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Inquiry on the transmission of U.S. aggregate shocks to Mexico: A SVAR approach ☆

Julio A. Carrillo [a], Rocio Elizondo [a], Luis G. Hernández-Román [b]

Bank of Mexico, Directorate General of Economic Research, 18 Av. Cinco de Mayo, 06059, Mexico City, Mexico
University of Warwick, Warwick Business School, Coventry CV4 7AL, UK

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ABSTRACT

We analyze the business cycle co-movement between Mexico and the US. We identify two shocks affecting US aggregate supply, three affecting its demand, and two types of monetary policy surprises with different financial implications. US shocks explain about 75% of expected output fluctuations in Mexico at a three-year horizon, with US demand shocks driving half of these variations alone. In turn, Mexican output responses to a monetary policy surprise in the US depend on the reaction of investors' sentiment to said surprise. Finally, for the sample period studied, financial-market interconnections are as important as goods-demand linkages for the international transmission of US shocks.

1. Introduction

In this paper we ask which macro shocks explain the co-movement observed in real activity, monetary-policy interest rates, and investors' sentiment in the U.S. and Mexico. To answer this question, we distinguish between two transmission mechanisms: the *financial-contagion channel*, and the *foreign-demand channel* (see also Fink and Schüler, 2015). The former focuses on the role of financial markets in the propagation of foreign economic conditions to the domestic economy; the latter refers to the direct effects that variations in trade patterns exert on domestic demand. In a nutshell, we find that shocks that affect U.S. spending, such as shocks to aggregate demand, economic policy uncertainty, and firms' cost of external finance, explain the co-movement of output, the policy rate, and the term premium in the U.S. and Mexico. In addition, both the financial-contagion and the foreign-demand channels are equally important for the propagation of U.S. spending shocks to real activity in Mexico.

To reach these findings, we use a structural vector autorregresion (SVAR) model with macro variables of the U.S. and Mexico in the sample period 2002M1-2018M3. We estimate the model using Bayesian techniques assuming that the U.S. economy is block-exogenous to the Mexican economy. This assumption is supported by the data, since Mexican variables do not seem to Granger cause U.S. variables. In turn, the sample period covers a single monetary-policy regime in Mexico, namely inflation targeting, which has reached lower and stable values of inflation and other nominal variables in this country.

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E-mail addresses: jcarrillo@banxico.org.mx (J.A. Carrillo), melizondo@banxico.org.mx (R. Elizondo), phd19lh@mail.wbs.ac.uk (L.G. Hernández-Román)

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We identify the following seven U.S. structural disturbances: two types of monetary policy shocks, which we call *type I* and *type II*; an aggregate demand shock; a productivity shock; a shock to the degree of uncertainty in the formulation of economic policies; a cost-push shock; and a financial shock affecting firms' cost of external funding. Our strategy to identify the monetary policy shocks is in line with recent studies that highlight a *signaling channel* of monetary policy (see Cesa-Bianchi and Sokol, 2017; Melosi, 2017; Jarociński and Karadi, 2018). Namely, an unexpected monetary-policy decision may bring about not only a change of the policy instrument, but also the central bank's assessment about the economy. In this vein, our identification strategy addresses the observation that the U.S. term premium has responded differently to unexpected monetary policy decisions. Such responses seem, in turn, to depend on whether investors perceive the information contained in the policy surprises as positive or negative to the economy.

The term premium contains information about the compensation that long-term bond holders ask for bearing the risk that short-term interest rates evolve differently than expected. For example, if investors believe that the economy's future is bright and stable, the term premium tends to decrease; instead, if they think that the economy is set to a downturn, the term premium tends to rise.[1] Accordingly, after a *type I* monetary policy unexpected loosening, the term premium falls. The intuition is that investors perceive that the new policy stance contributes to economic stability. In contrast, after a *type II* monetary policy unexpected loosening, the term premium rises. In this case, investors may perceive that the news accompanying the monetary policy decision indicates a bad economic outlook. In Section 3.1.1 we provide some examples of the differentiated investors' responses to U.S. monetary policy surprises, where we count at least ten *type I* shocks, and nine *type II* shocks in our sample.

We find three main results. The first one is that the seven identified U.S. macro shocks seem to explain nearly 75% of the fluctuations of the Mexican output gap after a three-year horizon. In fact, for every one percentage point change observed in this variable, 40 basis points belong to shocks that affected U.S. spending (i.e., shocks that hit directly economic policy uncertainty, aggregate demand, and the cost of external finance). From these 40 basis points, our calculations suggest that the financial-contagion and foreign-demand channels each transmitted 20 basis points. Our results echo those by Cushman and Zha (1997), Justiniano and Preston (2010), Mumtaz et al. (2011), International Monetary Fund (2014), who also find similar explanatory power of U.S. shocks on the real activities of other small open economies. However, our decomposition of the transmission into two separate channels delves into the propagation mechanism of U.S. shocks to small open economies.

The second result is that the response of Mexican real activity to a U.S. monetary policy shock depends crucially on the type of this shock. Accordingly, Mexican real activity might be prompted to a persistent slowdown after a *type I* unexpected tightening, while it might not be significantly affected after a *type II* unexpected policy stance. These predictions follow from the response of the Mexican term premium to a U.S. monetary policy shock. Namely, after a *type I* shock, the U.S. and Mexican term premia co-move; in contrast, after a *type II* shock, these spreads display a zero correlation. Therefore, the financial-contagion channel is stronger after a *type I* shock than after a *type II* shock, which suggests that in the case of the former long-term bond holders in the two countries seem to share a similar sentiment. Nonetheless, since monetary policy shocks are rare events, their contributions to output fluctuations in the U.S. and Mexico are relatively small.

Finally, the third result suggests that the co-movement in the policy interest rates of the U.S. and Mexico seems to be accounted for by shocks affecting U.S. spending. In other words, if only these shocks were observed, the correlation between the Fed funds rate and Banco de México's target rate would be close to 90%, as opposed to 70% in the data. The lower correlation follows from the fact that the Mexican policy rate reacts to domestic and other non-U.S. shocks that affect real activity and inflation. Therefore, both domestic and foreign factors are important for the determination of the Mexican policy interest rate.

The literature studying the propagation of shocks across borders is vast and growing. Some recent examples include Cushman and Zha (1997), Canova (2005), Uribe and Yue (2006), Justiniano and Preston (2010), Mumtaz et al. (2011), International Monetary Fund (2014), García-Verdú and Ramos-Francia (2014), Takáts and Vela (2014), Mohanty (2014), Fink and Schüler (2015), Caceres et al. (2016), Cesa-Bianchi and Sokol (2017), Dedola et al. (2017), Vicondoa (2019), among others. Our results complement their findings, as we emphasize the key role played by domestic financial markets in the dissemination of foreign shocks.[2]

In contrast to these studies, we consider broader measures of the opportunity cost of funds in a small open economy. While most papers only include the so-called *country interest rate* (i.e., the difference between the yields of dollar-denominated long-term domestic bonds and 10-year U.S. Treasury bonds),[3] we have also included the domestic policy rate and the term premium of domestic-currency denominated bonds. We argue that the country interest rate is an incomplete indicator of the cost of funds, since it excludes both the short- and long-term end of the domestic yield curve. In this paper, we show that it is important to include domestic financial market rates in order to dissect the propagation mechanism of U.S. shocks into the domestic economy.

This paper also differentiates from the majority of the aforementioned papers as we identify several shocks simultaneously, rather than one or two shocks at a time. Our strategy allows us to disentangle the relative importance of shocks in

[1] For further details, see this entry of the NY Fed's Liberty Street Economics Blog: http://libertystreeteconomics.newyorkfed.org/2013/04/do-treasury-term-premia-rise-around-monetary-tightenings-.html.

[2] Notably, Caceres et al. (2016) suggest that changes in U.S. interest rates transmit to long-term interest rates in small open economies mainly through unexpected changes in the U.S. term premium. In this regard, Chen et al. (2014) find that the U.S. monetary policy significantly affected capital flows and asset prices in EMEs during the implementation of unconventional monetary policy.

[3] See for instance Uribe and Yue (2006) or Vicondoa (2019).

driving the business cycle in the U.S. and Mexico. In this sense, our work is similar to Canova (2005), who studies the transmission of shocks to U.S. monetary policy, aggregate demand, and aggregate supply to Latin American countries, including Mexico, in the sample period 1990Q1-2002Q2. In contrast to our results, Canova (2005) finds that U.S. demand and supply shocks did not explain output fluctuations in Latin America, whereas U.S. monetary policy shocks and global factors did. With regard to Mexico, the difference between his results and ours can be attributed to at least two factors. First, it may be due to the different sample period. For example, exports and financial activities had a lower weight and importance in the 1990's Mexico in comparison to today's. And second, the monetary policy regime was very different. In the late 1990s, Mexico transitioned from a fixed-exchange-rate regime to a floating regime, in which monetary policy became independent and started focusing on price stability.

In addition, Fink and Schüler (2015) analyze the effect of U.S. systemic financial stress shocks in eight emerging market economies (EMEs), including Mexico. They find that real activity in EMEs and the U.S. falls for several months after an increase in the stress of U.S. financial markets. These authors also emphasize the importance of the financial-contagion channel for the propagation of the shock. We find akin results in our analysis, since the shock to the cost of external finance of U.S. firms is similar to Fink and Schüler (2015)'s financial stress shock. Accordingly, we find that real activity in the U.S. and Mexico persistently decrease after the shock, while the term premium rises in both countries. However, we find contrasting results concerning the exchange rate. In our case, the peso depreciates against the dollar after the shock; in theirs, it appreciates. The discrepancy may be due to the sample period chosen by Fink and Schüler (2015), since nominal variables followed a decreasing path triggered by a disinflationary process from the late 90s till 2003 in Mexico. Our sample excludes this period.

The rest of the paper is organized as follows. Section 2 reviews stylized facts between Mexican and U.S. variables. Section 3 describes the VAR model, along with the identification strategy of U.S. structural shocks. Section 4 describes the data and estimation algorithms. Section 5 discusses the variance decomposition, selected impulse-response functions, a co-movement analysis based on the model's historical decomposition, and robustness checks. Section 6 concludes.

2. Some stylized macro facts between Mexico and the U.S.

Mexico is an emerging market and a small open economy, whose major trading partner is the U.S. On average, from 2002 to 2017, about 80% of all Mexican exports were bought by the U.S. In contrast, U.S. exports to Mexico were only 14% of the total volume.[4] The difference is partly explained by the fact that the U.S. economy is 8 times larger than the Mexican one, as measured by GDP in PPP terms. In addition, for the period mentioned, about 50% of foreign direct investment received in Mexico originated in the U.S. Strong relationships in trade and investment sectors partially explain the co-movement between economic activity in Mexico and that of its northern neighbor. Panel (a) in Fig. 1 shows the output gaps for Mexico and the U.S. in the period spanning from 2002 to early 2018. These variables present a 65% correlation during this period.

A second determinant of the co-movement in real activity relates to tight financial links between Mexico and the U.S. Panel (b) in Fig. 1 displays two measures of external finance premium: one is the difference between *Moody*'s Baa corporate bond yield averages and the 10-year Treasury security yield in the U.S., while the other is *J.P. Morgan*'s EMBI + spread for Mexico, which measures the sovereign premium in bonds issued by the Mexican government in U.S. dollars (all of which are long-term bonds). The EMBI + spread can be interpreted as a floor of the cost of external financing in international markets for Mexican corporate firms.[5] In the literature, the EMBI + spread is also known as the *country-interest-rate* spread (see Uribe and Yue, 2006). The correlation between the two measures of the cost of external finance is close to 75% for the aforementioned period. In addition, panel (c) shows the difference between the yields of 10-year and the 2-year government bonds denominated in domestic currency in Mexico and the U.S. This spread summarizes the term premium in both countries, which is an anti-cyclical indicator that helps to predict future economic activity. The correlation of the *10-year-2-year* spread between the two countries amounts to 67% for the period of interest.

In sharp contrast with real and financial activities, price setting in both countries seems to be independently determined. Panel (d) in Fig. 1 shows the monthly inflation rates of core *personal consumption expenditures* (PCE) for the U.S. (i.e., excluding food and energy), and the core *consumer price index* (CPI) for Mexico. The correlation between these two variables is virtually zero.

Despite the lack of co-movement in the inflation rates, the short-run nominal interest rates of the two countries co-move with a correlation of 68% for the period shown in panel (e) of Fig. 1. These rates are highly influenced by the monetary policy target or reference rate, which we simply call the *policy rate*. The correlation goes up to 73% if we approximate U.S. monetary policy with a *shadow* Fed funds interest rate. This indicator is not constrained by the zero lower bound, and serves as a proxy of the degree of accommodation induced by unconventional monetary policies implemented by the Fed.[6]

[4] In what follows, the ratios presented in this section are computed with data from Mexico's *Instituto Nacional de Estadística y Geografía* (INEGI) and *Ministry of Economics*, the U.S. *Bureau of Economic Analysis* (BEA), and the *World Bank*.

[5] Notice also that fluctuations in the EMBI + spread may also capture deviations in the uncover interest rate parity condition, since this spread is computed as the difference between the yield of dollar-denominated long-term bonds issued by the Mexican government and the yield of 10-year U.S. Treasury bonds. As both bonds are settled in dollars, exchange-rate risk is absent, so what remains is a UIP deviation containing sovereign risk.

[6] In panel (e) of Fig. 1, we use the average of three different shadow measures proposed in the literature: one by Krippner (2015), another by Wu and Xia (2016), and another one by Lombardi and Zhu (2018). Our average shadow measure takes negative values during the Fed funds rate's zero lower bound period, from 2009M1 to 2015M12, and it is equal to the Fed funds rate outside this period.

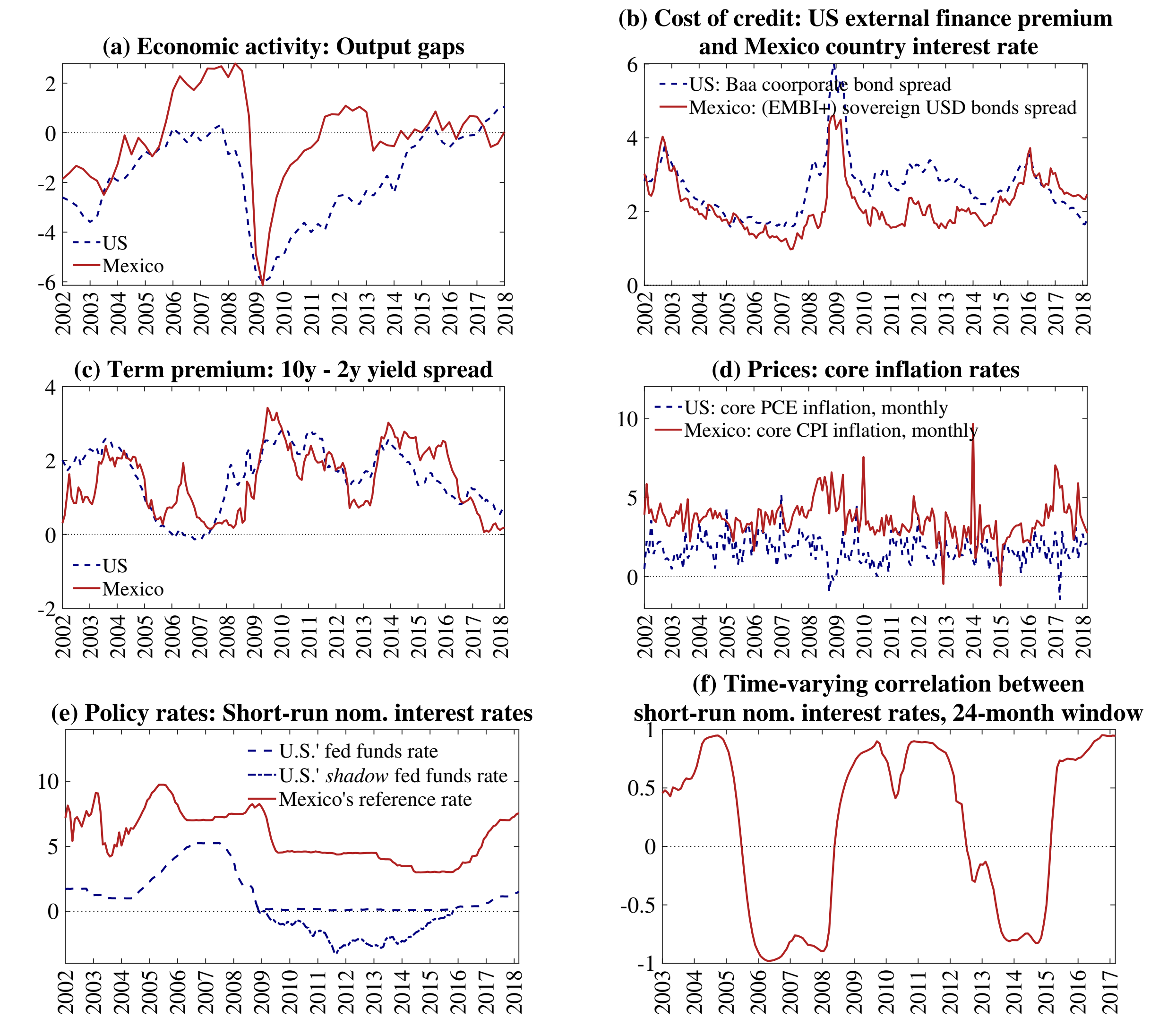

Fig. 1. Selected U.S. and Mexican variables. **Note:** The average shadow Fed funds rate measure takes negative values during the Fed funds rate's zero lower bound period, from 2009M1 to 2015M12, and is equal to the Fed funds rate outside that period. Details of the data are provided in the online appendix, Section A. **Source:** Mexican data are provided by *INEGI*, *Banco de México*, *Valmer*, and *Bloomberg*. U.S. data are provided by the *St. Louis Fed* FRED database, Krippner (2015),Wu and Xia (2016), and Lombardi and Zhu (2018).

Although the policy rates show similar trends north and south of the *Río Bravo/Grande*, there are episodes of absent co-movement. Panel (f) of Fig. 1 presents a time-varying correlation between the policy rates of Mexico and the U.S., which we compute through a centered 24-month rolling window. Two subperiods stand out with a clear negative correlation: 2006–2008 and 2013–2015.

In light of these stylized facts, we prompt the following questions:

1. *What type of macro shocks explain the co-movement registered in real and financial activities in the two countries?, and*
2. *Which US shocks raise the probability of observing co-movement in the policy rates of the two countries at short-run horizons (and which ones reduce this probability)?*

In the following sections, we aim to provide plausible answers to these questions.

3. Empirical model

We now describe the VAR model and the set of assumptions that help us identify various aggregate shocks originating in the U.S. These shocks stem from exogenous variations in the degree of economic policy uncertainty, productivity, aggregate demand, financial sector, production costs, and monetary policy in the U.S. economy. Since we aim to measure the propagation of these shocks into Mexico, we resort to a two-country VAR model with block-exogeneity, which we describe next.

3.1. A structural VAR model for a small open economy

Let $X_{1,t}$ be a vector of n_1 foreign (U.S.) variables and $X_{2,t}$ a vector of n_2 domestic (Mexican) variables at time t. We assume that $X_{1,t}$ is block exogenous to the domestic variables, i.e. neither the dynamics of $X_{2,t}$, nor structural shocks to this vector, affect any of the values in $X_{1,t+k}$ for any $k \geqslant 0$. Indeed, we assume that the domestic economy is too small in comparison to the large foreign economy.[7] Let $X_t = \left[X'_{1,t}, X'_{2,t}\right]'$ and $n \equiv n_1 + n_2$, so the structural aggregate dynamics of the domestic and foreign economies are represented by [8]

$$A^{-1}X_t = C + \sum_{\ell=1}^{p} B_\ell X_{t-\ell} + \upsilon_t, \tag{1}$$

where C is an $n \times 1$ vector of constants, A^{-1} is an invertible $n \times n$ matrix of contemporaneous structural relations between the variables of the system, B_ℓ are $n \times n$ matrices of persistence, and υ_t is a vector of structural innovations with mean zero, not autocorrelated, and with a variance–covariance matrix equal to $\mathrm{E}\{\upsilon_t \upsilon'_t\} = \mathbf{I}_n$, i.e. the identity matrix.

The reduced-form version of this model results from pre-multiplying the above system by A, which leads to

$$X_t = \tilde{C} + \sum_{\ell=1}^{p} \tilde{B}_\ell X_{t-\ell} + \xi_t, \tag{2}$$

where $\tilde{C} \equiv AC, \tilde{B} \equiv AB$, and $\xi_t \equiv A\upsilon_t$, which is a vector of reduced-form innovations with a variance–covariance matrix equal to $\mathrm{E}\{\xi_t \xi'_t\} = \mathbf{\Omega}$. It follows that matrix A satisfies $AA' = \mathbf{\Omega}$.

To satisfy block exogeneity, matrices A and $\tilde{B}_\ell$ contain a block of zeros, such that

$$A \equiv \begin{bmatrix} A_{zz} & \mathbf{0} \\ A_{zy} & A_{yy} \end{bmatrix} \quad \text{and} \quad \tilde{B}_\ell \equiv \begin{bmatrix} \tilde{B}_{zz,\ell} & \mathbf{0} \\ \tilde{B}_{zy,\ell} & \tilde{B}_{yy,\ell} \end{bmatrix}$$

where A_{zz} denotes the impact effect of U.S. shocks on U.S. variables, A_{zy} indicates the said effect on Mexican variables, and A_{yy} represents the impact effect of domestic and non-U.S. shocks on Mexican variables. Similarly, $B_{zz,\ell}$ expresses the effect of U.S. variables lagged ℓ periods on current U.S. variables, $B_{zy,\ell}$ is the said effect on current Mexican variables, and $B_{yy,\ell}$ carries the effect of Mexican variables lagged ℓ periods on current Mexican variables.

3.1.1. Structural restrictions

We set a strategy to identify 7 different U.S. aggregate structural disturbances: a productivity shock, a cost-push shock, an economic policy uncertainty shock, two types of monetary policy shocks, an aggregate demand shock, and a financial shock affecting the cost of credit. To achieve a full-ranged identification, we include 7 indicators of the U.S. economy which measure (1) the degree of economic policy uncertainty, epu_t^{us}, (2) TFP growth, Δtfp_t^{us}, (3) the output gap, $\hat{y}_t^{us}$, (4) the inflation rate, π_t^{us}, (5) the policy rate, i_t^{us}, (6) the term premium, tp_t^{us}, and (7) a measure of the cost of external finance, efp_t^{us}. In turn, for the Mexican economy we include indicators related to (1) the output gap, $\hat{y}_t^{mx}$, (2) the inflation rate, π_t^{mx}, (3) the policy rate, i_t^{mx}, (4) the depreciation of the nominal exchange rate, Δs_t^{mx}, (5) the country-interest-rate spread, $embi_t^{mx}$, and (6) the term premium, tp_t^{mx}. Further details on these variables are displayed in Section 4.1 and Appendix A.

To identify U.S. structural shocks, we apply a set of sign and zero restrictions to the responses of U.S. variables to U.S. shocks, and we impose no restrictions to the responses of Mexican variables. Table 1 summarizes the structural restrictions at two specific horizons: the impact period, and 6 months after the shock. Together, these restrictions ensure that the 7 structural shocks are indeed different from each other, as the analysis of impulse-response functions shows in Section 5 and Appendix B. In the table, a '+' sign indicates that the response of variable x is expected to be positive h periods after a positive innovation in structural shock $\upsilon_{i,t}$, a '−' sign describes the opposite, and a '0' states that variable x does not respond to the shock. We impose '0' restrictions only on the impact period, following the rationale that some variables need time to adjust to certain shocks. A blank space in Table 1 means that the response of variable x is unrestricted at the specified horizon.

We first review the zero restrictions imposed, and after that we explain the rationale behind the assumed sign restrictions. At the end of this section, we illustrate the reasons behind the separation of the monetary policy shocks into two types in further detail.

[7] In Appendix D we formally test this assumption using the results of our Bayesian estimation.

[8] We omit non-U.S. world variables from the model to keep the analysis simple. However, it would be worth studying this dimension in an extension of this work.

Table 1
Structural effects of U.S. shocks on U.S. variables.

Shock	epu^{us}_{t+h}	Δtfp^{us}_{t+h}	$\hat{y}^{us}_{t+h}$	π^{us}_{t+h}	i^{us}_{t+h}	tp^{us}_{t+h}	efp^{us}_{t+h}
RESTRICTIONS ON IMPACT, $h = 0$							
Econ. policy uncertainty	+						
Productivity	0	+					
Aggregate demand	0	0	+				
Cost-push	0	0		+			
Monetary policy, type I	0	0	0		+	+	
Monetary policy, type II	0	0	0		+	–	
Financial	0	0	0		0		+
RESTRICTIONS AT $h = 6$ MONTHS AFTER THE SHOCK							
Econ. policy uncertainty							
Productivity							
Aggregate demand			+	+	+		
Cost-push			–	+			
Monetary policy, type I			–	–	+		
Monetary policy, type II			–	–	+		
Financial					–		+

Note: The sign refers to the impact response of the variable of interest to a structural shock. A '+' (or '–') indicates that $\partial x_{t+h}/\partial \upsilon_{i,t} > 0$ (or < 0), where x_{t+h} is the response of the variable of interest h periods after the shock, and $\upsilon_{i,t}$ is the structural shock studied. Blank spaces mean that no restriction is imposed on the variable.

Zero restrictions. We assume that the measure of economic policy uncertainty and TFP growth respond contemporaneously to their own innovation, rather than to any other structural aggregate shock.[9] The exception is TFP growth, which may respond on impact to the economic policy uncertainty shock. We allow this response since we are constrained by the number of zeros we can impose in the rows of matrix A. Indeed, we cannot impose $n - 1$ zeros in two rows of A because the identification strategy becomes over identified. Nonetheless, the assumption that productivity may respond to uncertainty is not economically ungrounded. Productivity may fall after an increase in uncertainty because firms could follow a *wait-and-see* strategy, pausing their investment, production and hiring decisions until the uncertainty clears.[10]

In turn, we allow output to respond on impact to real shocks, such as uncertainty, productivity, costs of inputs, and aggregate demand. In contrast, output is not expected to respond contemporaneously to shocks that hit financial markets first, such as a shock to the cost of external finance or to monetary policy. In these cases, we assume that output needs more than one period to react to these innovations. Finally, we also assume that the U.S. policy rate does not react on impact to the financial shock. This assumption helps us to differentiate the financial shock from the monetary policy shocks. Fink and Schüler (2015) also adopt this identification strategy.

Sign restrictions: aggregate demand, cost-push, and financial shocks. The sign restrictions imposed on output, prices and the policy rate after an aggregate demand and cost-push shocks are motivated by a vast class of DSGE models with nominal rigidities, such as Smets and Wouters (2007). These signs allow us to differentiate the referred shocks from one another. Six months after an aggregate demand shock, we assume that output, prices, and the policy rate move in the same direction. In contrast, six months after a cost-push shock we assume that output and prices move in opposite directions, while the policy rate bears no restriction.

For the financial shock, we require that the external finance premium rises after a positive innovation of this shock and six months after. We remain agnostic about the responses of the rest of U.S. variables, with the exception of the U.S. policy interest rate, which we assume displays a fall six months after the shock. Indeed, similar to Fink and Schüler (2015), we assume that the Federal Reserve answers to a tightening in financial conditions by loosing monetary policy, although this response may happen with some delay. This assumption, along with the zero restriction of the U.S. policy rate at the impact period, ensures that the financial shock is uniquely identified.

Sign restrictions: monetary policy shocks. We assume that six months after a monetary policy shock, prices and output co-move and head to the direction opposite to that of the policy rate. However, we split the monetary policy shocks as follows:

- ***Type I***: *in the impact period, the policy rate and the term premium move in the* <u>*same*</u> *direction, and*
- ***Type II***: *in the impact period, the policy rate and the term premium move in* <u>*opposite*</u> *directions.*

[9] We adopted zero restrictions on impact on these variables because preliminary exercises suggest that the EPU index and TFP are primarily determined by their own innovations.

[10] Despite this negative correlation, we impose no sign on the response of TFP growth to an economic policy uncertainty shock, as can be seen in Table 1. It is worth mentioning that in robustness exercises, we have changed the order between epu^{us}_t and Δtfp^{us}_t, so that EPU answers to TFP shocks on impact. All our results are quantitatively very similar.

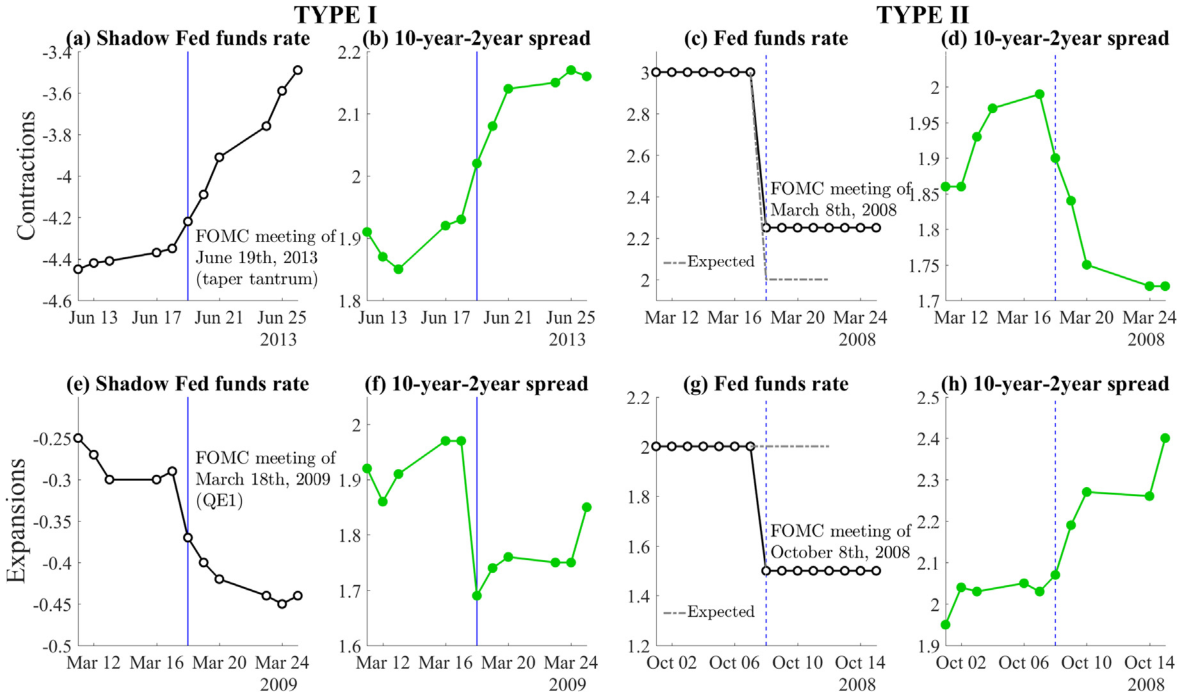

Fig. 2. Specific monetary policy surprises in the U.S. **Source:** Federal Reserve Board and Krippner (2015).

We differentiate between the two types of monetary policy shocks because we noticed that the U.S. term premium reacted qualitatively different to monetary policy surprises. In Appendix C we delve into the motivation of this new identification strategy. To gain intuition, notice that long-term interest rates can be viewed as the sum of two components: (i) investors' expectations about the future course of short-term interest rates, and (ii) a compensation investors ask for bearing the risk that short-term interest rates may evolve differently than expected. This component is also called *term premium*. The difference between the yields of long-term bonds and medium-term bonds, such as the 10-year-2-year spread reviewed in the previous section, partially removes the expectations component of the long-term bond interest rate. Therefore, the term premium, computed as the difference between long- and medium-term yields, can be interpreted as a rough measure of the term premium. If investors perceive favorable economic conditions in the future, the term premium tends to decrease, while if they anticipate adverse or uncertain economic conditions ahead, the term premium tends to rise.[11]

In particular, we assume that under a *type I* contractionary monetary policy shock, the term premium increases on impact. The interpretation is that investors judge that the restrictive policy stance may destabilize the economy. An example of such a shock is the *taper tantrum* in mid 2013. Fig. 2 shows the reaction of financial markets to the FOMC's policy decision statement of June 19, 2013. This decision was preceded by comments of former Fed's chairman Ben Bernanke pointing out that the FOMC could start tapering its QE programs *soon*. When the FOMC statement confirmed slight improvements in economic projections, investors interpreted that the tapering was imminent, and therefore too early relative to prevalent economic indicators. In consequence, the 10-year-2-year spread rose sharply in the days following the FOMC statement, as shown in panel (b) of Fig. 2. In turn, panel (a) shows the monetary policy restriction through an increase in Krippner (2015)'s daily measure of the notional *shadow* Fed funds rate.[12]

In contrast, we assume that under a *type II* contractionary monetary policy shock, the term premium falls on impact. In this case, we interpret that investors judge the surprise restriction in the policy stance as an action consistent with economic stability. For example, on March 18, 2008, at the onset of the global financial crisis, markets expected a cut in the Fed funds rate of about 100 basis points (bp), according to the Fed funds rate futures a day before the policy announcement (see panel (c) in Fig. 2). As a surprise to the markets, the Fed cut the policy rate by 75 bp, i.e. 25 bp less than market expectations, resulting in a lower-than-expected expansion *or* unexpected contraction. Investors interpreted the policy decision as a signal that economic conditions were not as adverse as previously thought. As a result, panel (d) shows that the term premium fell on the days that followed the policy announcement.

The data also support *type I* and *type II* unexpected monetary policy expansions. For instance, on March 18, 2009, the FOMC announced its first comprehensive quantitative easing program, also known as *QE1*. Although the program was

[11] There are alternative measures of the U.S. term premium based on affine term-structure models. In particular, the methods proposed by Kim and Wright (2005) and Adrian et al. (2013) are quite popular. However, given that there is no a consensus over which method is more reliable, we decided to include rough data of the U.S. term premium into the VAR model, i.e. the 10-year-2 year spread.

[12] We use Krippner (2015)'s shadow measure in Figs. 2 and 3 because it is the only shadow indicator available on a daily basis.

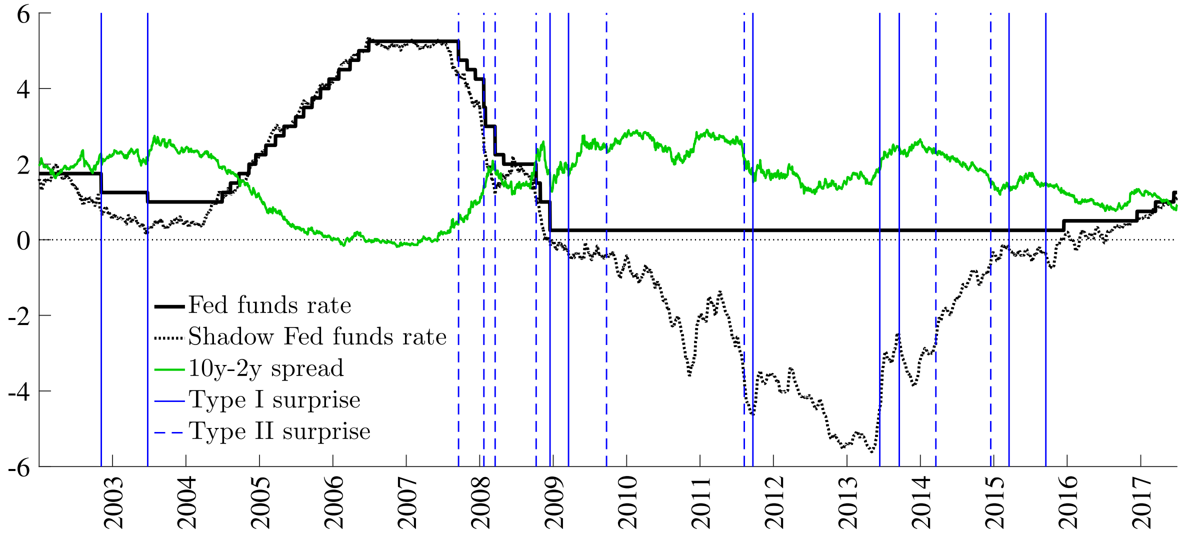

Fig. 3. Monetary policy surprises in the U.S. **Source:** *Federal Reserve Board* and Krippner (2015).

expected, it contained details about the intended large-scale asset purchases that cleared uncertainty and surpassed investors' expectations. As a result, the 10-year-2-year spread fell by more than 30 bp following the expansion announcement, resulting in a type I unexpected monetary policy expansion (see panels (e) and (f) in Fig. 2). A *type II* expansion occurred on October 8, 2008, when the FOMC surprised the markets by cutting the policy rate by 50 bp in an unscheduled decision. Markets interpreted the movement as a clear signal that economic conditions were worsening faster than expected, and so the 10-year-2-year spread rose 40 bp in less than a week (see panels (g) and (h) in Fig. 2).

Fig. 3 marks the FOMC decisions that contain surprises since 2002, according to some studies in the literature. Up to December 2008, the monetary policy surprises are captured through Kuttner (2001)'s methodology. From 2009 onwards, the surprises are taken from Swanson (2016) and Gupta et al. (2017), who use high-frequency asset-price data to identify the monetary policy shocks. We classify each of these surprises as *type I* or *type II* according to the behavior of the 10-year-2-year spread in the day of the surprise. We marked with a vertical line type I surprises (10 in total), and with a dashed line *type II* surprises (9 in total). The balanced distribution of these surprises justifies the separation between the two types of shocks. In Appendix F we document each of the surprises marked in the figure.

4. Estimation

In this section we present the data used in our model. Then, we describe the procedure to impose exogeneity in the reduced-form VAR model. Finally, we present the estimation procedure of the structural-form VAR model.

4.1. Data

The data used in the estimation, shown in Fig. 4, spans from the period January 2002 to March 2018, at a monthly frequency.[13] For the U.S., we use the economic policy uncertainty (EPU) index of Baker et al. (2016), the utilization-adjusted TFP growth series of Basu et al. (2006), a monthly linear interpolation of the *CBO* output gap, the inflation rate of the core PCE price index, a shadow measure of the Fed funds rate (see footnote 6), the term premium, and a proxy for the cost of external financing computed by the difference between *Moody*'s Baa corporate bond yield averages and the 10-year Treasury bond yield. For Mexico, we use and a Bayesian monthly interpolation of *Banco de México*'s output gap official estimate, following Elizondo (2012)'s methodology, the core CPI inflation, the overnight bank funding rate (which is tightly related to *Banco de México*'s policy rate), the percent change in the peso/dollar exchange rate, *J.P. Morgan*'s EMBI + spread for Mexico as a measure of the country-interest-rate spread, and the term premium computed as the difference between the 10-year and 2-year zero-coupon Treasury bond yields issued by the Mexican government in domestic currency. All of these variables are explained in detail in Appendix A. U.S. data are obtained from the *St. Louis Fed* FRED database and the *San Francisco Fed*, except for the shadow Fed funds rate. Mexican data are obtained from *INEGI*, *Banco de México*, *Valmer*, and *Bloomberg*.

[13] We start our sample in 2002 due to a change in the monetary policy regime that took place in that year and to the transition from a fixed to floating forex regime that was a direct consequence of the 1995 Tequila crisis. As a result of these two changes, the time series properties of inflation, the short-run nominal interest rate, and the exchange rate changed dramatically. In this regard, Chiquiar et al. (2010) show that Mexican inflation became stationary around the year 2001, once Banco de México started pursuing annual inflation targets and prompted a disinflationary process.

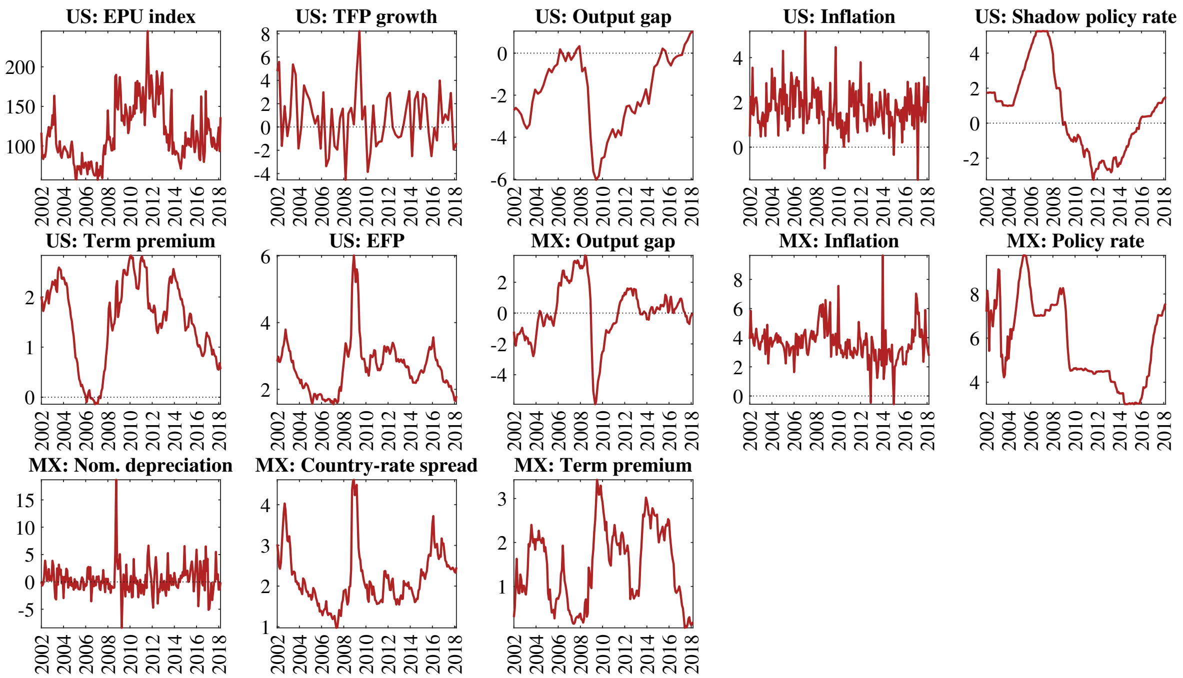

Fig. 4. Data used in the estimation. **Source:** Data for the U.S. are obtained from the *St. Louis Fed* FRED database and the *San Francisco Fed*, except for the shadow Fed funds rate which we computed as stated in footNote 6. Data for Mexico are obtained from *INEGI*, *Banco de México*, *Valmer*, and *Bloomberg*.

4.2. Reduced-form VAR estimation

It is convenient to represent the VAR model in Eq. (2) in terms of the demeaned vector $Z_t \equiv X_t - \mu$, where $\mu = \left(I_n - \sum_{\ell=1}^{p} \tilde{B}_\ell\right)^{-1} \tilde{C}$ is the vector of the steady-state values of the variables in X_t. Therefore, the reduced-form VAR model can be rewritten as

$$Z_t = \tilde{\mathbf{B}} W_t + \xi_t,$$

where $\tilde{\mathbf{B}} \equiv \left[\tilde{B}_1, \tilde{B}_2, \ldots, \tilde{B}_p\right]$ is an $n \times pn$ matrix, and $W_t \equiv [Z_{t-1}, Z_{t-2}, \ldots, Z_{t-p}]'$ is a $pn \times 1$ vector. Further, the compact version of the above system of equations is given by

$$\mathbf{Z} = \mathbf{W}\tilde{\mathbf{B}}' + \xi,$$

where $\mathbf{Z} \equiv [Z_1, Z_2, \ldots, Z_T]'$ is a $T \times n$ matrix of observations, $\mathbf{W} \equiv [W_1, W_2, \ldots, W_T]'$ is a $T \times np$ matrix of regressors, $\boldsymbol{\xi} \equiv [\xi_1, \xi_2, \ldots, \xi_T]'$ is a $T \times n$ matrix of residuals, and T is the sample size.

The reduced-form VAR is estimated through Bayesian methods with an *independent Normal-inverse-Wishart* prior for $\tilde{\mathbf{B}}$ and $\boldsymbol{\Omega}$, and a long-run or steady-state prior à *la* Villani (2009) for μ. We included 12 lags in the estimation, which ensures that the estimated reduced-form innovations are white noise.

4.2.1. Prior and posterior distributions

The prior for $\tilde{\mathbf{B}}$ is a multivariate normal such that $p\left(\text{vec}(\tilde{\mathbf{B}})\right) \sim N\left(\tilde{\mathbf{B}}_0, \mathbf{H}\right)$. The prior mean and variance follow a structure similar to a Minnesota prior, with the difference that we enforce block exogeneity by setting a tight prior in the block of zeros of matrices $\tilde{B}_\ell$. These assumptions indeed presume that most of parameters of the BVAR model are equal to zero, except for the first order autoregressive parameters (i.e., we expect that *at least* 13 parameters will be statistically different from zero).[14] In particular, the prior mean of $\tilde{\mathbf{B}}$ is set according to

$$(\tilde{\mathbf{B}}_{0,\ell})_{ij} = \begin{cases} \rho_i, & \text{for } j = i, \ell = 1 \\ 0, & \text{otherwise}, \end{cases}$$

[14] In fact, we obtained 83 parameters statistically different than zero at one-standard deviation confidence band.

where $(\tilde{\mathbf{B}}_{0,\ell})_{ij}$ is the element in position (i,j) in the partition matrix $\tilde{\mathbf{B}}_{0,\ell}$, which corresponds to the ℓ-th lag matrix in $\tilde{\mathbf{B}}_0$, and ρ_i is the autoregressive coefficient in the OLS regression $(Z_t)_i = \alpha_i + \rho_i (Z_{t-1})_i + \eta_{i,t}$, where α_i is the intercept and $\eta_{i,t}$ is an innovation with standard deviation equal to σ_i. If ρ_i approaches 1, the prior sets the belief that variable $(Z_t)_i$ is very persistent, while if ρ_i approaches 0, the prior states that the variable behaves similar to a white noise.

Further, the variance of the prior mean, $\tilde{\mathbf{B}}_0$ follows the sequence

$$(\mathbf{H})_{ij} = \begin{cases} \left(\lambda_1/\ell^{\lambda_3}\right)^2, & \text{for } j = i \\ \left((\sigma_i\lambda_1)/(\ell^{\lambda_3}\sigma_j\lambda_2)\right)^2, & \text{for } j \neq i \\ 0, & \text{for } i \leqslant n_1, j > n_1, \end{cases}$$

where λ_1 controls the overall tightness of the prior, λ_2 weighs the relative importance of lags of variable $(Z_t)_j$ on variable $(Z_t)_i$, and λ_3 measures the speed of decay of the effect of the ℓ-th lag on current variables values. We impose block exogeneity in the third line of the above expression, where we specify that there is a tight prior that $\partial X_{1,t}/\partial X_{2,t-\ell} = 0$ for any $\ell > 0$, i.e. that the lags of domestic variables do not affect current values of the foreign variables. The hiper-parameters λ_i are set to achieve a good forecasting performance of the model. They take the values $\lambda_1 = 0.2$, $\lambda_2 = 0.4$, and $\lambda_3 = 1$.[15]

The prior for $\mathbf{\Omega}$ is an inverse Wishart distribution such that $p(\mathbf{\Omega}) \sim iW(\bar{S}, \varphi)$, where $\bar{S}$ is the scale matrix and φ denotes the degrees of freedom. We set $\bar{S} = \text{diag}(\sigma_1, \sigma_2, \ldots, \sigma_n)$ and $\varphi = n + 1$.

Following Villani (2009), we set a normal prior for μ such that $p(\mu) \sim N(\mu_0, \mathbf{H}_\mu)$, where μ_0 contains the prior steady-state values of the variables in the model, and $\mathbf{H}_\mu$ is the variance–covariance matrix of this prior. We set the prior that in the long run the output gaps in the U.S. and Mexico reach zero, the inflation rates are equal to their targets (2% in the U.S. and 3% in Mexico), and the measure of the external finance premium, the 10-year-2-year spread, and the peso/dollar nominal exchange rate percent change equal their historical means. Further, we assume that the Fed funds rate long-run convergence level is 3%, which equals the upper bound of the FOMC's March 2018 long-run projection for this variable. Finally, we assume that the long-run value of *Banco de México*'s policy rate equals 5.5%, which is the medium point of the estimated range of the nominal neutral rate of interest in Mexico.[16] In turn, $\mathbf{H}_\mu$ is equal to $\bar{S}\lambda_\mu$, where λ_μ is the tightness of the steady-state prior, which we set to 0.5.[17]

The conditional posterior distributions of the estimated parameters are equal to

$$\pi\left(\text{vec}(\tilde{\mathbf{B}})|\mathbf{\Omega}, \mu, \mathbf{X}\right) \sim N(M^*, V^*) \tag{3}$$

$$\pi\left(\mathbf{\Omega}|\tilde{\mathbf{B}}, \mu, \mathbf{X}\right) \sim iW(\bar{\mathbf{\Omega}}, T + \varphi) \tag{4}$$

$$\pi\left(\mu|\tilde{\mathbf{B}}, \mathbf{\Omega}, \mathbf{X}\right) \sim N\left(M^*_\mu, V^*_\mu\right) \tag{5}$$

where $\mathbf{X} \equiv [X_1, X_2, \ldots, X_T]'$ is a matrix containing all variables and observations. The conditional posterior moments are given by

$$\begin{aligned} V^* &= \left(H^{-1} + \mathbf{\Omega}^{-1} \otimes \mathbf{W}'\mathbf{W}\right)^{-1}, \\ M^* &= V^*\left(H^{-1}\text{vec}(\tilde{\mathbf{B}}_0) + \mathbf{\Omega}^{-1} \otimes \mathbf{W}'\mathbf{W}\text{vec}(\tilde{\mathbf{B}}_{ols})\right), \\ \bar{\mathbf{\Omega}} &= \bar{S} + \xi'\xi, \\ V^*_\mu &= \left(H_\mu^{-1} + U'(D'D \otimes \mathbf{\Omega}^{-1})U\right)^{-1}, \\ M^*_\mu &= V^*_\mu\left(H_\mu^{-1}\mu_0 + U'\text{vec}(\mathbf{\Omega}^{-1}\mathbf{Y}'D)\right), \end{aligned}$$

where $\tilde{\mathbf{B}}_{ols} = (\mathbf{Z}'\mathbf{Z})^{-1}(\mathbf{Z}'\mathbf{W})$, $U = [I_n, \tilde{B}_1, \tilde{B}_2, \ldots, \tilde{B}_p]'$, $D = [\mathbf{1}_{T\times 1}, -\mathbf{1}_{T\times p}]$, $\mathbf{Y} = \mathbf{X} - (\mathbf{W} + \mathbf{1}_{T\times p} \otimes \mu')\tilde{\mathbf{B}}'$, and $\mathbf{1}_{k\times h}$ is a matrix of ones with k rows and h columns.

We estimate the joint posterior distribution through a Gibbs sampling algorithm that follows the order of Eqs. (3)–(5). The algorithm shows good convergence properties after 20,000 iterations, from which we save the last 1,000 for inference and discard the rest. In Appendix D we present the convergence diagnostics of the Gibbs sampler.

[15] The triplet $(\lambda_1, \lambda_2, \lambda_3)$ belongs to the neighborhood where the marginal likelihood is maximal. To find the triplet, we evaluated the likelihood in a grid of values in the interval $[0.1, 1] \times [0.1, 1] \times 0, 1, 2$. The marginal likelihood was maximized with values of λ_1 closed to 0.1. Since this parameter controls the overall tightness of the prior, we decided to lose a bit this prior to allow the model to learn more from the data.

[16] See Banco de México Quarterly Report October - December 2017, pp. 65. The report provides the range of the neutral rate in real terms, which is 1.7% to 3.3%, with a medium point of 2.5%. We add the 3% inflation target to compute the mid point of the neutral rate in nominal terms. More details can be seen in Carrillo et al. (2018).

[17] The choice of λ_μ was set to provide the model with a fair chance to learn from the data about the long-run values of variables such as inflation and nominal interest rate.

4.3. Structural VAR estimation

Structural VAR models based on sign restrictions are not uniquely identified, since there is a set of matrices A that satisfy both the structural restrictions and $AA' = \mathbf{\Omega}$. We closely follow the procedure of Binning (2013) to jointly impose sign and zero restrictions to set-identified models.[18] We extend his procedure using the method of Arias et al. (2014), which allows us to properly draw from the posterior distribution of structural parameters. Both Binning (2013) and Arias et al. (2014) make use of the properties of orthonormal matrices, where rotation matrices are used to ensure that both zero and sign restrictions are satisfied. The algorithm to compute the SVAR models, i.e. the set of matrices A, is summarized as follows:

1. *Select randomly a natural number k from the discrete uniform distribution [1,2, ..., 1,000]. Then, draw from the posterior distribution the triplet* $(\mu_k, \tilde{\mathbf{B}}_k, \mathbf{\Omega}_k)$.
2. *Draw a random* $n \times n$ *orthonormal matrix* $\mathcal{Q}$ *from the QR decomposition, such that* $\mathcal{Q}\mathcal{Q}' = I_n$. *Set the candidate model* $A_{k,0} = \mathrm{chol}(\mathbf{\Omega}_k)\mathcal{Q}$, *where* $\mathrm{chol}(\mathbf{\Omega}_k)$ *is the lower-triangular Cholesky decomposition of* Ω_k *.Verify that* $A_{k,0}A'_{k,0} = \mathbf{\Omega}_k$.
3. *Find a rotation matrix* $\mathcal{P}$ *with* $\mathcal{P}\mathcal{P}' = I_n$, *such that* $A_{k,1} = A_{k,0}\mathcal{P}$ *fulfills the zero restrictions. Verify that the rotated candidate model fulfills* $A_{k,1}A'_{k,1} = \mathbf{\Omega}_k$.
4. *Verify whether matrix* $A_{k,1}$ *satisfies the sign restrictions. Keep it if it does, discard it otherwise.*
5. *For draw* $(\mu_k, \tilde{\mathbf{B}}_k, \mathbf{\Omega}_k)$, *repeat 20 times steps 2 to 4.*
6. *Repeat steps 1 to 5 until collecting 5,000 models* $A_{k,1}$ *that satisfy both zero and sign restrictions.*

5. Results

In this section, we present first a central metric of the forecast error variance decomposition that is computed from the set-identified SVAR models. Then, we describe the impulse-response functions of selected shocks and analyse which shocks explain the co-movement observed in real and financial variables in the two countries. Finally, we present a number of robustness checks.

5.1. Variance decomposition

We now present the contribution of the identified shocks to the expected fluctuations of U.S. and Mexican variables. As a first approximation, we grouped the shocks into three principal categories: *supply shocks*, composed by productivity and cost-push shocks; *demand shocks*, which contain shocks to economic policy uncertainty, aggregate demand, and financial markets (i.e. to the external finance premium); and *monetary policy shocks*, which include shocks of *types I* and *II*. Tables 2 and 3 display the forecast-error (or f.e.) variance decomposition of U.S. and Mexican variables to the three shocks categories. The numbers correspond to the mean of the distribution of all identified models $\{A_{k,1}\}_{k=1}^{5,000}$.[19] Further, Fig. 5 and Tables B1 and B2 in the appendix show in detail the specific shock contribution to the f.e. variance of the variables at different horizons.

We discuss first the U.S. variables. According to Table 2, the f.e. variance of the output gap, the policy rate, the term premium, and the external finance premium responds mainly to shocks contained in the *demand* category. The only exceptions are the f.e. variances of the term premium for horizons shorter than six months, and of inflation for horizons shorter than three years. The monetary policy shocks seem to be major contributors to these variances. On the one hand, it seems natural that the fluctuations in the term premium, especially in the short run, respond strongly to monetary policy surprises, given the no-arbitrage conditions existing between financial instruments. But, on the other hand, it seems odd that the f.e. variance of inflation also depends strongly on these surprises. We consider that this is perhaps a dimension that the proposed identification strategy does not handle well, and it requires an analysis that goes beyond the scope of this paper.

Regarding the contributions of specific shocks, from Fig. 5 and Table B1 it can be inferred that the expected fluctuations of the EPU index and TFP growth are largely explained by their own shocks in all horizons. Further, the aggregate demand shock is the largest contributor to the expected fluctuations in the output gap in all horizon. In turn, the contributions of financial and economic policy uncertainty shocks to the f.e. variance of the output gap increase after horizons longer than 10 months. This can be interpreted as a signal that the effects of these shocks are more persistent than the rest of identified shocks. Finally, shocks to monetary policy and productivity seem to have a moderate contribution to the f.e. variance of the output gap.[20]

With respect to financial variables, Fig. 5 shows that the financial shock is the major contributor to the f.e. variance of the external finance premium in all horizons. In turn, both types of monetary policy shocks account for the largest proportion of

[18] Binning (2013)'s method is in itself an extension of Rubio-Ramírez et al. (2010)'s algorithm for exactly identified models.

[19] We prefer to present the results of the mean of the distribution rather than the median. As discussed in Fry and Pagan (2011), the statistics based on the median come from different models, so there is no guarantee that a variance decomposition based on medians sums one across all shocks. However, the sum of the means of the variance decomposition distribution does sum one by construction.

[20] With respect to productivity shocks, notice that, according to the business-cycle theory, this type of shocks affect mainly the natural level of output, or potential growth. Therefore, the output gap, a measure that normalizes potential growth to zero, is not an appropriate metric to account for the effects of productivity shocks on economic activity.

Table 2
Forecast error variance decomposition for U.S. variables.

Shock	epu_t^{us}	Δtfp_t^{us}	$\hat{y}_t^{us}$	π_t^{us}	i_t^{us}	tp_t^{us}	efp_t^{us}
			6 MONTHS HORIZON				
Supply shocks	4.4	91.9	34.3	14.2	12.6	23.3	12.1
Demand shocks	93.4	5.9	61.5	40.6	64.0	35.3	68.4
Monetary policy shocks	2.1	2.2	4.1	45.2	23.4	41.4	19.5
			1 YEAR HORIZON				
Supply shocks	8.1	82.3	23.1	15.0	15.9	21.8	15.1
Demand shocks	88.5	13.2	68.0	40.9	67.9	44.6	68.2
Monetary policy shocks	3.4	4.5	8.9	44.1	16.3	33.6	16.7
			3 YEARS HORIZON				
Supply shocks	11.8	75.5	19.3	15.6	19.6	21.0	18.2
Demand shocks	82.3	16.9	67.4	42.2	70.3	57.6	65.2
Monetary policy shocks	5.9	7.6	13.3	42.2	10.1	21.4	16.6
			LONG RUN				
Supply shocks	13.6	72.8	19.8	16.0	20.3	21.0	18.7
Demand shocks	77.9	18.9	63.8	43.4	67.7	57.0	64.2
Monetary policy shocks	8.5	8.3	16.4	40.7	12.0	21.9	17.1

Note: Each entry corresponds to the mean contribution of the shock category listed in the rows to the forecast error variance decomposition of the variables listed in the columns. The distribution of all identified models corresponds to the set $\{A_{k,1}\}_{k=1}^{5,000}$.

Table 3
Forecast error variance decomposition for Mexican variables.

Shock	$\hat{y}_t^{mx}$	π_t^{mx}	i_t^{mx}	Δs_t^{mx}	$embi_t^{mx}$	tp_t^{mx}
			6 MONTHS HORIZON			
Mexican and non-U.S. shocks	53.1	80.2	84.8	56.8	25.0	61.0
All U.S. shocks	46.9	19.8	15.2	43.2	75.0	39.0
Supply shocks	7.8	7.0	4.7	5.1	8.5	9.6
Demand shocks	30.2	8.6	7.7	26.8	48.8	18.0
Monetary policy shocks	8.9	4.2	2.9	11.2	17.6	11.4
			1 YEAR HORIZON			
Mexican and non-U.S. shocks	34.7	75.6	70.6	55.3	20.2	47.1
All U.S. shocks	65.3	24.4	29.4	44.7	79.8	52.9
Supply shocks	8.7	8.6	7.3	6.2	13.8	9.5
Demand shocks	43.2	10.3	17.1	26.9	49.6	32.0
Monetary policy shocks	13.4	5.4	5.0	11.5	16.4	11.4
			3 YEARS HORIZON			
Mexican and non-U.S. shocks	21.5	65.4	46.9	53.1	16.3	30.1
All U.S. shocks	78.5	34.6	53.1	46.9	83.7	69.9
Supply shocks	17.8	10.7	11.4	7.7	21.7	13.3
Demand shocks	45.4	16.7	32.2	27.3	45.1	43.2
Monetary policy shocks	15.3	7.3	9.6	12.0	16.9	13.4
			LONG RUN			
Mexican and non-U.S. shocks	18.6	53.9	35.0	50.0	14.0	25.2
All U.S. shocks	81.4	46.1	65.0	50.0	86.0	74.8
Supply shocks	18.9	12.9	15.1	8.8	21.8	15.7
Demand shocks	45.3	23.9	37.5	28.5	45.5	44.0
Monetary policy shocks	17.1	9.3	12.3	12.6	18.7	15.1

Note: Each entry corresponds to the mean contribution of the shock category listed in the rows to the forecast error variance decomposition of the variables listed in the columns. The distribution of all identified models corresponds to the set $\{A_{k,1}\}_{k=1}^{5,000}$.

the f.e. variance of the term premium for horizons shorter than three years, followed in importance by the financial shock. Finally, the economic policy uncertainty shock and the financial shock contribute largely to the f.e. variance of the policy rate for horizons longer than 6 months.

We now turn to the Mexican variables. Table 3 shows that, for horizons shorter than or equal to 6 months, Mexican and non-U.S. shocks are major contributors to the f.e. variance of the output gap, inflation, the policy rate, the nominal exchange rate, and the term premium. In contrast, for the same horizon interval, the f.e variance of the country-interest-rate spread (i.e. the EMBI + spread) is mainly explained by U.S. shocks. The table also signals that U.S. shocks appear to be more persistent than Mexican and non-U.S. shocks, since for horizons longer than 1 year the former explain the largest proportion of the f.e. variance of the Mexican output gap, the country-interest-rate spread, and the term premium.

In this regard, at a long-run horizon (10 years), our results suggest that the contribution of U.S. shocks to the f.e. variance of the output gap in Mexico could reach a share of 80%. This result echoes the previous findings in the literature for small open economies. For instance, Justiniano and Preston (2010) find that, at a long-run horizon, U.S. shocks seem to explain 75% of the variance of the output gap in Canada. Mumtaz et al. (2011) find that, at a long-run horizon and for the sample period 1985–2007, 75% of the variance of output growth in Mexico is due to regional shocks (i.e. common shocks affecting

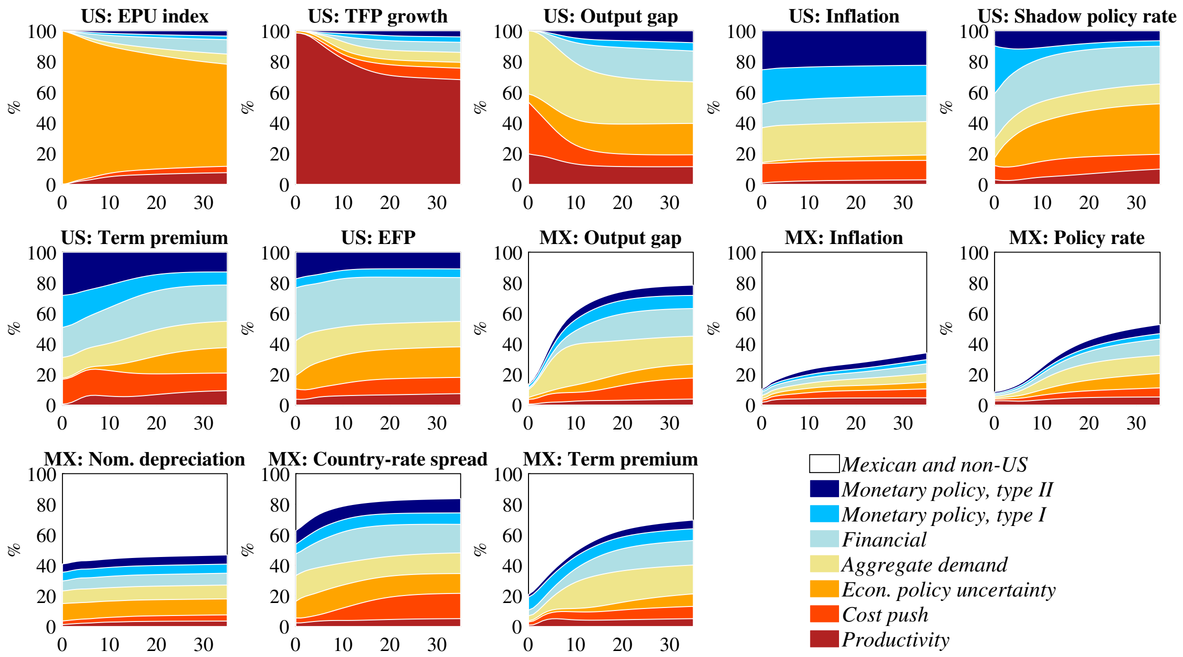

Fig. 5. Forecast error variance decomposition.

Latin American countries). Likewise, the IMF's *World Economic Outlook* of April International Monetary Fund (2014) reports that at a 5-year horizon, U.S. shocks explain 67% of the f.e. variance of output growth in Mexico and Russia, 50% in Malaysia, and 20% in Brazil and South Africa. Our findings are consistent with the numbers retrieved for Mexico. Finally, Fig. 5 shows that the U.S. aggregate demand shock is the largest contributor to the f.e. variance of the output gap in Mexico, followed by the financial shock for horizons longer than one year.

In turn, Table 3 shows that inflation is mainly driven by Mexican and non-U.S. shocks in all horizons. Further, U.S. shocks explain about half of the f.e. variance of the nominal exchange rate at long horizons. In turn, the f.e. variance of the policy rate remains mainly attached to Mexican and non-U.S. shocks for horizons shorter than 3 years.

Regarding the contribution of specific shocks, Fig. 5 and Table B2 in the appendix show that U.S. aggregate demand shocks have the highest contribution to the fluctuations of the output gap and the term premium in Mexico. The financial shock is the second one in importance for these variables. In contrast, uncertainty shocks to U.S. economic policy seem to have a moderate impact on the fluctuations of the Mexican output gap. The reason is that part of the shock effects seem to be absorbed by financial variables, such as the nominal exchange rate and the country-interest-rate spread. Table B2 shows that the economic policy uncertainty shock is the major U.S. shock contributor to the f.e. variance of the exchange rate, while it is one of the top three contributors to the f.e. variance of the country-interest-rate spread in any horizon.

5.2. Selected impulse-response functions

In this section, we present the impulse responses of the variables in the SVAR model to the shocks contained in the *demand* category, and to the two types of monetary policy shocks. We focus on the *demand* category given its large contribution to the f.e. variance of both U.S. and Mexican macro variables. In turn, the two types of monetary policy shocks are novel to the literature. We show that distinguishing between these two shocks matters for the understanding of the international transmission of U.S. monetary policy.[21]

We compute the IRFs to one-standard-deviation innovations, and we keep the same variable scale across Figs. 6,7,8,9 and 11 to ease the comparison between shocks. Finally, we cumulate the impulse responses of TFP growth, the percent change in the nominal exchange rate, and the U.S. and Mexican inflation rates in order to clearly distinguish changes in the levels of these variables. We denote cumulated variables as tfp^{us}, s^{mx}, P^{us}, and P^{mx}, respectively.

5.2.1. Shocks in the demand category

Shock to economic policy uncertainty. Fig. 6 shows the impulse responses of U.S. and Mexican variables to an EPU shock. Notice that we do not impose sign restrictions for this shock, so the impulse responses recover correlations present in the data.

[21] The IRFs of the *supply* category are shown in Appendix B.

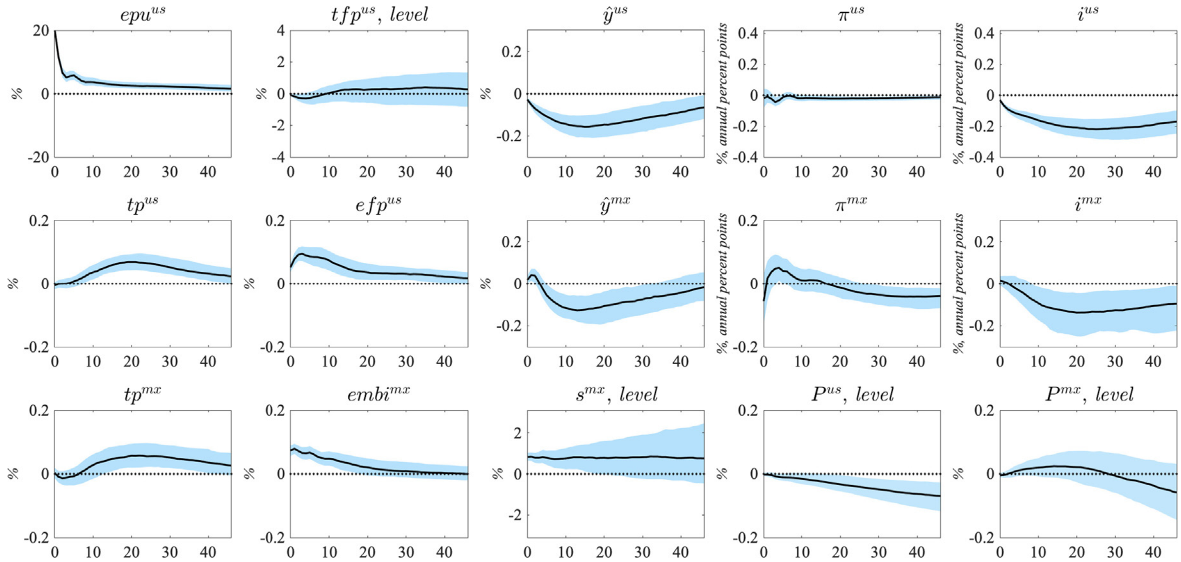

Fig. 6. Impulse responses to a sudden increase in U.S. economic policy uncertainty. **Note:** The *x*-axis denotes months after the shock, while the *y*-axis are percent points away from the long-run equilibrium value of a variable. The uncertainty band covers 68% of the distribution, and the continuous line is the median of the distribution. The IRF correspond to a 1 standard deviation innovation.

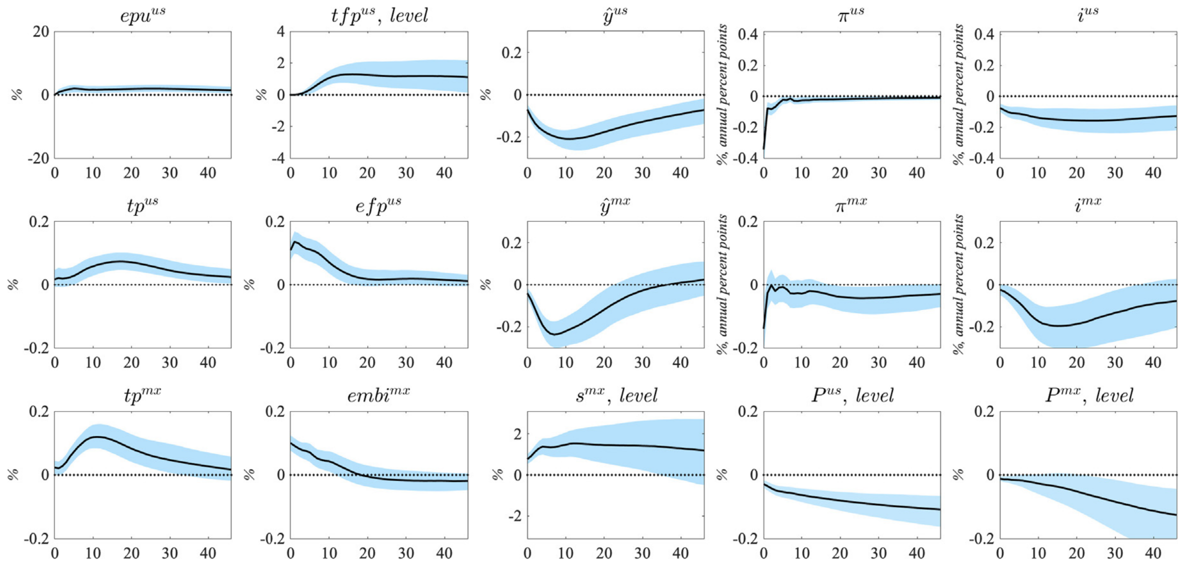

Fig. 7. Impulse responses to a negative shock in U.S. aggregate demand. **Note:** The *x*-axis denotes months after the shock, while the *y*-axis are percent points away from the long-run equilibrium value of a variable. The uncertainty band covers 68% of the distribution, and the continuous line is the median of the distribution. The IRF correspond to a 1 standard deviation innovation.

We start with the description of the responses of U.S. variables. Similar to Baker et al. (2016), we find that economic activity shrinks temporarily after the uncertainty in the formulation of U.S. economic policy increases. In addition, prices also seem to slow down, falling below trend. In turn, the policy rate falls protractedly for several months after the shock. In financial markets, the cost of external finance and the term premium rise. The latter may reflect an increase in the risk and the term premia demanded by credit suppliers and investors due to the contractionary economic outlook. Finally, productivity does not seem to be affected by this shock.

In Mexico, the output gap also decreases after an EPU shock, although at a lesser magnitude. The median peak effect of the output gap in Mexico equals -0.1%. In contrast, in the U.S. the median peak effect on the output gap is -0.14%. In both cases, this effect is reached one year after the shock. The data also suggest that, after an increase in economic policy uncertainty in the U.S., there is a co-movement between the policy rates and long-term interest rates in the two countries, while Mexico's country-interest-rate spread also co-moves with the U.S. external finance premium. This is again a signal that the Mexican

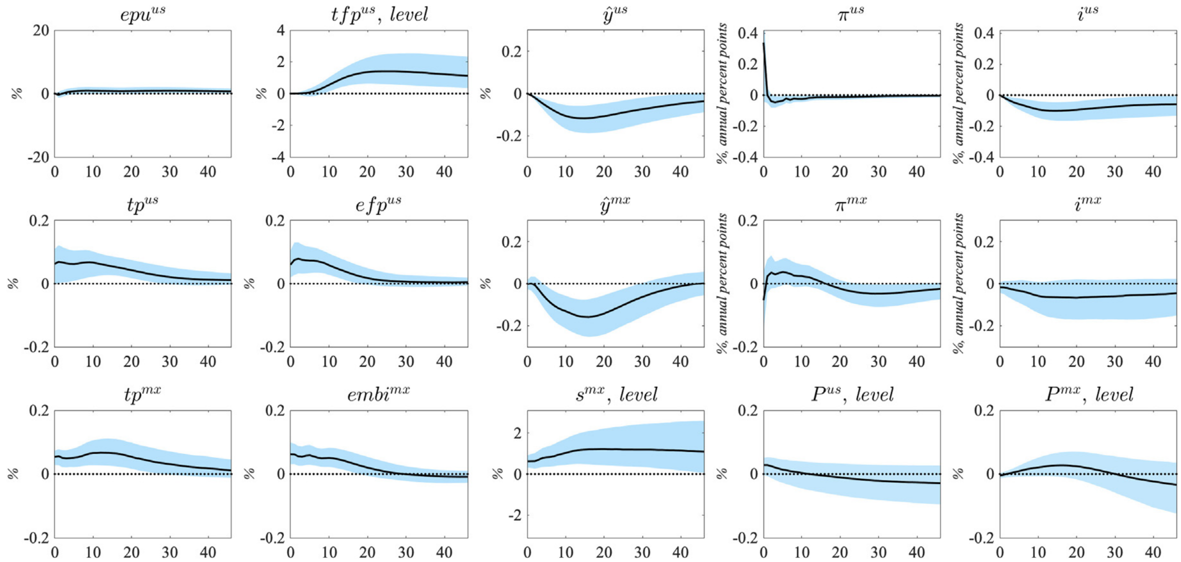

Fig. 8. Impulse responses to a sudden increase in the U.S. external finance premium. **Note:** The *x*-axis denotes months after the shock, while the *y*-axis are percent points away from the long-run equilibrium value of a variable. The uncertainty band covers 68% of the distribution, and the continuous line is the median of the distribution. The IRF correspond to a 1 standard deviation innovation.

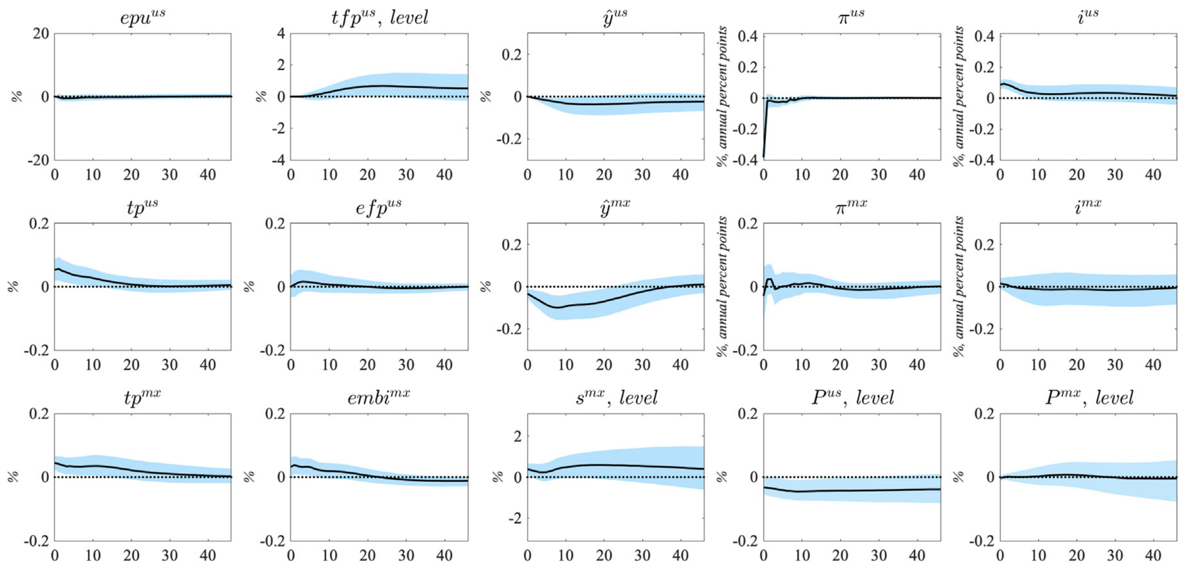

Fig. 9. Impulse responses to a contractionary type I monetary policy shock in the U.S. **Note:** The *x*-axis denotes months after the shock, while the *y*-axis are percent points away from the long-run equilibrium value of a variable. The uncertainty band covers 68% of the distribution, and the continuous line is the median of the distribution. The IRF correspond to a 1 standard deviation innovation.

financial market is tightly linked with that of the U.S. In addition, the shock seems to permanently depreciate the peso against the dollar by about 0.80% in terms of the median responses. Finally, prices in Mexico seem to increase somewhat for 12 months after the shock, probably due to the peso depreciation. For longer horizons, the Mexican price level decelerates, falling back to trend.

Aggregate demand shock. Fig. 7 shows the effects of a negative innovation in U.S. aggregate demand. In the U.S., output and prices fall for several months following the shock. In this regard, it is noteworthy that the sign restriction on inflation was imposed six months after the shock, rather than at the impact period, when we observe the peak effect of this variable. Following the deceleration of activity and prices, the U.S. policy rate falls below its long-term equilibrium level for several months. Interestingly, productivity presents a median increase of about 1% one year after the recessionary shock hits the economy. This result is consistent with the observation that labor productivity has grown in the U.S. during the last three recessions.[22] In turn, economic policy uncertainty registers a slight but persistent increase in the wake of the shock. In U.S. financial markets, following the economic deterioration, the cost of external finance surges on impact and returns to its mean level slowly, while the term premium becomes steeper. Again, this could reflect a rise in the risk and term premia demanded by credit suppliers and investors.

In Mexico, the output, the prices, the policy rate, the country interest rate, and the term premium co-move with their U.S. counterparts. The output and the policy rate are, however, less persistent than in the U.S., since they return to their mean in shorter periods. In turn, the peso persistently depreciates against the dollar, with a median peak effect of 1.2% twelve months after the shock. The depreciation is nonetheless transitory and eventually vanishes at a very slow pace. This might explain that prices do not experience a persistent upward pressure after the shock.

Financial shock. Fig. 8 displays the effects of a shock that persistently increases the cost of credit in the U.S. In this economy, aggregate variables respond similarly to how they do after a negative aggregate demand shock, but there are three important differences.

First, the external finance premium reverts to its mean approximately one year and a half after the financial shock, while it takes only a year after an aggregate demand shock. Second, the responses of output are remarkably more persistent after a financial shock than after an aggregate demand shock: the median peak effect is reached after fourteen months for the former, as opposed to eight months for the latter.[23] This result is consistent with the findings of Reinhart and Rogoff (2014), who document that financial recessions feature slower recoveries than non-financial recessions. And third, the median response of U. S. prices is positive on the impact period of a financial shock, while it is remarkably negative after an aggregate demand shock. However, there is an important amount of uncertainty surrounding the impact response of U.S. prices to a financial shock, as the wide confidence band indicates. This evidence is consistent with the results of Gilchrist et al. (2017), who find that liquidity-constrained firms tend to raise prices when financial conditions worsens, while unconstrained firms cut them down.[24] Therefore, the overall effect of a financial shock on prices is ambiguous.

In Mexico, the co-movement with U.S. variables is preserved for the output gap, the policy rate, the country interest rate, and the term premium, but not for prices. In turn, the peso permanently depreciates against the dollar, falling in value about 1.4% in median terms. This may explain the persistent upward pressure on prices following the shock. Further, the median peak effect of the output gap also arrives later after a U.S. financial shock than after a U.S. aggregate demand shock (i.e., fifteen months for the former and seven months for the latter).[25]

5.2.2. Monetary policy shocks

Type I shock. Fig. 9 presents the impulse responses to a *type I* contractionary monetary policy shock in the U.S. In this economy, the term premium steepens at the same time as the policy rate increases. The one-standard-deviation innovation translates into an unexpected rise in the Fed funds rate of roughly 10 bp in the impact period. The small size of the increase signals that monetary policy shocks are rare events. The mild change in the policy rate brings about moderate fluctuations in the rest of U.S. variables in comparison to shocks in the *demand* category. For instance, the median peak effect in the output gap is reached a year after the shock and it levels at -0.05%, i.e. three times smaller than the median peak effect of this variable after shocks in the *demand* category. The economic policy uncertainty index seems unaffected while productivity increases moderately. Finally, despite the negative sign restriction imposed on the six-month response of inflation, there is no clear effect on U.S. prices.

In Mexico, this shock has a particular amplification effect. Notably, economic activity falls twice as much as in the U.S., with a median peak effect of -0.1%, a level that is reached eight months after the shock. In turn, the peso persistently depreciates 0.75% against the dollar in median terms. Further, similar to the seminal work of Uribe and Yue (2006), and more recently Vicondoa (2019), we find that Mexico's country interest rate increases after the innovation in the U.S. policy rate. In addition, for a type I monetary policy shock, we find that the Mexican term premium follows its U.S. counterpart, so the yield of peso denominated long-term bonds also increases.[26] In contrast, the responses of the policy rate and prices in Mexico are ambiguous. On the one hand, the very persistent peso depreciation pushes prices upwards but, on the other hand, the slower economic activity pushes prices downwards. The monetary authority thus faces a dilemma, with a falling economic activity and mixed risks on prices. This scenario may explain why the policy rate in Mexico does not co-move with the U.S. policy rate after a

[22] Lazear et al. (2016) notice that from 2007 to 2009 output per capita in the U.S. non-farm business sector dropped by 7.2%, but the aggregate hours worked fell by 10% in the same sector. Since input fell more than output, productivity increased. These authors argue that their data favor the hypothesis that workers displayed more effort during the last recession, probably because they were afraid of losing their jobs.

[23] In turn, the size of the median effects is -0.16% for the financial shock and -0.18% for the aggregate demand shock.

[24] Gilchrist et al. (2017) build a model in which financial distortions motivate firms to increase prices in response to adverse financial shocks. In addition, they find supporting empirical evidence of this hypothesis using a database with information on U.S. firms' balance sheets and prices during the global financial crisis.

[25] Our results echo those found by Cesa-Bianchi and Sokol (2017), who study the transmission of shocks to U.S. monetary policy and corporate firms excess bond premia to the U.K. economy. One important difference between their paper and ours is the identification strategy. Cesa-Bianchi and Sokol (2017) use only sign restrictions to differentiate a financial shock from an aggregate-demand shock. In contrast, we differentiate these shocks by imposing a zero restriction in the output gap at the impact period of a financial shock (see Table 1).

[26] Notice that Uribe and Yue (2006) and Vicondoa (2019) do not include either policy rates or domestic currency denominated long-term bond yields into their analysis.

monetary policy shock. In this regard, Dedola et al. (2017) find that the mean responses of the policy rate and inflation rates of eighteen advanced economies and eighteen emerging economies are close to zero after a monetary policy shock in the U.S.[27] In the same way, Crespo Cuaresma et al. (2018) find that the mean responses of the policy rate and inflation are heterogenous between 12 advanced and emerging economies analyzed.[28]

The amplification effect observed in the Mexican output gap and the lack of co-movement between the two policy rates after this shock deserve further exploration. To analyze these features, we build two *counterfactual scenarios*, in which we increasingly disconnect the Mexican economy from the U.S. economy. These scenarios are:

- *C1*: *Mexican output does not respond to U.S. real and financial variables (i.e. productivity, output, long-term rates, and the external finance premium), and*
- *C2*: *C1* plus *Mexican output is not affected by Mexican financial variables (i.e. long-term interest rates and the country interest rate).*

For the first counterfactual scenario, we set the coefficients associated to $\Delta tfp, y^{us}, tp^{us}$ and efp^{us} equal to zero in the VAR equation for y^{mx}. In the second counterfactual scenario, we further impose in the same VAR equation zero values for the coefficients associated to tp^{mx} and $embi^{mx}$. The results are shown in Fig. 10, where we have normalized the IRFs to a surprise increase of 25 bp in the U.S. policy rate.

The figure presents the median responses of Mexican variables in 3 scenarios: the baseline, which corresponds to the median responses of Fig. 9, and the counterfactuals *C1* and *C2*. The difference between the baseline and *C1* uncovers the direct influence that U.S. variables have on goods demand in Mexico, either through the current account (e.g. exports), foreign direct investment, or other. We call this transmission mechanism the *foreign-demand channel*. As expected, the Mexican output gap moderates its fall when we turn this channel off in scenario *C1*. As a result, the Mexican inflation rate and the policy rate in this scenario are higher than their baseline levels, while the peso depreciation is also smaller (see dashed line in Fig. 9).

The difference between scenarios *C2* and *C1* unveils the crucial importance of the link between financial markets in Mexico and the U.S. We call this transmission mechanism the *financial-contagion channel*, and it operates through arbitrage in international capital markets. In scenario *C2*, we isolate Mexican real activity further by assuming that output does not respond to the Mexican term premium or the country interest rate, in addition to the conditions of scenario *C1*. The impact of this assumption is huge, since it yields stability in real activity. Therefore, in scenario *C2* the only risk for inflation comes from the peso depreciation and, as a result, the policy rate increases in response to the inflationary threat. Therefore, this counterfactual scenario features again a co-movement between the U.S. and Mexican policy rates in the medium run.

The financial-contagion channel is driven by investors' sentiment about the Mexican economy. When this channel is active, after a *type I* U.S. monetary policy shock, domestic and international investors demand a higher premia for holding Mexican long-term bonds, either denominated in pesos or dollars. Consequently, the cost of funding increases in the Mexican economy, which explains the slowdown of real activity.

Type II shock. Finally, Fig. 11 presents the responses to a *type II* contractionary monetary policy shock in the U.S. In this economy, the aggregate dynamics look similar to those after a type I shock, with the crucial difference that the U.S. term premium flattens in the period of the monetary policy surprise. The median peak effect of the U.S. policy rate after a one-standard-deviation innovation is of merely 7 bp, which is again a signal that this shock is a rare event.

Mexican aggregate dynamics display important differences with respect to a *type I* shock. First, output does not seem to fall as much (-0.07% at peak in median terms) and, more importantly, it reverts to zero faster. The quick recovery can be explained by the responses of the Mexican term premium to a *type II* shock. Notably, there is no evidence that investors ask a higher premium for holding peso denominated long-term bonds after a *type II* shock. In contrast, the country interest rate increases similarly regardless of the type of the monetary policy shock realized. Overall, the financial-contagion channel is relatively weaker after a *type II* shock than after a *type I* shock, and, therefore, loanable funds are less constrained in the former than in the latter. In addition, the peso does not depreciate as much after a *type II* shock, so upward risks to inflation are smaller. Finally, the Mexican policy rate seems to follow the path of inflation with a lag.

Summary of U.S. monetary policy shocks. There are four noteworthy results about the Mexican aggregate dynamics after a U.S. monetary policy shock: (1) the yield of dollar-denominated long-term bonds issued by the Mexican government (i.e. the country interest rate) rises after an unexpected increase in the Fed funds rate; (2) the response of peso-denominated, long-term bond yields depends on the reaction of the U.S. term premium to said shock; (3) domestic real activity may fall persistently if the domestic term premium steepens after the shock; and (4) the responses of domestic inflation and the policy rates are ambiguous. To the best of our knowledge, results (2) and (3) are novel to the literature, while results (1) and (4) confirm the previous findings.

[27] It is noteworthy that Dedola et al. (2017)'s results present significant heterogeneity on policy rate responses.

[28] In particular, in the case of Mexico the responses of policy rate and inflation are near zero.

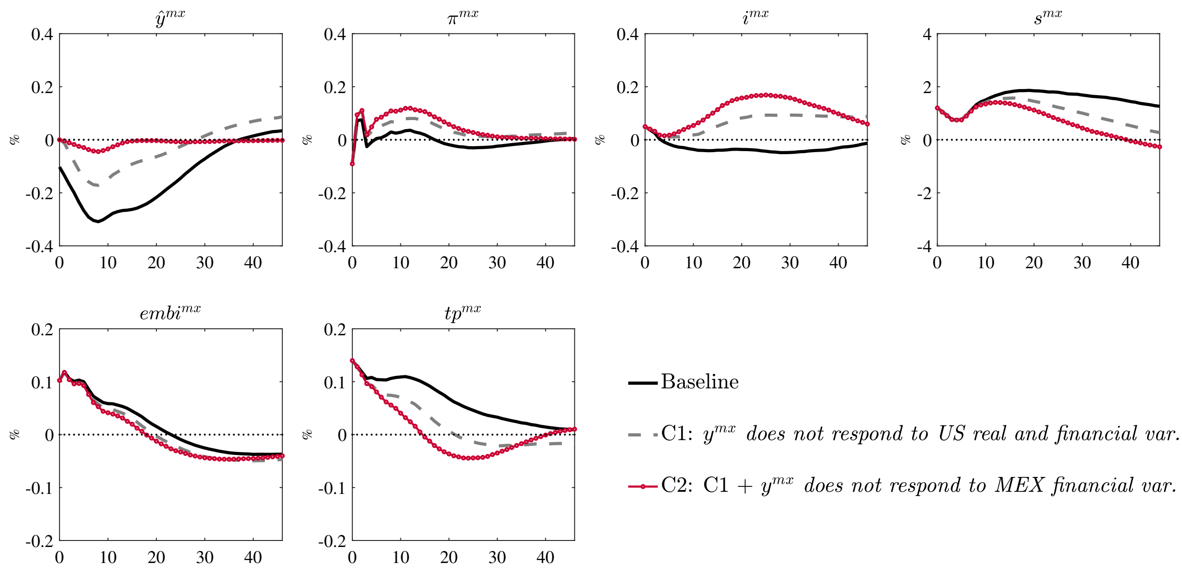

Fig. 10. Transmission channels of a U.S. type I monetary policy shock into Mexico. **Note:** The *x*-axis denotes months after the shock, while the *y*-axis are percent points away from the long-run equilibrium value of a variable. The IRFs correspond to the median of the distribution and are normalized to an increase in the Fed funds rate of **25 bp**.

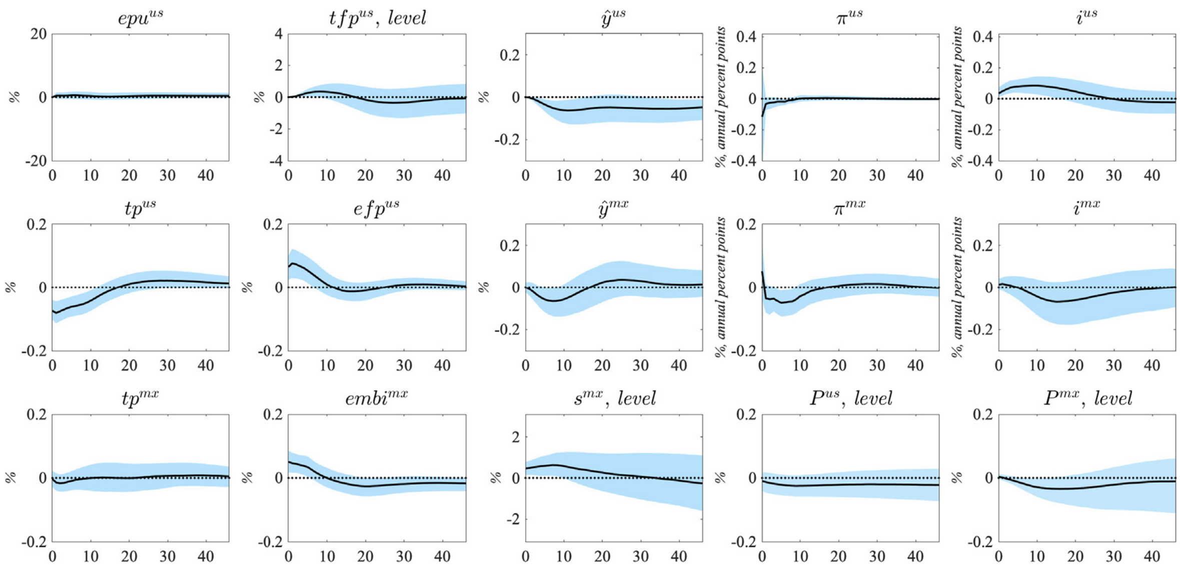

Fig. 11. Impulse responses to a U.S. type II monetary policy shock. **Note:** The *x*-axis denotes months after the shock, while the *y*-axis are percent points away from the long-run equilibrium value of a variable. The uncertainty band covers 68% of the distribution, and the continuous line is the median of the distribution. The IRF correspond to a 1 standard deviation innovation.

5.3. Co-movement analysis

In this subsection, we use the historical decomposition of the set-identified SVAR models to answer the two questions that we posed at the end of Section 2. We paraphrase them below:

1. *Which U.S. aggregate shocks account for the co-movement observed in the output gap, the term premium, and the cost of finance in the U.S. and Mexico?, and*
2. *Which of these shocks yield a co-movement in the policy rates of the U.S. and Mexico, and which ones reduce this co-movement?*

Fig. 12 decomposes the observed fluctuations in the output gap, the term premium, the external finance premium in the U.S., and the country-interest-rate spread in Mexico into the contributions attributed to, on the one hand, shocks generated

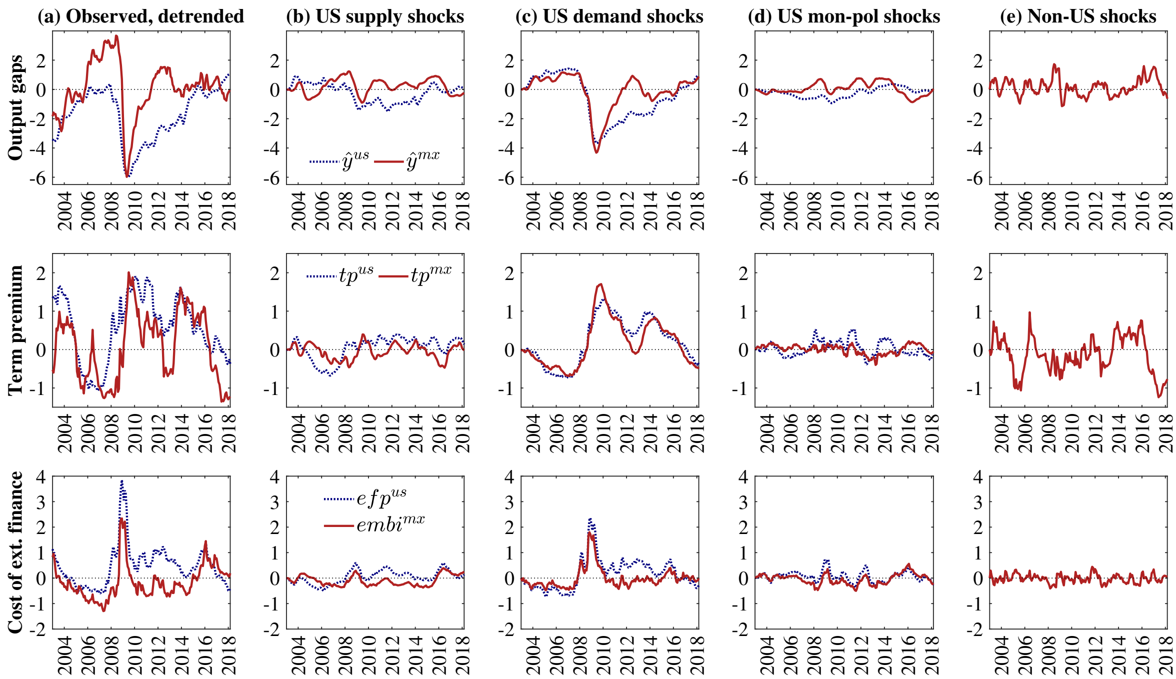

Fig. 12. Detrended variables and shock contributions. **Note:** The historical decomposition of a variable corresponds to the mean of the distribution of the posterior estimates. Column (b) keeps U.S. shocks contained in the supply category on, while it turns the rest of the shocks off. Column (c) keeps U.S. shocks contained in the demand category on, while it turns the rest of the shocks off, and so on. None of the columns from (b) to (e) include the initial conditions of each variable, i.e. the value of vector X_{-1} in the notation of Eq. (2). Notice that the effect of the initial condition of a variable inherits the persistence of this variable.

in the U.S. contained in the supply, demand, and monetary policy categories and, on the other hand, shocks not generated in the U.S. (i.e. Mexican and non-U.S. shocks). The figure illustrates the results already shown in the variance decomposition of Section 5.1 in a different way. That is, for the sample period studied, we have that shocks in the demand category explain the co-movement between real and financial variables in the U.S. and Mexico. In turn, shocks to U.S. supply and monetary policy play a relatively minor role in the determination of these variables. Finally, it is noteworthy that Mexican and non-U.S. shocks appear to be relatively more important for the determination of the Mexican term premium than for the country interest rate.

In Table 4, we have computed contrafactual correlations of the variables in Fig. 12 assuming that only a subset of shocks was observed. The first column in the table displays the correlation present in the data. In column (2), we assume that only U.S. demand shocks were observed while we turn off the rest of shocks. As a result, the correlation between the output gap in Mexico and in the U.S. increases with respect to the one observed in the data. The same applies for the term premium, and the costs of external finance in both countries. In column (3), we assume that Mexican and non-U.S. shocks are also observed in addition to U.S. demand shocks, which results in a reduction in the counterfactual correlations with respect to the previous case. Notice that if we further added the rest of U.S. shocks into the simulation along with the initial condition of each variable, i.e. the vector X_{-1} in the notation of Eq. (2), the model-based correlations and those observed in the data would match perfectly. Therefore, shocks that explain the co-movement between real and financial activities in the U.S. and Mexico relate to those contained in the U.S. demand category, while the rest of U.S. and non-U.S. shocks cannot explain the co-movement.

Table 4
Conditional correlations.

	(1): Data	(2): *Only* U.S. demand shocks	(3): Column (2) *plus* non-U.S. shocks
Output gaps	0.64	0.82	0.78
Term premia	0.71	0.94	0.78
Costs of external finance	0.74	0.84	0.75
Policy rates	0.73	0.93	0.72
Inflation rates	0.06	0.26	0.09

Note: The correlations are computed from the mean historical decomposition of each variable. Column (2) keeps U.S. shocks contained in the demand category on, while it turns the rest of shocks off. Column (3) adds the non-identified shocks to the previous case. Notice that if we include the rest of U.S. shocks and the initial condition of each variable, i.e. X_{-1} in the notation of Eq. (2), we obtain the correlation computed in the data.

Fig. 13 shows the historical decomposition of the policy rates and inflation rates in Mexico and the U.S. For the ease of interpretation, we have computed the annual inflation rates from the monthly rates that we used as input in the model. Remarkably, the evidence suggests that if only U.S. demand shocks were observed, the co-movement between the two policy rates would have been almost perfect (see Table 4, where the contrafactual correlation for the policy rates is 0.93 if only U.S. demand shocks were observed). However, domestic and non-U.S. factors tend to weaken the co-movement between these two variables. To illustrate this result, consider the period December 2015 to March 2018, in which the policy rate in Mexico increased the cumulated amount by 500 bp. According to the set-identified SVAR models, nearly 440 bp of this increase were triggered by shocks not generated in the U.S. (see column (e) in Fig. 13). Notably, during the same period, annual core inflation in Mexico registered an increase of 200 bp due to Mexican and non-U.S. shocks. Therefore, conditional to these shocks, the policy rate in Mexico rose on a basis of two-to-one with the inflationary risk. In sum, the takeaway from this exercise is that, during the cycle, the policy rate in Mexico is affected substantially by U.S. shocks contained in the demand category, but domestic factors are as important for the determination of this variable too.

Decomposition of the transmission channels. To conclude this section, we disentangle the transmission of the shocks contained in the demand category on Mexican real activity into the financial-contagion and the foreign-demand channels. To compute these channels, we follow a strategy similar to the one taken in Section 5.2.2. In particular, to unveil the financial-transmission channel we build a counterfactual scenario in which the output gap, the inflation, and the policy rate in Mexico do not respond to any U.S. variable, i.e. we set the coefficients associated to U.S. variables equal to zero in the VAR equations of $\hat{y}^{mx}$, π^{mx} and i^{mx}. In this scenario, any change in the Mexican output gap is due to the responses of Mexican financial variables (i.e. the exchange rate, the term premium, and the country interest rate) to the developments in the U.S. economy. To compute the foreign-demand channel, we assume the opposite. In particular, we build a counterfactual scenario in which the aforementioned Mexican financial variables do not respond to U.S. variables. Therefore, in this scenario any changes in the output gap are due to the direct responses of Mexican output, inflation, and policy rate to changes in U.S. variables.

Panel (a) in Fig. 14 shows that the sum of the foreign-demand and financial-contagion channels adds up to the total contribution of U.S. shocks in the demand category to the Mexican output gap. In turn, panel (b) depicts the two channels in isolation, and shows that the contribution of the channels is of similar importance, especially at the onset of the global financial crisis.

Finally, in panels (c) and (d) of Fig. 14, we show the specific contributions to the Mexican output gap of the U.S. shocks contained in the demand category. This decomposition shows that the three shocks in the category contributed to the fall of Mexican real activity in the period 2008–2010.

5.4. Robustness analysis

In this section we present a number of robustness checks. In particular, we specify different identification strategies for monetary policy shocks following Uhlig (2005) and Arias et al. (2019). In addition, we investigate the sensitivity to the usage of alternative measures of economic uncertainty in the U.S. (VIX), the data frequency, the identification strategy, alternative measures of economic activity, and the similarities between central bank information shocks and our monetary policy shocks. We show that our results are robust across alternative specifications.

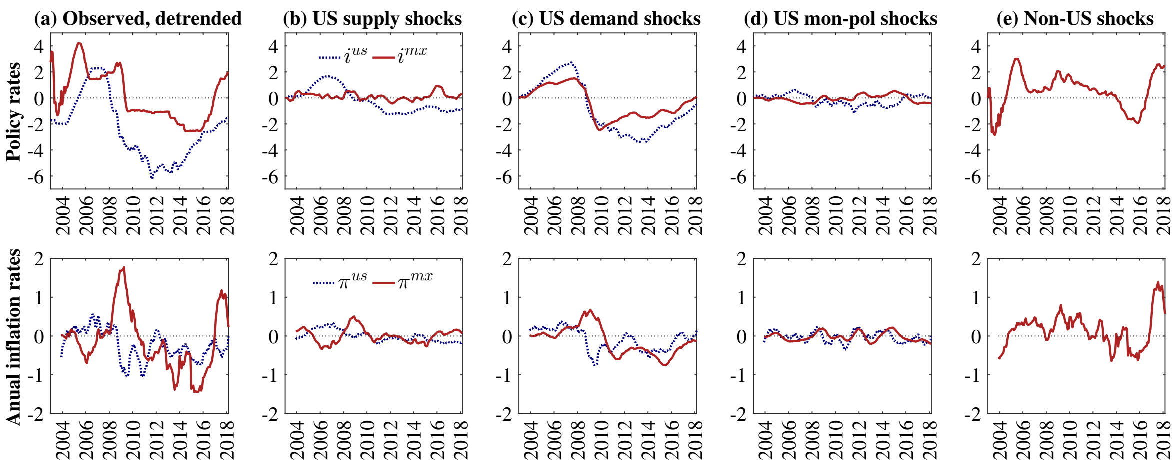

Fig. 13. Co-movement of monetary policy decisions and inflation rates. **Note:** The historical decomposition of a variable corresponds to the mean of the distribution of the posterior estimates. Column (b) keeps U.S. shocks contained in the supply category on, while it turns the rest of shocks off. Column (c) keeps U.S. shocks contained in the demand category on, while it turn the rest of shocks off, and so on. None of the columns from (b) to (e) include the initial conditions of each variable, i.e. the value of vector X_{-1} in the notation of Eq. (2). Notice that the effect of the initial condition of a variable inherits the persistence of this variable.

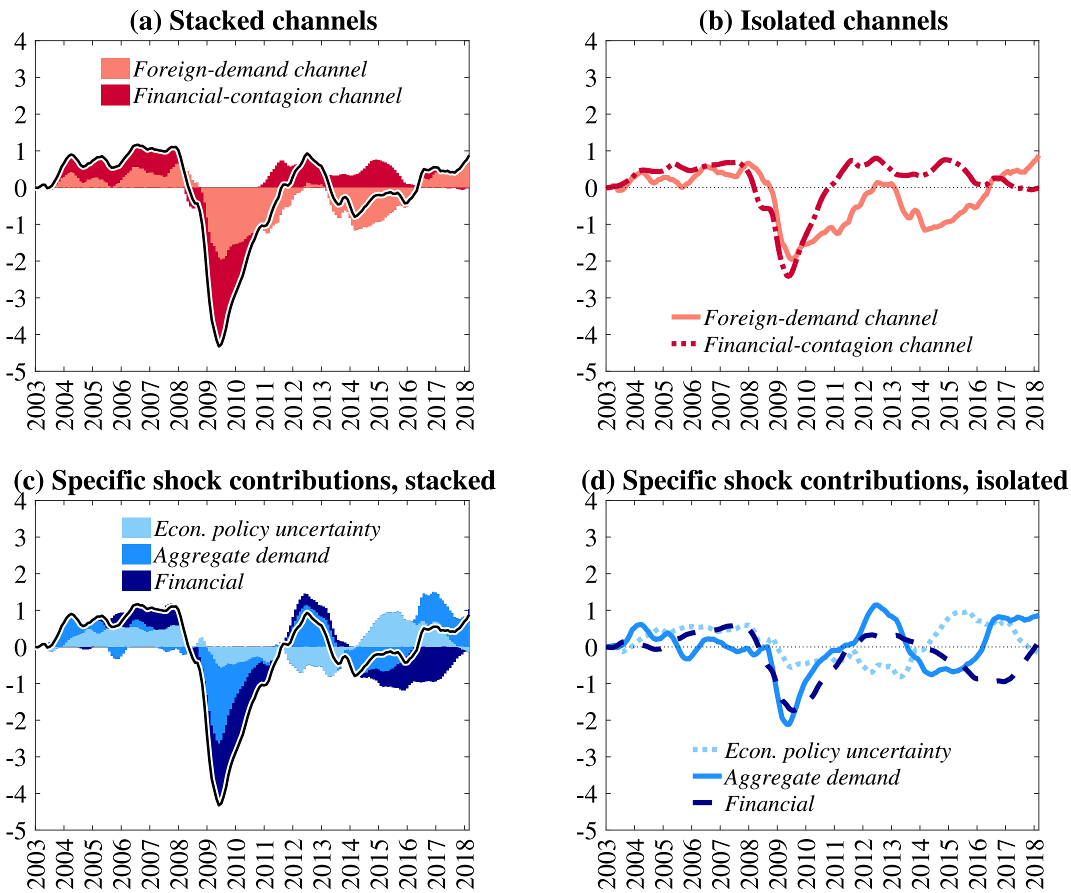

Fig. 14. Decomposition of U.S. demand shocks contributions to the Mexican output gap. **Note:** The historical decomposition of a variable corresponds to the mean of the distribution of the posterior estimates. The foreign-demand channel represents the responses of Mexican output to the U.S. shocks in the demand category assuming that Mexican financial variables do not respond *directly* to U.S. variables. In turn, the financial-contagion channel represents the responses of Mexican output to the U.S. shocks in the demand category assuming that Mexican output, inflation, and policy rate do not respond *directly* to U. S. variables.

5.4.1. Alternative identification strategy for monetary policy shocks

Our identification strategy for the two monetary policy shocks imposes signs on the responses of the policy rate, output, and inflation, which is a standard practice in the literature (see Arias et al., 2019, and the references included therein). However, this conventional strategy has been questioned by Uhlig (2005). He argues that the contractionary effects of monetary policy shocks are the result of imposing a zero restriction in the contemporaneous response of output. Thus, he follows an agnostic approach and imposes no restriction on the impact response of output, while imposing a negative response of prices, commodity prices, and nonborrowed reserves in his identification strategy both on impact and five months after the shock. As a result, Uhlig (2005) finds that a monetary policy shock has no clear effect on output in the U.S.

In order to handle this critique, we follow an identification strategy similar to the one in Uhlig (2005). We therefore impose no zero restriction on the impact response of output, while we maintain the required negative response of prices both on impact and during the six moths after our two contractionary monetary policy shocks. The restrictions in the remaining variables remain the same, as well as for the other five shocks. This exercise yields three results: (1) a *type I* monetary policy shock induces a co-movement between the U.S. and the domestic term premia, and a contraction in domestic activity; (2) a *type II* monetary policy shock has no clear effects on both domestic activity and the term premium; and (3) both monetary policy shocks have an ambiguous effect on U.S. output. Even when we do not impose a zero restriction on the impact response of U.S. output, our central results maintain qualitatively (Figs. E3 and E4 in Appendix E show the responses to a *type I* and *type II* monetary policy shocks).

Arias et al. (2019) identify a monetary policy shock using SVARs by imposing sign and zero restrictions on the contemporaneous structural parameters of their monetary equation and show that the identification strategy in Uhlig (2005) does not satisfy their restrictions. Moreover, they find that an increase in the federal funds rate leads to a contraction in U.S. output. Following a similar approach, in addition to the previous strategy identification following Uhlig (2005), we impose positive sign restrictions on the contemporaneous structural parameters associated to the U.S. output gap and inflation in our monetary equation. Interestingly, we qualitatively recover the central results from our baseline identification (Figs. E5 and E6 in Appendix E show the responses to a *type I* and *type II* monetary policy shocks).

5.4.2. Alternative measure of economic uncertainty

There are different measures to quantify economic uncertainty. In this section, we additionally employ the well-known CBOE Volatility Index (or VIX) commonly used in the literature to identify an uncertainty shock in the US and compare the IRFs with the ones obtained using the EPU index. In general, the responses of key variables implied by both indicators are qualitatively similar as is shown in Fig. E1 in Appendix E. Nonetheless, the responses implied by the VIX tend to be larger (in absolute terms) than those implied by the EPU index. This may be due to the fact that the VIX captures both economic uncertainty and stock market volatility. As pointed out by Caldara et al. (2016), using this indicator is likely to confound the respective effects of economic uncertainty and financial shocks. In our study the correlation between the VIX and the EPU index is 0.50, while the corresponding one between the VIX and the external finance premium is 0.80.

5.4.3. Quarterly frequency

Results from our baseline identification strategy are robust to using data at a quarterly frequency (results are available upon request). However, we impose only impact responses on IRFs to achieve identification. The justification is that one would expect to see our variables react to shocks within the quarter. Again, we find that U.S. shocks drive the business cycle co-movement between Mexico and the U.S. and the effects of monetary policy shocks on the domestic economy depend on the market sentiment.

5.4.4. Alternative identification strategy between financial and cost-push shock

From the restrictions in Table 1, it may appear that the financial shock and the cost-push shock are not fully identified, since the external finance premium may rise in both shocks in the impact period. Figs. 8 in the main text and in Appendix E show that this possibility does not materialize, since the external finance premium falls after a cost-push shock, while it rises after a financial shock. However, to enforce the identification, in this robustness exercise we add an extra zero restriction on the external finance premium at the impact period of the cost-push shock (see Table E1 in Appendix E). The impulse responses from this exercise, displayed in Figs. E7 and E8, show virtually no changes with respect to the baseline identification strategy.

5.4.5. Alternative measures of economic activity

In this exercise, we replace the output gap variables in the baseline estimation with two alternatives measures: the log of real GDP in levels, and an alternative output gap gathered from a model that computes the trend of real GDP for each country. For the latter, we use the fitted value from the time-varying parameter regression $\ln(GDP_t) = \alpha + \beta_t t + \varepsilon_t$. The estimated values of β_t correspond to the one-side, Kalman filter value of the Bayesian estimation of said regression (ran using data from 1960 to 2018 for the U.S., and from 1980 to 2018 for Mexico). The alternative trend and gap measures are shown in Fig. E9 in Appendix E. The impulse responses of this exercise are displayed in Figs. E10 and E11 in said Appendix. In general, the responses to U.S. spending shocks are consistent with those of the baseline estimation. For other shocks, however, there are some important differences. For instance, the responses of economic activity in the U.S. and Mexico to a cost push shock are very persistent, which is at odds with standard general-equilibrium macro models. In these models, a temporary rise in inflation delivers a short-lasted response in output. The baseline estimation delivers transitory responses of U.S. output, while the alternative economic measures do not.

5.4.6. Central Bank information shocks

Jarociński and Karadi (2018) use the reaction of the stock market in the 30 min window that follows a FOMC statement release to distinguish between two types of monetary policy shocks: a *pure* monetary policy tightening, and a central bank information shock. In the former, the stock market index and the Fed funds rate display a negative correlation, while in the latter both variables move in the same direction. According to the authors, their strategy helps to purge monetary policy surprises from positive information revealed by the central bank about the economy. Our identification strategy for *types I* and *II* monetary policy shocks exploits a similar feature, namely, the reaction of investors' sentiment to a monetary policy shock as reflected in the term premium of long-term government bonds.

In this exercise, we replace the U.S. term premium in the baseline estimation by the percent change in the *Standard & Poor's 500* stock market index. As such, after a *type I* (*II*) monetary policy tightening, the stock market falls (rises) in response to negative (positive) news to the economy. The impulse responses of this exercise are displayed in Fig. E12 in Appendix E. The results confirm our baseline findings: type I monetary policy tightenings, which are accompanied by a negative investors's sentiment, seem to have a larger effect on Mexican output than type II tightenings.

6. Conclusions

In this paper, we study the transmission mechanism of U.S. aggregate shocks to the Mexican economy. In a SVAR model for a small open economy, we used a combination of sign and zero restrictions to identify seven types of U.S. shocks: an economic policy uncertainty shock, a productivity shock, an aggregate demand shock, a cost-push shock, a financial shock, and two types of monetary policy shocks. We study the transmission of these shocks over the period 2002M1-2018M3.

Our approach provides results that are qualitatively in line with impulse responses from a vast class of DSGE models with nominal rigidities. The identified U.S. macro shocks explain close to 75% of the fluctuations in the Mexican output gap, led by shocks to U.S. aggregate demand and the cost of external finance, with around 20% each. Additionally, the response of Mexican real activity to unexpected changes in the U.S. monetary policy depends on the behavior of the financial markets in both economies. Thus, if the term premia in both economies co-moves with the Fed funds rate after a U.S. monetary policy shock, then Mexican real activity might be significantly affected, in a statistical sense. Nonethless, the contribution of this shock to output fluctuations in the U.S. and Mexico is relatively small. Finally, the co-movement observed in the policy interest rates of the U.S. and Mexico seems to be mainly explained by shocks affecting the spending side of the U.S. economy, i.e. by shocks affecting U.S. aggregate demand, the cost of external finance, and the degree of uncertainty in the formulation of the U.S. economic policy.

Finally, we would like to acknowledge some avenues for future research. Among them, it would be worth to differentiate the propagation mechanisms of different productivity disturbances, such as foreign neutral and investment-specific technology shocks, which may impact differently the domestic economy (see Altig et al., 2011). Also, we have simplified the identification of cost-push shocks, but it might also be interesting to unfold this identification to address diverse types of shocks generated in foreign labor markets, such as labor-supply shocks or wage bargaining shocks (Foroni et al., 2018). These topics are left for future research.

Appendix A. Supplementary material

Supplementary data associated with this article can be found, in the online version, at https://doi.org/10.1016/j.jimonfin.2020.102148.

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