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What Determines the Neutral Rate of Interest in an Emerging Economy?*

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Abstract: Evidence suggests that potential growth and the neutral rate co-move in advanced economies. In contrast, this co-movement is not observed in emerging economies. We argue that capital flows may explain this behavior. We focus on Mexico, a benchmark emerging economy, and find that capital inflows may account for a temporary reduction in the Mexican neutral rate after the global financial crisis. These inflows surged during the implementation of unconventional monetary policies in advanced economies. In turn, low-frequency changes in the neutral rate may be attributed to increasing domestic savings, demographics, and a decreasing global long-run real interest rate. These results are largely consistent with other studies showing that the neutral rate has decreased in the last 25 years in advanced and emerging economies.
Keywords: Neutral rate of interest, emerging market economies, transitory and structural factors
JEL Classification: C10, E43, E52

Resumen: La evidencia sugiere que el crecimiento potencial y la tasa neutral se mueven conjuntamente en economías avanzadas. En contraste, este co-movimiento no es observado en economías emergentes. Argumentamos que los flujos de capitales pueden explicar este comportamiento. Nos enfocamos en México, una economía emergente de referencia, y encontramos que entradas de capitales pueden explicar una reducción temporal en la tasa neutral Mexicana después de la crisis financiera global. Estas entradas aumentaron durante la implementación de políticas monetarias no convencionales en economías avanzadas. A su vez, cambios de baja frecuencia en la tasa neutral pueden atribuirse a un ahorro doméstico creciente, a cambios demográficos y a una tasa de interés real de largo plazo global decreciente. Estos resultados son ampliamente consistentes con otros estudios que señalan que la tasa neutral ha disminuido en los últimos 25 años en economías avanzadas y emergentes.
Palabras Clave: Tasa neutral de interés, economías emergentes, factores transitorios y estructurales

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1 Introduction

The neutral rate of interest, which we call interchangeably natural interest rate, neutral rate, or simply $r^\star$, can be defined as the level of the short-run real interest rate that is consistent with output near its potential, and stable inflation near its target (see Laubach and Williams, 2003). The neutral rate is determined in the domestic market of loanable funds, so factors that affect this market prompt changes in the neutral rate.[1] We can classify these factors into structural (such as potential growth, demographics, financial-markets development, etc.) and transitory (such as macroeconomic shocks; see Section 2 for further details). Since these factors are exogenous to central banks, $r^\star$ is not a policy choice.

In contrast, $r^\star$ is relevant for central banks because it helps them to determine the stance of monetary policy.[2] Despite its importance, the neutral rate is an elusive indicator for monetary policy because: (1) it is not observable, and must be inferred using quantitative methods that are subject to an important statistical uncertainty; and (2) it may vary due to changes in both structural and transitory factors.

Recent studies estimating $r^\star$ in both advanced and emerging economies attempt to reveal the levels of interest rates that support a *normal* pace of economic activity. After the global financial crisis (GFC, henceforth), the *new normal* for interest rates is expected to be lower than prior to the crisis because of the persistently low levels of economic activity in advanced economies since the GFC, a shift in demographics, and a decreasing global long-term real interest rate in the last three decades (see Rachel and Smith, 2015). In this context, the results found in these studies are remarkably similar: most estimates show a downward trend in $r^\star$ since before the GFC, but the downturn is especially evident during the crisis.[3]

[1] In this market, $r^\star$ is the price, and traded loans the quantity. *Desired savings*, both from domestic and foreign parties, determine the supply, while *investment demand*, composed by public and private debt, determines the demand.

[2] This stance is *neutral* if the short-run real interest rate equals $r^\star$, and it is contractionary (expansionary) if the short-run real rate locates above (below) $r^\star$. If the stance is contractionary, monetary policy slows down aggregate demand by setting an opportunity cost of funds for consumption and investment higher than it would *normally* be. The opposite happens if the stance is expansionary. If we add a medium-term measure of inflation expectations to $r^\star$, we get the level of the policy interest rate at which monetary policy is neutral.

[3] See Annex B for a non-exhaustive review of the literature.

Holston, Laubach and Williams (2017) find that the estimated neutral rates and trend growth rates of four advanced economies (AEs), namely the U.S., Canada, the U.K., and the Euro Area, have co-moved tightly for the last 25 years. These authors suggest that global factors may largely explain this behavior. In sharp contrast, the same co-movement does not hold in emerging market economies (EMEs), since most of these countries grow at relatively high rates, while at the same time their neutral rates have fallen (see Annexes A and B).[4]

A dimension that has not been fully explored in the neutral-rate literature is the role of capital flows in shaping $r^\star$. Indeed, sustained capital flows could have a long-lasting effect on the supply of loanable funds of an EME, affecting its neutral rate. This channel is potentially more important for EMEs than for AEs, given the exposure of the former in international markets. Therefore, in an EME, factors other than potential growth seem to have a relatively large importance in the determination of $r^\star$. In this paper, we seek to illustrate this point quantitatively. We focus on Mexico, a prototype EME with an important volume of international trade, and a financial market friendly to international investors. Mexico can be viewed as a benchmark study for EMEs, since the techniques used for this economy can serve other similar countries. In particular, the availability of shorter time series data for EMEs than for AEs poses significant methodological challenges, which we tackle in this paper. Our sample period spans from January 2000 to December 2017.[5]

Unlike the typical approach in the neutral-rate literature, we consider two different frequency domains to study the dynamics of $r^\star$. At a low frequency, we acknowledge that the neutral rate is determined exclusively by structural factors, which change slowly through

[4]The IMF's WEO of April 2018, Box 1.3, presents potential growth estimates for selected AEs and EMEs, and finds that potential growth has persistently decreased for the former, while it follows an inverted U-shaped pattern for the latter. In particular, for the group of AEs, trend growth fell from 2.5% in 2001 to 1.5% in 2017, while for the group of EMEs, trend growth located at 4% in both years, with a peak at 5% in 2007. In Annex A, we review the growth rates and short-run real interest rates for a larger set of AEs and EMEs. The data for output growth are consistent with the IMF's results. In addition, the data for the short-run real interest rate show a decreasing trend in both AEs and EMEs since at least 1993. These trends are confirmed for the neutral rate for AEs and EMEs in Annex B.

[5]Anchoring trend inflation through the adoption of an inflation-targeting regime in 2001 propelled a drastic change in the time-series properties of inflation in Mexico (see Chiquiar, Noriega and Ramos-Francia, 2010). This fact, along with a formal change in the monetary-policy instrument during the 2000s, make it difficult to combine different monetary-policy regimes using a single estimation method (Banco de México gradually changed its targeting instrument from the monetary base to the short-run nominal interest rate during the 2000s).

time. At a high frequency, we assume that transitory factors may temporarily divert the neutral rate from its fundamental value. We call the low-frequency measure of $r^\star$ as its *long-run convergence level*, while we refer to the high-frequency measure as *short-run* $r^\star$. In Section 2, we provide a formal description of these two concepts.

To achieve a robust estimate of short-run $r^\star$, we consider five different approaches: averages and filters, a simple Taylor rule estimated recursively, affine term-structure models, the Laubach and Williams (2003) model adapted for a small open economy, and a BVAR model with time-varying intercepts (or TVI-BVAR for short). Some of these estimates are clearly affected by high-frequency shocks, while other filter these shocks but still feature the effects of very persistent transitory factors. Therefore, we refer to these estimates, without distinction, as short-run or medium-run measures of $r^\star$. Conversely, to compute the long-run convergence level of the neutral rate, we estimate an augmented Taylor rule which includes a control for a very persistent transitory factor, and open-economy RBC model, and the 10-year expectation of the short-run nominal interest rate computed from an affine model.

All medium-run measures exhibit a similar path: $r^\star$ has trended downwards, in general, at least since 2001. The exception is at the onset of the GFC, where $r^\star$ followed a U-shaped pattern, falling to record low levels by 2012, and partially reverting to trend since 2014. We claim that persistent transitory factors explain the U-shaped pattern of $r^\star$, while its downward trend may be attributed to changes in structural factors.

On the *drivers of short- and medium-run* $r^\star$, we argue that both domestic and foreign transitory factors pushed the neutral rate downwards in the aftermath of the GFC. This is the case because the estimate of potential growth appears relatively steady in comparison to the estimates of $r^\star$. From the domestic dimension, slack conditions prevalent in the Mexican economy following the GFC implied a demand for loanable funds lower than normal, which depressed $r^\star$. From the foreign dimension, we find two important transitory factors: (1) persistent slack conditions in the U.S. after the crisis, and (2) the implementation of unconventional monetary policies (UMPs, henceforth) by central banks in some AEs, notably the Federal Reserve of the U.S. (or the Fed).

Concerning the first foreign transitory factor, the evidence points that the U.S. business cycle correlates not only with the Mexican business cycle, but also with the Mexican neutral rate. Therefore, transitory factors affecting the neutral rate in the U.S. may also impact the neutral rate in Mexico, given the strong commercial and financial links between Mexico and its northern neighbor.

Regarding the second foreign transitory factor, a growing amount of evidence suggests that UMPs triggered a boom in capital flows towards EMEs, and Mexico was no exception.[6] In the Mexican case, most of the increase in capital inflows between 2009 and 2014 was directed towards portfolio investment, in particular in government debt securities. The data suggest that the surge in government bond holdings by non-residents affected both ends of the Mexican yield curve, but it was the short end that experienced an unprecedented momentum precisely when UMPs were expanding. Then, when the Fed signaled for the first time the tapering its QE programs in mid 2013, the momentum vanished and non-residents scaled back their holdings of short-term Mexican debt. The rise and fall of these holdings, which increased and then decreased the supply of loanable funds in the country, may explain why the medium-run estimates of $r^\star$ reached a minimum between 2009 and 2012 in Mexico.

Our estimates of the Taylor rule for Mexico support the hypothesis that capital inflows depressed the Mexican neutral rate when UMPs were expanding. In Section 3.2, we recursively estimate a simple Taylor rule, from which the intercept corresponds to the estimate of $r^\star$. Notably, such estimate displays a U-shaped pattern as we add observations to the estimation. In contrast, when we include an indicator of the Fed's UMPs in the Taylor rule in Section 5.1, the time-variation in the estimate of $r^\star$ disappears. UMPs seem, thus, to capture a very persistent transitory factor affecting the Mexican neutral rate.

On the *drivers of the long-run convergence level of* $r^\star$, we argue again that both domestic and foreign structural factors account for the apparent fall registered from the 2000s to present. From the domestic side, we observe (1) a sustained growth of national savings as a percentage of GDP, (2) an increase of the proportion of the working-age population, (3) a de-

[6] See Fratzscher, Lo Duca and Straub (2016), Tillmann (2016), Anaya, Hachula and Offermanns (2017), Chari, Stedman and Lundblad (2017), Hernández-Vega (2017), Ramírez and González (2017), and references cited therein.

clining outlook in the growth rate of the labor force, and (4) a flat trend productivity. All four factors imply a lower long-run convergence level of $r^\star$.[7] From the foreign side, the sustained reduction in the global long-term real interest rate may have pushed international long-term credit towards the Mexican market, lowering the domestic long-term real interest rate through no-arbitrage conditions. The latter has contributed to increase the supply of loanable funds in the economy, putting downward pressure on $r^\star$.

In this paper, we make two contributions to the already vast literature on the neutral rate.[8] On the methodological side, we distinguish between high and low frequency determinants of $r^\star$, which complements the discussion on the implications for central banking. A major consequence is that a central bank's monetary-policy stance should be assessed not only regarding a relatively static object (the long-run convergence level of $r^\star$), but also with respect to a dynamic object that is affected by transitory factors (short-run $r^\star$).

The notion of short-run $r^\star$ and its longer-term value also appears in Laubach and Williams (2016). To distinguish between the two variables, the authors include a mean-reverting unobserved factor in their model for the U.S. that affects short-run $r^\star$. However, they do not interpret what this latent factor is. Interestingly, the only episode in which short-run $r^\star$ persistently diverts from its longer-term level is during the GFC. The unprecedented persistent transitory factors experienced during the crisis makes more relevant the distinction between high and low-frequency measures of the neutral rate. In this paper, we formalize the concept of short-run $r^\star$ and its long-run convergence level using economic theory. This decomposition helps central bankers to assess the factors that may alter the monetary-policy stance at short- and long-term horizons.[9]

[7] In the market of loanable funds, the first two factors raise the supply, while the last two reduce the demand (through their influence on investment).

[8] See Annex B for a literature review.

[9] As an illustration of this point, some participants at the FOMC meeting of January 30-31, 2018, distinguished between current estimates of $r^\star$ in the U.S. and its longer-term value: *"[b]y most estimates, the neutral level of the federal funds rate had been very low in recent years, but it was expected to rise slowly over time toward its longer-run level. However, the outlook for the neutral rate was uncertain [] For example, the neutral rate, which appeared to have fallen sharply during the Global Financial Crisis when financial headwinds had restrained demand, might move up more than anticipated as the global economy strengthened. Alternatively, the longer-run level of the neutral rate might remain low in the absence of fundamental shifts in trends in productivity, demographics, or the demand for safe assets"* (sic).

On the technical side, we propose a TVI-BVAR model explicitly designed for a small open economy (or SOE). This tool addresses some of the recent criticisms to the Laubach and Williams's approach.[10] The advantage of the TVI-BVAR model is that it can include a broader set of foreign and domestic variables to account for the dynamics of a SOE. Therefore, the TVI-BVAR model estimates a more stable pattern of the neutral rate in comparison to models with a more rigid structure. The TVI-BVAR model is similar in spirit to the full-fledged time-varying parameter (or TVP) BVAR models of Lubik and Matthes (2015), and Johannsen and Mertens (2016) for the U.S. A similarity between the TVI and TVP models is that they use their period-by-period trend component as a medium-run estimate of the neutral rate. The difference is that, given the broader set of variables considered in the TVI, this model permits time-variation only for a subset of its parameters. Considering all points, including the complex international interactions that a typical EME faces, the TVI-BVAR model is an appropriate and flexible tool to estimate $r^\star$ in this type of economies.

The remainder of the paper is organized as follows. Section 2 distinguishes structural and transitory factors in the determination of $r^\star$ using a simple general-equilibrium model. Section 3 presents the medium-run estimates of $r^\star$ for Mexico. Section 4 discusses the transitory factors explaining changes in $r^\star$ at business-cycle frequencies. Section 5 focuses on the long-run convergence level of the neutral rate in Mexico. The final section concludes.

2 A Primer on the Neutral Rate and its Determinants

The neutral rate of interest is the level of the short-run real interest rate that is consistent with economic activity near its potential or efficient level, and inflation near its long-run target. In other words, if the short-run real interest rate equals $r^\star$, the output gap is closed. The neutral rate may vary over time not only due to changes in structural factors, but also to the appearance of transitory factors. The former typically change very slowly (e.g., demographics, trend productivity, markets structure), while the latter reflect macro shocks that are expected to disappear eventually. The distinction between these two types of factors is relevant because it

[10] Kiley (2015) and Taylor and Wieland (2016) have emphasized the role of omitted factors in Laubach and Williams's model. See Sections 3.5 and 5.1 for details.

helps to disentangle the implications for monetary policy of changes in $r^\star$. A central bank may take a different course of actions depending on whether it expects changes in $r^\star$ to be permanent or temporary.

We now illustrate the effect of both structural and transitory factors on $r^\star$ with the help of a simple RBC model. Although we do not use this model in the quantitative analysis, it helps to fix ideas on the differences between short-run $r^\star$, and its long-run convergence level.

2.1 Closed-Economy Determinants

Assume a representative-agent economy, in which the agent chooses consumption C_t, bond holdings B_t, labor hours n_t, investment I_t, capital purchases K_t, and production Y_t in order to maximize his or her expected discounted lifetime utility subject to budget, technology and resource constraints. The objective of the agent is thus

$$\max_{C_\tau, B_\tau, n_\tau, I_\tau, K_\tau, y_\tau} \mathrm{E}_t \left\{ \sum_{\tau=t}^{\infty} \beta^{\tau-t} \left[\log C_\tau - h\left(n_t\right) \right] \right\} \ \forall\, \tau \geq t$$

subject to:

$$\begin{aligned}
0 &\leq Y_\tau + B_{\tau-1} - C_\tau - I_\tau - \frac{B_\tau}{\left(1+r_\tau\right)\exp\left(\varepsilon_\tau\right)}, && \text{(budget const.)}\\
Y_\tau &= K_{\tau-1}^{\alpha}\left(A_\tau n_\tau\right)^{1-\alpha}, && \text{(production tech.)}\\
K_\tau &\leq \left(1-\delta\right)K_{\tau-1} + I_\tau - \frac{\vartheta}{2}\left(\frac{I_\tau}{K_{\tau-1}} - \left(\delta+\gamma\right)\right)^2 K_{\tau-1}, && \text{(capital accum.)}\\
Y_\tau &= C_\tau + I_\tau. && \text{(resource const.)}
\end{aligned}$$

Variable r_t is the real interest rate, while ε_t is a mean-zero transitory stochastic disturbance that affects savings returns. The economy grows at a constant rate γ, which enters the model through the labor-neutral deterministic trend $A_t = A_{t-1}\left(1+\gamma\right)$. In turn, $\beta < 1$ is the subjective discount factor, $1-\alpha$ is the labor-income share, $\vartheta > 0$ measures the intensity of capital adjustment costs, and δ is the depreciation rate of capital. Further, notice that capital adjustment costs incorporate the fact that the economy is growing at rate γ, so the investment-to-capital ratio in the steady state is equal to $\delta+\gamma$.[11]

[11] To see this, write the law of motion of capital in terms of the de-trended variables for capital and investment, i.e. $k_{t-1} \equiv K_{t-1}/A_t$ and $i_t \equiv I_t/A_t$, so that $k_t A_{t+1} \leq \left(1-\delta\right)k_{t-1}A_t + i_t A_t -$

The equilibrium dynamics of the de-trended economy can be summarized by the following two relations (small-case letters denote growth-detrended variables, such that $x_t \equiv \frac{X_t}{A_t}$ for $X_t \in \{C_t, I_t, Y_t\}$ and $k_{t-1} \equiv \frac{K_{t-1}}{A_t}$):

$$1 + r_t = \mathrm{E}_t \left\{ \frac{c_{t+1}}{c_t} \right\} \frac{1+\gamma}{\beta} \frac{1}{\exp(\varepsilon_t)}, \tag{1}$$

$$1 + r_t = \mathrm{E}_t \left\{ \frac{1}{q_t \exp(\varepsilon_t)} \left[\alpha \frac{y_{t+1}}{k_t} + q_{t+1} (1 - \delta - \kappa_{t+1}) \right] \right\}, \tag{2}$$

where $q_t = \left[1 - \vartheta \left(\frac{i_t}{k_{t-1}} - (\delta + \gamma) \right) \right]^{-1}$ is the relative price of capital, and κ_t is a proportion of last-period capital that is lost due to adjustment costs when producing new capital.[12] Equation (1) is the first-order condition for bond holdings, and determines the *supply of loanable funds* in the economy, or the financial *desired savings* schedule. Since $c_t = y_t - i_t$, it follows that, everything else held constat, the supply of loanable funds is upward sloping in the (i_t, r_t)-space. Equation (2) results from combining the first-order conditions for capital and investment, and it denotes *aggregate investment demand*, which in this case is also the economy's *demand for loanable funds*.[13] Replacing the price of capital in that equation, it can be shown that the demand for loanable funds is downward sloping in the (i_t, r_t)-space.

To illustrate the effects of structural factors on $r^\star$, assume that the stochastic shock ε_t is equal to zero. From the supply of loanable funds, it can be seen that the long-run convergence level of the neutral rate, which we denote by $\bar{r}^\star$, is determined by γ and β, such that

$$1 + \bar{r}^\star = \frac{1+\gamma}{\beta}. \tag{3}$$

A lower potential economic growth γ or a more patient agent (i.e. with a higher β), will tend to lower the level at which $r^\star$ will converge in the long run. This is the case because in both scenarios the agent is willing to re-balance his portfolio towards bonds, so the supply

$\frac{\vartheta}{2} \left(\frac{i_t A_t}{k_{t-1} A_t} - (\delta + \gamma) \right)^2 k_{t-1} A_t$. At the steady state, adjustments costs are zero and capital and investment, $\bar{k}$ and $\bar{\imath}$ resp., solve the equation $\bar{k} A_{t+1} = (1 - \delta) \bar{k} A_t + \bar{\imath} A_t$, which yields $\bar{\imath}/\bar{k} = \delta + \gamma$.

[12] It can be shown that $\kappa_t = \vartheta \left(\frac{i_t}{k_{t-1}} - (\delta + \gamma) \right) \left[\frac{1}{2} \left(\frac{i_t}{k_{t-1}} - (\delta + \gamma) \right) - \frac{i_t}{k_{t-1}} \right]$.

[13] Notice that the demand for loanable funds may also include public debt if a government uses this instrument to finance public spending. Kocherlakota (2015) and Winter (2017) argue that the neutral rate may increase by the issuance of new public debt in a non-Ricardian environment (e.g. with financially constraint agents). Such a policy would effectively displace rightwards the demand for loanable funds, while non-Ricardian agents will not increase proportionally their desired savings.

of loanable funds shifts to the right, crossing investment demand at a lower equilibrium real interest rate.[14] Factors affecting γ relate to trend productivity, and population growth. In turn, changes in the demographic composition of a country could be represented, in reduced form, by changes in β. For instance, a higher proportion of the working-age population tends to increase the national savings rate, since this group is typically characterized by a stronger savings profile than others (i.e., the β of the representative agent will tend to be higher; see also Rachel and Smith, 2015). In addition, financial inclusion, by allowing a larger proportion of the population to access financial markets, could also raise β.

In order to understand the effect of transitory factors on $r^\star$, suppose there is a positive innovation of ε_t that raises the returns of bonds. After the shock, there are two alternatives to satisfy equation (1). In the first one, the real interest rate does not change, and current consumption falls below its long-run equilibrium level. In the second one, consumption stays put while the real interest rate drops in the same proportion as the shock increases. This latter scenario defines the neutral rate at higher frequencies or *short-run* $r^\star$. A sufficient drop in the real interest rate ensures that consumption does not deviate from its long-run equilibrium level, keeping the output gap closed. In this simple model, the relationship between short-run $r^\star$ and its long-run convergence level $\bar{r}^\star$ is given by

$$\begin{aligned} 1 + r_t^\star &= \mathrm{E}_t \left\{ \frac{c_{t+1}^\star}{c_t^\star} \right\} \frac{1+\gamma}{\beta} \frac{1}{\exp(\varepsilon_t)} \\ &= \frac{1+\bar{r}^\star}{\exp(\varepsilon_t)}, \end{aligned} \tag{4}$$

where $c_t^\star$ is the detrended-steady-state value for consumption (notice that $c_t^\star = c^\star$ for all t). Figure 1 displays graphically this relationship. In the figure, we assume low frequency changes for potential growth and time preferences, so $\bar{r}_t^\star$ depends on γ_t and β_t. In turn, ε_t displays temporary fluctuations at a high frequency.

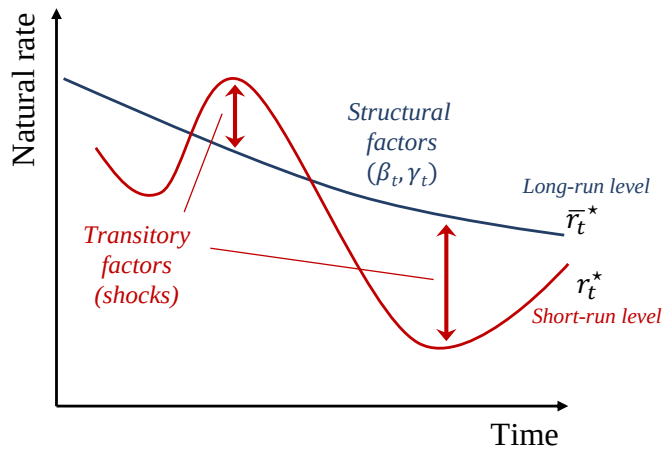
The distinction between short-run $r^\star$ and its long-run convergence level is not only interesting from an academic perspective, but also for policymakers when they instrument mone-

[14]Changes in other deep parameters, like α or δ, have an impact on the long-run level of total savings, investment and output, but leave the neutral rate unchanged. The reason is that changes in these parameters shift both the supply and demand for loanable funds in similar proportions so $\bar{r}^\star$ does not change.

tary policy. The following quote from the Fed's October 2015 FOMC Minutes illustrates clearly this point:

> *"Estimates derived using a variety of empirical models of the U.S. economy [...] indicated that short-run $r^\star$ fell sharply with the onset of the 2008–09 financial crisis and recession, quite likely to negative levels. [...] [On the] discussion of the potential use of $r^\star$ in monetary policy deliberations, policymakers [...] [indicated that real GDP growth] has, on average, exceeded growth of potential GDP, [... which] suggested that the actual level of short-term real interest rates has been below [...] the equilibrium real rate. [...] [S]hort-run $r^\star$ [is expected] to rise as the economic expansion continue[s], but probably only gradually. Moreover, [...] the longer-run normal level to which the nominal federal funds rate might be expected to converge in the absence of further shocks to the economy [...] would likely be lower than was the case in previous decades."*

Figure 1: Short-Run $r^\star$ and its Longer-Run Convergence Level



This quote is informative about the Fed's decision-making process. Not only the Fed expected a lower normal level of its policy instrument in the long run, but it also acknowledges how changes in short-run $r^\star$ affected its current monetary-policy stance.

2.2 Open-Economy Determinants

EMEs might be more vulnerable than AEs to adjustments in international capital markets. A simple way to incorporate this dimension into the previous model is to assume that the representative agent can borrow from abroad. In such a case, the budget constraint of the

agent is:

$$C_\tau + I_\tau + \frac{B_\tau}{(1+r_\tau)\exp(\varepsilon_\tau)} - \frac{\mathcal{E}_\tau B^f_\tau}{1+r^f_\tau} \leq Y_\tau + B_{\tau-1} - \mathcal{E}_\tau B^f_{\tau-1},$$

where B^f is foreign debt, $\mathcal{E}$ is the real exchange rate expressed as domestic goods for a unit of the foreign good, and r^f is the real interest rate at which foreign investors agree to buy home-issued claims. B^f represents capital inflows to the home economy, which is why it appears with an opposite sign to that of home bonds in the budget constraint. In turn, we assume that $r^f_t = r^w_t + \phi\left(\mathcal{E}B^f_t/Y_t\right)$, so r^f is a function of the world interest rate r^w, and the country risk premium ϕ, which is an increasing function of the foreign-debt-to-GDP ratio. Under the assumption of free capital flows, the equilibrium condition of foreign debt yields a typical uncovered interest rate parity condition:

$$r_t = r^w_t + \mathrm{E}_t\Delta_\%\mathcal{E}_{t+1} + \varphi_t,$$

where $\Delta_\%\mathcal{E} \equiv (\mathcal{E}-\mathcal{E}_{-1})/\mathcal{E}_{-1}$ is real depreciation, and $\varphi_t = \phi(\cdot)-\varepsilon_t$. This equation regulates capital flows in the economy. For example, a persistently low r^w may increase the demand for domestic assets. The capital inflow will, in turn, increase the supply of loanable funds in the economy, pushing downwards $r^\star$. In the opposite case, an expected reversal of r^w towards more normal levels could pull capital out of the economy, decreasing the supply of loanable funds.

2.3 Disambiguation of $r^\star$ Definitions

In an RBC model, like the one described above, prices and wages are sufficiently flexible to allow output to always reach its efficient level, and so the short-run real interest rate is always equal to short-run $r^\star$. New Keynesian DSGE models refer to this setup when they define the neutral rate, i.e. the short-run real interest rate that would prevail in the absence of nominal rigidities (see for instance Cúrdia, Ferrero, Ng and Tambalotti, 2015).

New Keynesian DSGE models have advantages and shortcomings when measuring $r^\star$. When taken to the data, these models pin down changes in short-run $r^\star$ to particular transitory shocks. An obvious limitation is that these interpretations are model-dependent. A more important shortcoming is that the long-run convergence level of the neutral rate is assumed

constant, and determined by the researcher. This is the case because the solution methods involved when estimating DSGE models are ill-suited to capture low-frequency changes in structural factors affecting $\bar{r}^\star$.[15] In contrast, more flexible reduced-form models, such as the semi-structural framework of Laubach and Williams (2003), may capture the trend of $\bar{r}^\star$ but may face problems to filter very persistent shocks from the estimation. Other time-series models, such as the TVP-BVAR model of Johannsen and Mertens (2016), suffer from the same mis-specification. For this reason, $r^\star$ estimates from reduced-form models could be understood at best as medium-run estimates.

In this paper, we focus on reduced-form models to estimate $r^\star$ in Mexico. Therefore, we provide medium-run measures of this variable. Some of these estimates are closer to the concept of short-run $r^\star$ since they are clearly affected by high-frequency shocks. In contrast, other estimates remain closer to the concept of $\bar{r}^\star$ to the extent that their estimation methods are able to filter (very) persistent transitory factors.

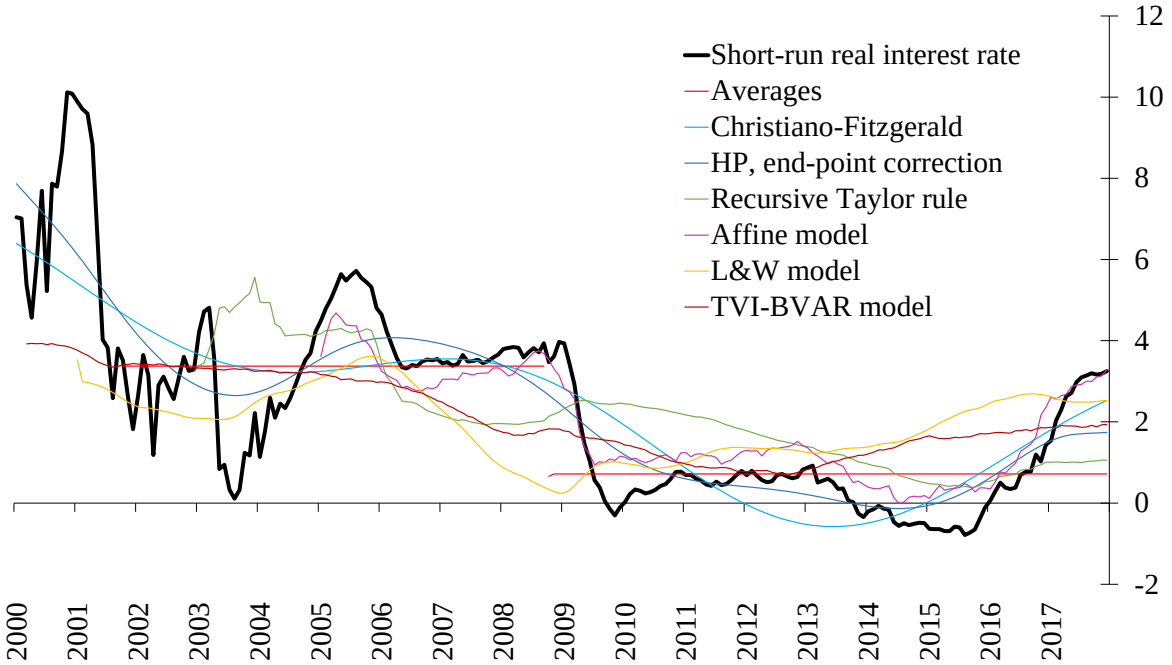
3 Short-run and Medium-run Measures of $r^\star$ in Mexico

We consider six different methodologies to study the evolution of $r^\star$ in the medium run: business-cycle averages, univariate filters, a simple Taylor rule estimated recursively, two affine term-structure models, an adaptation of the Laubach and Williams (2003)'s model for a SOE, and a BVAR model for the U.S. and Mexico with time-varying intercepts. The point estimates of these models are shown in Figure 2. Despite their differences, all methodologies suggest an important reduction in $r^\star$ in the aftermath of the 2008 GFC, and a robust increase since 2014. The degree of uncertainty surrounding these estimates depends on how well each methodology is able to differentiate between transitory factors, and changes in structural factors.

For all estimates, we use the ex-ante short-term real interest rate, measured as the overnight interbank nominal interest rate minus the one-year ahead expectations of headline inflation.

[15] The Bayesian inference used to estimate empirically suitable DSGE models make global solution methods very costly. Therefore, most Bayesian estimations of these models rely on solution methods based on linearization around a fixed steady-state equilibrium.

Figure 2: Summary of Results for Short- and Medium-Run $r^\star$



Source: Own estimates with data from INEGI, Banco de México, Valmer, PiP, the NY Fed, and St. Louis Fed's FRED database.

We extract the latter from Banco de México's survey of private professional forecasters.[16] For multivariate models, we discuss the data used therein when we present each method. The period of study spans at large from January 2000 to December 2017 at a monthly and quarterly frequency, depending on data availability, and the model used.

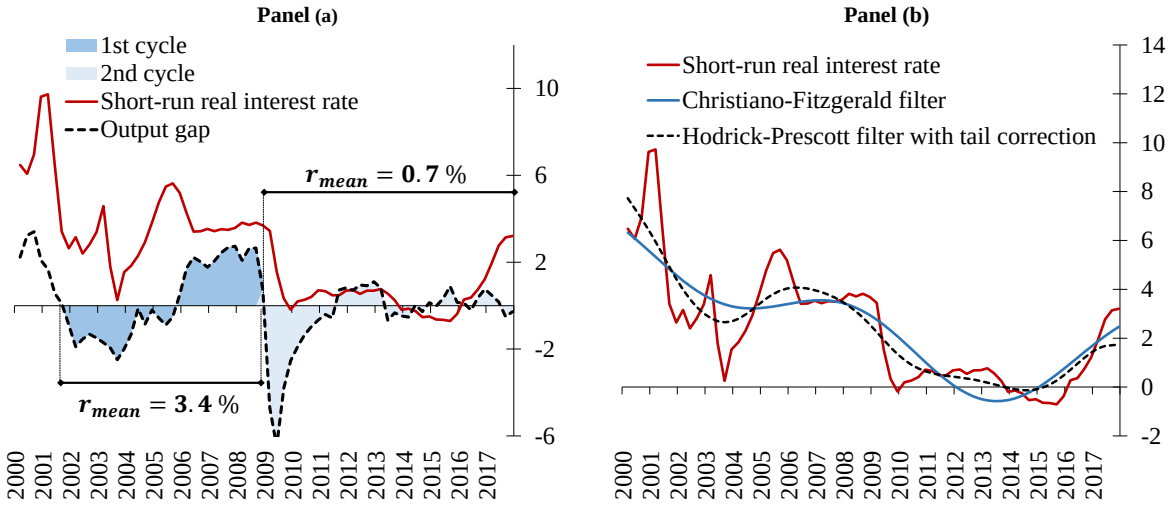
3.1 Averages and Filters

A simple indicator of $r^\star$ in the medium run is the average of the ex-ante real interest rate during a full business cycle, which we define as a completed downturn and upturn of output with respect to its trend. In our sample, we have one full business cycle from 2001 to 2008, while we assume that the second one is still ongoing, starting from 2009 up to present. These

[16]See the *Encuesta sobre las expectativas de los especialistas en economía del sector privado* by Banco de México (http://www.banxico.org.mx/informacion-para-la-prensa/comunicados/resultados-de-encuestas/expectativas-de-los-especialistas/index.html). We use survey-based measures of inflation expectations rather than market-based measures because the latter are model-dependent, so there is uncertainty over these indicators attributed to the modelling choices of the researcher.

cycles are shown in panel (a) of Figure 3 using Banco de Mexico's output gap estimate.[17] The figure also shows the ex-ante real interest rate at a quarterly frequency. According to this methodology, $r^\star$ decreased substantially from cycle to cycle, falling from 3.4% to 0.7%.[18]

Figure 3: Average and Trends of the Short-Run Real Interest Rate



Source: Own estimates with data from INEGI and Banco de México.

An alternative to measure changes in $r^\star$ is through univariate statistical filters. The estimates from the Christiano-Fitzgerald filter and the Hodrick-Prescott filter with tail correction are presented in panel (b) of Figure 3.[19] The smoothed series capture a downward trend in the ex-ante real interest from 2000 to 2008, and a U-shaped pattern from 2009 to 2017.

Because of their simplicity, the averages and smoothed series tell us nothing about the drivers behind the apparent fall in $r^\star$.

3.2 Simple Taylor Rule

Next, we assess the behavior of $r^\star$ by approximating a standard reaction function for the central bank in response to deviations of inflation and output from desired levels. The resulting

[17] See Banco de México (2009), pp. 69, for more details about this estimate.

[18] These numbers do not change much if instead we assume that the second cycle ends in 2014.

[19] For the Christiano and Fitzgerald (2003)'s filter, we use an asymmetric band-pass filter, isolating the cyclical components between 2 and 96 months (which is the usual assumption about the length of a business cycle). In the case of the Hodrick-Prescott filter, we use a smoothing parameter of 14,400 and a tail correction as suggested by St-Amant and van Norden (1997).

interest rate rule should not be taken as the policy directive of the central bank, but rather as a particular lens to interpret the systematic behavior of monetary policy. We estimate the following Taylor rule with interest-rate smoothing:

$$R_t = (1-\rho)[\bar{r}^\star + \bar{\pi} + \delta(\pi_t - \bar{\pi}) + \theta\hat{y}_t] + \rho R_{t-1} + \varepsilon_t, \tag{5}$$

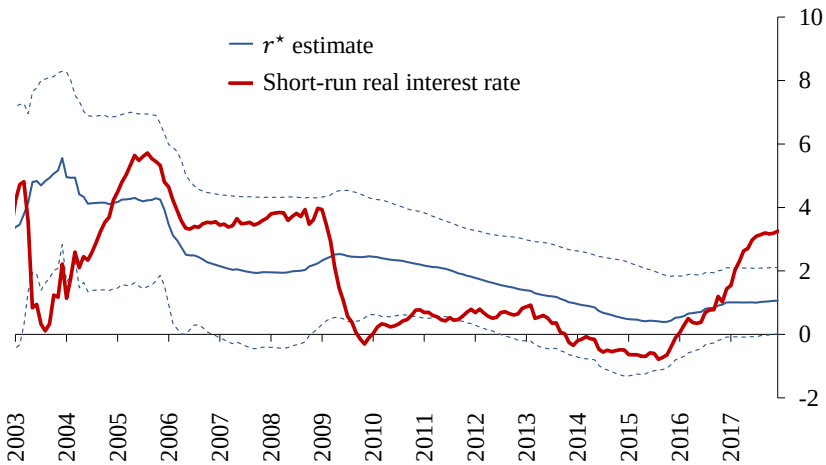
where R is the overnight interbank nominal interest rate, π is inflation, $\hat{y}$ is the output gap, and ε captures any change in R not explained by the rule. In addition, the rule includes a lag of the nominal interest rate to capture gradual adjustments in this variable induced by the central bank. Finally, the neutral rate $\bar{r}^\star$ denotes the level of the real interest rate that should prevail when inflation equals the inflation target $\bar{\pi}$, and the output gap equals zero. Notice that we have added an upper *bar* to this estimator to denote the long-run convergence level of the neutral rate (see equation 3).

In order to capture changes in $\bar{r}^\star$ over our sample period, we estimate equation (5) recursively at a monthly basis. In particular, we first estimate the rule from January 2000 to December 2002 through OLS, and then we re-estimate it adding an observation month-by-month until we reach the end of the sample. We use headline annual inflation measured by the CPI index, and estimate the output gap using Mexico's Global Indicator of Economic Activity (or *IGAE*, by its Spanish acronym), published monthly by INEGI.[20] The results of this exercise are shown in Figure 4. Similar to the evidence from averages and filters, the neutral rate seems to have fallen from 2008 to 2015, partially reverting its trend afterwards.

A question that remains unanswered from this exercise is *why does $\bar{r}^\star$ seem to change over our sample period, in particular when it is supposed to be a long-term indicator of the neutral rate?* The answer to this question may be given by factors omitted in equation (5). We come back to this point in Subsection 5.1, where we revisit the estimation of the Taylor rule for Mexico.

[20]To compute a measure of economic slack from IGAE, we used its percent deviation from trend, which we estimated using the Hodrick-Prescott filter with tail correction.

Figure 4: Taylor Rule Intercept and Short-Run Real Interest Rate



Note: The range corresponds to two standard deviations of the estimates.
Source: Own estimates with data from INEGI and Banco de México.

3.3 Affine Term Structure Models

In this exercise, we estimate two affine term structure models with Mexican data, one similar to Adrian, Crump and Moench (2013) and another to Kim and Wright (2005), henceforth ACM and KW, respectively.[21] The models assume no-arbitrage conditions in financial markets to compute an expected average of the nominal interest rate of a nth-month-maturity bond for a horizon of k periods. The structure of both models is similar, and it can be written in state-space form as

$$X_t = \underbrace{\mu + \phi X_{t-1} + \vartheta_{t+1},}_{\text{TRANSITION EQUATION}} \tag{6}$$

$$i_t^{(n)} = \underbrace{A_n + B_n^{'} X_t,}_{\text{MEASUREMENT EQUATION}} \tag{7}$$

where X is a vector of factors or state variables, $i^{(n)}$ is the nominal interest rate of a bond with maturity of n months, ϑ are white-noise state innovations, ϕ and B_n are coefficient matrices, and μ and A_n are coefficient vectors. The ACM model has five *observable* factors in vector X, each as a proxy for the following yield curve characteristics: (1) level, (2) slope, (3)

[21] Both models deliver market-based expectations of the short-run nominal interest rate. However, the level of these expectations may differ because the ACM model seems to be affected relatively more by transitory factors.

curvature, (4) implied excess returns, and (5) term premia. We estimate this model through OLS, and principal components. In contrast, vector X in the KW model contains only three *latent* factors, which correspond to the first three characteristics of the yield curve listed above. We estimate this model through maximum likelihood, and the Kalman filter. In both models, we use available data from 2004 to 2017 on the yields of government zero-coupon bonds with maturities of 1 month to 120 months.[22]

From each model, we obtain the average expected path of the nominal interest rate of 1-month maturity bonds for horizons running from 1 to 60 months ahead.[23] To obtain a measure of medium-run $r^\star$ in each model, we subtract inflation expectations to the expected nominal rate in each horizon, such that

$$r_t^{\star,m} = \frac{1}{60}\sum_{k=1}^{60}\left[\mathrm{E}_t\left\{i_{t+k}^{(1)}\right\} - \bar{\pi}_{t+k}^{e}\right], \tag{8}$$

where $m = ACM, KW$, and $\bar{\pi}_{t+k}^{e}$ is the expectation of headline inflation at horizon k as estimated by Aguilar-Argáez, Elizondo and Roldán-Peña (2016).[24] Our estimate of $r^\star$ is computed through the average of both trajectories,

$$r_t^\star = \frac{r_t^{\star,ACM} + r_t^{\star,KW}}{2}$$

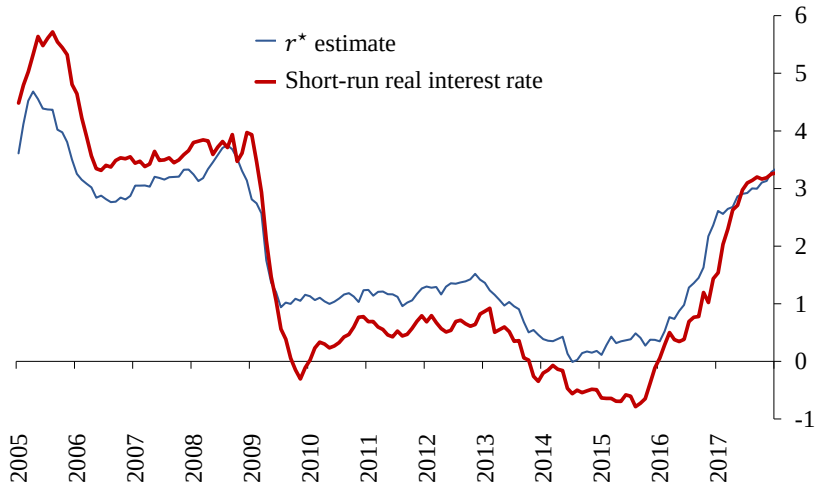
The estimate of $r^\star$ reflects the average level of the short-run real interest rate that investors expect in 1 month to 5 years ahead. Figure 5 displays the term-structure estimate of $r^\star$. Similar to our previous results, the affine models suggest that investors expected the short-run real interest rate to plunge at the onset of the GFC.

[22]More details about these methodologies can be found in Adrian *et al.* (2013), and Kim and Wright (2005). In particular, the coefficients A_n and B_n are estimated recursively and depend on risk parameters. When these parameters are equal to zero, we obtain the risk-free coefficients A_n^{RF} and B_n^{RF}. Using these coefficients, we can compute the average expectation at time t of short-term interest rates over the next k periods, since $E_t[i_{t+1,t+k}^{(1)}] = -(1/n)(A_n^{RF} + B_n^{RF}X_t)$.

[23]The mean squared error between the observed and fitted nominal interest rate from the ACM and KW models is 17 basis points and 1 basis point, respectively, for a 5-year horizon.

[24]The authors follow Adrian and Wu (2009), and Melo-Velandia and Granados-Castro (2010) to estimate inflation expectations implicit in financial instruments, i.e. the long-term break-even inflation. The model uses three factors, of which two are latent, and one is observed inflation. It is worth mentioning that $i_{t+k}^{(1)}$ and $\bar{\pi}_{t+k}^{e}$ are not estimated jointly, since the model presented in this section, and the one in Aguilar-Argáez *et al.* (2016) have different factor structures, parameters, and inputs.

Figure 5: Average Expectation of the Short-Run Real Interest Rate Implicit in Financial Instruments



Source: Own calculations with data from PiP and Valmer.

3.4 Laubach and Williams Model Adapted for a Small Open Economy

In this exercise, we adapt Laubach and Williams (2003) model, henceforth LW, for a small open economy. Similar to the original setting, we include reduced-form representations of aggregate supply and demand, i.e. an IS curve, and a Phillips curve, respectively. In addition, a system of transition equations drives the dynamics of the unobserved variables of the model, such as the neutral rate, and potential output. The model is thus represented by[25]

MEASUREMENT EQUATIONS

$$y_t - y_t^\star = a_y \left(y_{t-1} - y_{t-1}^\star\right) + a_r \left(r_{t-1} - r_{t-1}^\star\right) + a_y \hat{y}_t^{US} + \sum_{\ell=1}^{4} a_{q,\ell} \hat{q}_{t-\ell} + \epsilon_{y,t} \tag{9}$$

$$\pi_t = b_0 + b_\pi \pi_{t-1} + \frac{b_y}{2} \sum_{\ell=1}^{2} \left(y_{t-\ell} - y_{t-\ell}^\star\right) + b_s \left(\Delta s_{t-1} + \pi_{t-1}^{US}\right) + \epsilon_{\pi,t} \tag{10}$$

[25]To select the number of lags in the model, we run regressions of preliminary measures of the neutral rate and potential growth on their own lags. We select the number of lags following the AIC and BIC criteria, and keep those lags with statistically significant coefficients. The preliminary measures are computed using the Hodrick-Prescott and the Christiano-Fitzgerald filters.

TRANSITION EQUATIONS

$$r_t^\star = \gamma_t + z_t \tag{11}$$

$$z_t = z_{t-1} + \epsilon_{z,t} \tag{12}$$

$$y_t^\star = y_{t-1}^\star + \gamma_{t-1} + \epsilon_{y^\star,t} \tag{13}$$

$$\gamma_t = \gamma_{t-1} + \epsilon_{\gamma,t} \tag{14}$$

where y is output, $y^\star$ is potential output, r is the short-run real interest rate, $\hat{y}^{US}$ is a measure of the U.S. output gap, $\hat{q}$ is the percentage deviation of the real exchange rate from trend, π is inflation, Δs is the percent change in the nominal exchange rate, and π^{US} is the U.S. inflation rate. The IS and Phillips curves, expressed in equations (9) and (10), are affected by transitory shocks ϵ_y and ϵ_π, which we assume are white noise. The law of motion of $r^\star$ is given by equation (11), and is a function of potential output growth γ and a time-varying latent component z, which captures model-omitted factors. It is worth noticing that while γ is clearly a structural factor, z can contain both structural and transitory factors. Similar to Laubach and Williams, we assume that z and γ are random walks, while $y^\star$ is a random walk with drift. Similarly, we assume that the state innovations ϵ_z, ϵ_γ, and $\epsilon_{y^\star}$ are white noise.

Our SOE adaptation of the LW model includes the U.S. economic activity, and the real exchange rate in the IS curve to take into account foreign drivers of aggregate demand. In the same vein, the Phillips curve takes on board that home inflation may be affected by the relative purchasing power parity condition, so it includes nominal depreciation, and the inflation of U.S. prices.

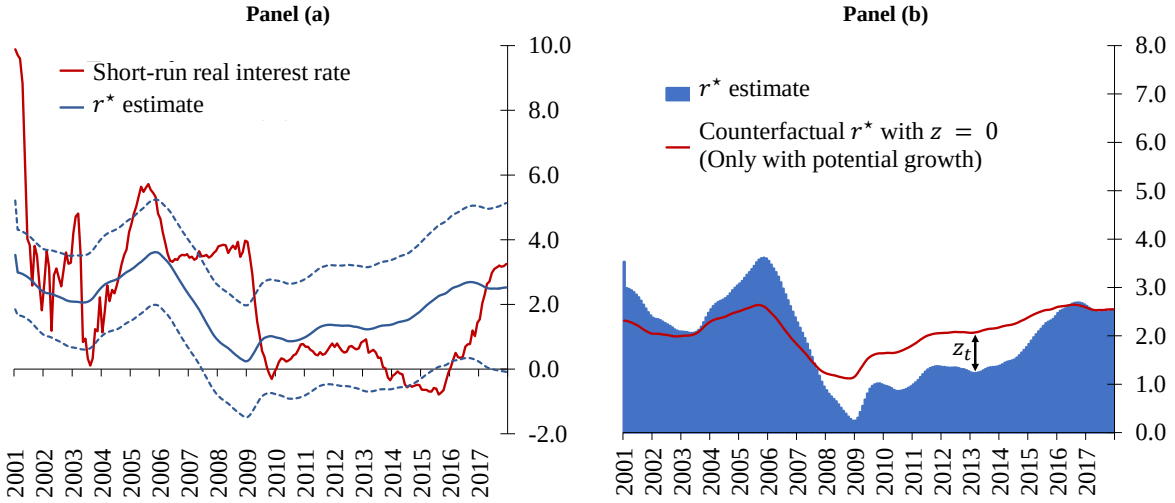
We estimate the model through maximum likelihood, and the Kalman filter using monthly data from January 2001 to December 2017.[26] As a measure of Mexico's output, we use the indicator of Elizondo (2012), who approximates monthly GDP from the IGAE index using a

[26] As it is well known, this estimation procedure suffers from the so-called "pile-up problem", which biases the estimation of the variances of ϵ_γ and ϵ_z towards zero, i.e. σ_γ and σ_z, respectively. For this reason, it is necessary to calibrate these parameters, so that $\lambda_\gamma = \frac{\sigma_\gamma}{\sigma_{y^\star}}$ and $\lambda_z = \frac{\sigma_z}{\sigma_y}$. As a strategy, we chose the couple $(\lambda_\gamma, \lambda_z)$ that minimize the distance between the model's estimated output gap and the output gap estimates by Banco de México (more details on this estimate can be found in Banco de México (2009), pp. 69). Further information about the ML estimation procedure of these parameters can be found in Laubach and Williams (2003), Mesonnier and Renne (2007), Magud and Tsounta (2012), and Pescatori and Turunen (2015), among others.

mixed-frequency Kalman filter (see Section 3.2 for the definition of IGAE). For the inflation and nominal depreciation rates, we use annual rates of headline CPI inflation for both Mexico and the U.S., and the peso-dollar nominal exchange rate. Finally, we use the U.S. industrial production index as a proxy for the U.S. economic activity, and the bilateral real exchange rate between the U.S. and Mexico.[27]

The left panel of Figure 6 shows the $r^\star$ estimate along with the short-run real interest rate. Consistent with the exercises above, this estimate declined sharply at the onset of the GFC, reached a minimum in 2009, and then reverted to higher levels at a very slow pace. The right panel of the figure redraws the estimated $r^\star$ (blue area) together with a counterfactual of this estimate assuming that the latent factor z is equal to zero during all periods (red line). In other words, the red line shows the proportion of the neutral rate that is explained by estimated potential growth alone. As it can be seen, the two measures have diverged persistently from 2006 to 2016, which suggests that model-omitted factors contained in z, and no potential growth may explain the sharp fall in $r^\star$ during the crisis.

Figure 6: Short-Run Real Interest Rate, Neutral Rate and Its Determinants in LW Model



Note: The confidence intervals in the left panel are of 90 percent significance. *Source:* Own estimates made with data from Banco de México, and the St. Louis Fed's FRED database.

[27] We compute gaps from these variables using the Hodrick-Prescott filter with a λ-parameter equal to 14,400 (because of their monthly frequency).

3.5 BVAR with Time-Varying Intercepts

For the final exercise of this section, we consider a joint Bayesian vector autoregression model for Mexico and the U.S. We include a SOE assumption in the model, meaning that U.S. variables influence the Mexican business cycle, but not vice versa (i.e. the U.S. is block exogenous to Mexico). Also, we assume time-variation in the BVAR's intercepts, which allows to capture changes in $r^\star$ for the medium term. We call this model a TVI-BVAR. It is worth noticing that in addition to block exogeneity, we do not impose further restrictions on the dynamics of the model.

Let $X = \begin{bmatrix} X_f' & X_h' \end{bmatrix}'$ be the joint vector of foreign (U.S.) variables X_f, and home (Mexican) variables X_h so that the VAR model reads

$$X_t = C_t + A_1 X_{t-1} + A_2 X_{t-2} + \xi_t, \tag{15}$$

$$C_t = C_{t-1} + \upsilon_t, \tag{16}$$

where C is a vector of time-varying intercepts that follow random walk processes, A_ℓ are conformable matrices of parameters, and ξ and υ are white-noise innovations. Notice that in the absence of shocks, the variables in the system converge to

$$\bar{X}_t = (I - A_1 - A_2)^{-1} \times C_t. \tag{17}$$

We take the element of vector $\bar{X}$ that corresponds to the short-run real interest rate as the estimate of the neutral rate. Notice that vector C may capture both changes in structural factors, and very persistent transitory factors. The main difference between the TVI-BVAR model and the LW model is that the former includes a wider set of foreign and home variables, within a more flexible structure. Therefore, the TVI-BVAR estimate contains more information about the joint dynamics of foreign and home variables that may help to the inference of $r^\star$. Finally, notice that since our object of interest is $\bar{X}$, it is not necessary to include more lags into the model, which is also costly given the short time span of the data.

We estimate the model through Bayesian techniques, and a Kalman smoother following the Carter-Kohn algorithm.[28] Vector X_f contains the U.S. PCE inflation rate, the growth

[28] Kim and Kim (2013) and Kiley (2015) note that the pile-up problem is much less severe with Bayesian methods than with a maximum-likelihood estimation. The priors of the model are set as follows. For the

of the U.S. industrial production index, a shadow measure for the fed funds rate that is not constrained by the zero lower bound (and serves as a proxy for the unconventional monetary policies implemented by the Fed),[29] the average of the 10-year term premia of U.S. bonds as estimated by Kim and Wright (2005) and Adrian *et al.* (2013), the VIX index and the TED spread to control for financial markets volatility, and the price of oil to control for commodity prices. In turn, vector X_h includes core inflation, a monthly approximation of GDP growth using the methodology proposed by Elizondo (2012),[30] the short-term real interest rate, the nominal peso-dollar exchange rate, and the J. P. Morgan's EMBI+ index for Mexico to control for country risk.[31]

Figure 7 shows the results of this methodology. Similar to all previous exercises, the $r^\star$ estimate started declining around 2006, reaching its minimum value by the end of 2012, and reverting its trend afterwards, although not yet to the pre-crisis levels. In the right panel of the figure we have plotted the estimate of potential growth according to the TVI-BVAR. Similar to the results of the LW model for Mexico, estimated potential growth cannot explain the persistent fall in the estimated neutral rate.

3.6 Summary

Table 1 displays the average of the point estimates of $r^\star$ for the periods 2001Q1-2008Q4 and 2009Q1-2017Q4. The table shows that all methodologies find consistent results, namely that the estimates of short-run $r^\star$ in Mexico fell during the GFC, from an average of 3% to around

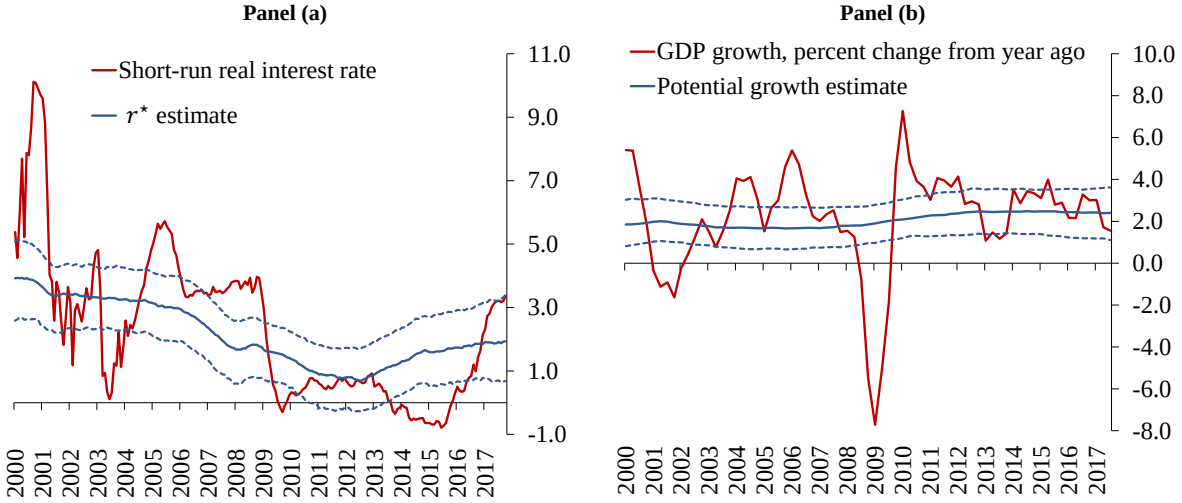
set of time-invariant parameters, we use a Minnesota-style prior for A_ℓ, where we select the hyperparameters that maximize the forecasting performance of the model, while we impose block exogeneity in the upper-right quadrant of these matrices. We opt for maximizing the forecasting performance rather than the marginal likelihood as a selection criterion for the prior since computing the latter turned challenging for the hybrid mode. Further, we use a diffuse inverse Wishart prior for the variance-covariance matrix $\mathrm{E}\{\xi\xi'\}$. For the time-varying intercepts, we use a diffuse prior for the initial state vector C_0, and an adjusted inverse Wishart prior for $\mathrm{E}\{\upsilon\upsilon'\}$ to minimize the distance between the model's estimated output gap and the output gap estimate by Banco de México. Further details about the estimation of the TVI-BVAR are available upon request.

[29] In particular, we use the average of the measures proposed by Lombardi and Zhu (2014), Krippner (2015), and Wu and Xia (2016). These series take negative values during the fed funds rate's zero lower bound period, from 2009M1 to 2015M12, and are equal to the fed funds rate outside that period.

[30] See also Section 3.4.

[31] Given the broad set of home and foreign variables in the model, it soon became cumbersome to estimate it with a fully-fledged model with time-varying parameters and stochastic volatility, like that proposed by Primiceri (2005).

Figure 7: Short-Run Real Rate, Neutral Rate, and Trend Growth from the TVI-BVAR Model



Note: The confidence intervals reflect the 16th and 68th percentiles. *Source:* Own estimates made with data from Banco de México, the NY Fed, St. Louis Fed's FRED database.

1.3% in real terms for the periods indicated. If we translate these results into nominal terms, using the average of the 12-months ahead inflation expectations for each period, we find that the neutral nominal interest rate decreased on average from 7.1% to 5.1%.

Table 1: Summary of Quantitative Results for the Short and Medium Run

Methods	Real neutral rate, $r_t^\star$		Nominal neutral rate, $r_t^\star + \pi_t^e$	
	2001Q1-2008Q4	2009Q1-2017Q4	2001Q1-2008Q4	2009Q1-2017Q4
Averages and trends	3.44	0.74	7.45	4.58
Standard Taylor rule	3.30	1.39	7.31	5.23
Affine model	3.42	1.19	7.43	5.03
Laubach and Williams model	2.26	1.59	6.27	5.43
TVI-BVAR model	2.82	1.35	6.83	5.19
Average	3.05	1.25	7.06	5.09

Note: To compute the nominal neutral rate, we add the average of headline inflation expectations for 12-months ahead to $r^\star$. We extract the former from Banco de México's survey of professional forecasters. According to this survey, for the period 2001Q1-2008Q4, the average inflation expectation was 4.01%, while for the period 2009Q1-2017Q4, it reached 3.84%.

4 Drivers of $r^\star$ in Mexico at business-cycle frequencies

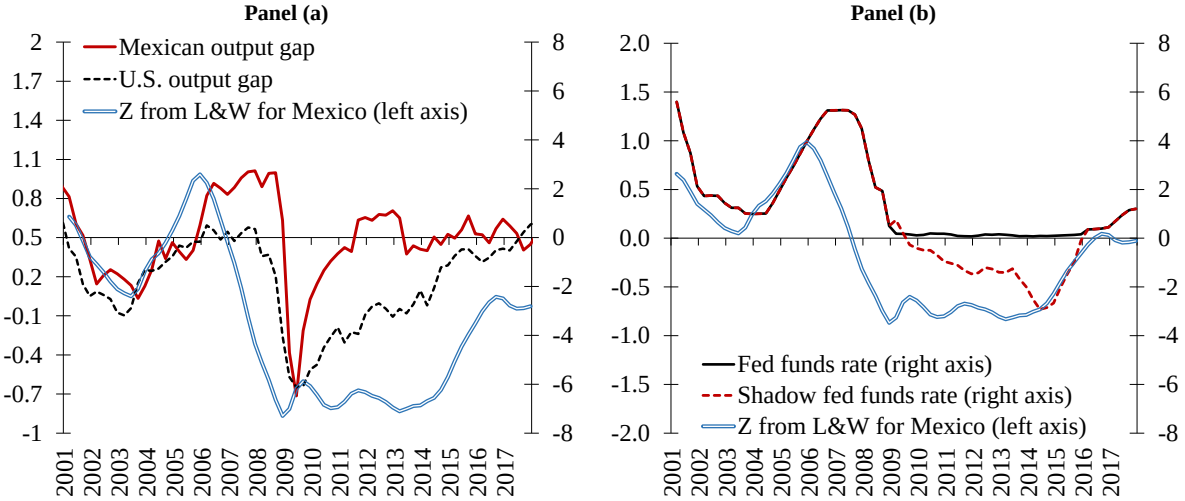
The results from the LW model, and the TVI-BVAR model are consistent with the claim that trend growth alone cannot explain the dynamics of short- and medium-run $r^\star$ in Mexico. In terms of LW's framework, these results imply that other factors contained in the latent variable z are driving the dynamics of estimated $r^\star$ (see equation 11). In this subsection, we analyze some of the transitory factors that may have affected $r^\star$, disentangling them between foreign and domestic.

4.1 Medium-run Drivers through the Lens of the LW Model

In Section 2, we argue that transitory factors, such as recessionary shocks, can depress short-run $r^\star$ for several periods. In the sample studied, we find at least two of these shocks: one relates to the recession that followed the burst of the 2001's *dot-com* bubble, and the other is the global financial crisis. Panel (a) in Figure 8 shows that these events slowed down economic activity in both the U.S. and Mexico. The red line depicts the Mexican output gap (as computed by Banco de México), the black dashed line is the U.S. output gap (as computed by the CBO), and the blue line is the latent variable z extracted from the LW model (whose scale is on the left vertical axis). From 2001 to 2008, z seems to co-move with the output gaps of both countries, which suggests that aggregate demand shocks could explain some of the variation in the estimated $r^\star$ from the LW model. In particular, z reaches its minimum at the same time than the output gaps in 2003, and dips again during the GFC (although it starts falling somewhat earlier, since 2006). A low aggregate demand at home pushes downwards short-run $r^\star$ because investment demand falls, and some households increase desired savings. For a small open economy, such as Mexico, aggregate demand conditions abroad also matter because they influence exports dynamics, financial flows, and economic activity at home.

From 2009 to 2015, z seems to capture other factors besides those related with economic slack. Panel (b) in Figure 8 suggests that the Fed's unconventional monetary policies (or UMPs) implemented since 2009 may have affected z's dynamics. In the picture, the black line is the observed fed funds rate, the red dashed line is Wu and Xia (2016)'s shadow fed

Figure 8: z from the LW Model, Output Gap in Mexico and the U.S., and the Shadow Fed Funds Rate



Note: Own calculations with results from the LW model for Mexico, the shadow fed funds rate of Wu and Xia (2016), and the CBO.

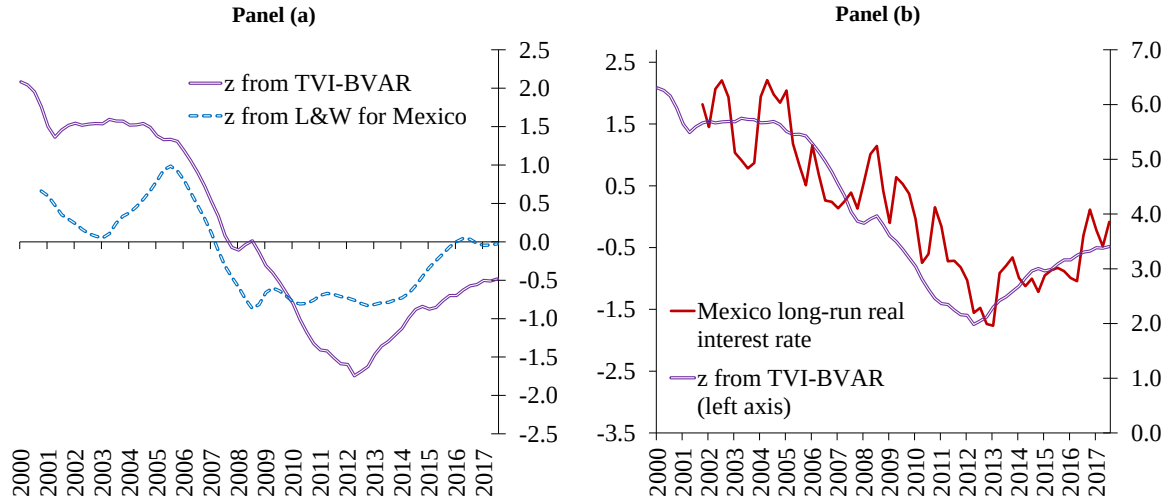
funds rate, a proxy for the degree of monetary policy accommodation of UMPs,[32] and the blue line is the z from the LW model, whose scale is again on the left vertical axis. Variable z remains at low levels from 2009 until 2014, moment in which it starts rising. Remarkably, z lands at 0 precisely a quarter after the FOMC started to normalize the fed funds rate at the end of 2015. The dynamics of z from the LW model allow us to posit the following hypothesis: UMPs might have pressured downwards short-run $r^\star$ in Mexico through its influence on capital flows. Accordingly, several investors might have re-balanced their portfolios away from the U.S. and other AEs with low returns, to favor relatively safe EMEs with higher returns, such as Mexico. If so, capital inflows would have increased the supply of loanable funds in the domestic market, pushing $r^\star$ downwards. We explore in detail this hypothesis next.

[32]The negative values in the shadow rate aim to measure the degree of monetary policy accommodation achieved by the Fed's UMPs. A larger negative value in absolute terms implies a larger monetary policy accommodation.

4.2 Medium-run Drivers through the Lens of the TVI-BVAR Model

Kiley (2015) criticizes the LW model because it might omit important "demand shifters" in the IS curve, i.e. persistent transitory factors in our framework. This omission may affect $r^\star$ estimates in important ways. For instance, Kiley shows that including credit spreads into the IS curve makes the LW's $r^\star$ estimate for the U.S. more stable. It follows that if persistent transitory factors are not included in the IS curve, the latent variable z, and the estimate of $r^\star$ may implicitly capture them. In this context, the TVI-BVAR model, which covers a broader set of variables, seems to deliver a more stable estimate of $r^\star$ compared to the LW estimate, as can be seen in Figures 6 and 7. Despite its greater stability, the TVI-BVAR's $r^\star$ estimate still persistently declines around the GFC, and has not reverted to the pre-crisis levels ever since.

Figure 9: z from L&W Model and TVI-BVAR, and Mexico's Long-Run Real Interest Rate

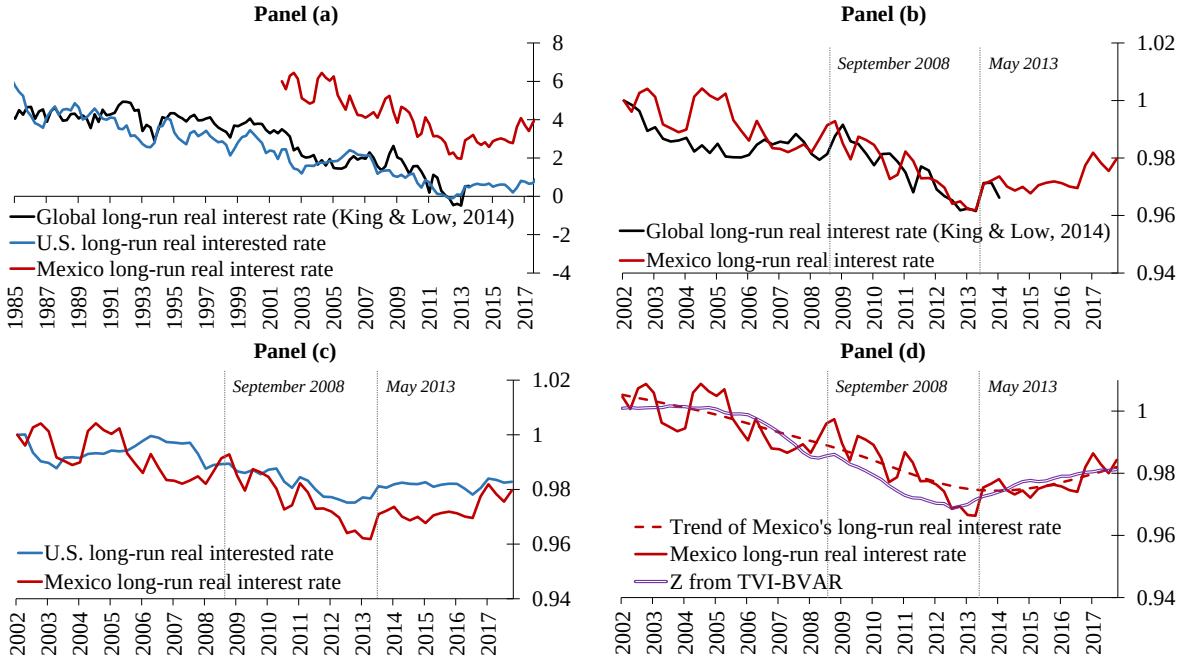


Note: Own calculations with results from the LW model and the TVI-BVAR. For Mexico's long-term real interest rate, we use the 10-year coupon bonds rate minus 10-year inflation expectations from Aguilar-Argáez *et al.* (2016). The nominal interest rate of coupon bonds was obtained from PiP for the period 2003Q3 to 2016Q2. To extrapolate the coupon rate for 2002Q1 to 2003Q2, we ran a regression between the coupon rate and the zero-coupon rate from Valmer for the period 2003M7-2016M6, then we use the zero-coupon rate to extrapolate the coupon-rate for the missing sample. It is noteworthy that the correlation between the coupon rate and the zero-coupon rate is 99.5 percent.

Panel (a) in Figure 9 compares the LW model's z (blue dashed line) with a similar measure computed from the TVI-BVAR model (purple line, generated as the difference between the estimates of $r^\star$ and potential growth). Two features are noteworthy. First, although the

TVI-BVAR delivers a z with fewer fluctuations than the LW model, it shows a more pronounced downward trend. And second, as shown in Panel (b) in the figure, the trend of the TVI-BVAR's z seems to be correlated with the trend observed in the long-run real interest rate of Mexico (red line, computed as the difference between the 10-year nominal yield of government bonds minus a measure of 10-year inflation expectations).[33] Therefore, the TVI-BVAR model seems to succeed at cleaning the estimate of $r^\star$ from mildly persistent shocks, such as the 2001's *dot-com* recession, but it is still sensitive to other persistent transitory factors, such as the GFC, and the array of UMPs that followed.

Figure 10: Long-Run Real Interest Rates and the z from the TVI-BVAR



Note: Source: Own calculations with data from King and Low (2014), the Federal Reserve, PiP, Valmer and Aguilar-Argáez *et al.* (2016).

Panel (a) in Figure 10 compares the long-run real interest rates of the world (as computed by King and Low, 2014), the U.S., and Mexico, whose available data start in 2002. The downward path in the global long-run real interest rate is an issue widely discussed in academic and policymaking forums. We discuss the drivers behind this trend in more detail in Annex C. For the time being, it is important to notice that Mexico does not seem to be

[33]For further details about this measure of inflation expectations, see Section 3.3 and footnote 24.

insulated from such a path. Panel (b) in Figure 10 explores similarities in the trend of the long-run real interest rates of Mexico and the world. The series are normalized to be equal to one in 2002Q1. Up to 2013Q4, which is the quarter where King and Low's series stops, the two rates seem to share a common trend, a fact that is especially evident after September 2008, at the onset of the GFC. In turn, in panel (c) the same comparison is done between the Mexican and the U.S. rates. Although the trends are similar, the Mexican rate decreased faster than the U.S. rate between September 2008 and May 2013. This period corresponds to the rapid expansion of the UMPs implemented by the Fed and other central banks in AEs.[34]

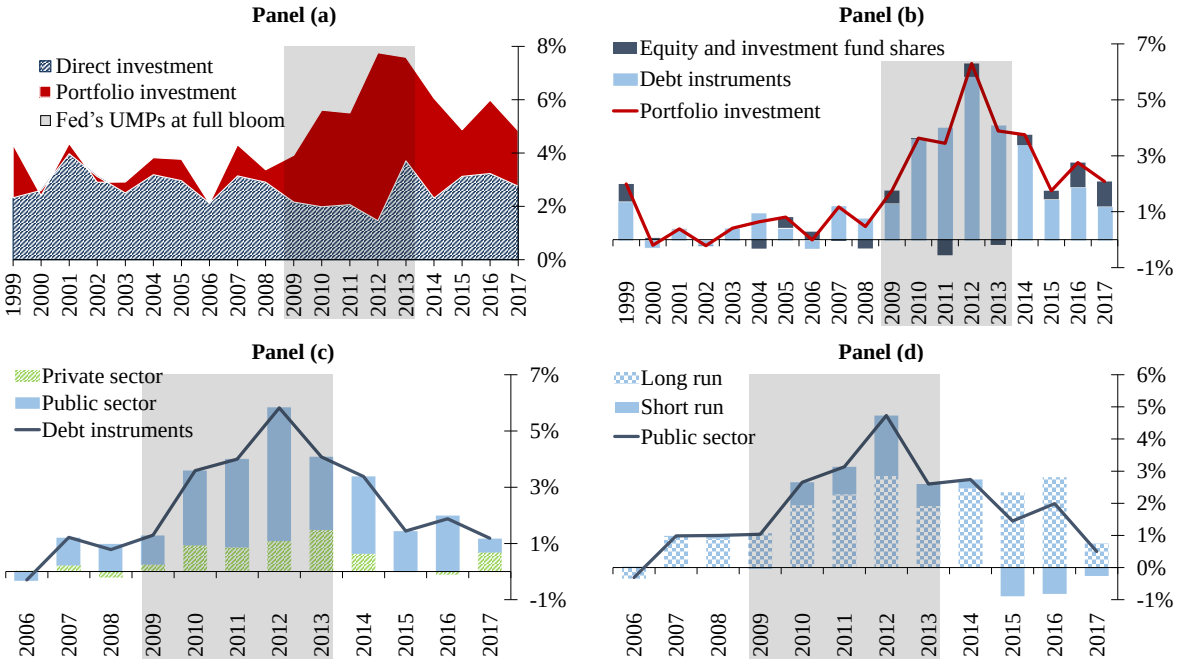
The faster decrease of the Mexican long-term rate as compared to its U.S. counterpart can be interpreted as an indication that financial capitals flowed towards the higher-yield Mexican market. These inflows would have therefore pushed downwards $r^\star$ in Mexico by increasing the supply of loanable funds in the economy. Interestingly, panel (d) in Figure 10 provides evidence that seems to support this hypothesis. The picture compares the Mexican long-run real interest rate with the constructed variable z from the TVI-BVAR estimation; this time, though, we have included the trend of the Mexican long-run real interest rate as computed by an HP filter (dashed red line), and we have normalized the series again to be equal to one in 2002Q1. Remarkably, the trend of the Mexican long-run rate, and the TVI-BVAR's z are almost identical from 2002 until September 2008. From that month onwards, z falls below the long-run real interest rate, and the gap does not seem to close until the second half of 2013, during the *taper tantrum*. Recall that it is precisely between 2008 and 2013 that the TVI-BVAR model recovers the lowest estimates of $r^\star$(see Figure 7).

Capital inflows may explain the accelerating reduction in both the Mexican long-run real interest rate and the TVI-BVAR's $r^\star$ estimate between 2009 and 2013. If that is the case, such inflows would have affected both ends of the yield curve, but with a more drastic impact at its short end. Figures 11 and 12 provide strong support to this hypothesis. The figures mark in gray the period where the Fed's UMPs were in full expansion. Panel (a) in Figure 11 shows that between 2009 and 2013 gross capital inflows spiked in Mexico. These flows

[34]In May 2013, the Fed for the first time mentioned the tapering of their QE programs, an event known as the *taper tantrum*.

directed towards portfolio investment, which peaked at 6.3% of GDP in 2012 from an average of 0.5% of GPD between 1999 and 2008. From the *taper tantrum* in mid-2013 until the end of 2017, portfolio flows stabilized around 2% of GDP. Panel (b) in the figure shows that almost the entire portfolio inflows went into domestic debt instruments, in particular to the public sector, as shown in panel (c). Finally, panel (d) decomposes the flow towards public instruments into short-run debt (with maturity of one year or less), and long-term debt (with maturity larger than a year). Notably, there is an increase in the purchases of public debt by foreign investors at both ends of the yield curve in 2010. Part of this increase may be due to the inclusion of Mexican peso-denominated debt to Citigroup's World Government Bond Index (WGBI) in October of that year.[35] This index is used as a benchmark by institutional investors who aim to buy highly-rated long-term debt. Interestingly, from 2013 onwards, there is a strong reversal in short-term debt purchases by non-residents, whereas long-term debt purchases remain relatively steady.

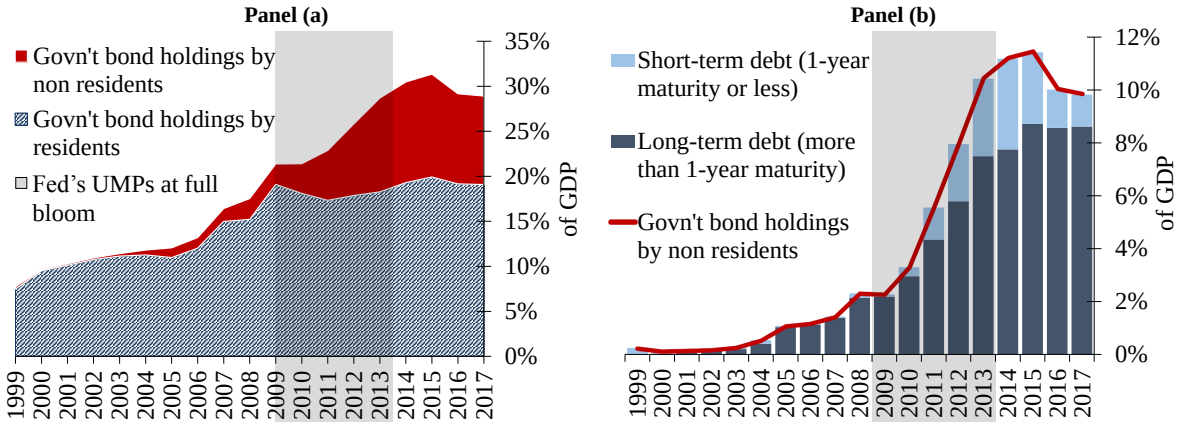
Figure 11: Capital Inflows



Note: The data are presented at an annual frequency and as percentage of GDP. *Sources:* IMF, INEGI and Banco de México.

[35]The index includes fixed-rate bonds with remaining maturity of one year or longer, from 22 countries with highly developed and liquid markets.

Figure 12: Composition of Government Bonds



Note: The data is presented at an annual frequency, and as percentage of GDP. *Sources:* INEGI and Banco de México.

Figure 12 presents two different decompositions of the stock of holdings of public debt. Panel (a) displays the holdings by residents and non-residents as a proportion of GDP. Notably, between 2009 and 2014, holdings by non-residents increased substantially, from 2.3% of GDP in 2009 to 11.2% of GDP by the end of 2014. Panel (b) shows that non-residents raised their holdings of both long- and short-term debt, but it was the latter that had a notorious momentum during the expansion of the Fed's UMPs. Government short-term debt holdings by non-residents went from 0.1% of GDP in 2009 to 3.4% of GDP in 2014. In the last three years of the sample, the short-term debt momentum by non-residents receded, and by the end of 2017 their holdings fell to just 1.2% of GDP. In turn, panel (b) exhibits that non-residents holdings of government long-term debt also increased importantly during the same period, rising from 2.2% of GDP in 2009 to about 9% of GDP by 2014. These holdings have stabilized at that level for the last three years of the sample.

In sum, the data are consistent with the hypothesis that capital inflows have increased the supply of loanable funds in the country in recent years, especially from 2009 to 2013. In particular, the rise and fall of the short-term debt momentum by non-residents provides a rationale on why all models estimates of $r^\star$ plunged during the GFC, and reverted afterwards.

5 Long-Run Convergence Level of $r^\star$ in Mexico

In this section, we present three different quantitative methods that estimate the convergence level of the neutral rate, $\bar{r}^\star$. We first estimate an augmented Taylor rule that controls for the Fed's UMPs. Second, we apply an open-economy RBC model to Mexico to get a long-run average of the equilibrium real interest rate. And third, we compute the implicit long-term expectation of the short-run policy rate that emerges from an affine term-structure model. Finally, we present the summary of an heuristic analysis of structural factors affecting $\bar{r}^\star$.

5.1 Taylor Rule Revisited: Augmented Version

In Section 3.2, we present a simple Taylor rule whose $\bar{r}^\star$ estimate falls between 2008 and 2014. Moreover, in Section 4.2 we argue that the Fed's UMPs affected the Mexican neutral rate during this period through their effects on capital flows. In this context, Taylor and Wieland (2016) argue that omitting important information in the estimating reaction function of the central bank may result in a noisy estimate of the neutral rate, one that misleadingly absorbs the fluctuations of the omitted factors. Motivated by the latter, we include an indicator of the Fed's UMPs as an additional regressor in an augmented Taylor rule, which now reads

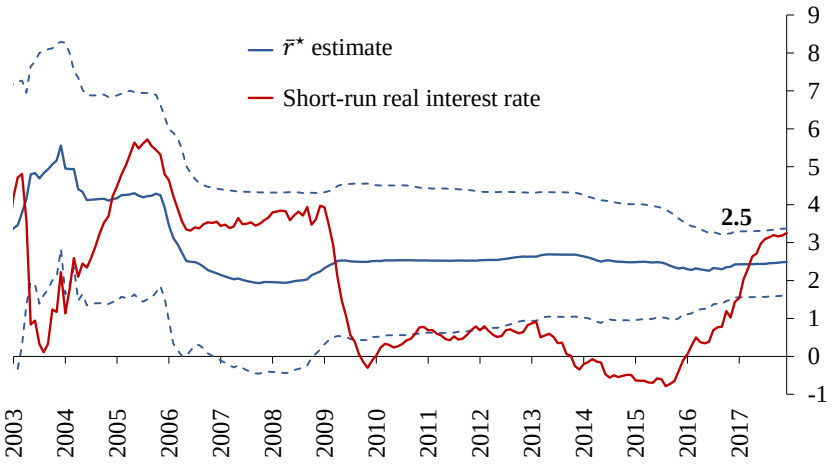
$$R_t = (1-\rho)[\bar{r}^\star + \bar{\pi} + \gamma\left(\mathbb{1} \times R_t^{US,shadow}\right) + \beta(\pi_t - \bar{\pi}) + \theta\hat{y}_t] + \rho R_{t-1} + \varepsilon_t, \qquad (18)$$

where $R_t^{US,shadow}$ is the shadow fed funds rate of Wu and Xia (2016), and the indicator variable $\mathbb{1}$ takes the value of zero when $R_t^{US,shadow}$ is positive, and one when $R_t^{US,shadow}$ is negative (i.e. from July 2009 to December 2015). We include only the information of the shadow rate during the ZLB period as a proxy for the Fed's UMPs. Therefore, we explicitly assume that these policies capture a very persistent transitory factor, and not a structural factor.[36] The augmented Taylor rule is estimated recursively from 2002 onwards, akin to the estimation presented in Section 3.2.

Figure 13 suggests that the omitted-variable critique of Taylor and Wieland (2016) is important for the case of Mexico. Controlling for the Fed's UMPs yields a relatively steady

[36]Note that if the long-run value of the Fed's UMPs is not zero, the interpretation of the intercept as an estimator of $\bar{r}^\star$ in the Taylor rule changes.

Figure 13: Augmented Taylor Rule



Note: The confidence intervals are of 90 percent significance. *Source:* Own estimates made with data from Banco de México.

estimate of $\bar{r}^{\star}$, close to 2.5% since 2008. The latter translates into a neutral nominal policy rate of 5.5%, if we add Banco de México's inflation target of 3%.[37]

5.2 Open-Economy RBC Model

As an alternative to measure $\bar{r}^{\star}$, we use a neoclassical growth model for a small open economy. We follow the business-cycle model of Lama (2011) who, similar to Chari, Kehoe and McGrattan (2007), includes four sources of macroeconomic fluctuations into the model: an efficiency wedge (or TFP), a labor wedge, a capital wedge, and a bond wedge. These wedges allow the model to perfectly match the fluctuations of output, consumption, investment, and hours worked. Our estimate of $\bar{r}^{\star}$ is the average over a long period of time of the equilibrium real rate of capital returns, r^k, which is a model-consistent measure of the actual macro

[37] We have considered alternative measures of the shadow fed funds rate, as those described in footnote 29. The results remain quantitatively similar. Also, in a robustness exercise, we estimate the augmented Taylor rule through a rolling-window OLS regression. However, the lack of variability in the Mexican nominal interest rate in certain parts of the sample generates a quite volatile estimation. In a different robustness exercise, we estimate the monetary rule through a state-space model via maximum likelihood, and the Kalman filter. The measurement equation is given by the Taylor rule, and the state equation by a random walk process for the neutral rate. In this case, the estimate of the neutral rate is very similar to the recursive regression. The results of these robustness exercises are available upon request.

dynamics. We consider a long-period average of r^k since the aforementioned wedges are reduced-form distortions that may capture both structural and transitory factors.[38]

Lama (2011)'s model contains a competitive firm, and a representative household with an increasing number of members. The firm chooses labor l_t, and capital services k_t to maximize profits:

$$\max_{l_t,k_t} \quad A_t k_t^{\alpha} \left((1+\gamma)^t l_t \right)^{1-\alpha} - w_t l_t - z_t k_t,$$

where A_t is TFP, w_t is the real wage, z_t is the rental rate of capital, α is the share of capital income on GDP, and γ is the growth rate of technological progress. A representative household chooses consumption per capita c_t, international debt b_{t+1}, investment i_t, the next period's capital stock, and the labor supply, in order to maximize its expected discounted utility, subject to a budget constraint, the law of motion for capital accumulation, and a supply of funds for international borrowing:

$$\max_{c_t,b_{t+1},l_t} E_0 \left\{ \sum_{t=0}^{\infty} N_t \beta^t \left[\log c_t + \psi \log (1-l_t) \right] \right\},$$

subject to

$$\begin{aligned}
(1+n)\, b_{t+1} + c_t + i_t &\leq (1-\tau_{lt})\, w_t l_t + (1-\tau_{kt})\, z_t k_t \\
&\quad + (1+\tau_{bt}) \left(1 + r_t^W\right) b_t + \Upsilon_t, \\
(1+n)\, k_{t+1} &\leq (1-\delta)\, k_t + i_t - \phi \left(\frac{i_t}{k_t} \right) k_t, \\
1 + r_t^W &= \left(1 + r^W\right) \left(\frac{b_t}{b^W} \right)^{\upsilon},
\end{aligned}$$

where β is the subjective discount factor, ψ is a normalizing constant, n is the growth rate of population, δ is the depreciation rate of capital, and $\upsilon > 0$ is the elasticity of the supply of international borrowing. In turn, N_t is the size of the population, r_t^W is the world real interest

[38]Recently, Caballero, Farhi and Gourinchas (2017) notice that for the case of the U.S. there is a growing divergence between the return on productive capital, and the return of safe assets. For the case of Mexico, it is not clear that such divergence is as secular as in the U.S. Nonetheless, we bear in mind that even a long-period average of r^k might be a poor approximation of $\bar{r}^{\star}$. We decided to keep the neoclassic analysis for two reasons. First, Dorich, Reza and Sarker (2017) perform a similar exercise for Canada, and notice that potential output growth plays a prominent role in the determination of $\bar{r}^{\star}$. And second, there are not many methods available in the literature to estimate $\bar{r}^{\star}$.

rate, Υ_t represent government transfers, and $\phi\left(i_t/k_t\right) = \vartheta/2 \times \left(i_t/k_t - \tilde{\delta}\right)^2$ measures capital adjustment costs, where $\tilde{\delta} = \delta + n + \gamma + n\gamma$. Finally, $(1-\tau_{lt})$ is the labor wedge, $(1-\tau_{kt})$ is the capital wedge, and $(1+\tau_{bt})$ is the bond wedge. These wedges enter the model as taxes, and multiply each price in the economy to reflect market distortions in the otherwise efficient-allocation conditions. The supply of international funds is upward sloping in order to ensure that the model economy does not display a unit root (see Schmitt-Grohé and Uribe, 2003).

The wedges evolve according to $x_t = x^{1-\rho_x} x_{t-1}^{\rho_x} \exp\left(\varepsilon_{xt}\right)$ for $x \in \{A, 1-\tau_l, 1-\tau_k, 1+\tau_b\}$, where $\varepsilon_{xt} \sim N\left(0, \sigma_x\right)$ are normally-distributed, white-noise innovations. The dynamics of the de-trended economy are given by the law of motion for capital, the wedges processes, and the following market-clearing conditions:

$$\tilde{y}_t - \tilde{c}_t - \tilde{\imath}_t = (1+n)(1+\gamma)\tilde{b}_{t+1} - \left(1 + r_t^W\right)\tilde{b}_t, \tag{19}$$

$$\psi \frac{\tilde{c}_t}{1-l_t} = (1-\tau_{lt})(1-\alpha)\frac{\tilde{y}_t}{l_t}, \tag{20}$$

$$\frac{1}{\tilde{c}_t} = \frac{\beta}{1+\gamma} E_t \left\{ \frac{1}{\tilde{c}_{t+1}} (1+\tau_{bt+1}) \left(1 + r_{t+1}^W\right) \right\} \tag{21}$$

$$\frac{1}{\tilde{c}_t} = \frac{\beta}{1+\gamma} E_t \left\{ \frac{1}{\tilde{c}_{t+1}} \left(1 + r_{t+1}^k\right) \right\}, \tag{22}$$

$$1 + r_t^k \equiv \frac{1}{q_{t-1}} \left[(1-\tau_{kt})\alpha \frac{\tilde{y}_t}{\tilde{k}_t} + q_t \left(1 - \delta - \phi\left(\frac{\tilde{\imath}_t}{\tilde{k}_t}\right) + \phi'\left(\frac{\tilde{\imath}_t}{\tilde{k}_t}\right)\frac{\tilde{\imath}_t}{\tilde{k}_t} \right) \right], \tag{23}$$

where $\tilde{x}_t$ denotes a detrended variable, such that $\tilde{x}_t \equiv x_t/(1+\gamma)^t$ for $x \in \{y, k, i, c\}$, and $q_t = \left(1 - \phi'\left(\tilde{\imath}_t/\tilde{k}_t\right)\right)^{-1}$ is Tobin's Q. Equation (19) denotes the economy's resource constraint; equations (20)-(22) are the household's first-order conditions; and equation (23) describes the evolution of the real rate of capital returns. The estimate of $\bar{r}^\star$ is given by

$$\bar{r}^\star = \frac{1}{T}\sum_{t=1}^{T} r_t^k.$$

Similar to Lama (2011), we calibrate the deep parameters of the model, while we estimate the parameters governing the dynamics of the wedges through maximum likelihood, using time series for output, consumption, investment, and hours worked (see Table 2). In contrast with Lama, we use quarterly frequency data instead of annual, and we focus on the recent

Table 2: Calibrating and Estimating Parameters for the Neoclassical Model

Calibration			Estimation		
Parameter	Symbol	Value	*Wedge*	ρ_x	σ_x
Population growth	n	1.84% *app*	TFP	0.99 (0.002)	0.013 (0.001)
Exogenous tech. progress	γ	0.86% *app*	$1-\tau_{lt}$	0.99 (0.003)	0.016 (0.002)
Depreciation rate	δ	5.00% *app*	$1-\tau_{kt}$	0.70 (0.229)	0.151 (0.163)
Discount factor	β	0.99	$1+\tau_{bt}$	0.95 (0.050)	4×10^{-4} (2×10^{-4})
Leisure weight	ψ	2.80			
Capital adjustment costs	ϑ	12.98			
Labor income share	$1-\alpha$	0.30			
International real rate	r^W	4.00% *app*			
Supply of international funds	υ	1×10^{-4}			

Note: The acronym *app* stands for annual percent points. For the estimated parameters, the numbers in parenthesis are the standard deviation of the estimated value.

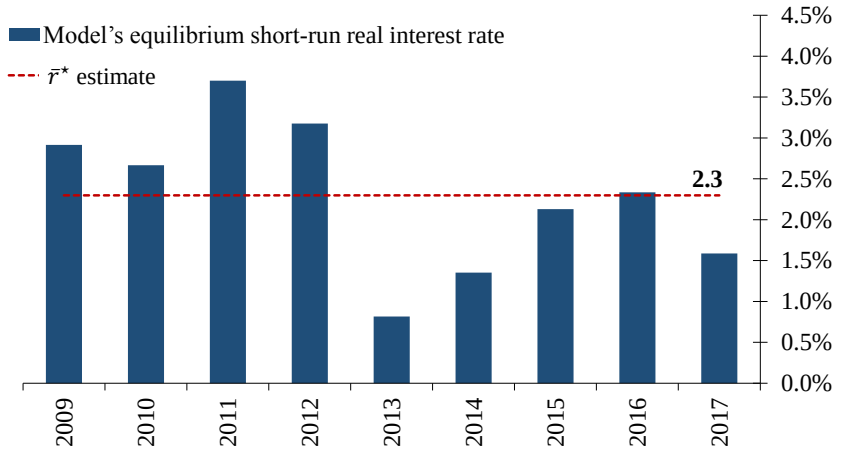
period, from 2006Q1 to 2017Q4 (we start late our sample because quarterly data for hours worked is only available from 2006).[39] The latter implies that we need to adjust certain calibrating parameters for the quarterly frequency and the different time period. We assume that potential growth is 2.7% in annual terms, which is consistent with the estimation results from the LW and TVI-BVAR models in Sections 3.4 and 3.5.[40] Since population growth averaged 1.84% annually during this period, it turns out that the exogenous technological progress must equal 0.86% at a annual basis. The international real rate r^W equals 4%, similar to Lama (2011). Given these numbers, we adjusted the discount factor β so that it satisfies equation (21) at the steady state. The leisure parameter ψ is set to match the average of hours worked per day in Mexico, which equals 41.23 hours per week for the time period studied. For the rest of parameters, we followed closely the strategy of Lama. We used standard values for the depreciation rate δ, the labor income share $1-\alpha$ for a Latin American

[39] Lama (2011) use similar data for Mexico for the period 1991 to 2006, at an annual basis.

[40] We have also performed the exercise assuming a more conservative potential growth, i.e. 2.4% instead of 2.7%. The results in terms of the estimated $\bar{r}^\star$ are quite similar.

economy, and the inverse of the elasticity of supply of international funds υ (further details can be found in Lama, 2011). Similar to Bernanke, Gertler and Gilchrist (1999), the value for the adjustment cost parameter ϑ is consistent with a price elasticity of capital with respect to the investment-capital ratio η equal to 0.25. Using Tobin's Q to compute this elasticity, we impose that at the steady state it must hold that $\eta = \vartheta\tilde{\delta}$, and solve this expression to find ϑ.

Figure 14: Real Rate of Capital Returns and Long-Run $r^\star$



Source: Own estimates made with data from Banco de México and INEGI.

Figure 14 shows that the estimated of $\bar{r}^\star$ equals 2.3% from 2009 to 2017, which corresponds to the time period of the second business cycle considered in Section 3.1. This estimate is located in an one-standard-deviation confidence interval of [1.2%, 3.2%]. Finally, the neutral nominal policy rate becomes 5.3%, if we add Banco de México's 3% inflation target to the above estimate, while the interval becomes [4.2%, 6.2%]. These results are similar to those obtained from the estimation of the augmented Taylor rule.

5.3 Financial Markets

To compute an alternative estimate of $\bar{r}^\star$, we use the long-run expectation of the short-run nominal interest rate that is derived from financial-markets information. We retrieve this expectation from an affine model similar in structure to the KW model (see Section 3.3).[41]

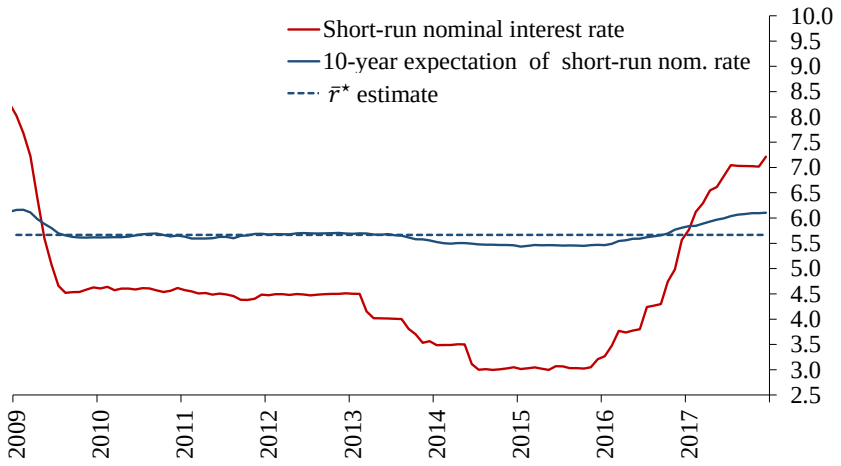
[41] We decided to consider only the estimate of $\bar{r}^\star$ coming from the KW model because such a model seems to filter better the effects of transitory factors on the neutral rate long-run estimates. The KW model at long

We use a horizon of 10 years since in that time period it is quite likely that even the most persistent transitory factors would have faded away. In particular, the estimate of $\bar{r}^{\star}$ is given by

$$\bar{r}^{\star} = \mathrm{E}_t \left\{ i_{t+10}^{(1)} \right\} - \bar{\pi}, \tag{24}$$

where $\bar{\pi}$ is the inflation target.

Figure 15: Long-run Expectation of the Short-Run Nominal Interest Rate Implicit in Financial Instruments



Source: Own calculations with data of PiP and Valmer.

Figure 15 shows that the long-run expectation of the short-run nominal interest rate averaged a level of 5.7% from 2009 to 2017, the time period that corresponds to the second business cycle studied in Section 3.1. During this period, the minimum value of the long-run expectation of the short-run nominal interest rate is 5.4%, while the maximum value is 6.1%. In real terms, $\bar{r}^{\star}$ becomes 2.7% if we subtract Banco de México's 3% inflation target, while the variation interval translates to [2.4%, 3.1%]. These results are again consistent with those from previous methods.

horizons seems to capture the trend of $r^{\star}$, which is the object we look for. In contrast, the ACM model is more susceptible to transitory shocks, since even at long-term horizons the neutral rate estimate displays a lot of fluctuations.

5.4 Summary of Quantitative Methods for $\bar{r}^\star$ and Outlook

Table 3 summarizes the results of the methodologies we use to compute plausible values for $\bar{r}^\star$. The range for this rate, calculated from the average of the minimum and maximum levels obtained with each method, suggests that $\bar{r}^\star$ could be located between 1.7% to 3.3% in real terms, and from 4.7% to 6.3% in nominal terms, with mid points at 2.5% and 5.5%, respectively. To compute the latter, we simply added Banco de México's 3% inflation target.

Table 3: Summary of Quantitative Results for the Long Run

Methods	Real neutral rate, $r_t^\star$		Nominal neutral rate, $r_t^\star + \bar{\pi}$	
	Central point	Range	Central point	Range
Augmented Taylor rule	2.49	1.60 - 3.37	5.49	4.60 - 6.37
Neoclassical growth model	2.30	1.16 - 3.19	5.30	4.16 - 6.19
Affine model	2.70	2.40 - 3.10	5.70	5.40 - 6.10
Average	2.50	1.72 - 3.22	5.50	4.72 - 6.22

Note: We compute the long-run nominal neutral rate by adding to the estimated long-run real neutral rate the inflation target of Banco de México, which equals 3%.

5.5 Outlook for the Long-Run Convergence Level of the Neutral Rate

Structural factors determine the long-run convergence level of the neutral rate. Risks to this variable depend on how these factors affect the supply of loanable funds, and investment demand in the economy. In Annex C, we review relevant structural factors affecting the Mexican neutral rate in detail. In sum, we find that downside risks to $\bar{r}^\star$ in Mexico are given by an slowdown in the growth rate of the labor force, a higher proportion of the working-age population, a flat trend productivity, and a secular reduction in the global long-run real interest rate. Upside risks, in turn, relate to a potential increase in productivity generated by recent structural reforms implemented in the country.

6 Concluding Remarks

We find that potential growth cannot account for the changes in the neutral rate of a benchmark emerging market economy such as Mexico. We show that different medium-run estimates of the Mexican neutral rate followed, in general, a downward trend from 2000 to 2017. The exception is in the aftermath of the global financial crisis, were $r^\star$ followed a U-shaped pattern from 2009 to the end of the sample.

We argue that persistent transitory factors, such as conditions of economic slack in Mexico and the U.S., and the implementation of unconventional monetary policies in advanced economies, affected the Mexican neutral rate at business-cycle frequencies. Specifically, UMPs, by lowering government bond yields in advanced economies, dispatched capital flows towards government bond markets in EMEs. In Mexico, non-residents holdings of short-term public debt spiked from 2009 to 2012, and then scaled back during the *taper tantrum* in 2013. The boom-and-bust of these holdings seems to explain the estimated U-shaped pattern of the neutral rate obtained around the same time period. Finally, we argue that structural factors, such as increasing domestic savings, demographic shifters, and a decreasing global long-run real interest rate, appear to explain the estimated downward trend in the Mexican neutral rate.

An important caveat of the research agenda on the neutral rate is that all quantitative methods available to measure this variable are subject to a considerable degree of statistical uncertainty. This implies that a central bank must continue to observe a wide set of economic indicators in order to set the monetary policy stance consistent with achieving its objectives.

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A Output Growth and Money-Market Rates in AEs and EMEs

The IMF's WEO of April 2018, Box 1.3, presents potential growth estimates for selected AEs and EMEs, and finds that potential growth has persistently decreased for the former, while it follows an inverted U-shaped pattern for the latter. In particular, for the group of AEs, trend growth fell from 2.5% in 2001 to 1.5% in 2017, while for the group of EMEs, trend growth located at 4% in both years, with a peak at 5% in 2007. In the IMF's study, AEs include Australia, Canada, France, Germany, Italy, Japan, Korea, Spain, the U.K., and the U.S., while EMEs include Brazil, India, Mexico, Russia, and Turkey. When China is included in the EMEs sample, average trend growth of these subgroup is even stronger.

In this section, we present complementary evidence to Box 1.3 using the Funds's IFS data. Table 4 presents long-run averages of annual output growth rates and money-market real interest rates for a wider set of AEs and EMEs, 17 for the former and 30 for the latter. Money-market rates refer to the interest rate of assets with maturity of one year or less. These rates are therefore closely related to short-term government bond rates, such as T-bills. We compute ex-post real interest rates using annual inflation in each country. In the table, AEs include Australia, Belgium, Canada, Denmark, France, Germany, Italy, Japan, Korea, the Netherlands, New Zealand, Singapore, Spain, Sweden, Switzerland, the U.K., and the U.S. In turn, EMEs include Algeria, Angola, Argentina, Brazil, Bulgaria, Chile, Colombia, Cote d'Ivoire, Hungary, India, Indonesia, Kuwait, Malaysia, Mexico, Morocco, Pakistan, Peru, the Philippines, Poland, Romania, Russia, Saudi Arabia, Serbia, South Africa, Thailand, Tunisia, Turkey, Ukraine, Venezuela, and Vietnam.[42] The observation of each country is weighted by its proportion in world GDP. The sample starts in 1993 due to issues with data availability, especially for EMEs. The average weight of AEs' GDP in the sample is 50.9% of the world production, while that of EMEs is 28%. To compute average real interest rates, we excluded year observations in which the inflation rate is higher than 25%. The trimmed sample avoids, thus, distorted measures of real interest rates due to super-inflationary periods. In AEs there

[42] China is not included since data for its mainland money markets are not available for most of the period of interest.

are zero episodes with such characteristics, while in EMEs there are 84, from which 56 are located between 1993 and 1999.

The statistics shown in Table 4 are consistent with the IMF's results. Notably, the long-run average of output growth in AEs decreases from the first period considered to the last, while for EMEs this statistic fluctuates between 4% and 5%. Long-run averages cover 7 or 8 years, which is the typical length assumed for a business cycle. We opt for excluding the years 2008 and 2009 from the sample, since these years were severely affected by the GFC.[43]

In addition, the data show a decreasing trend in the long-run averages of the short-run real interest rate in both AEs and EMEs since at least 1993. The table shows that there is a clear positive co-movement between average growth rates and short-run real rates in AEs. In EMEs, this co-movement is fairly weak.

Table 4: Output Growth and Short-Run Real Interest Rate Statistics

Time period	Annual output growth rate		Money-market real interest rate	
	AEs	EMEs	AEs	EMEs
1993-2000	3.0	4.2	2.4	7.5
2001-2007†	2.3	4.9	0.7	3.3
2010-2017†	1.9	4.2	-1.0	0.9

Note: The statistics considers 17 advanced economies, and 30 emerging economies. Each country-observation is weighted according to the proportion of the country's production on global GDP. Money-market real interest rates are computed with realized inflation in a given year. *Source:* Own computations with data from the International Financial Statistics of the IMF.
† The years 2008 and 2009, where the effects of the GFC reached their peak, were removed from the sample.

B Recent Estimates of $r^\star$ Around the World

This section non-exhaustively surveys the recent evidence related to $r^\star$ in AEs and EMEs. The main takeaway is that almost all studies capture a downward trend in $r^\star$ that started around the 90s, and that sharpened in the wake of the 2008 global financial crisis.

[43] Including these years into the calculation reduces the long-run average of output growth of AEs, but not so much in EMEs.

B.1 Advanced Economies

For the U.S., Yellen (2015) presents a set of estimates of short-run $r^\star$ obtained from New-Keynesian DSGE models developed by the Fed's staff, and shows that this variable plunged towards negative levels at the onset of the GFC, and reached zero by the end of 2015. These models interpret the reduction in short-run $r^\star$ as a response to persistent macro shocks to aggregate demand, such as tighter financing conditions and lesser access to credit, de-leveraging by households, lower global growth, and greater uncertainty.[44] More flexible methodologies, such as state-space models with a time-varying structure, find similar results. For the case of the U.S., Laubach and Williams (2016), and Johannsen and Mertens (2016) estimate a clear downward trend in $r^\star$ that has started at least since the 80s, but has deepened since the financial crisis. Laubach and Williams (2016) relate the fall in $r^\star$ to a decreasing potential growth. In contrast, Del Negro, Giannone, Giannoni and Tambalotti (2017), using both time-series and a DSGE model, attribute the fall in $r^\star$ to a rising premia for the liquidity and safety of Treasury bonds, also known as *convenience yield*. Their findings add to the literature showing that Treasury bonds are valued not only by their pecuniary return, but also by the safe and liquidity services they offer.

The evidence of a downward trend of $r^\star$ is not exclusive to the U.S. Holston *et al.* (2017) find evidence that $r^\star$ and potential growth in Canada, the Euro Area, and the U.K. have followed a downward trend for several decades.[45] Additionally, they find that these estimates and those for the U.S. have a considerable amount of co-movement over time. Thus, the authors suggest that global factors play an important role in explaining the trends of $r^\star$, and potential growth in these economies. Similarly, Bouis, Rawdanowicz, Renne, Watanabe and Christensen (2013) find that for seven OCDE economies $r^\star$ has generally fallen since 1980.[46]

[44]The estimates of these DSGE models assume the existence of nominal rigidities and other frictions to capture transitory macroeconomic shocks. To estimate short-run $r^\star$, the models compute the real interest rate that would prevail if prices and wages were flexible. Therefore, the estimated short-run $r^\star$ in this type of models is a counterfactual measure, not observable, and highly volatile, since it is subject to a wide set of transitory shocks.

[45]See also Berger and Kempa (2014) for Canada.

[46]The countries are the U.S., Japan, the Euro Area, the U.K., Canada, Sweden, and Switzerland. The last two countries are the exceptions, since their estimates of $r^\star$ have remained stable, and relatively high, since the financial crisis.

They argue that the fall of $r^\star$ is likely the result of a lower potential growth. In addition, they mention that, according to OECD projections, $r^\star$ may converge to a lower level than before the GFC. For Japan, Fujiwara, Iwasaki, Muto, Nishizaki and Sudo (2016) show that $r^\star$ has followed a downward trend since the 90s, and relate this trend to a slowdown in potential growth. Similarly, the European Central Bank (2004) finds that $r^\star$ in the Euro Area has decreased since the mid 90s, and argues that this trajectory may reflect the slowdown in productivity and population growth in the region.[47] For Norway, Bernhardsen and Gerdrup (2007) find that $r^\star$ has fallen since at least 1990, and explain that one of the reasons is partly a lower inflationary risk premia, since inflation and its expectations stabilized towards low levels. For New Zealand, Basdevant, Björksten and Karagedikli (2004) find evidence that suggests a downward trend in $r^\star$ since 1992, while Björksten and Karagedikli (2003) conclude that the reduction in $r^\star$ can be partly attributed to a worldwide decline in natural rates, and to local factors. Richardson and Williams (2015) find similar evidence for New Zealand. Schmidt-Hebbel and Walsh (2009) present more evidence on $r^\star$ in other advanced economies.[48] Although they do not find clear evidence of a downward trend in $r^\star$ in all cases, they show that the neutral rates of these economies are highly correlated.

B.2 Emerging Market Economies

The evidence for EMEs is not very different from that for AEs. Trends of natural rates in EMEs have also declined. In particular, Magud and Tsounta (2012), using different methodologies, document some stylized facts for $r^\star$ in ten Latin American countries:[49] (i) $r^\star$ tends to be lower in countries with stronger fundamentals; (ii) wider ranges in $r^\star$ estimates are associated with weaker monetary policy frameworks and higher inflation risk premia, although the

[47] See also Cuaresma, Gnan and Ritzberger-Gruenwald (2004), Mesonnier and Renne (2007), Garnier and Wilhelmsen (2009), and Fries, Mésonnier, Mouabbi and Renne (2018).

[48] The countries covered are the U.S., the Euro Area, Japan, and some inflation targeting countries, such as Australia, Canada, New Zealand, Norway, the U.K., Sweden, and Chile.

[49] The countries are Brazil, Chile, Colombia, Costa Rica, the Dominican Republic, Guatemala, Mexico, Paraguay, Peru, and Uruguay. The methodologies used by the authors include: the HP filter, an implicit common stochastic trend using short- and long-term interest rates, dynamic Taylor rules, expected-inflation augmented Taylor rules, the Laubach and Williams model, consumption-smoothing models, and the uncovered interest rate parity (UIP) condition. Their sample spans from 2000 to 2012.

dispersion could be also caused by short samples and unavailable data; and (iii) $r^\star$ presents a downward trend in the last decade for most of the countries studied. Magud and Tsounta argue that this trend is possibly due to stronger economic fundamentals in the region, as well as more accommodative global financing conditions that would have increased the supply of loanable funds in the region.

In the same vein, Perrelli and Roache (2014) find a downward trend in the estimates of $r^\star$ in a wider set of EMEs.[50] These authors focus on the experience of Brazil, and find that the fall in its neutral rate can be explained by both domestic and foreign factors. Regarding the former, they argue that financial deepening, a declining public debt, and a lower sovereign risk premium have contributed to increase the desired savings in the country. Concerning the latter, they find evidence suggesting that the global real interest rate has also contributed to the decrease in Brazil's neutral rate.

In other individual-country analyses, Fuentes and Gredig (2008) and González, Ocampo, Pérez and Rodríguez (2012) study the cases of Chile and Colombia using a battery of models to estimate plausible paths for $r^\star$.[51] In the case of Chile from 1980 to 2007, all models find that the estimated $r^\star$ presents a downward trend. For Colombia, the estimates of $r^\star$ vary significantly.

Finally, Zhu (2016) also finds that, with the exceptions of China and Thailand, estimates of $r^\star$ have declined substantially since 2005 in a group of countries in the Asia-Pacific region.[52] Consistent with the existing evidence, the author finds that for some economies (e.g. the U.S., Japan, Korea, and Singapore), the downward trend in $r^\star$ started in the 1980s. Also,

[50]They include the following countries: Brazil, Chile, China, Colombia, the Czech Republic, Egypt, Hungary, India, Indonesia, Israel, Korea, Malaysia, Mexico, Peru, the Philippines, Poland, Russia, South Africa, Taiwan, Thailand, Turkey, and Uruguay. The authors use statistical filters to document the decline of $r^\star$ in a sample spanning from 2002 to 2013. Further, using a principal components analysis, the authors find that two common factors may explain about 45% of the common fluctuations in real policy rates of the analyzed countries. The first of these factors represents the common trend, while the second one is the common cycle.

[51]The models used in these papers can be classified in three categories: (i) economic theory (traditional consumption model, uncovered parity interest rate condition, general equilibrium reduced-form models); (ii) implicit expectations of $r^\star$ in financial instruments (forward rates, state-space models with common stochastic trend in short-run and long-run interest rates, and yield curve models); and (iii) statistical models (filters).

[52]These countries are: Australia, China, Hong Kong, India, Indonesia, Japan, Korea, Malaysia, New Zealand, the Philippines, Singapore, Thailand, and the U.S. The sample spans from 1950 to 2014. The author exploits the spectral density of the data to find low-frequency changes.

Zhu finds that low-frequency movements in the neutral rate seem to be strongly related to demographics and global factors (e.g. trade and capital flows, global liquidity), while the relationship with potential growth appears to be weaker.

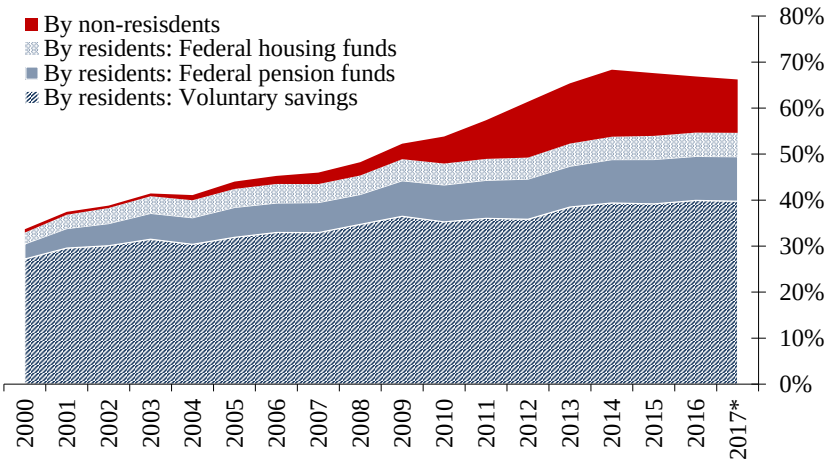
C Heuristic Analysis of Structural Factors in Mexico

The outlook of $\bar{r}^{\star}$ depends on how structural factors are expected to change, and how they will affect the supply of loanable funds, and investment demand in the economy. We now review trends of some important structural factors.

Savings. Domestic savings have increased robustly as a percentage of GDP since the beginning of the 2000s in Mexico. Voluntary savings by residents, distributed along public and private instruments, reached 40% of GDP in November 2017 as compared to 27% in 2000. Also, federal pension and housing funds, a compulsory type of savings, reached 15% of GDP in November 2017 relative to 5.7% in 2000. In addition, domestic assets holdings by non-residents became important only after 2008. Overall, the trends signal that the supply of loanable funds in the economy will continue to grow, which represents a downward pressure on $\bar{r}^{\star}$ in the future.

Population. Demographics have also played a role in the determination of $\bar{r}^{\star}$ in at least two dimensions. First, changes in the distribution of the Mexican population may have favored an environment conducive to strengthening the savings profile of the country. And second, a slower growth of the labor force might have negatively affected potential output growth. With respect to the former, the National Population Council (or CONAPO, by its Spanish acronym) estimates that the proportion of the working-age population in Mexico (those between 16 and 65 years old) increased from 59.3% of the total population in 2000 to 64.7% in 2018. This subgroup of the population has the highest ability to save in comparison to other subgroups. CONAPO expects the working-age population to peak by 2025, at 65.4%. Regarding the labor force, CONAPO estimates that its growth rate diminished from 1.7% in 2000 to 1.4% in 2016, and that it might reach 0.6% by the end of the 2020s. If capital and labor are

Figure 16: Holdings of Domestic Assets

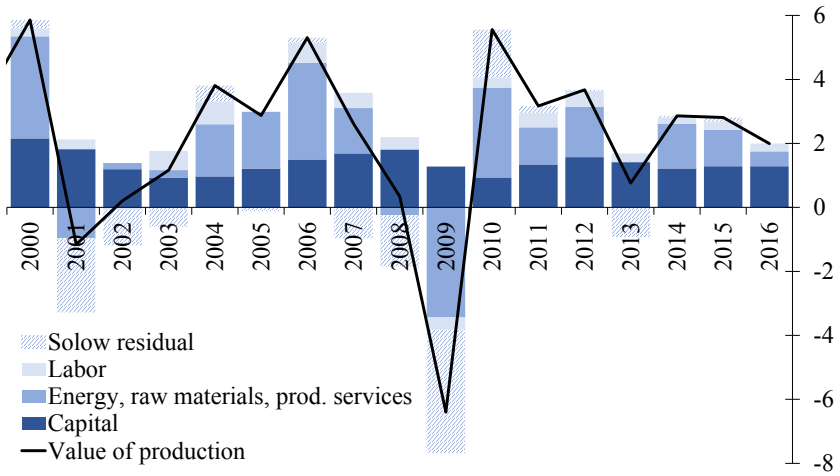


Note: The data is presented at an annual frequency and as percentage of GDP. The year 2017 covers January to November due to availability. *Sources:* INEGI and Banco de México.

complements, this pattern for the labor force represents a poorer outlook for the marginal product of capital and investment returns, which implies that investment demand might also grow slowly. Demographics have, thus, posed downside risks to $\bar{r}^{\star}$ in recent years, and the outlook going forward does not seem to be different.

Productivity and growth. INEGI's Total Factor Productivity statistics decompose GDP growth into the contributions proceeding from capital, labor, energy, raw materials and production services from 2000 to 2016, the latest available year. The difference between total growth, and the sum of contributions of each factor is total factor productivity (TFP), or the Solow residual. This taxonomy of growth shows that capital services are the most stable contributors, while TFP is the most unstable. Since TFP does not show a clear pattern in the data, it is difficult to assess its possible impact on $\bar{r}^{\star}$. However, the latter might be reverted if the structural reforms recently implemented in Mexico boost productivity in the coming years. Part of these reforms encourage competition in sectors such as telecommunications, and energy production (oil and electricity), while a deeper long-term reform seeks to substantially upgrade the quality of elementary education in public schools.

Figure 17: Growth Accounting



Note: The data is presented at annual frequency. *Sources:* INEGI.

Global cost of money. Section 4.2 shows that the global long-run real interest rate has presented a clear downward trend for at least 25 years. Academics and policymakers have hotly debated about the drivers behind this trend.[53] Rachel and Smith (2015) have recently argued that at least 400 basis points of the fall in the global long-run real interest rate registered between 1985 and 2015 may be ascribed to secular factors affecting global desired savings and global investment demand. The structural factors pushing outwards global desired savings are an increase in the proportion of the working-age population, higher inequality, and, to a lesser extent, the glut of precautionary savings by emerging markets. In turn, structural factors that have affected negatively global investment demand are a falling relative price of capital, lower public investment, and an increase in the spread between the risk-free rate and the rate of capital returns. In contrast, Rachel and Smith argue that economic growth seems not to have affected negatively the global long-run real interest rate until 2008. After that year, the prospect of a lower global growth could have contributed to a fall of 100 basis points in the global long-run interest rate.

[53]For instance, as early as 2005 the former Fed's Chairman Ben Bernanke expressed concerns about the growing global savings glut, i.e. a situation in which global desired savings exceeds global investment demand.

The global factors just described may also affect the Mexican long-run interest rate through international arbitrage. Therefore, we should not expect to see structural upside risks to $\bar{r}^{\star}$ coming from international capital markets in the near future.

Outlook summary. Downside risks to the long-run convergence level of the neutral rate in Mexico are given by an expected slowdown in the growth rate of the labor force, a higher proportion of the working-age population, and a secular reduction in the global long-run real interest rate. Upside risks, in turn, relate to a potential increase in productivity generated by recent structural reforms in the country.