve private information and the stock market. *Journal of Banking & Finance* **106**(3), 34-49.

[43] Litterman, R.B., 1979. Techniques of forecasting using vector autoregressions. Federal Reserve Bank of Minneapolis, Working Paper 115.

[44] Litterman, R. B. (1986). Forecasting with Bayesian vector autoregressions—five years of experience. *Journal of Business & Economic Statistics* **4**(1), 25-38.

[45] Lucca, D. O., and Moench, E. (2015). The pre-FOMC announcement drift. *The Journal of Finance* **70**(1), 329-371.

[46] Maćkowiak, B. (2007). External shocks, US monetary policy and macroeconomic fluctuations in emerging markets. *Journal of Monetary Economics* **54**(8), 2512-2520.

[47] Melosi, L. (2017). Signalling effects of monetary policy. *The Review of Economic Studies* **84**(2), 853-884.

[48] Mertens, K., and Ravn, M. O. (2013). The dynamic effects of personal and corporate income tax changes in the United States. *American Economic Review* **103**(4), 1212-47.

[49] Miranda-Agrippino, S., and Rey, H. (2020). US monetary policy and the global financial cycle. *The Review of Economic Studies* **87**(6), 2754-2776.

[50] Miranda-Agrippino and Ricco, G. (2023). Identification with external instruments in structural VARs. *Journal of Monetary Economics* **135**, 1-19.

[51] Nakamura, E., and Steinsson, J. (2018). High-frequency identification of monetary non-neutrality: the information effect. *The Quarterly Journal of Economics* **133**(3), 1283-1330.

[52] Paul, P. (2020). The time-varying effect of monetary policy on asset prices. *Review of Economics and Statistics* **102**(4), 690-704.

[53] Ramey, V. A. (2016). Macroeconomic shocks and their propagation. *Handbook of Macroeconomics* **2**, 71-162.

[54] Rubio-Ramirez, J. F., Waggoner, D. F., and Zha, T. (2010). Structural vector autoregressions: Theory of identification and algorithms for inference. *The Review of Economic Studies* **77**(2), 665-696.

[55] Stock, J. H., and Watson, M. W. (2012). Disentangling the channels of the 2007–2009 recession. *Brookings Papers on Economic Activity* **42**(1), 81-135.

[56] Swanson, E. T., and Williams, J. C. (2014). Measuring the effect of the zero lower bound on medium-and longer-term interest rates. *American Economic Review* **104**(10), 3154-3185.

[57] Uhlig, H. (2005). What are the effects of monetary policy on output? Results from an agnostic identification procedure. *Journal of Monetary Economics* **52**(2), 381-419.

[58] Uribe, M., and Yue, V. Z. (2006). Country spreads and emerging countries: Who drives whom? *Journal of International Economics* **69**(1), 6-36.

[59] Vicondoa, A. (2019). Monetary news in the United States and business cycles in emerging economies. *Journal of International Economics* **117**, 79-90.

[60] Wu, J. C., and Xia, F. D. (2016). Measuring the macroeconomic impact of monetary policy at the zero lower bound. *Journal of Money, Credit and Banking* **48**(2-3), 253-291.

A Appendix: Bayesian estimation

The reduced form representation of the VAR model is as follows:

$$\begin{pmatrix} \boldsymbol{M} & \widetilde{\boldsymbol{Y}} \end{pmatrix} = \boldsymbol{X} \begin{pmatrix} 0 & \boldsymbol{B} \end{pmatrix} + \begin{pmatrix} \boldsymbol{U}^{M} & \boldsymbol{U}^{\widetilde{Y}} \end{pmatrix} \tag{1}$$

or

$$\begin{pmatrix} \boldsymbol{m}_t \\ \boldsymbol{y}_t^* \\ \boldsymbol{y}_t \end{pmatrix} = \begin{pmatrix} 0 \\ \boldsymbol{c}_{y^*} \\ \boldsymbol{c}_y \end{pmatrix} + \sum_{p=1}^{P} \begin{pmatrix} 0 & 0 & 0 \\ B_{y^*m}^p & B_{y^*y^*}^p & 0 \\ B_{ym}^p & B_{yy^*}^p & B_{yy}^p \end{pmatrix} \begin{pmatrix} \boldsymbol{m}_{t-p} \\ \boldsymbol{y}_{t-p}^* \\ \boldsymbol{y}_{t-p} \end{pmatrix} + \begin{pmatrix} \boldsymbol{u}_t^m \\ \boldsymbol{u}_t^{y^*} \\ \boldsymbol{u}_t^y \end{pmatrix} \tag{2}$$

where $\boldsymbol{M}$ represents a vector of high-frequency surprises, denoted by $\boldsymbol{m}_t$, $\widetilde{\boldsymbol{Y}}$ is a vector of monthly variables, $\boldsymbol{y}_t^*$ and $\boldsymbol{y}_t$, $\boldsymbol{X}$ is a matrix that collects the right-hand-side variables, $\boldsymbol{B}$ is a matrix that contains the coefficient matrix at lag p, $\boldsymbol{B}^p$, and a vector of constants $\boldsymbol{c}$. In turn, $\boldsymbol{U}^M$ and $\boldsymbol{U}^{\widetilde{Y}}$ are vectors of white noise residuals following the normal distribution $N(0, \boldsymbol{\Sigma})$. In particular, $\boldsymbol{m}_t$ is a vector of N_m surprises in US financial instruments, $\boldsymbol{y}_t^*$ is a vector of N_{y^*} US variables, and $\boldsymbol{y}_t$ is a vector of N_y domestic (Mexican) variables. We assume that the high-frequency surprises are exogenous with respect to the monthly variables in Eq. (2). Thus, we set the corresponding coefficients of m_t in the coefficient matrices and in the vector of constants to zero as in Jarociński and Karadi (2020) and Breitenlechner et al. (2021). Following Fink and Schüler (2015) and Carrillo et al. (2020), we assume that $\boldsymbol{y}_t^*$ is block exogenous with respect to the domestic variables, i.e. the dynamics of $\boldsymbol{y}_t$ does not affect any of the values in $\boldsymbol{y}_t^*$.

We estimate the model using a Bayesian approach, as it is standard in the sign restrictions literature. In particular, we follow the literature on Bayesian VAR models and use the Minnesota prior of Litterman (1979), which is an independent normal inverted Wishart distribution, $p(\boldsymbol{B}, \boldsymbol{\Sigma}) = p(\boldsymbol{B})p(\boldsymbol{\Sigma})$, such as:

$$p(\boldsymbol{\Sigma} \mid \underline{\boldsymbol{S}}, \underline{\boldsymbol{v}}) = IW(\underline{\boldsymbol{S}}, \underline{\boldsymbol{v}}) \tag{3}$$

$$p(vec\boldsymbol{B} \mid \underline{\boldsymbol{B}}, \underline{\boldsymbol{Q}}) = N(vec\boldsymbol{B}, \underline{\boldsymbol{Q}}) \tag{4}$$

where *IW* denotes the Inverted Wishart distribution. Similar to Jarociński and Karadi (2020), the coefficient of the first own lag of each variable in the matrix $\underline{\boldsymbol{B}}$ is fixed at 1, and the remaining entries are zero, reflecting the prior assumption that these variables follow a random walk.[37] In turn, $\underline{\boldsymbol{Q}}$ is a diagonal matrix, such that the standard deviation of lag p of variable j in equation i is $\lambda_1^{-1}\sigma_i/\sigma_j p^{-\lambda_2}$. Following Litterman (1986), we set $\lambda_1 = 5$, $\lambda_2 = 1$. σ_i (σ_j) represents the standard error in the autoregression of order P of variable i (j). As in Kadiyala and Karlsson (1997), we set $\underline{\upsilon} = N+2$, which is the minimum value that guarantees the existence of the prior mean of Σ. Finally, $\underline{\boldsymbol{S}}$ is a diagonal matrix with σ_i^2, $i = 1,...,N_m + N_{y^*} + N_y$ on the main diagonal. We include one lag in our baseline specification, as is common in studies analyzing the international transmission of US monetary policy.

We use a Gibbs sampler to compute the posterior. In particular, this algorithm consists of drawing $\boldsymbol{\Sigma}$ and $\boldsymbol{B}$ from their conditional posteriors $p(\boldsymbol{\Sigma} \mid \widetilde{\boldsymbol{Y}}, \boldsymbol{B})$ and $p(vec\boldsymbol{B} \mid \widetilde{\boldsymbol{Y}}, \boldsymbol{\Sigma})$, respectively, until the sampler converges. The conditional posteriors of these parameters are standard.[38] Specifically, the conditional posterior of $\boldsymbol{\Sigma}$ is as follows:

$$p(\boldsymbol{\Sigma} \mid \widetilde{\boldsymbol{Y}}, \boldsymbol{M}, \boldsymbol{B}) = IW(\overline{\boldsymbol{S}}, \overline{\upsilon}) \tag{5}$$

where

$$\overline{\boldsymbol{S}} = \left(\begin{pmatrix} \boldsymbol{M} & \widetilde{\boldsymbol{Y}} \end{pmatrix} - \boldsymbol{X}\begin{pmatrix} \boldsymbol{0} & \boldsymbol{B} \end{pmatrix}\right)' \left(\begin{pmatrix} \boldsymbol{M} & \widetilde{\boldsymbol{Y}} \end{pmatrix} - X\begin{pmatrix} \boldsymbol{0} & \boldsymbol{B} \end{pmatrix}\right) + \underline{\boldsymbol{S}} \tag{6}$$

$$\overline{\upsilon} = T + \underline{\upsilon} \tag{7}$$

In turn, the conditional distribution of $\boldsymbol{B}$ is given by:

$$p(vec\boldsymbol{B} \mid \widetilde{\boldsymbol{Y}}, \boldsymbol{M}, \boldsymbol{\Sigma}) = N(\overline{\boldsymbol{B}}, \overline{\boldsymbol{Q}}) \tag{8}$$

where

$$\overline{\boldsymbol{Q}} = (\underline{\boldsymbol{Q}}^{-1} + \boldsymbol{\Sigma}^{-1}_{\widetilde{Y}\widetilde{Y}.1} \otimes \boldsymbol{X}'\boldsymbol{X})^{-1} \tag{9}$$

$$vec\overline{\boldsymbol{B}} = \overline{\boldsymbol{Q}}(\underline{\boldsymbol{Q}}^{-1} vec\underline{\boldsymbol{B}} + (\boldsymbol{\Sigma}^{-1}_{\widetilde{Y}\widetilde{Y}.1} \otimes \boldsymbol{X}')vec(\widetilde{\boldsymbol{Y}} + \boldsymbol{M}\boldsymbol{\Sigma}^{-1}_{MM}\boldsymbol{\Sigma}_{M\widetilde{Y}})) \tag{10}$$

[37] It is worth mentioning that this assumption has no impact on the final results.

[38] Further details on the derivation of the conditional posteriors can be found in Jarociński and Karadi (2020).

and

$$\boldsymbol{\Sigma} = \begin{pmatrix} \boldsymbol{\Sigma}_{MM} & \boldsymbol{\Sigma}_{MY} \\ \boldsymbol{\Sigma}_{\widetilde{Y}M} & \boldsymbol{\Sigma}_{\widetilde{Y}\widetilde{Y}} \end{pmatrix} \tag{11}$$

$$\boldsymbol{\Sigma}_{YY.1} = \boldsymbol{\Sigma}_{\widetilde{Y}\widetilde{Y}} - \boldsymbol{\Sigma}_{\widetilde{Y}M}\boldsymbol{\Sigma}_{MM}^{-1}\boldsymbol{\Sigma}_{M\widetilde{Y}} \tag{12}$$

Following Jarociński and Karadi (2020) and Breitenlechner et al. (2021) the estimation is implemented with 10,000 iterations of the Gibbs sampling procedure of which the first 2000 are dropped as burn-in and every fourth iteration is saved afterwards.

B Appendix: Supplemental Results

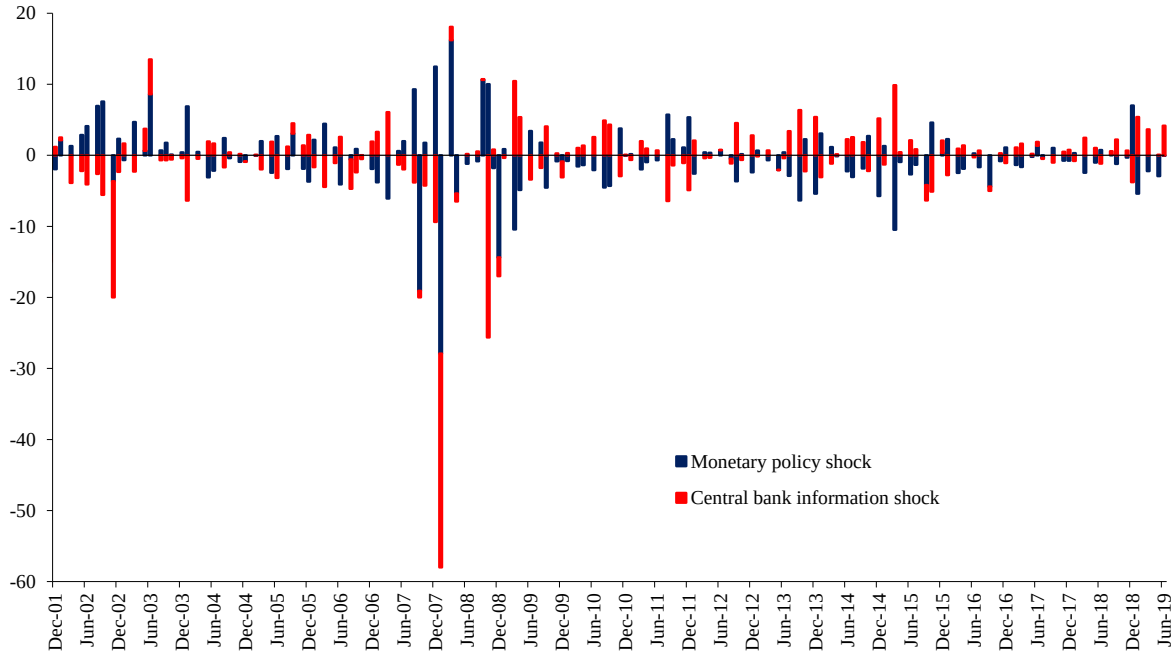

Figure B.1: Contributions of Shocks to the Surprises in the One-Month Fed Funds Futures

Notes: Aggregated to the monthly frequency. The shocks are scaled in terms of the one-month fed funds futures surprises, in basis points, and summarized by their posterior means.

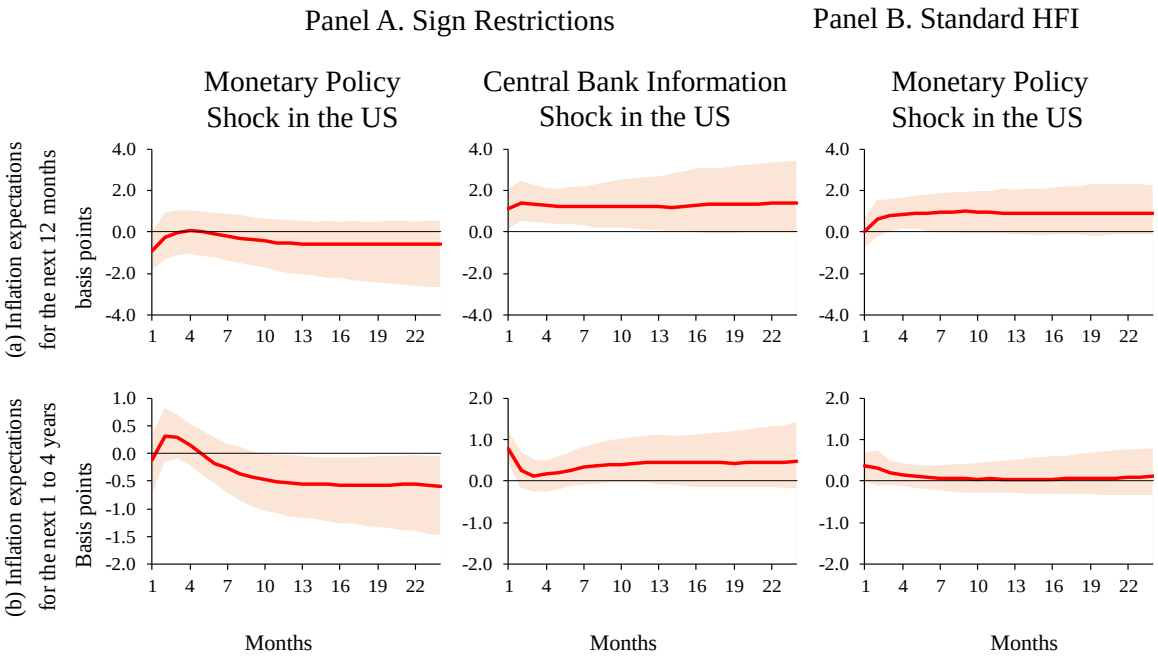

Figure B.2: Impulse Response Functions of Additional Variables

Notes: Responses are presented for a 24-month horizon with the associated 68% highest posterior density intervals (HPDIs).

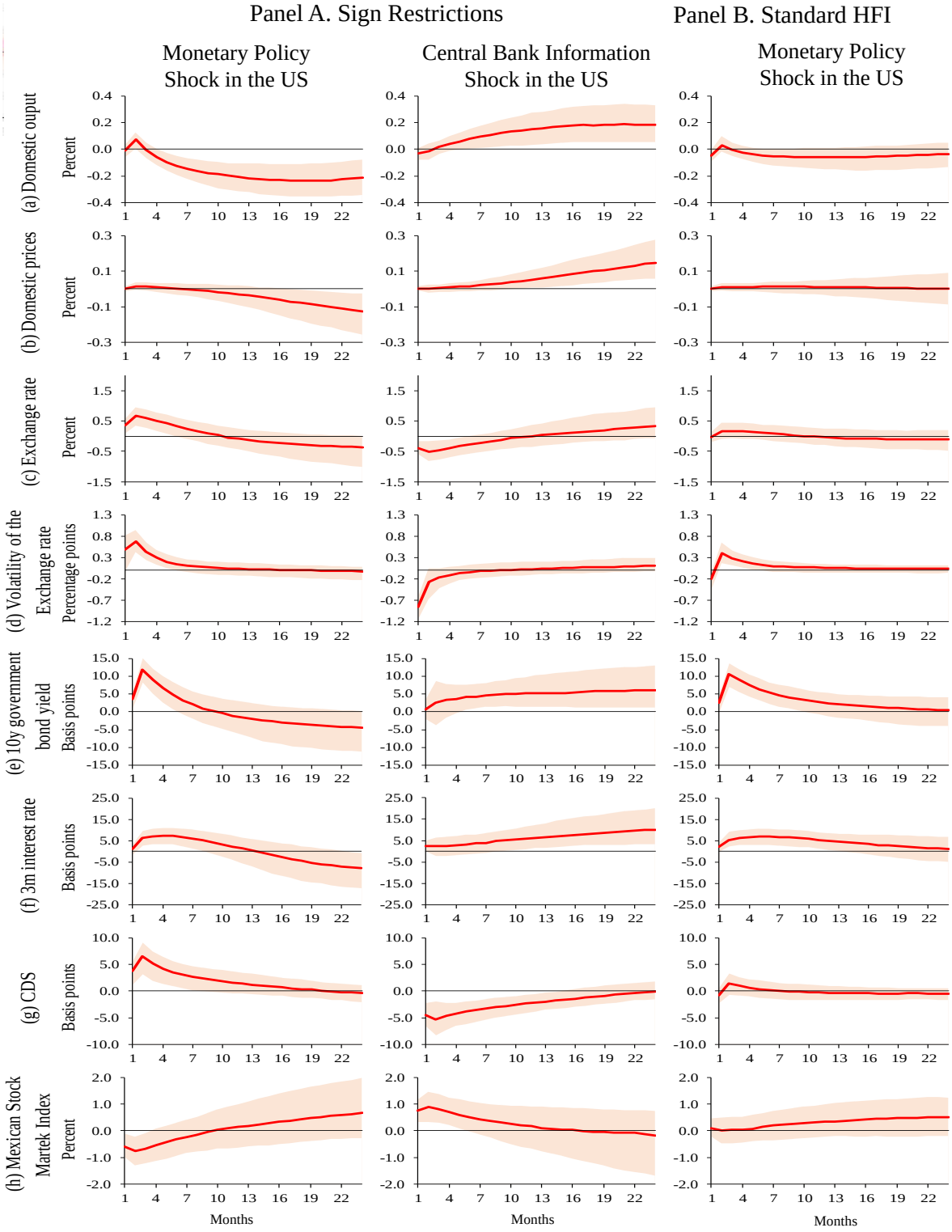

Figure B.3: Impulse Response Functions by using the change in the Three-Month Fed Funds Future

Notes: Responses are presented for a 24-month horizon with the associated 68% highest posterior density intervals (HPDIs).

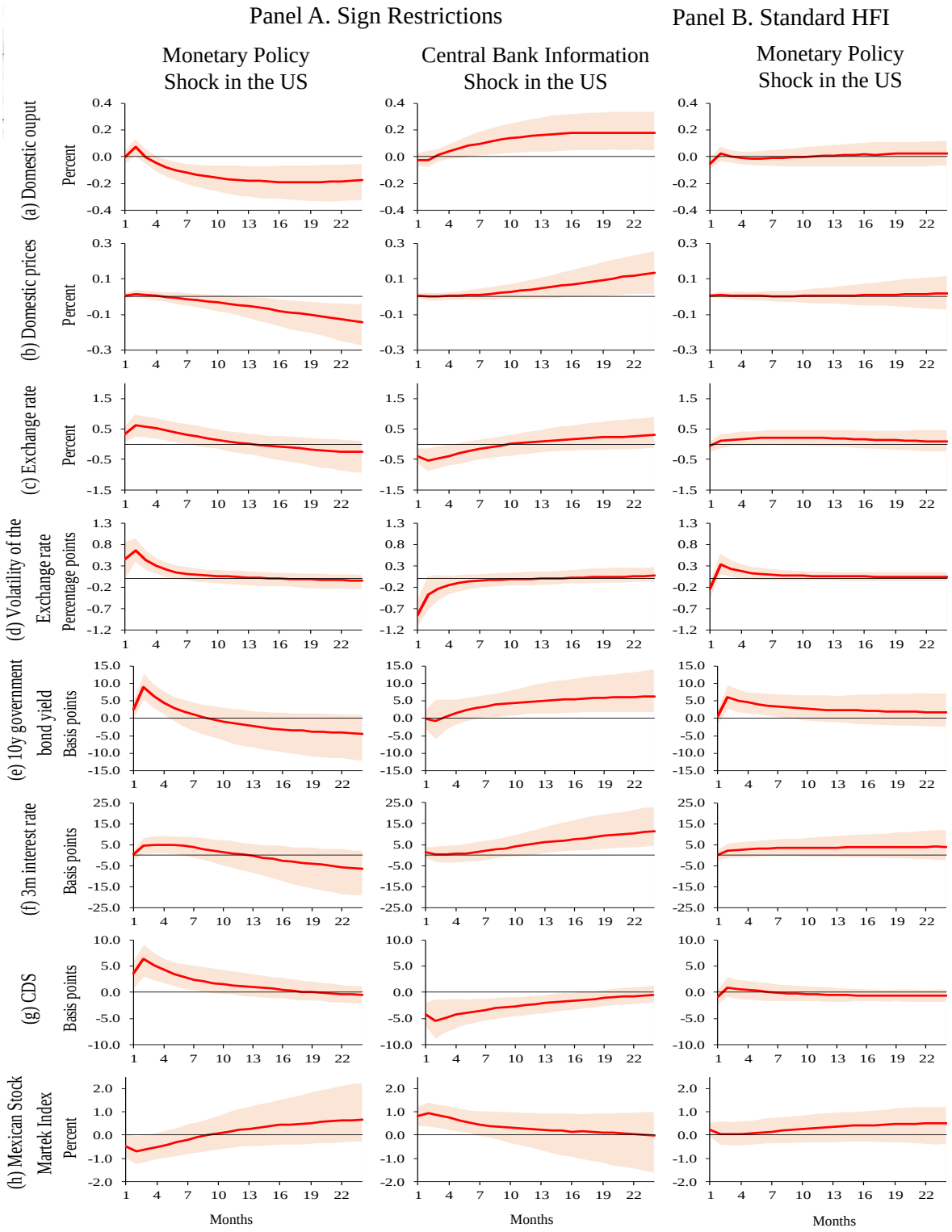

Figure B.4: Impulse Response Functions by using the Target Factor of Gurkaynak et al. (2005)

Notes: Responses are presented for a 24-month horizon with the associated 68% highest posterior density intervals (HPDIs).

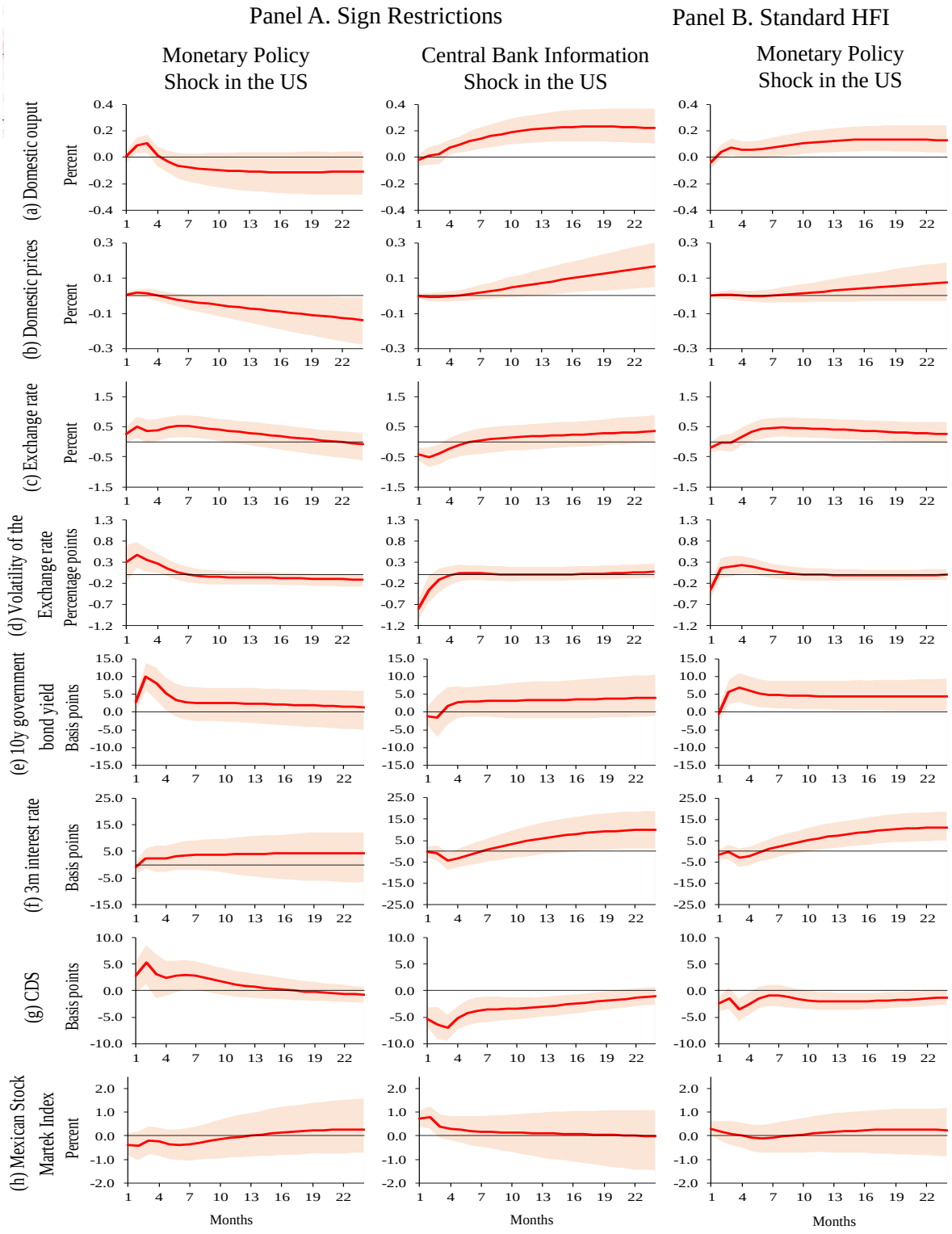

Figure B.5: Impulse Response Functions: VAR Model with 2 Lags

Notes: Responses are presented for a 24-month horizon with the associated 68% highest posterior density intervals (HPDIs).

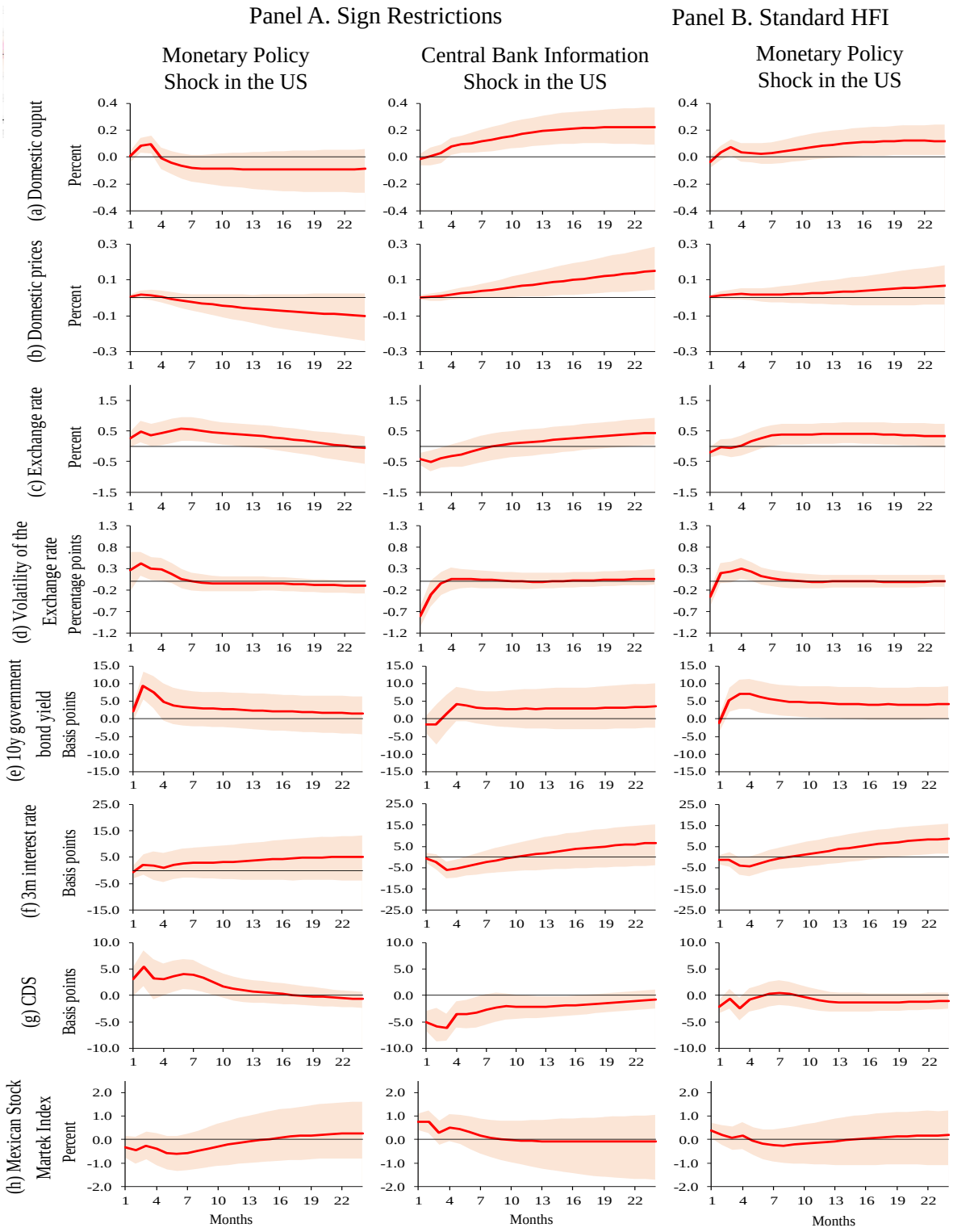

Figure B.6: Impulse Response Functions: VAR Model with 3 Lags

Notes: Responses are presented for a 24-month horizon with the associated 68% highest posterior density intervals (HPDIs).

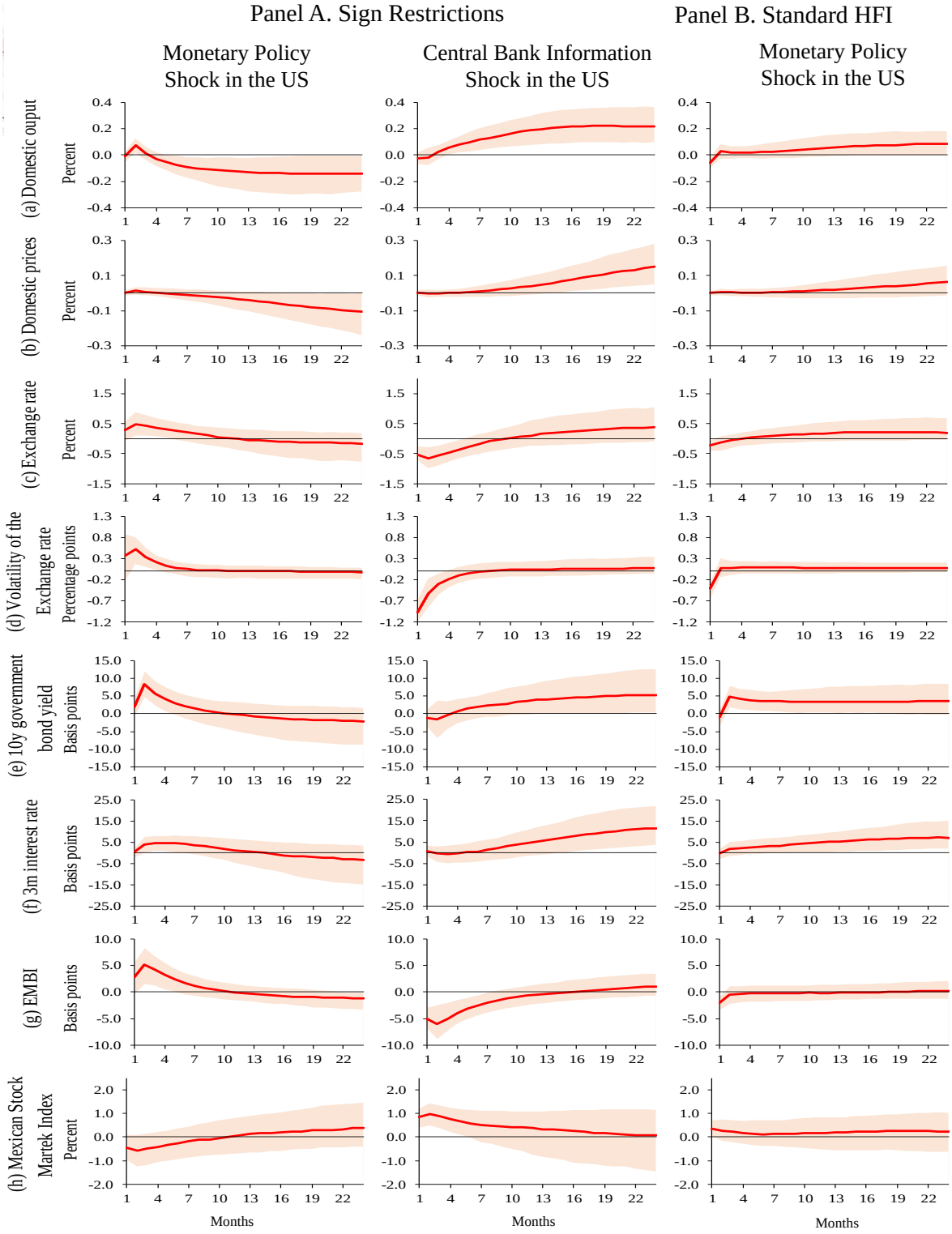

Figure B.7: Impulse Response Functions by using the EMBI+

Notes: Responses are presented for a 24-month horizon with the associated 68% highest posterior density intervals (HPDIs).

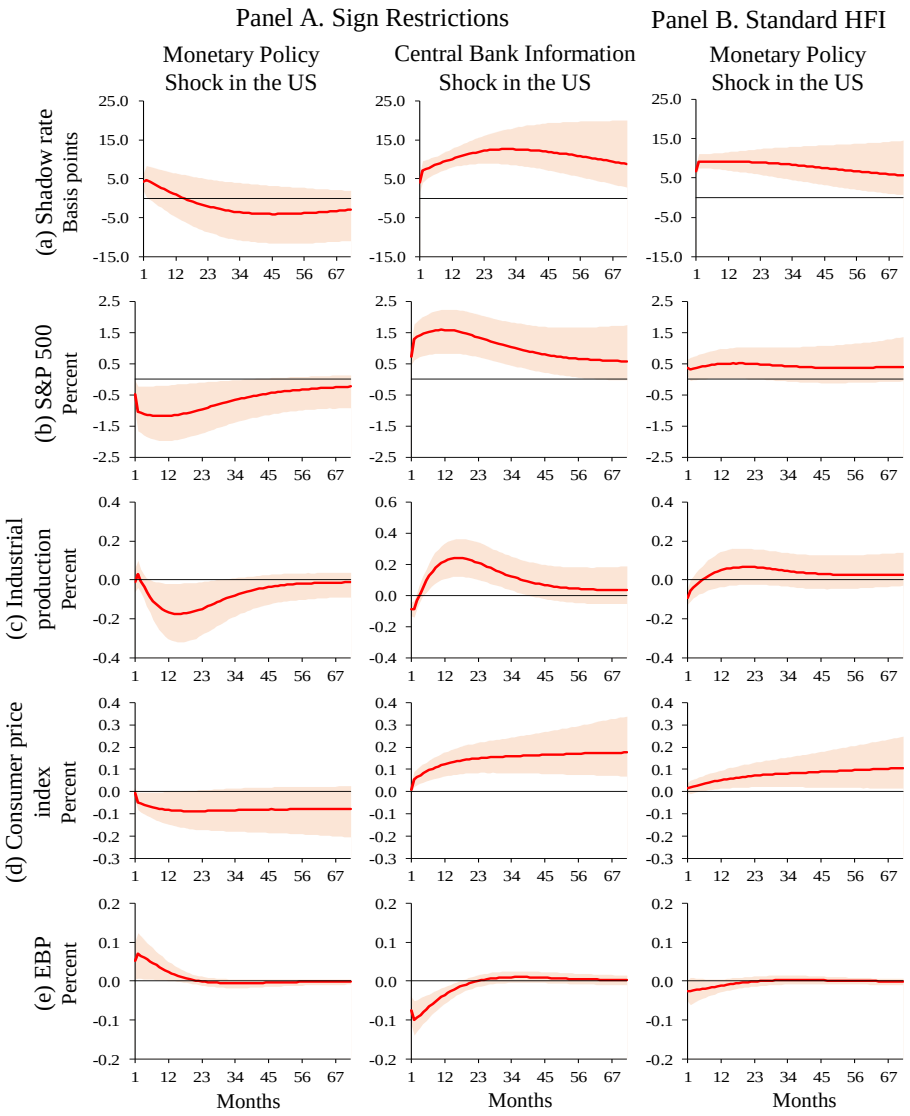

Figure B.8: Impulse Response Functions of US Variables to the Monetary Policy and Central Bank Information Shocks in the US for a 72-Month Horizon

Notes: Responses are presented for a 72-month horizon with the associated 68% highest posterior density intervals (HPDIs).

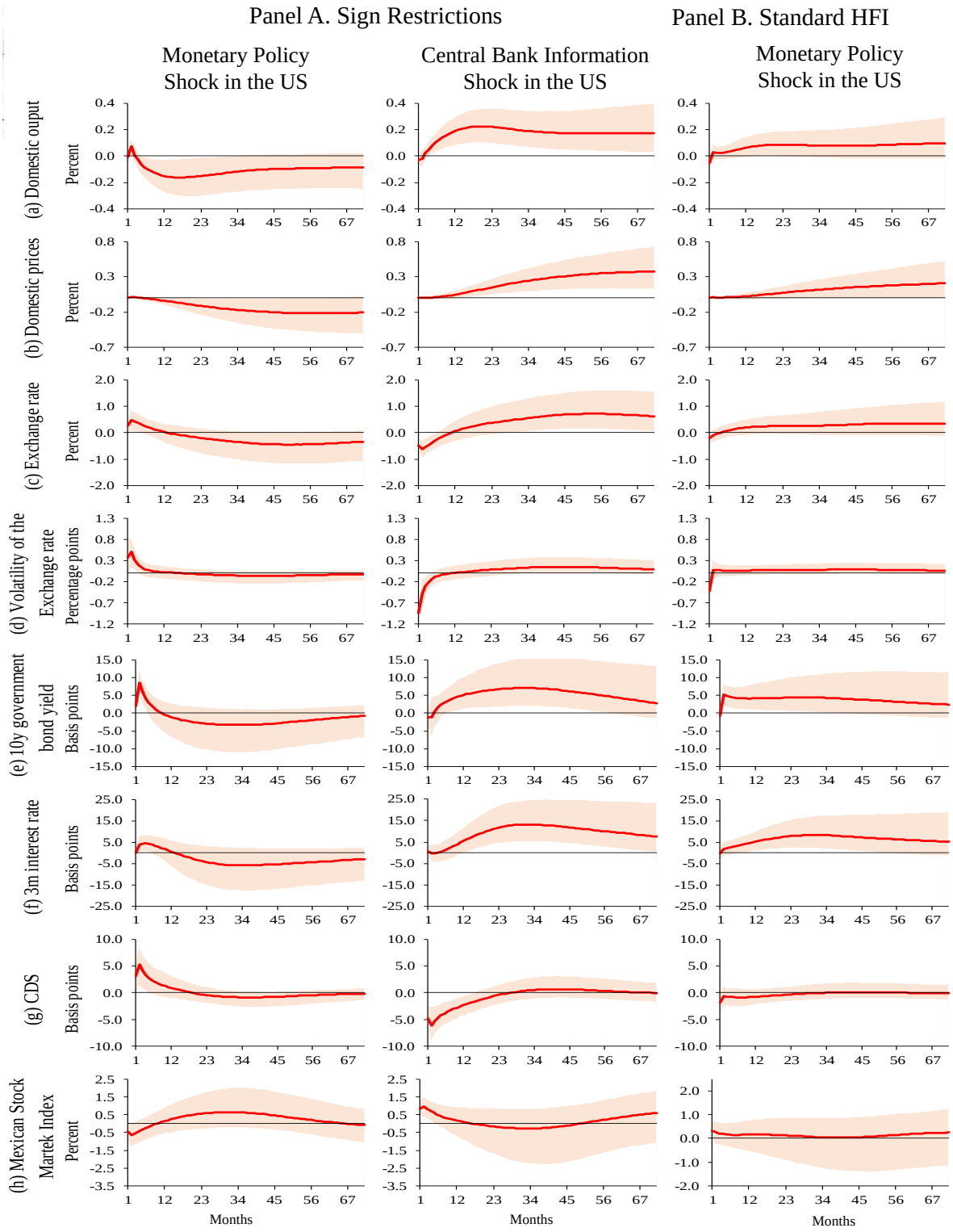

Figure B.9: Impulse Response Functions of Mexican Variables to the Monetary Policy and Central Bank Information Shocks in the US for a 72-Month Horizon

Notes: Responses are presented for a 72-month horizon with the associated 68% highest posterior density intervals (HPDIs).