Are Business Cycles in Emerging Market Economies Alike?*

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November 1, 2022

Abstract

This paper explores the predictions of real business cycle theory on the roles of total factor productivity (TFP) and financial frictions to explain business cycles in emerging market economies (EMEs). I obtain evidence about TFP, price of capital, risk premium, and collateral constraint shocks by estimating structural vector autoregressions (SVARs) on a Brazilian sample from 1999Q1 to 2018Q4 and a Mexican sample from 1997Q1 to 2018Q4. On each sample, two SVARs are estimated. One SVAR identifies shocks by imposing restrictions on their short-run impact. The other SVAR is grounded on restrictions that shocks have long-run effects on business cycles in EMEs. Estimates of the SVARs show the TFP shock is the main driver of business cycle movements in Brazil and Mexico. However, this evidence is produced by the SVAR under the long-run restrictions, which indicates the identification of shocks matters to the explanation of business cycles in EMEs. Next, the Brazilian and Mexican business cycles are markedly different, as the contributions of shocks to aggregate fluctuations vary across the two countries. Hence, findings of this paper suggest although not all business cycles are alike in EMEs, "the cycle is the trend" view on aggregate fluctuations in EMEs remains valid.

JEL Classification: E32; F44

Keywords: Business cycles; emerging market economies; structural vector autoregressions; short-run identification; long-run identification.

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*Acknowledgments: I am grateful and indebted to my advisor Jim Nason for his guidance and support. I extend my gratitude to the other members of my dissertation committee, Giuseppe Fiori, Daisoon Kim, and Xiaoyong Zheng, for their thoughtful comments and suggestions. I thank Givi Melkadze, Margaret Jacobson, and participants at the Business Fluctuations session at the 91st Southern Economics Association conference for helpful comments. My thanks also go to Tasha Bigelow for editorial assistance. All remaining errors are my own.

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1. Introduction

Business cycles in emerging market economies (EMEs) are distinguished from business cycles in developed economies. The key features are (a) larger volatility of consumption than output (see Aguiar and Gopinath 2007; Boz, Daude, and Durdu 2011; García-Cicco, Pancrazi and Uribe 2010), (b) countercyclical world real interest rates and trade balances (see Akinci 2013; Neumeyer and Perri 2005; Uribe and Yue 2006), and (c) infrequent periods of large contractions in output and investment accompanied by a spike in the trade balance. The latter are often referred as "sudden stop events" (see Calvo, Izquierdo, and Talvi 2006; Mendoza 2010).

There are two lines of research that seek to explain business cycles in EMEs. One line is "the cycle is the trend" started by Aguiar and Gopinath (2007). They show that permanent shocks to total factor productivity (TFP) dominate business cycles in EMEs. Their small open economy (SOE) real business cycle (RBC) model driven primarily by a permanent TFP shock is successful at matching the key features of business cycles in EMEs.

The other strand of studies states that EMEs face frictions in financial markets; this assigns financial shocks the responsibility for generating business cycle fluctuations. For example, Neumeyer and Perri (2005) and Uribe and Yue (2006) incorporate a working capital constraint as a financial friction into a SOE-RBC model.[1] They report that shocks to the

[1] The working capital constraint forces firms to borrow internationally to pay for labor in order to produce. As a result, the working capital constraint creates a transmission mechanism in that shocks to international financial markets affect output in a SOE.

international bond market account for about one-third of the business cycle movements in EMEs. Mendoza (2010) utilizes a SOE-RBC model, which includes a Kiyotaki and Moore (1997) type collateral constraint, to theorize that changes in price of capital and collateral requirement of foreign debt are origins of sudden stop events in Mexico. In spite of RBC theory's success on explaining business cycles in EMEs, it remains under debate whether the permanent TFP shock dominates, or whether financial shocks contribute largely to business cycle movements in EMEs.

This paper contributes to this debate by employing structural vector autoregressions (SVARs) estimated on the data from EMEs to explore the predictions of RBC theory on business cycles in EMEs. I estimate SVARs on a Brazilian sample from 1999Q1 to 2018Q4 and a Mexican sample from 1997Q1 to 2018Q4. Both samples consist of real GDP, a country-specific world real interest rate, Tobin's Q, and the trade balance-output ratio. The choice of these variables rests on two elements. First, dynamics of business cycles in EMEs are captured reasonably well by these variables, as documented in Mendoza (2010), Neumeyer and Perri (2005), and Uribe and Yue (2006). Second, they aid in identifying real and financial shocks. Variations in real GDP, the world real interest rate, Tobin's Q, and the trade balance-output ratio correspond to the TFP, risk premium, price of capital, and collateral constraint shocks, respectively.

I estimate two SVARs on each sample of Brazil and Mexico. The first SVAR is identified by recursive short-run restrictions. These restrictions rely on assuming whether shocks affect the variables at impact or one-quarter after. A Blanchard and Quah (1989) style recursive long-run restrictions are imposed to estimate the second SVAR. This identification strategy is motivated by the long-run properties of the data. Unit root tests cannot reject that real GDP, the world real interest rate, and Tobin's Q are integrated of order one but do reject the unit root null in the trade balance-output ratio for both samples. The presence of unit roots imply the TFP, risk premium, and price of capital shocks have permanent effects on the Brazilian and Mexican variables.

The recursive structure of short- and long-run restrictions follows the order of real GDP, the world real interest rate, Tobin's Q, and the trade balance-output ratio. I gather an assembly of restrictions from the EME RBC literature and neoclassical investment theory to guide the short- and long-run identifications. Ordering real GDP first is consistent with the SOE-RBC models in which a TFP shock hits output first. Studies on default and neoclassical investment theory offer short- and long-run restrictions for identifying the risk premium and price of capital shocks. Theories on financial frictions in EMEs provide insight on the identification of the collateral constraint shock.

This paper finds empirical evidence that supports the permanent TFP shock hypothesis of business cycles in EMEs. However, this evidence is sensitive to the identification of SVARs. The short-run restrictions render limited knowledge on business cycles in Brazil and Mexico. SVARs estimated with these restrictions result in real GDP, the world real interest rate, Tobin's Q, and the trade balance respond mostly to their own shocks. The long-run restrictions yield forecast error variance decompositions (FEVDs) that show the TFP shock is the primary driving force of business cycles in Brazil and Mexico. In this case, the TFP shock explains about three-fourths of the variation in Brazilian real GDP after the one-year horizon while also being responsible for about half of the fluctuations in the Brazilian trade balance from one- to twenty-quarter horizon. The TFP shock is the dominant shock in the SVAR identified on the long-run restrictions and estimated on the Mexican sample. This shock accounts for more than three-fourths of the movements of Mexican real GDP, about one-third of the variation in Tobin's Q, and nearly half of the fluctuations in the trade balance throughout the five-year horizon. These findings suggest the view of Aguiar and Gopinath (2007) on business cycles in EMEs is conditional on the long-run identification.

Next, FEVDs reveal distinct business cycles in Brazil and Mexico as shocks to TFP and financial markets have different roles in driving the aggregate fluctuations across the short- and long-run identifications. For example, the TFP shock pinned down by the long-run restrictions is negligible to the movement of Brazilian Tobin's Q, but causes large swings in the Mexican Tobin's Q throughout the five-year horizon. The price of capital shock identified on the short-run restrictions is insignificant to Brazilian real GDP, but generates about a quarter of the fluctuation in Mexican real GDP at the business cycle horizon.

The next section describes the data. Section 3 discusses the short- and long-run SVAR identifications. Estimates of the SVARs are presented in section 4 and section 5 concludes.

2. Data and Long-Run Properties

This section describes the data and discusses their long-run properties. I conduct augmented Dickey-Fuller tests to obtain evidence of whether unit roots exist in the Brazilian and Mexican variables.

2.1. Data

The Brazilian and Mexican samples include real GDP, a country-specific world real interest rate, Tobin's Q, and the trade balance-output ratio. The Brazilian sample runs from 1999Q1 to 2018Q4. The Mexican sample begins two years earlier in 1997Q1, and also ends with 2018Q4.[2]

Aggregate real GDP and trade balance data for Brazil and Mexico are drawn from International Financial Statistics (IFS) and Brazilian Institute of Geography and Statistics (IBGE) databases. These data are seasonally adjusted, quarterly, real, and measured in local currency units.

The Brazilian and Mexican world real interest rates are constructed as the ex-post U.S. real interest rate plus the emerging market bond index plus (EMBI+) spread.[3] I measure Tobin's Q with a ratio of the equity price index to the producer price index of Brazil and

[2]Longer Brazilian and Mexican samples are available. However, SVARs estimated on the longer samples produce badly behaved impulse response functions (IRFs) and FEVDs due to non-stationary endogenous variables. Discussions about the sample size are included in the appendix.

[3]Construction of the world real interest rate follows Neumeyer and Perri (2005) and Uribe and Yue (2006).

Mexico. The world real interest rate and Tobin's Q are quarterly, real, and obtained from the IFS, Global Economic Monitor (GEM), and Federal Reserve Economic Data (FRED) databases.[4]

Figures 1 and 2 display plots of the Brazilian and Mexican samples, respectively. The log levels of real GDP, the world real interest rate, and Tobin's Q appear non-stationary in both samples. It is difficult to tell whether the trade balance-output ratios are stationary. The first differences of log real GDP, the log world real interest rate, log Tobin's Q, and the trade balance-output ratio seem stationary in both samples. Nonetheless, visual inspections offer limited information on the degree of integration of the data. I gather evidence of unit roots by running augmented Dickey-Fuller tests on the Brazilian and Mexican samples.

2.2. Unit Root Tests

Dickey and Pantula (1987) point out there is a bias in the sequential testing of unit roots. In order to address the bias, they contend to start unit root tests at a high degree of differencing. In absence of the evidence of a unit root, unit root test should be conducted on one fewer difference of the data. This process continues until either the unit root null is rejected or the unit root test is applied to the series level of the data. As a result, I start augmented Dickey-Fuller tests on the first differences of the data.

The augmented Dickey-Fuller regression includes an intercept, a linear time trend, and

[4]See the appendix for details about the sources and construction of data.

lag differences of the variable. The length of lag differences is chosen according to Ng and Perron (1995). They begin the augmented Dickey-Fuller regression with eight lag differences. If the absolute value of the t-statistic for the eighth lag difference is greater than 1.6, the lag length is set at eight. If not, one lag is deleted and the process is repeated until the t-statistic for the last lag difference is greater than 1.6 or the lag length reaches zero.

Table 1 reports the results of augmented Dickey-Fuller tests. For the Brazilian sample, the tests reject the null of a unit root in the first differences of log real GDP, the log world real interest rate, log Tobin's Q, and the trade balance-output ratio at the 5% significance level. Next, I apply augmented Dickey-Fuller tests to log real GDP, the log world real interest rate, log Tobin's Q, and the trade balance-output ratio of the Brazilian sample. There is no evidence to reject the null of a unit root in these variables at the 5% significance level. However, the test rejects the null of a unit root in the Brazilian trade balance-output ratio at the 10% significance level.[5]

Similarly, for the Mexican sample, the augmented Dickey-Fuller tests are applied to the first difference levels. Next, I estimate the augmented Dickey-Fuller regression on the series levels of the variables. The test results show one unit root in log real GDP, the log world real interest rate, and log Tobin's Q but the trade balance-output ratio is stationary at the 5% significance level.

I conclude from the unit root tests that the first differences of log real GDP, the log world

[5]The small sample critical value for the Brazilian trade balance-output ratio equals -3.160 at the 10% significance level.

real interest rate, and log Tobin's Q plus the level of the trade balance-output ratio from the Brazilian and Mexican samples are stationary.

3. SVARs

This section presents reduced-form and structural VARs. Next, the short- and long-run identifications are discussed. Estimation of the SVARs is also reviewed.

3.1. Reduced-Form and Structural VARs

The covariance-stationary vector of endogenous variables is

$$\boldsymbol{y_t} = \left[\Delta lnY_t \quad \Delta lnR_t \quad \Delta lnQ_t \quad \frac{TB_t}{Y_t} \right]',$$

where ΔlnY_t is real GDP growth, ΔlnR_t is the first difference of the log world real interest rate, ΔlnQ_t is the first difference of log Tobin's Q, and $\frac{TB_t}{Y_t}$ is the trade balance-output ratio. Consider a reduced-form VAR to be estimated on $\boldsymbol{y_t}$,

$$\boldsymbol{y_t} = v + C(L)\boldsymbol{y_{t-1}} + \boldsymbol{u_t},\ \boldsymbol{u_t} \sim N(0, \Sigma_u), \tag{1}$$

where v is a 4×1 vector of intercepts, L is the lag operator defined such that $L\boldsymbol{y_t} = \boldsymbol{y_{t-1}}$, $C(L) = \sum_{i=1}^{p} C_i L^{i-1}$ are the 4×4 slope coefficient matrices associated with lags of the reduced-form VAR, p is the lag length, and Σ_u is the covariance matrix of the reduced-form errors, $\boldsymbol{u_t}$.

Build on the reduced-form VAR in equation (1) to construct a SVAR as

$$\boldsymbol{y_t} = v + C(L)\boldsymbol{y_{t-1}} + D\boldsymbol{\varepsilon_t},\ \boldsymbol{\varepsilon_t} \sim N(0, I), \tag{2}$$

by mapping the reduced-form errors into the structural shocks with

$$\boldsymbol{u_t} = D\boldsymbol{\varepsilon_t},$$

where D is the 4×4 nonsingular structural matrix, and $\boldsymbol{\varepsilon_t} = [\varepsilon_{y,t}, \varepsilon_{r,t}, \varepsilon_{p,t}, \varepsilon_{d,t}]'$ is the vector of orthogonalized unit variance structural shocks.

The short- and long-run identifications of structural shocks rest on imposing restrictions on the structural matrix, D, and the long-run structural matrix, $\Gamma(1)$, respectively. The long-run structural matrix, $\Gamma(1)$, is also mapped from the structural matrix as $\Gamma(1) = [I_{4\times4} - C(1)]^{-1}D$.

3.2. Short-Run Identification

Recursive short-run zero restrictions are imposed on the short-run structural matrix, D, to identify TFP, risk premium, price of capital, and collateral constraint shocks by the ordering of $\boldsymbol{y_t}$ shown above. This ordering is consistent with RBC theory as it is applied to EMEs.

Real GDP growth is ordered first in $\boldsymbol{y_t}$ because SOE-RBC models often assume a TFP shock hits output first.

Assumptions of placing the first difference of the log world real interest rate second in $\boldsymbol{y_t}$ are based on the literature on default in EMEs. The literature posits that the risk premium component of the world real interest rate hinges upon the probability of default.[6] An EME is likely to default when it has a large net foreign debt position and output is low.[7] The TFP shock, which as already noted drives output, influences the probability of default. As a result, the world real interest rate responds to the TFP shock at impact. The literature is silent on the role of Tobin's Q in EME defaults. I assume the price of capital shock does not affect the world real interest rate at impact because there is no compelling evidence to suggest otherwise. The collateral constraint shock alters the amount of collateral required to borrow foreign debt for an EME.[8] Collateral is essentially comprised of capital and therefore the collateral constraint shock shares the same assumption on the price of capital shock that it has no contemporaneous effect on the world real interest rate.

The neoclassical investment theory guides the placement of the first difference of log Tobin's Q in $\boldsymbol{y_t}$. Hayashi (1982) theorizes that the market value of capital, which is the numerator of Tobin's Q formula, is the present discounted value of future profits generated by an additional unit of investment. Computing that value involves the stochastic discount

[6]For reference, see Akinci (2013), Neumeyer and Perri (2005), and Uribe and Yue (2006).

[7]See, among others, Aguiar and Gopinath (2006), Arellano (2008), and Eaton and Gersovitz (1981).

[8]Mendoza (2010) and Schmitt-Grohé and Uribe (2016) are examples of studying the Kiyotaki and Moore collateral constraint in a SOE-RBC model.

rate. An EME actively borrows at the international bond market to finance its domestic investment. Moreover, it faces the world real interest rate on foreign debt. Thus, it is reasonable to assume that the world real interest rate influences the stochastic discount rate in an EME. This assumption gives a channel for the risk premium shock to affect Tobin's Q at impact. Since the value of an extra unit of investment responds to a TFP shock, it also moves Tobin's Q at impact. The collateral constraint shock does not influence Tobin's Q contemporaneously for the same reason that disturbances to the amount of collateral take one quarter to affect an EME.

The trade balance-output ratio is ordered last in $\boldsymbol{y_t}$. The TFP shock affects this ratio at impact, as real GDP is the denominator of that ratio. Disturbances to the collateral requirement and the world real interest rate elicit changes to the size of foreign loans (import) and interest payments to foreign investors (export), respectively. Therefore the collateral constraint and risk premium shocks alter the trade balance-output ratio contemporaneously. The trade balance-output ratio also responds to the price of capital shock at impact. The reason is EMEs often borrow internationally to fund investment. These investment decisions are reflected in Tobin's Q which connects to the trade balance.

3.3. Long-Run Identification

In terms of the long-run identification, I apply the Blanchard and Quah (1989) decomposition to restrict the long-run structural matrix, $\Gamma(1)$, as a lower triangular matrix. The long-run identification uses the same ordering of $\boldsymbol{y_t}$ discussed in the previous section. Note that the long-run identification restricts the TFP, risk premium, and price of capital shocks to have permanent effect. Discussion on the long-run identification centers around the six long-run zero restrictions.

The long-run neutrality of real GDP yields three long-run zero restrictions because only the TFP shock has a permanent effect on real GDP.

The previous assumption that the world real interest rate affects the stochastic discount rate in an EME is key to the long-run identification. According to this assumption, it is the world real interest rate moves Tobin's Q, not the converse. This gives one long-run zero restriction, meaning that the price of capital shock does not drive the world real interest rate in the long run.

The remaining two long-run zero restrictions are based on the transitory nature of the collateral constraint shock. The stationarity of trade balance-output ratio from the Brazilian and Mexican samples imply the collateral constraint shock has transitory effect on output and foreign debt. Since output and foreign debt are factors of the probability of default and the collateral constraint shock only affect them during the short run, I assume that this shock does not drive the world real interest rate in the long run. Moreover, the collateral constraint shock has no influence over the stochastic discount rate that there is zero channel for it to generate any impact on Tobin's Q during the long run.

3.4. Estimation of SVARs

Estimation of the SVAR under the short-run restrictions relies on ordinary least squares (OLS) and specification of the structural matrix, D. The slope coefficient matrices, $C(1)$, and the covariance matrix, Σ_u, are estimated using OLS. Since the short-run restrictions are recursive, D is computed by taking the Cholesky decomposition of the covariance matrix, $D = \Sigma_u^{0.5}$.

The long-run identification slightly complicates the SVAR estimation. I construct the structural infinite-order vector moving average (VMA) representation of the SVAR from equation (2) as

$$\boldsymbol{y_t} = \mu + \sum_{j=0}^{\infty} \Gamma_j \boldsymbol{\varepsilon_{t-j}}, \tag{3}$$

where $\mu = [I_{4\times4} - C(1)]^{-1}v$ is a 4×1 vector of unconditional mean of $\boldsymbol{y_t}$, and Γ_j are the 4×4 structural moving average coefficient matrices that follow

$$\Gamma_0 = D, \quad \boldsymbol{u_t} = \Gamma_0 \boldsymbol{\varepsilon_t}, \quad \Gamma(L) = \sum_{j=0}^{\infty} \Gamma_j L^j, \quad \Gamma(L)\boldsymbol{\varepsilon_t} = [I_{4\times4} - C(L)]^{-1}\boldsymbol{u_t}.$$

The goal is to compute the structural matrix, Γ_0, given the lower triangular long-run structural matrix, $\tilde{\Gamma}(1)$ after the Blanchard and Quah (1989) decomposition. This process starts with multiplying $\Gamma(1)\boldsymbol{\varepsilon_t}$ by $[\Gamma(1)\boldsymbol{\varepsilon_t}]'$ to obtain

$$\Gamma(1)\boldsymbol{\varepsilon_t}\boldsymbol{\varepsilon_t}'\Gamma(1)' = [I_{4\times4} - C(1)]^{-1}\boldsymbol{u_t}\boldsymbol{u_t}'\Big([I_{4\times4} - C(1)]^{-1}\Big)'.$$

Since $\boldsymbol{\varepsilon_t}\boldsymbol{\varepsilon_t}' = I_{4\times4}$ and $\boldsymbol{u_t}\boldsymbol{u_t}' = \Sigma_u$, rewrite the above equation as

$$\Gamma(1)\Gamma(1)' = [I_{4\times4} - C(1)]^{-1}\Sigma_u\Big([I_{4\times4} - C(1)]^{-1}\Big)'.$$

The long-run structural matrix, $\tilde{\Gamma}(1)$, is found by taking the Cholesky decomposition of $\Gamma(1)\Gamma(1)'$, $\tilde{\Gamma}(1) = \Big[\Gamma(1)\Gamma(1)'\Big]^{0.5}$. Because of $\Sigma_u = \Gamma_0\Gamma_0'$, the Cholesky decomposition of $\Gamma(1)\Gamma(1)'$ also yields

$$\Big[\Gamma(1)\Gamma(1)'\Big]^{0.5} = \Big\{[I_{4\times4} - C(1)]^{-1}\Gamma_0\Gamma_0'\Big([I_{4\times4} - C(1)]^{-1}\Big)'\Big\}^{0.5} \Rightarrow \tilde{\Gamma}(1) = [I_{4\times4} - C(1)]^{-1}\Gamma_0.$$

Finally, the structural matrix, Γ_0, is recovered by computing $\Gamma_0 = [I_{4\times4} - C(1)]\tilde{\Gamma}(1)$.

The impulse response functions (IRFs) and FEVDs are computed similarly under the short- and long-run restrictions. Estimates of the short-run structural matrix, $\hat{D}$, and the reduced-form slope coefficient matrices, $\hat{C}(1)$, pin down the IRFs and FEVDs identified on the short-run restrictions. Under the long-run restrictions, estimates of the structural matrix, $\hat{\Gamma}_0$, and $\hat{C}(1)$ are enough to calculate the IRFs and FEVDs.

4. Results

This section discusses the IRFs and FEVDs computed by the SVARs. The SVARs are estimated on the first differences of log real GDP, the log world real interest rate, and log Tobin's Q plus the trade balance-output ratio from 1999Q1 to 2018Q4 for Brazil and 1997Q1 to 2018Q4 for Mexico. The lag length of the SVARs equals two, $p = 2$, for both samples.[9] The IRFs display dynamic responses of real GDP, the world real interest rate, Tobin's Q, and the trade balance to TFP, risk premium, price of capital, and collateral constraint shocks.[10] The FEVDs address the importance of shocks for explaining variations in these four variables. Confidence bands of the IRFs and FEVDs are calculated using the bootstrap sup-t confidence bands proposed by Olea and Plagborg-Møller (2019).

4.1. IRFs

Figure 3 depicts the IRFs of real GDP estimated on the Brazilian sample from impact to a twelve-quarter horizon. Real GDP increases permanently in response to a positive TFP shock under the short- and long-run restrictions. The IRFs of real GDP with respect to the risk premium, price of capital, and collateral constraint shocks identified on both restrictions display considerable uncertainty because the confidence bands are wide and cover the zero axis at most horizons.

Figure 4 displays a selection of IRFs of the world real interest rate, Tobin's Q, and the trade balance estimated on the Brazilian sample. Seven out of twelve IRFs in figure 4 have broad confidence bands. The remaining five IRFs exhibit narrower confidence bands but are identified on own shocks. As a result, other than the importance of own shocks, there is little

[9] I follow the common practice in the EME empirical literature of estimating SVARs with two lags. See, for example, Akinci (2013), Bruno and Shin (2015), and Kim, Lim, and Sohn (2020).

[10] The appendix details the transformation from real GDP growth, the first difference of the log world real interest rate, the first difference of log Tobin's Q, and the trade balance-output ratio to real GDP, the world interest rate, Tobin's Q, and the trade balance.

information to assess RBC theory on the Brazilian business cycle from the IRFs estimated under the short- or long-run restrictions.

The Mexican sample produces several more economically interesting IRFs compared with the Brazilian sample. Figure 5 shows that the price of capital and collateral constraint shocks identified on the short-run restrictions produce positive hump-shaped responses in Mexican real GDP. The long-run restrictions, by contrast, generate negative IRFs of real GDP with respect to the price of capital and collateral constraint shocks with confidence bands mostly covering the zero axis from impact to twelve-quarter horizon.

Figure 6 presents the selected IRFs of the world real interest rate, Tobin's Q, and the trade balance estimated on the Mexican sample.[11] The IRF of Tobin's Q with respect to a TFP shock under the long-run restrictions matches a prediction from Schmitt-Grohé and Uribe (2016). Their SOE-RBC model, which is built with an occasionally binding Kiyotaki and Moore (KM) collateral constraint, predicts that Tobin's Q rises in response to a positive TFP shock in an EME.[12] Other notable findings from figure 6 are the IRFs of the world real interest rate and trade balance. In response to a TFP shock, the world real interest rate and trade balance increase under both identifications. This shows the Mexican world real interest rate and trade balance are procyclical, which questions the stylized fact from the literature that they are countercyclical in EMEs.

[11] The remaining IRFs are available in the appendix.

[12] In their SOE-RBC model, a positive TFP shock loosens the KM collateral constraint that enables a small open economy to borrow more foreign debt. As a result, the demand for capital increases its price.

In sum, the IRFs offer only incomplete evidence to evaluate competing business cycle theories for Brazil and Mexico. The evidence for both economies is limited by substantial uncertainties surrounding many of the IRFs. Moreover, the IRFs of the world real interest rate and trade balance with respect to a TFP shock are inconsistent evidence to assess which RBC theory might best describe the Mexican economy.

4.2. FEVDs

The FEVDs present more economically interesting results than the IRFs for learning about business cycles in Brazil and Mexico.[13] Table 2 contains the FEVDs estimated on the Brazilian sample from one-quarter to a five-year horizon under the short-run restrictions. These FEVDs show that fluctuations in real GDP, the world real interest rate, Tobin's Q, and the trade balance are attributed most often to their own shocks throughout the five-year horizon. The exceptions are that the collateral constraint shock explains between 1% and 74%, with a median estimate of about a quarter of the variation in real GDP, and also explains between 2% and 75%, with a median estimate of about one-third of the variation in Tobin's Q at the four- to five-year horizon. There are substantial uncertainties in these FEVDs, as the width of the confidence bands surrounding the point estimates of FEVDs often exceed 40%.

The FEVDs indicate the TFP shock identified using the long-run restrictions is the main

[13] I interpret point estimates of FEVDs greater than 20% as economically important and confidence bands surrounding the median estimate of a FEVD greater than 40% interval as large uncertainty.

source of business cycle fluctuations in Brazil. Table 3 displays these FEVDs. The TFP shock explains more than two-third of the variation in real GDP at the one- to five-year horizon. In addition, from two-third to half of the movement in the trade balance are generated by the TFP shock from one-quarter to the five-year horizon.

The long-run identification also reveals a rich business cycle dynamic in Brazil. The price of capital shock is important to the Brazilian business cycle. This shock drives one-third to a quarter of the fluctuation in real GDP at the one- to four-quarter horizon and about a quarter of the variation in the trade balance throughout the five-year horizon. These FEVDs indicate the presence of KM collateral constraint, which is a financial friction, in the Brazilian economy. The width of the confidence bands covering the median estimates of TFP and price of capital shocks for explaining the variation in real GDP and the trade balance are greater than 40% across the business cycle horizons.

The FEVDs estimated on the Mexican sample are markedly different. As reported in table 4, movements in the world real interest rate, Tobin's Q, and the trade balance often depend on own shocks under the short-run restrictions for Mexico. The only anomaly is almost half of the variation in Mexican real GDP is explained by the price of capital and collateral constraint shocks at the one- to five-year horizon. The associated confidence bands are relatively narrow from one-quarter to the five-year horizon. The Mexican price of capital and collateral constraint shocks matter for explaining the variation in real GDP under the short-run restrictions. Nevertheless, these are weak evidence to suggest the short-run identification offers sufficient information to understand better the business cycle in Mexico.

Table 5 presents the FEVDs estimated on the Mexican sample under the long-run restrictions. These FEVDs show the TFP shock dominates the Mexican business cycle. This shock is responsible for producing about three quarters to nearly all of the variation in real GDP, anywhere between one-third to a quarter of the movement in Tobin's Q, and more than one-third of the fluctuation in the trade balance throughout the five-year horizon. The tight confidence bands surrounding the median estimates of the FEVD of real GDP with respect to the TFP shock is evidence it makes economically important contributions to the aggregate fluctuations in Mexico.

In summary, FEVDs show the evidence on sources of business cycle movements in Brazil and Mexico is dependent on the identification. The FEVDs under the short-run restrictions yield business cycles in both countries in which the TFP shock has a limited role for generating fluctuations in asset prices and returns and the converse is true for the risk premium, price of capital, and collateral constraint shocks. By contrast, the long-run identification produces economically interesting FEVDs to learn about business cycles in Brazil and Mexico. These two business cycles have distinct features as the contributions of TFP and financial shocks to business cycle fluctuations differ in both economies across the identifications. Moreover, TFP shock remain dominant to aggregate fluctuations in both countries conditional on the long-run identification. This finding supports the permanent TFP shock view on business cycles in EMEs as in Aguiar and Gopinath (2007) .

5. Conclusion

This paper presents estimates of SVARs using the Brazilian and Mexican samples of real GDP, the world real interest rate, Tobin's Q, and the trade balance-output ratio in recent twenty years. Short- and long-run restrictions are applied on SVARs to identify shocks to TFP, risk premium, price of capital, and collateral constraint. I report IRFs and FEVDs computed by the SVARs to learn whether the Brazilian and Mexican business cycles are best explained by the permanent TFP shock, or whether financial shocks are the main driver of aggregate fluctuations in both countries.

The IRFs computed under the short- and long-run restrictions produce few economically interesting results about the business cycle dynamics in Brazil and Mexico. This is because substantial uncertainties pervade nearly half of the Brazilian and Mexican IRFs. For the few IRFs that display statistical significant movements, they often conflict with RBC theory as it relates to EMEs.

The FEVDs offer useful information about the sources of business cycle movements in Brazil and Mexico. I draw two conclusions from the FEVDs. First, the Brazilian and Mexican FEVDs show that TFP shock dominates business cycles in EMEs. This result is in line with the permanent TFP shock hypothesis of business cycles in EMEs. However, there is a caveat that SVAR estimates are sensitive to the identification of the structural shocks. The FEVDs conditional on the short-run restrictions do not yield sufficient information to decide whether business cycles in Brazil and Mexico are best explained by permanent TFP or financial shocks. The long-run identification serves to produce FEVDs that display rich dynamics of Brazilian and Mexican business cycles that are primarily driven by the TFP shock.

Second, the Brazilian and Mexican business cycles have more differences than similarities. For example, the TFP shock is important for generating fluctuations in the Mexican Tobin's Q, but that is not the case for Brazil under the long-run restrictions. The price of capital shock identified by the short-run restrictions contributes more to the variation in Mexican than Brazilian real GDP. But when the price of capital shock is pinned down by the long-run restrictions, it accounts for about a quarter of the variations in real GDP and trade balance in Brazil, but is negligible to the business cycle in Mexico.

This paper should server as a reminder that real factor shocks remain prominent to EMEs and there is rich heterogeneity in business cycles across EMEs. Future research should aim to build dynamic stochastic general equilibrium models with structures that can reproduce this heterogeneity. Furthermore, these models should focus to study the connection between the permanent TFP shock and incomplete financial market conditions.

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Table 1: Augmented Dickey-Fuller Tests for Unit Roots in the Brazilian and Mexican Samples

	Brazilian Sample	1999Q1 to 2018Q4	**Mexican Sample**	1997Q1 to 2018Q4
Variable	Length of Lag Differences	t-statistic	Length of Lag Differences	t-statistic
ΔlnY_t	1	-5.509 (-3.468)	1	-6.546 (-3.462)
ΔlnR_t	1	-7.841 (-3.468)	1	-7.441 (-3.462)
ΔlnQ_t	4	-4.671 (-3.468)	6	-4.876 (-3.462)
$\Delta\frac{TB_t}{Y_t}$	2	-4.975 (-3.468)	1	-7.699 (-3.462)
lnY_t	1	-0.487 (-3.467)	2	-2.939 (-3.461)
lnR_t	3	-0.999 (-3.468)	3	-1.026 (-3.462)
lnQ_t	5	-1.721 (-3.468)	7	-0.382 (-3.462)
$\frac{TB_t}{Y_t}$	8	-3.263 (-3.467)	6	-4.087 (-3.461)

Notes: Real GDP, the world real interest rate, Tobin's Q, and the trade balance-output ratio are denoted as Y_t, R_t, Q_t, and $\frac{TB_t}{Y_t}$, respectively. The numbers in parentheses are small sample critical values at the 5% significance level. Except for the trade balance-output ratio, augmented Dickey-Fuller tests are run on log levels and first differences of log levels of the data. For the trade balance-output ratio, the augmented Dickey-Fuller tests are run on the series level and the first difference of it. The first difference of the trade balance-output ratio is computed as $100*(\frac{TB_t}{Y_t}-\frac{TB_{t-1}}{Y_{t-1}})/\frac{TB_{t-1}}{Y_{t-1}}$.

Table 2: FEVDs Estimated on the Brazilian Sample from 1999Q1 to 2018Q4 Under Short-Run Restrictions

Variable	Shocks \ Horizons	Q1	Q2	Q4	Q8	Q16	Q20
Y_t	TFP	0.95 (0.87, 0.99)	0.90 (0.77, 0.97)	0.86 (0.66, 0.95)	0.79 (0.50, 0.93)	0.67 (0.26, 0.91)	0.64 (0.21, 0.91)
	Risk Premium	0.02 (0.00, 0.09)	0.05 (0.01, 0.15)	0.05 (0.01, 0.18)	0.04 (0.01, 0.17)	0.02 (0.01, 0.16)	0.02 (0.01, 0.15)
	Price of Capital	0.02 (0.00, 0.06)	0.04 (0.01, 0.13)	0.07 (0.01, 0.21)	0.07 (0.01, 0.23)	0.07 (0.01, 0.23)	0.07 (0.01, 0.23)
	Collateral Constraint	0.01 (0.00, 0.03)	0.01 (0.00, 0.06)	0.02 (0.00, 0.15)	0.10 (0.01, 0.39)	0.24 (0.01, 0.67)	0.27 (0.01, 0.74)
R_t	TFP	0.01 (0.00, 0.07)	0.03 (0.00, 0.14)	0.02 (0.00, 0.14)	0.01 (0.00, 0.15)	0.01 (0.00, 0.17)	0.01 (0.00, 0.18)
	Risk Premium	0.87 (0.76, 0.94)	0.83 (0.67, 0.93)	0.83 (0.63, 0.93)	0.80 (0.50, 0.93)	0.73 (0.29, 0.93)	0.72 (0.24, 0.93)
	Price of Capital	0.11 (0.05, 0.20)	0.13 (0.05, 0.26)	0.14 (0.04, 0.29)	0.14 (0.03, 0.30)	0.13 (0.02, 0.29)	0.12 (0.02, 0.29)
	Collateral Constraint	0.01 (0.00, 0.04)	0.01 (0.00, 0.06)	0.01 (0.00, 0.13)	0.05 (0.00, 0.32)	0.13 (0.00, 0.58)	0.15 (0.00, 0.64)
Q_t	TFP	0.02 (0.00, 0.09)	0.03 (0.00, 0.10)	0.08 (0.01, 0.25)	0.12 (0.01, 0.36)	0.13 (0.01, 0.41)	0.13 (0.01, 0.42)
	Risk Premium	0.05 (0.03, 0.13)	0.04 (0.02, 0.12)	0.03 (0.02, 0.12)	0.03 (0.01, 0.16)	0.04 (0.00, 0.20)	0.04 (0.00, 0.21)
	Price of Capital	0.92 (0.81, 0.96)	0.92 (0.79, 0.95)	0.83 (0.60, 0.93)	0.67 (0.36, 0.89)	0.49 (0.15, 0.83)	0.45 (0.12, 0.82)
	Collateral Constraint	0.01 (0.00, 0.05)	0.01 (0.00, 0.06)	0.06 (0.01, 0.20)	0.18 (0.01, 0.46)	0.34 (0.02, 0.70)	0.38 (0.03, 0.75)
TB_t	TFP	0.01 (0.00, 0.03)	0.01 (0.00, 0.06)	0.01 (0.00, 0.12)	0.02 (0.00, 0.19)	0.03 (0.00, 0.25)	0.03 (0.00, 0.26)
	Risk Premium	0.03 (0.00, 0.09)	0.04 (0.00, 0.13)	0.04 (0.00, 0.17)	0.05 (0.00, 0.21)	0.05 (0.00, 0.24)	0.05 (0.00, 0.25)
	Price of Capital	0.01 (0.00, 0.03)	0.01 (0.00, 0.04)	0.01 (0.00, 0.06)	0.01 (0.00, 0.08)	0.01 (0.00, 0.09)	0.01 (0.00, 0.10)
	Collateral Constraint	0.95 (0.90, 0.99)	0.94 (0.85, 0.99)	0.94 (0.77, 0.99)	0.92 (0.67, 0.99)	0.91 (0.61, 0.99)	0.91 (0.60, 0.99)

Notes: The table summarizes the contributions of TFP, risk premium, price of capital, and collateral constraint shocks to the fluctuations of real GDP, Y_t, the world real interest rate, R_t, Tobin's Q, Q_t, and the trade balance, TB_t. The parentheses contain 90% bootstrap sup-t confidence bands.

Table 3: FEVDs Estimated on the Brazilian Sample from 1999Q1 to 2018Q4 Under Long-Run Restrictions

Variable	Horizons / Shocks	Q1	Q2	Q4	Q8	Q16	Q20
Y_t	TFP	0.48 (0.01, 0.92)	0.52 (0.01, 0.94)	0.60 (0.03, 0.96)	0.72 (0.11, 0.98)	0.85 (0.35, 0.99)	0.87 (0.42, 0.99)
	Risk Premium	0.09 (0.00, 0.53)	0.05 (0.00, 0.47)	0.02 (0.00, 0.43)	0.02 (0.00, 0.37)	0.01 (0.00, 0.26)	0.01 (0.00, 0.23)
	Price of Capital	0.33 (0.02, 0.74)	0.29 (0.01, 0.70)	0.23 (0.01, 0.65)	0.15 (0.01, 0.55)	0.09 (0.00, 0.37)	0.07 (0.00, 0.32)
	Collateral Constraint	0.10 (0.00, 0.46)	0.14 (0.01, 0.50)	0.15 (0.01, 0.49)	0.11 (0.01, 0.36)	0.05 (0.01, 0.20)	0.05 (0.01, 0.16)
R_t	TFP	0.03 (0.00, 0.26)	0.04 (0.00, 0.33)	0.05 (0.00, 0.35)	0.08 (0.01, 0.41)	0.12 (0.01, 0.56)	0.13 (0.01, 0.60)
	Risk Premium	0.71 (0.12, 0.92)	0.68 (0.10, 0.92)	0.73 (0.14, 0.93)	0.77 (0.21, 0.95)	0.79 (0.24, 0.97)	0.79 (0.23, 0.97)
	Price of Capital	0.16 (0.01, 0.61)	0.16 (0.01, 0.59)	0.13 (0.01, 0.55)	0.09 (0.01, 0.46)	0.05 (0.00, 0.31)	0.05 (0.00, 0.26)
	Collateral Constraint	0.10 (0.02, 0.41)	0.12 (0.02, 0.43)	0.09 (0.01, 0.39)	0.06 (0.01, 0.29)	0.04 (0.00, 0.16)	0.03 (0.00, 0.13)
Q_t	TFP	0.07 (0.00, 0.33)	0.06 (0.00, 0.29)	0.05 (0.01, 0.29)	0.08 (0.01, 0.49)	0.13 (0.01, 0.73)	0.14 (0.01, 0.78)
	Risk Premium	0.13 (0.01, 0.41)	0.14 (0.01, 0.39)	0.13 (0.02, 0.37)	0.13 (0.01, 0.42)	0.12 (0.01, 0.50)	0.12 (0.01, 0.52)
	Price of Capital	0.06 (0.00, 0.43)	0.15 (0.01, 0.52)	0.34 (0.05, 0.67)	0.51 (0.06, 0.79)	0.62 (0.05, 0.88)	0.64 (0.05, 0.90)
	Collateral Constraint	0.74 (0.26, 0.94)	0.65 (0.23, 0.89)	0.48 (0.15, 0.76)	0.28 (0.09, 0.53)	0.13 (0.04, 0.26)	0.10 (0.04, 0.21)
TB_t	TFP	0.60 (0.01, 0.97)	0.56 (0.01, 0.97)	0.51 (0.01, 0.98)	0.47 (0.01, 0.98)	0.45 (0.00, 0.98)	0.44 (0.00, 0.98)
	Risk Premium	0.01 (0.00, 0.43)	0.01 (0.00, 0.43)	0.01 (0.00, 0.45)	0.01 (0.00, 0.47)	0.01 (0.00, 0.48)	0.01 (0.00, 0.48)
	Price of Capital	0.21 (0.00, 0.67)	0.23 (0.01, 0.68)	0.26 (0.01, 0.69)	0.28 (0.01, 0.70)	0.29 (0.01, 0.71)	0.29 (0.01, 0.71)
	Collateral Constraint	0.18 (0.00, 0.54)	0.20 (0.00, 0.55)	0.22 (0.00, 0.58)	0.24 (0.01, 0.60)	0.25 (0.01, 0.62)	0.26 (0.01, 0.62)

Notes: The table summarizes the contributions of TFP, risk premium, price of capital, and collateral constraint shocks to the fluctuations of real GDP, Y_t, the world real interest rate, R_t, Tobin's Q, Q_t, and the trade balance, TB_t. The parentheses contain 90% bootstrap sup-t confidence bands.

Table 4: FEVDs Estimated on the Mexican Sample from 1997Q1 to 2018Q4 Under Short-Run Restrictions

Variable	Shocks \ Horizons	Q1	Q2	Q4	Q8	Q16	Q20
Y_t	TFP	0.87 (0.78, 0.93)	0.71 (0.57, 0.83)	0.54 (0.37, 0.74)	0.51 (0.29, 0.73)	0.49 (0.25, 0.73)	0.48 (0.25, 0.73)
	Risk Premium	0.01 (0.00, 0.03)	0.01 (0.00, 0.04)	0.01 (0.00, 0.09)	0.02 (0.00, 0.14)	0.02 (0.00, 0.16)	0.02 (0.00, 0.16)
	Price of Capital	0.05 (0.01, 0.11)	0.12 (0.03, 0.24)	0.20 (0.05, 0.36)	0.21 (0.05, 0.41)	0.22 (0.05, 0.44)	0.23 (0.05, 0.44)
	Collateral Constraint	0.07 (0.03, 0.15)	0.16 (0.07, 0.29)	0.25 (0.09, 0.41)	0.26 (0.08, 0.46)	0.27 (0.08, 0.48)	0.27 (0.08, 0.48)
R_t	TFP	0.02 (0.00, 0.07)	0.05 (0.00, 0.16)	0.09 (0.01, 0.26)	0.11 (0.01, 0.28)	0.12 (0.01, 0.31)	0.12 (0.01, 0.31)
	Risk Premium	0.89 (0.77, 0.96)	0.83 (0.66, 0.93)	0.80 (0.58, 0.92)	0.81 (0.57, 0.93)	0.82 (0.55, 0.94)	0.82 (0.55, 0.94)
	Price of Capital	0.08 (0.01, 0.18)	0.09 (0.01, 0.23)	0.08 (0.01, 0.25)	0.06 (0.01, 0.25)	0.05 (0.01, 0.25)	0.05 (0.00, 0.25)
	Collateral Constraint	0.01 (0.00, 0.06)	0.03 (0.00, 0.12)	0.03 (0.00, 0.13)	0.02 (0.00, 0.13)	0.01 (0.00, 0.14)	0.01 (0.00, 0.14)
Q_t	TFP	0.01 (0.00, 0.03)	0.01 (0.00, 0.07)	0.01 (0.00, 0.13)	0.01 (0.00, 0.14)	0.01 (0.00, 0.15)	0.01 (0.00, 0.15)
	Risk Premium	0.02 (0.00, 0.07)	0.04 (0.00, 0.13)	0.06 (0.00, 0.20)	0.07 (0.00, 0.22)	0.07 (0.00, 0.24)	0.07 (0.00, 0.24)
	Price of Capital	0.94 (0.85, 0.98)	0.91 (0.77, 0.97)	0.89 (0.71, 0.97)	0.90 (0.69, 0.97)	0.90 (0.66, 0.97)	0.90 (0.66, 0.97)
	Collateral Constraint	0.03 (0.00, 0.10)	0.04 (0.00, 0.14)	0.04 (0.01, 0.13)	0.02 (0.00, 0.13)	0.02 (0.00, 0.14)	0.02 (0.00, 0.14)
TB_t	TFP	0.01 (0.00, 0.03)	0.01 (0.00, 0.05)	0.01 (0.00, 0.07)	0.01 (0.00, 0.09)	0.01 (0.00, 0.10)	0.01 (0.00, 0.11)
	Risk Premium	0.01 (0.00, 0.04)	0.01 (0.00, 0.05)	0.01 (0.00, 0.11)	0.02 (0.00, 0.15)	0.02 (0.00, 0.17)	0.02 (0.00, 0.17)
	Price of Capital	0.02 (0.00, 0.08)	0.05 (0.00, 0.16)	0.07 (0.00, 0.23)	0.08 (0.00, 0.27)	0.08 (0.00, 0.28)	0.08 (0.00, 0.29)
	Collateral Constraint	0.96 (0.90, 0.99)	0.93 (0.81, 0.98)	0.91 (0.72, 0.98)	0.89 (0.65, 0.98)	0.89 (0.62, 0.98)	0.89 (0.61, 0.98)

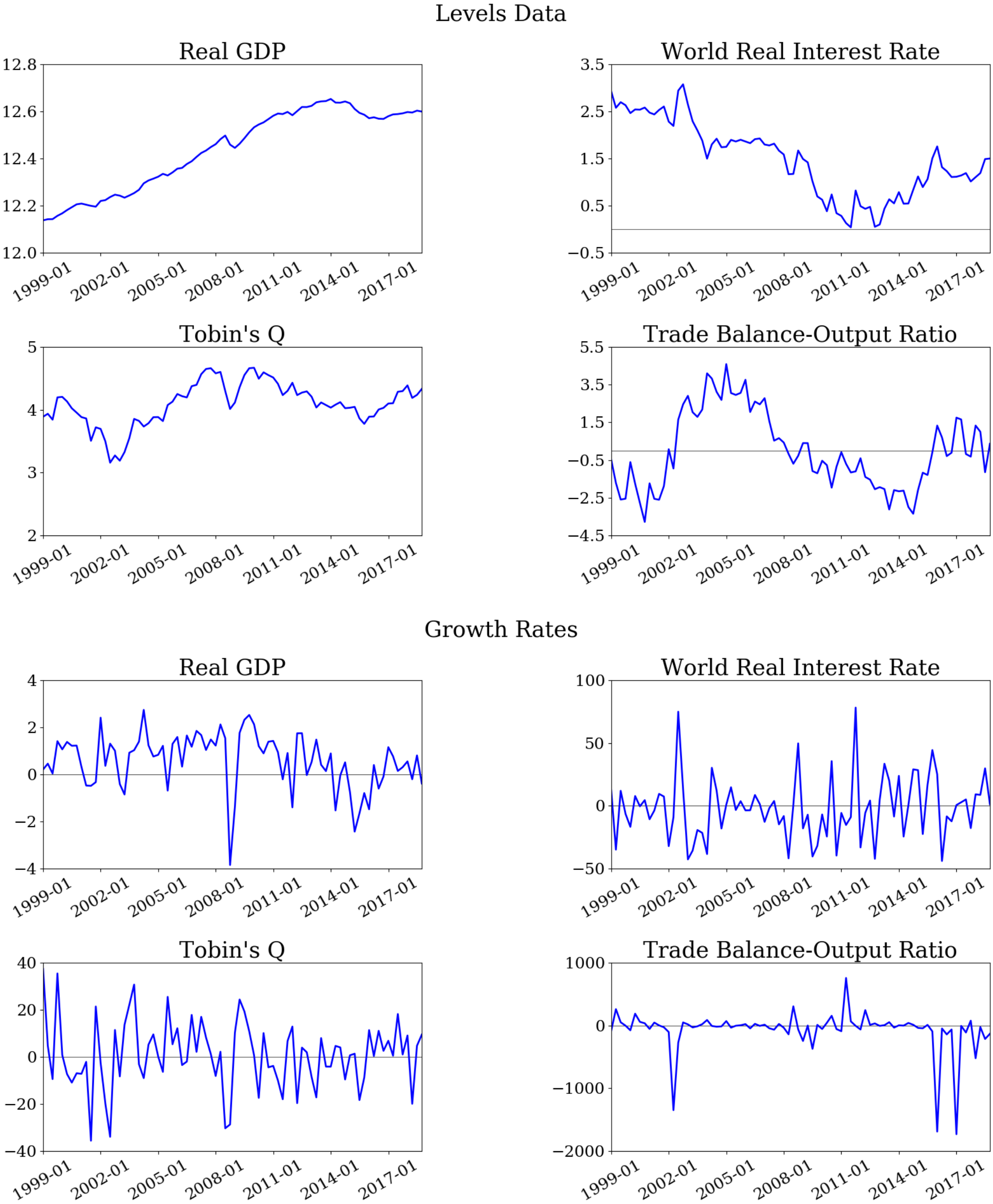
Notes: The table summarizes the contributions of TFP, risk premium, price of capital, and collateral constraint shocks to the fluctuations of real GDP, Y_t, the world real interest rate, R_t, Tobin's Q, Q_t, and the trade balance, TB_t. The parentheses contain 90% bootstrap sup-t confidence bands.

Table 5: FEVDs Estimated on the Mexican Sample from 1997Q1 to 2018Q4 Under Long-Run Restrictions

Variable	Shocks \ Horizons	Q1	Q2	Q4	Q8	Q16	Q20
Y_t	TFP	0.73 (0.48, 0.87)	0.85 (0.67, 0.93)	0.92 (0.83, 0.96)	0.96 (0.90, 0.98)	0.97 (0.95, 0.99)	0.97 (0.95, 0.99)
	Risk Premium	0.01 (0.00, 0.08)	0.01 (0.00, 0.07)	0.01 (0.00, 0.04)	0.01 (0.00, 0.02)	0.01 (0.00, 0.01)	0.01 (0.00, 0.01)
	Price of Capital	0.14 (0.03, 0.32)	0.07 (0.02, 0.19)	0.04 (0.01, 0.10)	0.02 (0.00, 0.06)	0.01 (0.00, 0.03)	0.01 (0.00, 0.03)
	Collateral Constraint	0.12 (0.03, 0.29)	0.07 (0.02, 0.17)	0.03 (0.01, 0.09)	0.01 (0.01, 0.05)	0.01 (0.00, 0.03)	0.01 (0.00, 0.02)
R_t	TFP	0.01 (0.00, 0.14)	0.01 (0.00, 0.15)	0.03 (0.01, 0.24)	0.04 (0.00, 0.31)	0.05 (0.00, 0.35)	0.05 (0.00, 0.35)
	Risk Premium	0.93 (0.73, 0.96)	0.93 (0.72, 0.96)	0.92 (0.65, 0.96)	0.92 (0.62, 0.97)	0.92 (0.59, 0.98)	0.92 (0.59, 0.98)
	Price of Capital	0.03 (0.00, 0.14)	0.04 (0.01, 0.19)	0.04 (0.01, 0.22)	0.03 (0.01, 0.22)	0.02 (0.00, 0.22)	0.02 (0.00, 0.23)
	Collateral Constraint	0.03 (0.00, 0.13)	0.02 (0.00, 0.09)	0.01 (0.00, 0.06)	0.01 (0.00, 0.04)	0.01 (0.00, 0.02)	0.01 (0.00, 0.02)
Q_t	TFP	0.34 (0.12, 0.58)	0.32 (0.10, 0.57)	0.28 (0.07, 0.55)	0.26 (0.05, 0.55)	0.25 (0.04, 0.56)	0.24 (0.03, 0.56)
	Risk Premium	0.03 (0.00, 0.13)	0.02 (0.00, 0.08)	0.01 (0.00, 0.05)	0.01 (0.00, 0.03)	0.01 (0.00, 0.02)	0.01 (0.00, 0.01)
	Price of Capital	0.61 (0.35, 0.82)	0.64 (0.38, 0.85)	0.70 (0.42, 0.89)	0.72 (0.43, 0.93)	0.73 (0.43, 0.95)	0.74 (0.43, 0.96)
	Collateral Constraint	0.02 (0.00, 0.11)	0.02 (0.01, 0.07)	0.01 (0.01, 0.05)	0.01 (0.00, 0.03)	0.01 (0.00, 0.02)	0.01 (0.00, 0.01)
TB_t	TFP	0.36 (0.12, 0.60)	0.41 (0.15, 0.63)	0.43 (0.15, 0.67)	0.45 (0.15, 0.70)	0.46 (0.15, 0.72)	0.46 (0.14, 0.72)
	Risk Premium	0.10 (0.01, 0.29)	0.08 (0.01, 0.25)	0.05 (0.00, 0.21)	0.04 (0.00, 0.19)	0.03 (0.00, 0.18)	0.03 (0.00, 0.18)
	Price of Capital	0.01 (0.00, 0.13)	0.01 (0.00, 0.11)	0.01 (0.00, 0.13)	0.01 (0.00, 0.14)	0.01 (0.00, 0.15)	0.01 (0.00, 0.16)
	Collateral Constraint	0.53 (0.28, 0.74)	0.50 (0.27, 0.73)	0.51 (0.27, 0.74)	0.50 (0.25, 0.75)	0.50 (0.24, 0.76)	0.50 (0.24, 0.76)

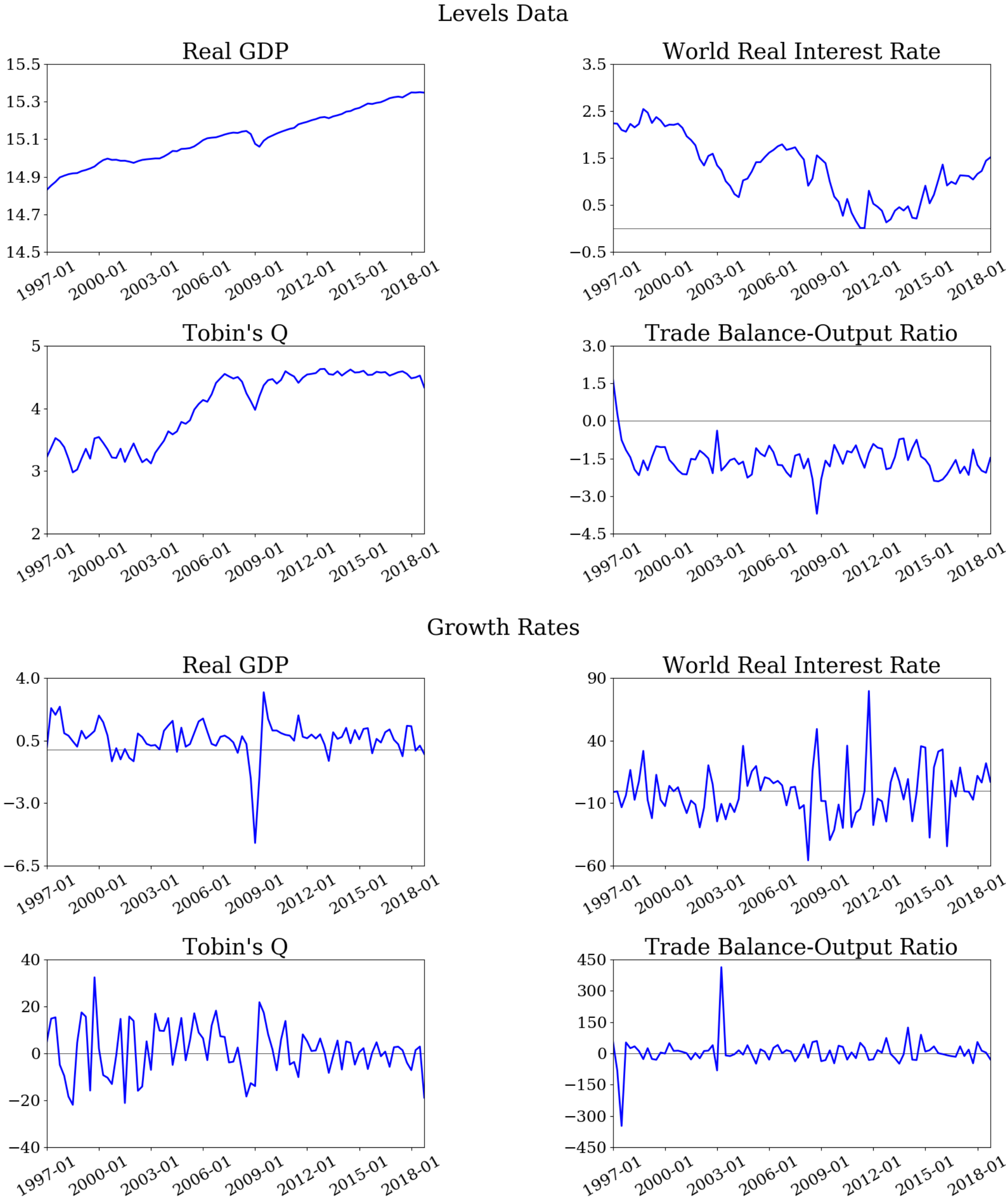
Notes: The table summarizes the contributions of TFP, risk premium, price of capital, and collateral constraint shocks to the fluctuations of real GDP, Y_t, the world real interest rate, R_t, Tobin's Q, Q_t, and the trade balance, TB_t. The parentheses contain 90% bootstrap sup-t confidence bands.

Figure 1: Brazilian Sample from 1999Q1 to 2018Q4



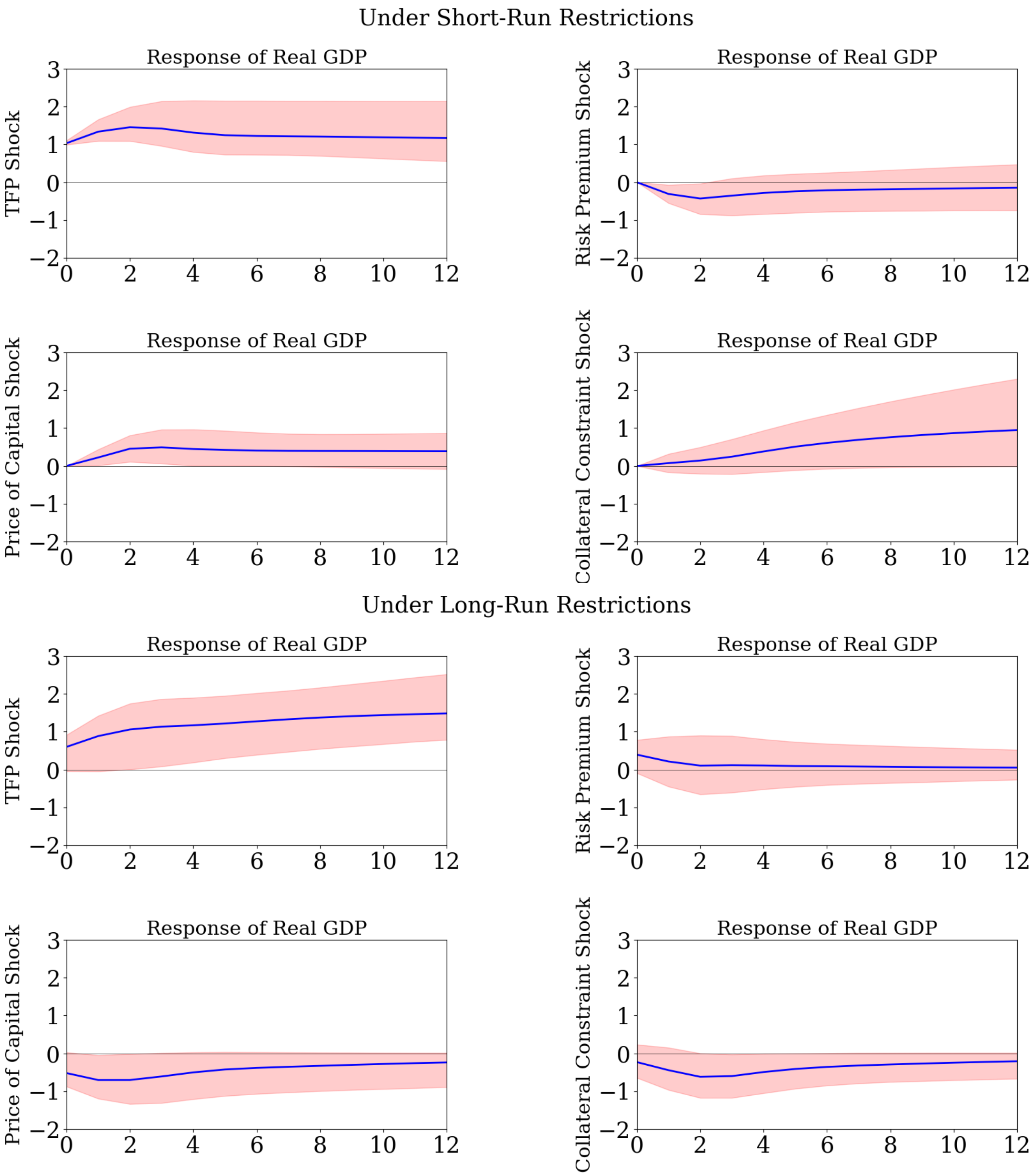
Notes: The top four graphs present log real GDP, the log world real interest rate, log Tobin's Q, and the trade balance-output ratio. Growth rates are the first differences of the log levels. The first difference of the trade balance-output ratio is computed as $100{*}(\frac{TB_t}{Y_t}\text{-}\frac{TB_{t-1}}{Y_{t-1}})/\frac{TB_{t-1}}{Y_{t-1}}$.

Figure 2: Mexican Sample from 1997Q1 to 2018Q4



Notes: The top four graphs present log real GDP, the log world real interest rate, log Tobin's Q, and the trade balance-output ratio. Growth rates are the first differences of the log levels. The first difference of the trade balance-output ratio is computed as $100*(\frac{TB_t}{Y_t}\text{-}\frac{TB_{t-1}}{Y_{t-1}})/\frac{TB_{t-1}}{Y_{t-1}}$.

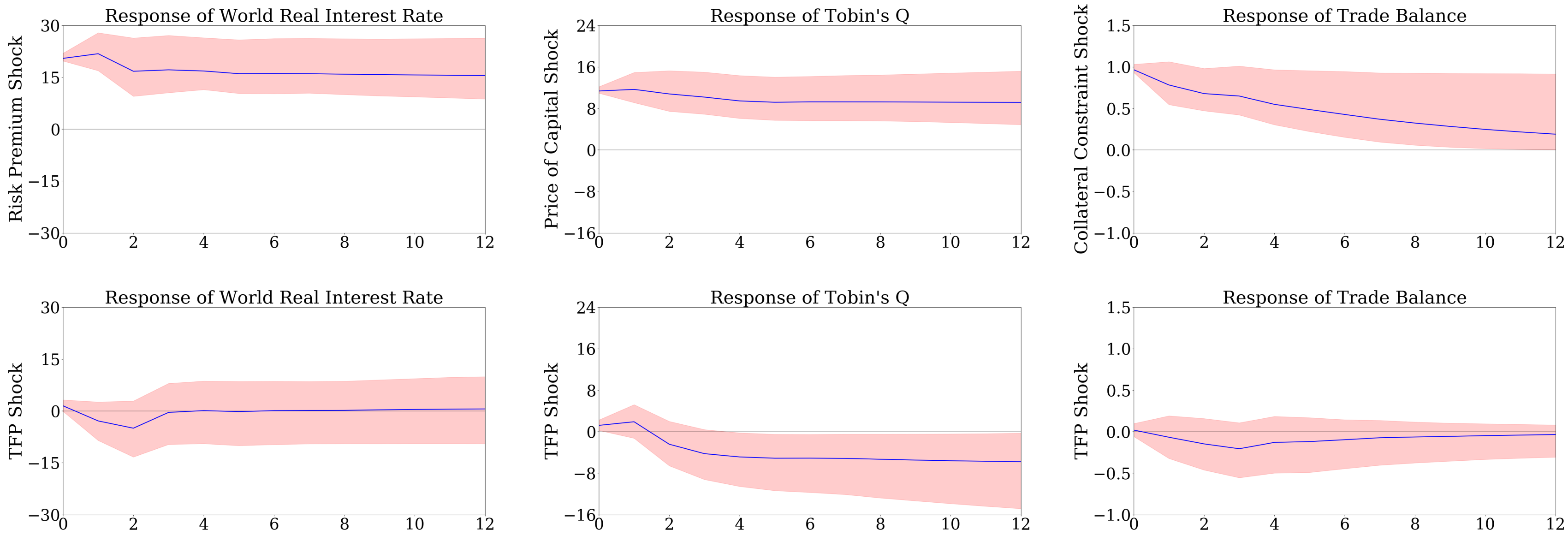
Figure 3: IRFs of Real GDP Estimated on the Brazilian Sample from 1999Q1 to 2018Q4



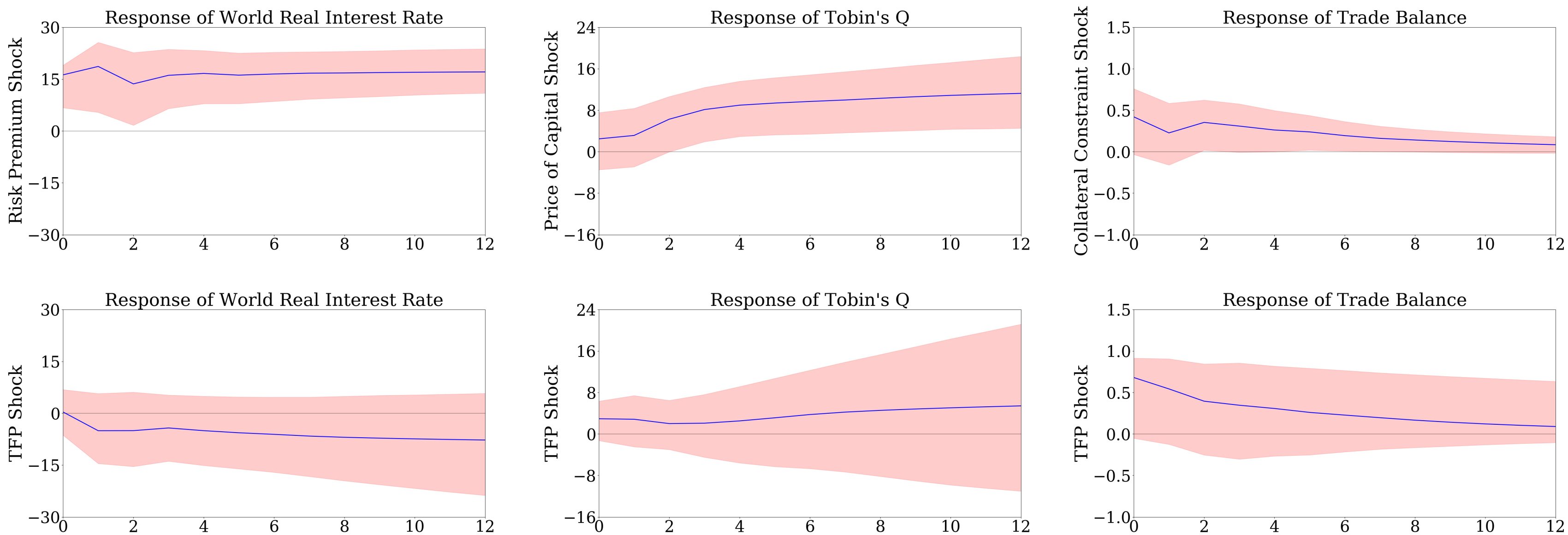
Notes: The blue solid lines trace the IRFs. The red shadings trace the 90% bootstrap sup-t confidence bands. The units of real GDP are in log levels.

Figure 4: Selected IRFs Estimated on the Brazilian Sample from 1999Q1 to 2018Q4

Under Short-Run Restrictions



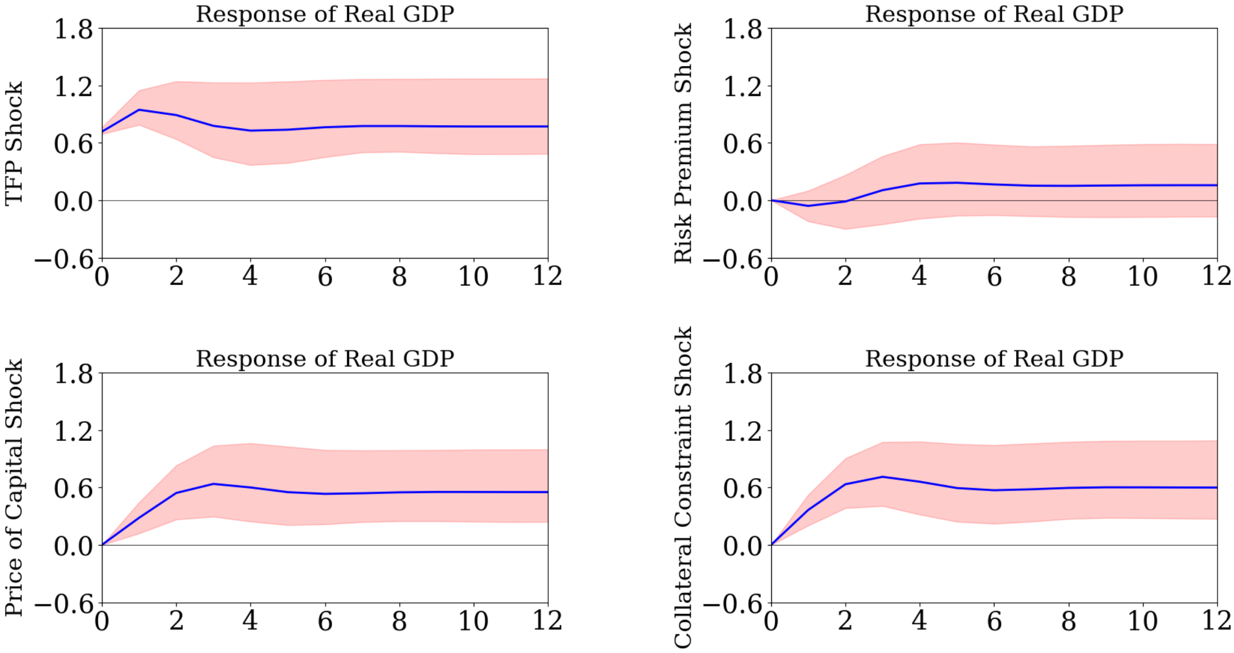
Under Long-Run Restrictions



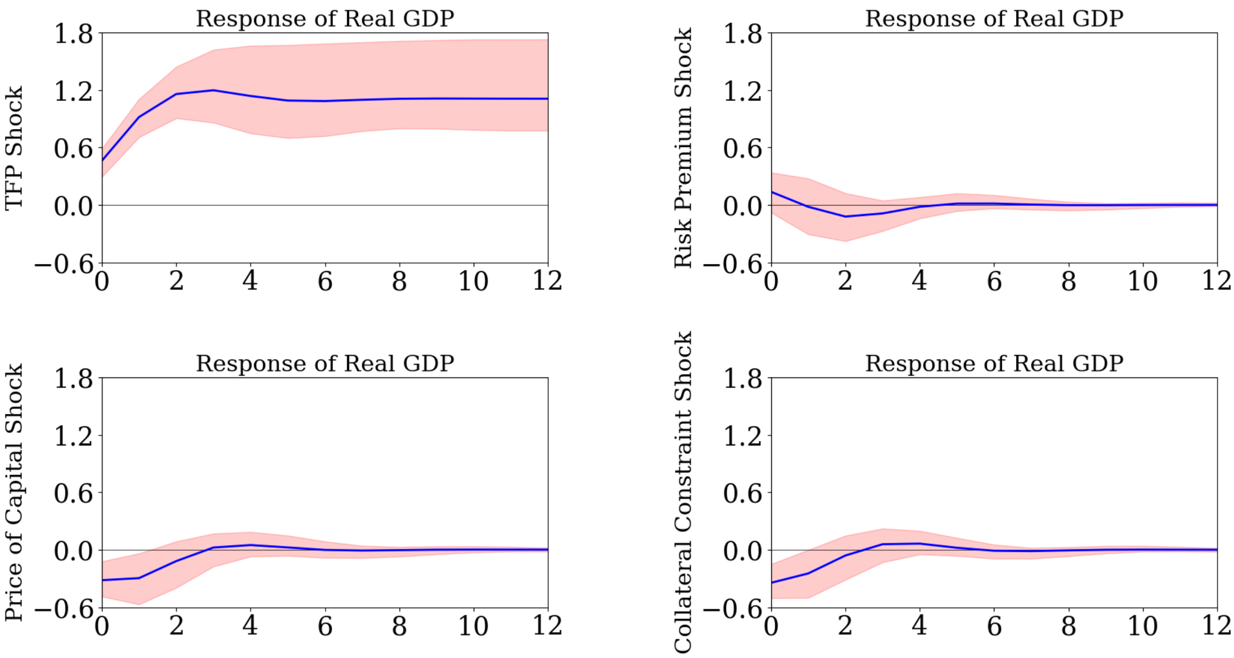
Notes: The blue solid lines trace the IRFs. The red shadings trace 90% bootstrap sup-t confidence bands. The units of the trade balance are in millions of local currency units; the world real interest rate and Tobin's Q are in percentages.

Figure 5: IRFs of Real GDP Estimated on the Mexican Sample from 1997Q1 to 2018Q4

Under Short-Run Restrictions

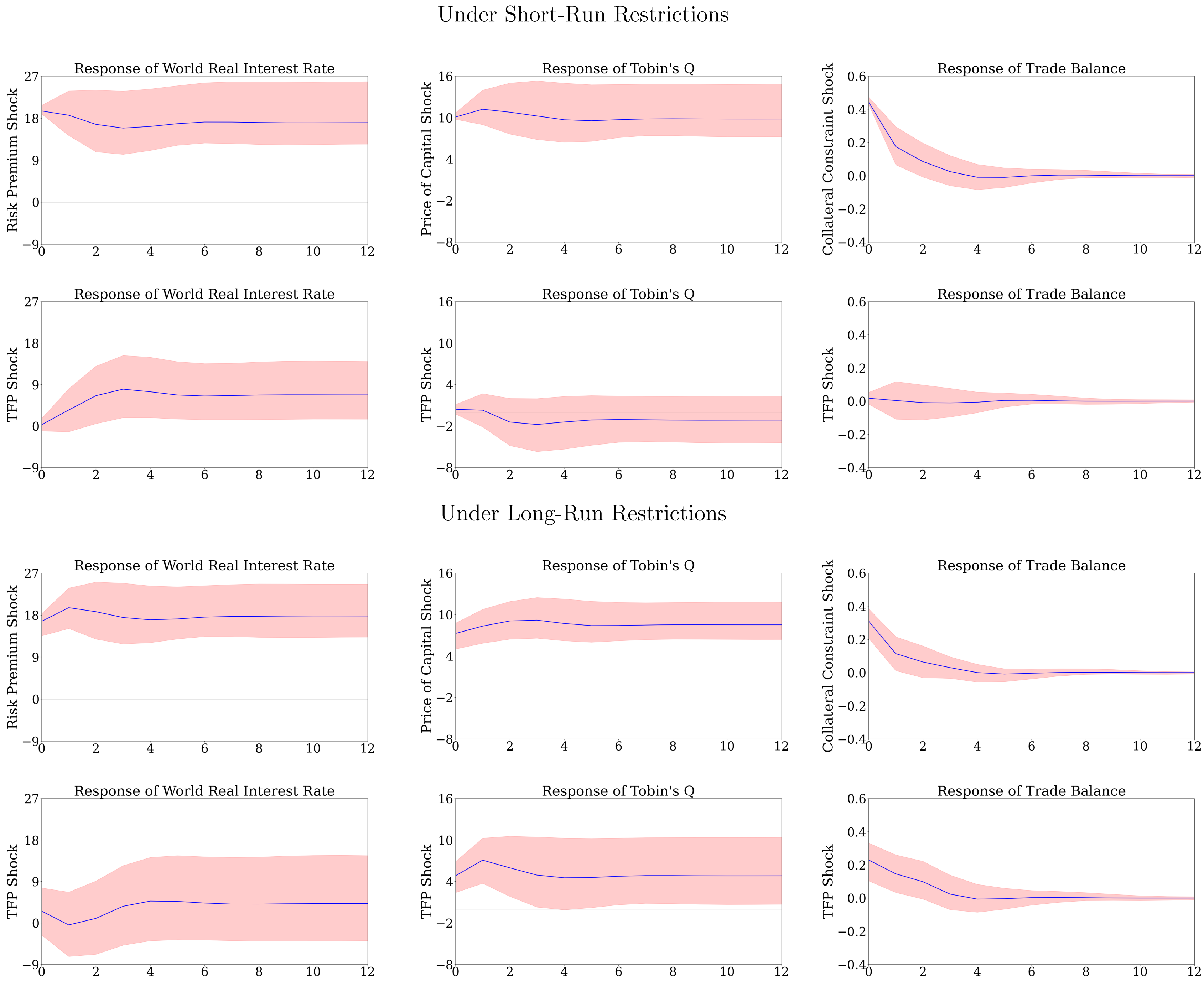


Under Long-Run Restrictions



Notes: The blue solid lines trace the IRFs. The red shadings trace the 90% bootstrap sup-t confidence bands. The units of real GDP are in log levels.

Figure 6: Selected IRFs Estimated on the Mexican Sample from 1997Q1 to 2018Q4



Notes: The blue solid lines trace the IRFs. The red shadings trace 90% bootstrap sup-t confidence bands. The units of the trade balance are in millions of local currency units; the world real interest rate and Tobin's Q are in percentages.