The Longer-Term Convergence Level of the Neutral Rate of Interest in Mexico

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Abstract

In this paper, we argue that foreign and domestic structural factors may explain the decline of the long-run convergence level of the neutral rate of interest in Mexico. In particular, we find that low-frequency changes in the neutral rate may be attributed to increasing domestic savings, demographics shifts, 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 alike.

Keywords: Neutral rate of interest, long-run determinants, Mexico.

JEL classification: C10, E43, E52.

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1. INTRODUCTION

The *neutral* or *natural* rate of interest is defined as the level of the short-run real interest rate that is consistent with output near its *potential* or *natural* level.[1] It is also common to refer to the neutral rate of interest as r^*. The domestic market for loanable funds determines r^*. In this market, *desired savings*, composed of foreign and home portfolios in fixed-income markets, firms' and households' bank deposits, and other types of savings, establish the supply, while *investment demand*, composed of public and private debt, determines the demand. The neutral rate helps to determine the stance of monetary policy,[2] but unfortunately this variable is not observed and must be estimated.

Recent evidence on r^* in both advanced and emerging economies (AEs and EMEs) yields remarkably similar results: most estimates show a downward trend in r^* over the past 25 years. Some commentators, such as Holston, Laubach and Williams (2017), have observed that potential growth and the neutral rate have co-moved in advanced economies during this time period. By contrast, this co-movement is not observed in emerging economies. A dimension that has not been fully explored in the literature is the role of capital flows in the determination of r^*. Sustained capital flows could have a long-lasting effect on the neutral rate of an EME since these flows affect the supply of loanable funds in such an economy. Although this channel is present in both AEs and EMEs, it could be more important in the latter given their greater sensitivity to external events.

In this document, we estimate the *long-run convergence level* of the neutral rate in Mexico, a prototype EME with a significant volume of international trade and a financial market highly integrated with the global market. We also refer to the long-run convergence level of the neutral rate as $\bar{r}^*$, i.e., with a *bar* accent over r^*. Following the FOMC Minutes of October 2015,

1 At this level of output, we also find the natural rate of unemployment, i.e., the level of unemployment at which inflation remains stable in the absence of shocks.

2 The monetary policy stance is neutral if the short-run real interest rate equals r^* and it is contractionary (expansionary) if the short-run real rate is located above (below) r^*. 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 measure of inflation expectations to r^*, we get the level of the monetary-policy interest rate at which the policy is neutral.

we formally define $\bar{r}^*$ as *"the longer-run normal level to which the* [short-run real interest] *is expected to converge in the absence of further shocks to the economy."* We perform our estimations for the sample period January 2000 to December 2017.

The long-run convergence level of r^* is determined exclusively by structural factors, such as potential growth, demographics, and financial-markets development.[3] To compute a robust measure of $\bar{r}^*$, we estimate three quantitative methods: an augmented Taylor rule which includes a control for a very persistent transitory (or non-structural) factor, an open-economy RBC model, and the 10-year expectation of the short-run nominal interest rate computed from an affine term- structure model. In contrast to studies focusing on higher-frequency measures of the neutral rate, there are a limited number of techniques that can be used to estimate $\bar{r}^*$. A suitable method for the low-frequency measure must be able to disentangle the effects of transitory and structural factors at the level of the neutral rate. This task becomes difficult when very persistent transitory factors are observed in the economy (e.g., the shocks that led to the global financial crisis, its aftermath, and the policy responses that followed). Since no method is perfect, we also review some structural factors that are informative about upward and downward risks for $\bar{r}^*$ in Mexico.

Our results are as follows. The evidence suggests that the long-run convergence level of r^* in 2017 is lower than the level that prevailed at the beginning of the 2000s, falling from around 3% in real terms to close to 2.5%.[4] As mentioned above, a downward trend for $\bar{r}^*$ has also been estimated in several AEs and EMEs.

We argue that both domestic and foreign structural factors account for the apparent fall in $\bar{r}^*$ registered from the 2000s to present. On the domestic side, we observe (1) sustained growth of national savings as a percentage of GDP, (2) an increase in working-age share of the population,

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[3] See Carrillo, Elizondo, Rodríguez-Pérez and Roldán-Peña (2018) for a formal definition. In that paper, we also considered higher-frequency measures of r^*, where we analyzed transitory factors (e.g. macro shocks) that diverted the neutral rate temporarily from its long-run fundamental value in Mexico. In Section 3, we briefly review the results and drivers of these higher-frequency measures of r^*.

[4] The former corresponds with the average of the estimates of r^* for the short-, and medium-run during the period 2001Q1-2008Q4, while the latter is the average of the estimates of $\bar{r}^*$ in Section 4. It is worth mentioning that the uncertainty surrounding these estimates is quite significant, so punctual results must be taken with caution.

(3) a declining outlook for the growth rate of the labor force, and (4) a flat trend in productivity. All four factors imply a lower long-run convergence level of the neutral rate. In the market for loanable funds, the first two factors drive up the supply, while the last two reduce the demand (through their influence on investment). On the foreign side, the sustained reduction in the global long-term real interest rate seems to have pushed international long-term credit toward the Mexican market. The latter could have lowered the domestic long-term real interest rate through no-arbitrage conditions. Indeed, sustained capital inflows seem to have contributed to permanent increases in the supply of loanable funds in the economy, putting downward pressure on $\bar{r}^*$.

The remainder of the paper is organized as follows. Section 2 reviews international evidence on trend output growth and money-market rates in a large set of AEs and EMEs. The section also reviews several studies that estimate r^* in several economies. Section 3 summarizes the findings on higher-frequency measures of the neutral rate in Mexico. In turn, Section 4 focuses on the long-run convergence level of the neutral rate in the country. The final section concludes.

2. INTERNATIONAL EVIDENCE

Holston *et al.* (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 over the past 25 years. These authors suggest that global factors may largely explain this behavior. In sharp contrast, this kind of co-movement does not hold for emerging market economies (EMEs) as most of these countries grow at relatively high rates, while at the same time their neutral rates have fallen. In this Section, we first review data on the growth rates and short-run real interest rates for a large set of AEs and EMEs that confirm the aforementioned trends. We then review some recent estimations of the neutral rate in different economies.

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

The IMF's *World Economic Outlook* of April 2018, Box 1.3, presents potential growth estimates for ten AEs and five EMEs;[5] the report finds that potential

[5] 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.

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 was 4% for both years, with a peak at 5% in 2007.[6]

We now present complementary evidence to Box 1.3 using the Fund's IFS data. Table 1 presents long-run averages of annual output growth rates and money-market real interest rates for a wider set of AEs and EMEs, seventeen for the former and thirty for the latter.[7] 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. To compute the mean for each category, we weighted the observation of each country by its proportion in world GDP. The average weight of AEs' GDP in the sample is 50.9% of the world production, while that of EMEs is 28%.[8] The sample starts in 1993 due to issues with data availability, especially for EMEs. 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 global financial crisis (*GFC*, for short).[9]

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6 When China is included in the EMEs sample, average trend growth of this subgroup is even stronger.

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

8 When computing the average money market real interest rate, we excluded observations for years 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 are zero episodes with such characteristics, while in EMEs there are 84, of which 56 occurred between 1993 and 1999.

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

Table 1

OUTPUT GROWTH AND SHORT-RUN REAL INTEREST RATE STATISTICS

	Annual output growth rate		*Money-market real interest rate*	
Time period	*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 consider 17 advanced economies and 30 emerging economies. Each country-observation is weighted according to the proportion of global GDP contributed by the country's production. 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, when the effects of the global financial crisis reached their peak, were removed from the sample.

The statistics shown in Table 1 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%. In addition, the data show a downward trend in the long-run average of the short-run real interest rate in both AEs and EMEs since at least 1993. The table shows a clear co-movement between average growth rates and short-run real rates in AEs but, notably, not in EMEs. This evidence suggests that a structural factor different than potential growth seems to drive the trend of the short-run real interest rate in an EME. We now provide further evidence on recent estimations of the neutral rate in different countries.

2.2 Recent Estimates of r* Around the World

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

2.2.1 Advanced Economies

For the U.S., Yellen (2015) presents a set of estimates of r^* obtained from New-Keynesian DSGE models developed by the Fed's staff; these show that this variable plunged toward 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^* as a response to persistent macro shocks to aggregate demand, such as tighter financing conditions and less access to credit, de-leveraging by households, lower global growth, and greater uncertainty.[10] 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^* that has started at least since the 80s but has deepened since the financial crisis. Laubach and Williams (2016) relate the fall in r^* 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^* 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 for their pecuniary return, but also for the safe and liquidity services they offer.

The evidence for a downward trend in r^* is not exclusive to the U.S. Holston *et al.* (2017) find evidence that r^* and potential growth in Canada, the Euro Area, and the U.K. have followed a downward trend for several decades.[11] 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 trends in r^* and potential growth in these economies. Similarly, Bouis, Rawdanowicz, Renne, Watanabe and Christensen (2013) find that for seven OECD economies r^* has generally fallen since 1980.[12] They argue that the fall of r^*

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

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

[12] 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^* have remained stable, and relatively high, since the financial crisis.

is likely the result of lower potential growth. In addition, they mention that, according to OECD projections, r^* may converge to a lower level than that before the GFC. For Japan, Fujiwara, Iwasaki, Muto, Nishizaki and Sudo (2016) show that r^* has followed a downward trend since the 90s, and they relate this trend to a slowdown in potential growth. Similarly, the European Central Bank (2004) finds that r^* 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.[13] For Norway, Bernhardsen and Gerdrup (2007) find that r^* has fallen since at least 1990, and they explain that one of the reasons is a lower inflationary risk premia, since inflation and its expectations stabilized toward low levels. For New Zealand, Basdevant, Björksten and Karagedikli (2004) find evidence that suggests a downward trend in r^* since 1992, while Björksten and Karagedikli (2003) conclude that the reduction in r^* can be partly attributed to a worldwide decline in natural rates as well as to local factors. More recently, Richardson and Williams (2015) confirm this evidence for New Zealand. Schmidt-Hebbel and Walsh (2009) present more evidence on r^* in other advanced economies.[14] Although they do not find clear evidence of a downward trend in r^* in all cases, they *do* show that the neutral rates of these economies are highly correlated.

2.2.2 Emerging Market Economies

The evidence for EMEs is not very different from that for AEs. Neutral rates in EMEs have also shown a downward trend. In particular, Magud and Tsounta (2012), using different methodologies, document some stylized facts for r^* in ten Latin American countries:[15] (i) r^* tends to be lower in countries with stronger fundamentals; (ii) wider ranges in r^* estimates are associated with weaker monetary policy frameworks and higher inflation risk

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

[14] 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.

[15] 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 Hodrick-Prescott 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.

premia, although the dispersion could be also caused by short samples and unavailable data; and (iii) r^* experienced a downward trend in the past 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 to more accommodative global financing conditions that would have increased the supply of loanable funds in the studied countries.

In the same vein, Perrelli and Roache (2014) find a downward trend in the estimates of r^* for a wider set of EMEs.[16] 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. For 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^*[17]. In the case of Chile from 1980 to 2007, all models find that the estimated r^* trends downward. For Colombia, the estimates of r^* vary significantly.

Finally, Zhu (2016) finds that, with the exceptions of China and Thailand, estimates of r^* have declined substantially since 2005 for a group of countries in the Asia-Pacific region.[18] Consistent with the existing evidence, the author finds that for some economies (e.g. the U.S., Japan,

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16 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^* in a sample spanning from 2002 to 2013. Furthermore, 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.

17 The models used in these papers can be classified into three categories: (i) economic theory (traditional consumption model, uncovered parity interest rate condition, general equilibrium reduced-form models); (ii) implicit expectations of r^* 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).

18 These countries are: Australia, China, Hong Kong, India, Indonesia, Japan, Korea, Malaysia, New Zealand, the Philippines, Singapore, Thailand, and the

Korea, and Singapore), the downward trend in r^* started in the 1980s. Additionally, 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.

3. SHORT-RUN NEUTRAL RATE IN MEXICO: A SUMMARY

Carrillo, Elizondo, Rodríguez-Pérez and Roldán-Peña (2018), henceforth CERR for short, conduct detailed analysis of higher-frequency measures of the neutral rate in Mexico. For convenience, CERR call these measures of the neutral rate short-run r^* or medium-run r^*. CERR consider that transitory factors (e.g. macro shocks) divert the neutral rate temporarily from its long-run fundamental value, which is determined exclusively by structural factors. In this section, we provide a brief summary of the results in CERR, and we invite the interested reader to review the paper for further details.[19]

The takeaway from CERR's analysis is that, the rise and fall of non-residents holdings of short- term Mexican debt in the aftermath of the GFC, which increased and then decreased the supply of loanable funds in the country, may explain the temporary dip of the neutral rate observed from 2010 to 2014 in Mexico. These capital inflows surged during the implementation of unconventional monetary policies in advanced economies, and started to reverse when the Fed signaled for the first time the tapering of its QE programs in mid-2013. Therefore, before the GFC, long-term averages of the estimates of short-run r^* seem to be good approximations of the long-run convergence level of the neutral rate in Mexico, since transitory factors were not very persistent. However, after 2008 this is no longer the case because short-run r^* was affected by very persistent transitory factors, such as slack economic conditions and the implementation of ultra-accommodative monetary policies in AEs.

To perform the short-, and medium-run estimations, CERR use the ex ante short-term real interest rate, measured as the overnight interbank

U.S. The sample spans from 1950 to 2014. The author exploits the spectral density of the data to find low-frequency changes.

[19] CERR also provides a formal disambiguation between the neutral in the short run and its long-run convergence level.

nominal interest rate minus the one-year ahead expectations of headline inflation.[20] The period of study in CERR spans from January 2000 to December 2017 at a monthly and quarterly frequency, depending upon data availability and the model used.

To achieve a robust estimate of r^* for the short-, and medium-run, CERR consider five different methodologies: business-cycle averages and filters, a simple Taylor rule estimated recursively, affine term-structure models, an adapted version of the Laubach and Williams (2003) model, and a BVAR model with time-varying intercepts. The average of the point estimates of these models is shown in Figure 1, where the range corresponds to the minimum and maximum values of the point estimates of all methodologies in every period.

Notably, all methodologies display similar paths. In particular, the results suggest that short-run r^* in Mexico decreased from 2009 onwards, persistently deviating from its long-run convergence value, falling to minimum levels in 2012, and partially returning to its trend by 2014.

CERR argue that foreign and domestic transitory factors pushed short-run r^* below trend between 2009 and 2014. Among the domestic transitory factors that depressed short-run r^*, CERR mention slack conditions that prevailed in the Mexican economy following the GFC, which implied a demand for loanable funds lower than normal. Concerning the foreign transitory factors, CERR analyze two: (1) the persistent slack conditions that prevailed in the U.S. after the crisis, and (2) the implementation of unconventional monetary policies (or UMPs) by central banks in some advance economies, and in particular the Federal Reserve in the U.S.

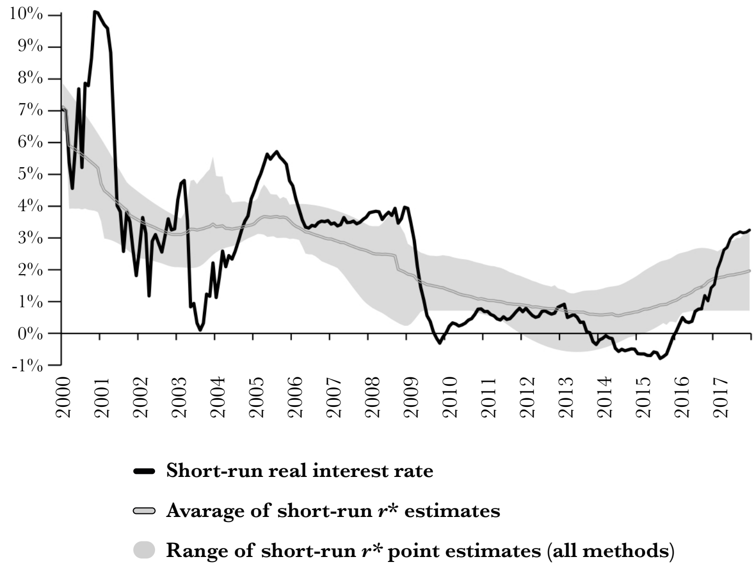
Concerning the first foreign factor, CERR argue that the U.S. business cycle co-moves not only with the Mexican business cycle, but also with the Mexican neutral rate. Therefore, transitory factors affecting the neutral rate in the U.S. may also impact the neutral rate in Mexico. For a small open economy such as Mexico, aggregate demand conditions abroad matter because they influence export dynamics, financial flows, and economic activity in general.

Regarding the second foreign factor, a growing literature finds that UMPs contributed to strong capital inflows toward EMEs, including Mexico. Accordingly, CERR argue that several investors might have re-balanced their portfolios away from the U.S. and other advanced economies with low returns to favor relatively safe emerging economies with higher

[20] Inflation expectations can be extracted from Banco de México's survey of private professional forecasters.

Figure 1

SUMMARY OF RESULTS FOR SHORT- AND MEDIUM-RUN r^*



Source: CERR's estimates with data from INEGI, Banco de México, Valmer, PiP, the NY Fed, and St. Louis Fed's FRED database. The range corresponds to the minimum and maximum values of the point estimates of all methodologies in every period.

returns, such as Mexico. In particular, the evidence suggests that the abundant liquidity in international financial markets could have persistently increased the supply of loanable funds in Mexico through foreign holdings of short-term Mexican debt, thus reducing short-run r^* from 2009 to 2014.

Table 2 displays the average of the point estimates of short-run r^* for the periods 2001Q1-2008Q4 and 2009Q1-2017Q4. The table shows that all methodologies considered in CERR find consistent results, namely that the estimates of short-run r^* in Mexico fell during the GFC from an average of 3% to around 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%.

4. LONG-RUN CONVERGENCE LEVEL OF r* IN MEXICO

In this section, we present three different quantitative methods that estimate the long-run convergence level of the neutral rate, i.e. $\bar{r}^*$. We resort to these methods after the observation that a long-term average of the estimated of short-run r^* might be a poor approximation of $\bar{r}^*$ due to the presence of very persistent transitory factors affecting the Mexican economy after 2009. 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 a heuristic analysis of structural factors affecting $\bar{r}^*$.

4.1 Augmented Taylor Rule

Taylor rules are commonly used tools that help to estimate the systematic behavior of a central bank's monetary policy. In particular, the estimated interest-rate rule aims to approximate the reaction function of the policy rate toward deviations of inflation from its target, and of output from its potential level. It is worth emphasizing that the resulting interest-rate rule should not be taken as the main policy directive of the central bank, but rather as a particular lens for interpreting the systematic behavior of monetary policy. Furthermore, notice that when inflation equals its target and the output gap equals zero, $\bar{r}^*$ is given by the difference between the rule's intercept and the inflation target. For example, consider the following Taylor rule with interest rate smoothing:

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$$R_t = (1+\rho)[\bar{r}^* + \bar{\pi} + \partial(\pi_t - \bar{\pi}) + \theta y_t] + \rho R_{t-1} + \varepsilon_t,$$

where R is the overnight interbank nominal interest rate, π is the inflation, $\hat{y}$ is the output gap, and ε captures any change in the nominal interest rate not explained by the rule. Additionally, the specification includes a lag of R to capture gradual adjustments in this variable induced by the central bank. CERR estimate equation (1) recursively over the period 2001 to 2017 in order to capture changes in $\bar{r}^*$. The results from this exercise show that the estimate of falls from 2008 to 2014, partially reverting to pre-GFC levels afterwards. In light of these results, and the evidence

Table 2

SUMMARY OF QUANTITATIVE RESULTS FOR THE SHORT AND MEDIUM RUN

Methods	Real neutral rate r_t^* 2001Q1-2008Q4	Real neutral rate r_t^* 2009Q1-2017Q4	Nominal neutral rate $r_t^* + \pi_t^e$ 2001Q1-2008Q4	Nominal neutral rate $r_t^* + \pi_t^e$ 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 William 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, CERR add the average of headline inflation expectations for 12-months ahead to r^*. See CERR for further details.

summarized in the previous section, CERR argue that the Fed's UMPs seem to have affected the Mexican neutral rate during this period through their effects on capital flows.

In this context, Taylor and Wieland (2016) show that the omission of relevant variables in the reaction function of the central bank can produce a bias in the estimation of $\bar{r}^*$. In particular, the authors argue that omitting important information in model estimation can result in a noisy estimate of the $\bar{r}^*$ one that misleadingly absorbs the fluctuations of the omitted factors. With this in mind, CERR include an indicator of the Fed's UMPs as an additional regressor in a modified version of Taylor's rule. Therefore, to obtain an estimate of $\bar{r}^*$. CERR estimate the following augmented Taylor rule:

2 $$R_t = (1+\rho)[\bar{r}^* + \bar{\pi} + \gamma(\mathbf{1x}R_t^{US,shadow}) + \beta(\pi_t - \bar{\pi}) + \theta y_t] + \rho R_{t-1} + \varepsilon_t,$$

where $R_t^{US,shadow}$ is the shadow fed funds rate of Wu and Xia (2016), and the indicator variable 1 takes the value of zero when $R_t^{US,shadow}$ is positive or the value of one when $R_t^{US,shadow}$ is negative (i.e. from July 2009 to December 2015).[21] CERR only include information about the shadow rate during the ZLB period as a proxy for the Fed's UMPs. Therefore, they explicitly assume that these policies capture a very persistent transitory factor, not a structural factor.[22]

CERR estimate equation (2) recursively on a monthly basis, using the short-run nominal interest rate, headline annual inflation measured by the CPI, and a measure of the output gap using Mexico's Global Indicator of Economic Activity (or IGAE, by its Spanish acronym), published monthly by INEGI. To compute a measure of economic slack from IGAE, CERR use its percent deviation from trend, which they estimate using the Hodrick-Prescott filter with tail correction.

Figure 2 presents the results from this exercise. The results suggest that $\overline{r}^*$ has fluctuated around the level of 2.5% since 2009. This number translates into a neutral nominal policy rate of 5.5%, if we add Banco de México's inflation target of 3%.

4.2 Open-Economy rbc Model

As an alternative means of measuring $\overline{r}^*$, we use a neoclassical growth model for a small open economy. We follow the business-cycle model of Lama (2011) who, in a manner similar to Chari, Kehoe and McGrattan (2007), includes four sources of macroeconomic fluctuations in 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 $\overline{r}^*$ is the long-term average of the equilibrium real rate of capital returns, r^k which is a model-consistent measure of the actual macro 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.[23]

[21] We have considered alternative measures of the shadow fed funds rate, such as those proposed by Lombardi and Zhu (2014), and Krippner (2015). The results remain quantitatively similar.

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

[23] Recently, Caballero, Farhi and Gourinchas (2017) noticed 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 it is in the U.S. Nonetheless, we bear in mind that even a long-period average of r^k might be a poor approximation of $\overline{r}^*$. We decided to keep

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

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

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:

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

subject to

$$(1+n)b_{t+1} + c_t + i_t \leq (1+\tau_{lt}) w_t l_t + (1+\tau_{kt}) z_t k_t + (1+\tau_{bt})(1+r_t^W) b_t + \Upsilon_t,$$

$$(1+n)k_{t+1} \leq (1+\partial)k_t \leq (1+\partial)k_t + i_t - \phi\left(\frac{i_t}{k_t}\right)k_t,$$

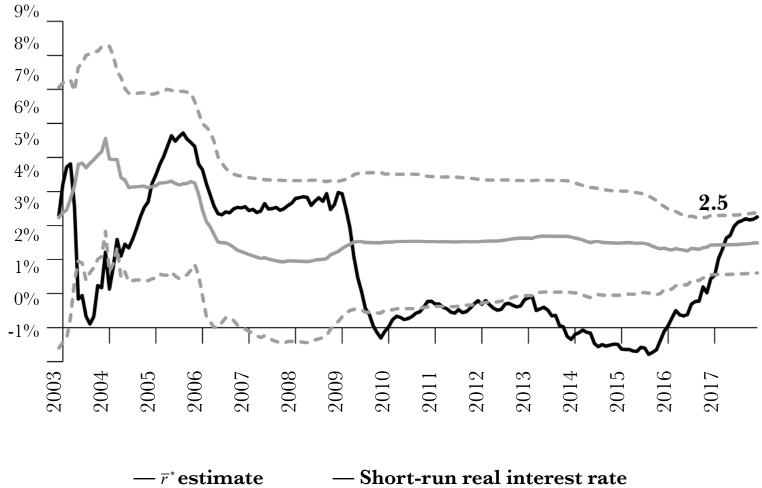
$$1 + r_t^W = (1+r^W)\left(\frac{b_t}{b^W}\right)^{\upsilon}$$

where β is the subjective discount factor, ψ is a normalizing constant, n is the population growth rate, ∂ is the capital depreciation rate, and $\upsilon > 0$ is the elasticity of the supply of international borrowing. In turn, N_t is the size of the population, r^W is the world real interest rate, Υ_t represent government transfers, and $\phi(i_t / k_t) = \vartheta / 2x(i_t / k_t - \tilde{\partial})^2$ measures capital

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 $\overline{r}^*$. Second, there are not many methods available in the literature to estimate $\overline{r}^*$.

Figure 2

AUGMENTED TAYLOR RULE



Note: The confidence intervals are of 90% significance.

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

adjustment costs, where $\tilde{\partial} = \partial + n + \gamma + n\gamma$. 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 they 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

$x_t = x^{1-\rho_x} x_{t-1}^{\rho_x} \exp(\varepsilon_{xt})$ for

$x \in \left\{A, 1-\tau_l, 1-\tau_k, 1-\tau_b\right\}$

where $\varepsilon_{xt} \sim N(0,\sigma_x)$ are normally-distributed, white-noise innovations. The dynamics of the detrended economy are given by the law of motion for capital, the wedge processes, and the following market-clearing conditions:

$$\tilde{y}_t - \tilde{c}_t - \tilde{l}_t = (1+n)(1+\gamma)\tilde{b}_{t+1} - (1+r_t^W)\tilde{b}_t, \tag{3}$$

$$\psi \frac{\tilde{c}_t}{1-\tilde{l}_t} = (1-\tau_{lt})(1-\alpha)\frac{\tilde{y}_t}{\tilde{l}_t}, \tag{4}$$

$$\frac{1}{\tilde{c}_t} = \frac{\beta}{1+\gamma} E_t \left\{ \frac{1}{\tilde{c}_{t+1}} (1+\tau_{bt+1})(1+r_{t+1}^W) \right\}, \tag{5}$$

$$\frac{1}{\tilde{c}_t} = \frac{\beta}{1+\gamma} E_t \left\{ \frac{1}{\tilde{c}_{t+1}} (1+r_{t+1}^k) \right\}, \tag{6}$$

$$1 + r_{t+1}^k \equiv \left[(1+\tau_{kt})\alpha \frac{\tilde{y}_t}{\tilde{k}_t} + q_t \left(1 - \partial \; - \phi\left(\frac{\tilde{l}_t}{\tilde{k}_t}\right) + \phi'\left(\frac{\tilde{l}_t}{\tilde{k}_t}\right)\frac{\tilde{l}_t}{\tilde{k}_t} \right) \right], \tag{7}$$

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 $q_t = (1-\phi'(\tilde{l}_t / \tilde{k}_t))^{-1}$ is Tobin's Q. Equation (3) denotes the economy's resource constraint; equations (4)-(6) are the household's first-order conditions; and equation (7) describes the evolution of the real rate of capital returns. The estimated of $\bar{r}^*$ is given by

$$\bar{r}^* = \frac{1}{T}\sum_{t=1}^{T} r_t^K$$

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 3). In contrast with Lama, we use quarterly frequency data instead of annual measurements, and we focus on the recent period, from 2006Q1 to 2017Q4 (the starting point of our sample is delayed because quarterly data for hours worked is only available from 2006).[24] The latter implies that we need to adjust certain calibrating

[24] Lama (2011) uses similar data for Mexico for the period 1991 to 2006 on an annual basis.

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 Carrillo *et al.* (2018).[25] Since population growth averaged 1.84% annually during this period, it turns out that the exogenous technological progress must equal 0.86% on an annual basis. The international real rate equals 4%, similar to Lama (2011). Given these numbers, we adjusted the discount factor β so that it satisfies equation (5) 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 closely followed the strategy of Lama. We used standard values for the depreciation rate ∂, the labor income shares 1-∂ for a Latin American economy, and the inverse of the elasticity of supply of international funds ν (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 $\eta = \vartheta^* \partial'$, and solve this expression to find ϑ.

Figure 3 shows that the estimated of $\overline{r}^*$ equals 2.3% from 2009 to 2017, which corresponds to the time period of the second business cycle considered in Carrillo *et al.* (2018). This estimate is located in 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.

[25] 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 $\overline{r}^*$ are quite similar.

Table 3

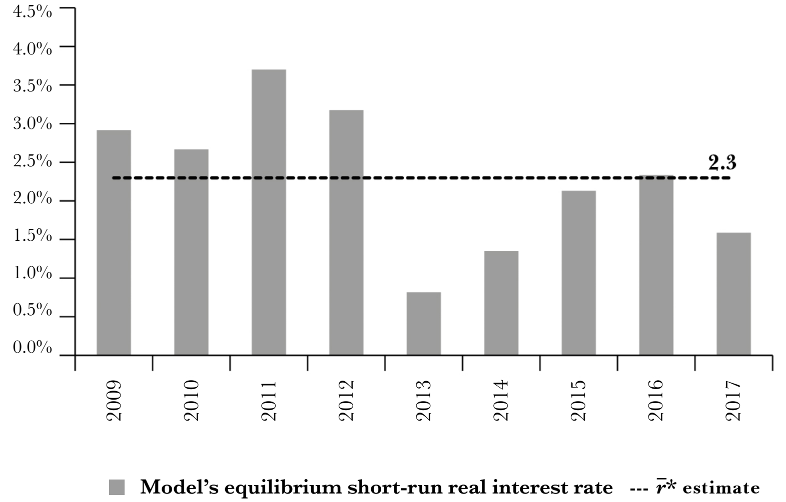
CALIBRATING AND ESTIMATING PARAMETERS FOR THE NEOCLASSICAL MODEL

	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	0.016
Depreciation rate	∂	5.00% *app*	$1-\tau_{kt}$	0.70	0.151
Discount factor	β	0.99	$1-\tau_{bt}$	0.95	4 × 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 int. 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.

Figure 3

REAL RATE OF CAPITAL RETURNS AND LONG RUN CONVERGENCE LEVEL OF r^*



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

4.3 Affine Term Structure Model of the Interest Rate

To compute an alternative estimate of $\overline{r}^*$ 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 Kim and Wright (2005) model, henceforth KW. We decided to use this model to estimate of $\overline{r}^*$ since such a model seems to better filter the effects of transitory factors than do alternative models, such as the one proposed by Adrian, Crump and Moench (2013). As a result, the KW model seems to capture the trend of r^* at long horizons, which is the object of our research.

The KW model assumes 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 the models is written in state-space form as

8 $$\mathbf{X}_t = \mu + \phi \mathbf{X}_{t-1} + \vartheta_{t+1}, \quad \text{[Transition Equation]}$$

9 $$i_t^n = \mathbf{A}_n + \mathbf{B}_n \mathbf{X}_t, \quad \text{[Measurement Equation]}$$

where $\boldsymbol{X}$ is a vector of factors or state variables, i^n is the nominal interest rate of a bond with maturity of n months, ϑ is white-noise state innovations, ϕ and $\mathbf{B}_n$ are coefficient matrices, and μ and $\mathbf{A}_n$ are coefficient vectors. The KW model includes three *latent* factors in vector $\boldsymbol{X}$, each as a proxy for the following yield curve characteristics: *i)* level, *ii)* slope, and *iii)* curvature. We estimate the coefficients of the model using maximum likelihood estimation and the Kalman filter. We use the 2004 to 2017 sample of the yields of government zero-coupon bonds with selected maturities of 1, 3, 6, 12, 24, 36, 60, 84, and 120 months.[26]

From the model, we can obtain the average expected path of the nominal interest rate of 1-month maturity bonds for horizons running from 1 month to n years ahead, namely,

10 $$E_{t,t+n}^i = \frac{1}{n} \sum_{k=1}^{n} E_t \left\{ i_{t+k}^{(1)} \right\}$$

Notice that form equation (10), we can recursively derive the long-run expectation of short-term nominal interest rate, i.e., $E_t \left\{ i_{t+k}^{(1)} \right\}$ where n corresponds to the future period in years. In particular, we have that:

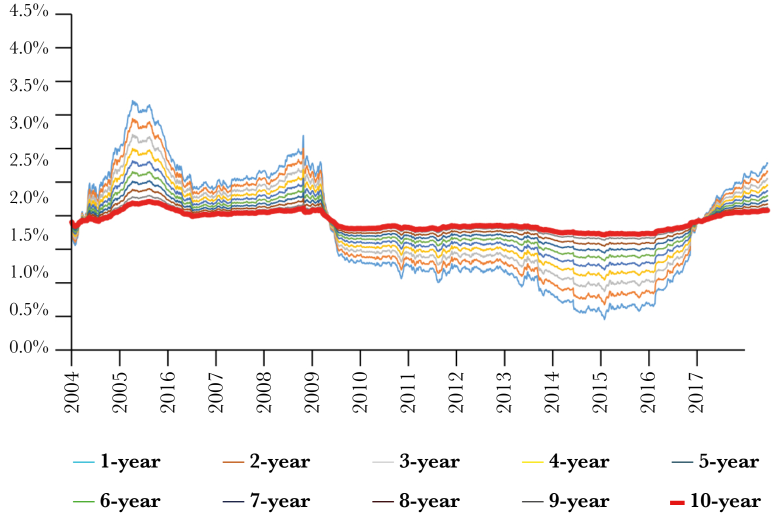
$$E_t \left[i_{t+1}^{(1)} \right] = E_{t,t+1}^i$$

$$E_t \left[i_{t+2}^{(1)} \right] = 2E_{t,t+2}^i - E_{t,t+1}^i$$

[26] More details about this methodology can be found in Kim and Wright (2005). In particular, the coefficients $\mathbf{A}_n$ and $\mathbf{B}_n$ are estimated recursively and depend on market risk parameters. When these parameters are equal to zero, we obtain the risk-free coefficients $\mathbf{A}_n^{RF}$ and $\mathbf{B}_n^{RF}$. Using these coefficients, we can compute the average expectation at time t of short-term interest rates over the next k periods.

Figure 4

ANNUAL EXPECTATIONS OF THE SHORT-RUN NOMINAL INTEREST IMPLICIT IN FINANCIAL INSTRUMENTS



Source: Own calculations with data of PiP and Valmer.

$$E_t\left[i^{(1)}_{t+3}\right] = 3E^i_{t,t+3} - 2E^i_{t,t+2}$$

$$\vdots$$

$$E_t\left[i^{(1)}_{t+10}\right] = 10E^i_{t,t+10} - 9E^i_{t,t+9}$$

Figure 4 shows annual expectations of the short-term nominal interest rate at different horizons. We use a horizon of $n = 10$ years as our measure of $\overline{r}^*$, since in that time period it is quite likely that even the most persistent transitory factors would have faded away (thick red line in Figure 4).[27]

Thus, the estimate of $\overline{r}^*$ in real terms, is given by

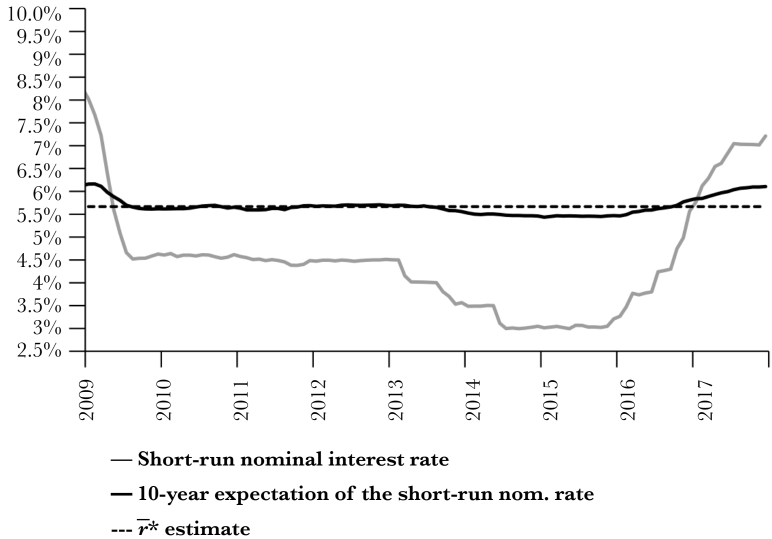
11 $$\overline{r}^* = E_t\left\{i^{(1)}_{t+10}\right\} - \overline{\pi},$$

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

[27] The mean square error between the 10-year nominal interest rate observed and estimated is 0.40.

Figure 5

LONG-RUN EXPECTATION OF SHORT-RUN NOMINAL INTEREST RATE IMPLICIT FINANCIAL INSTRUMENTS



Source: Own calculations with data of PiP and Valmer.

Figure 5 shows that the long-run expectation of the short-run nominal interest rate averaged 5.7% from 2009 to 2017, the time period that corresponds to the second business cycle considered Carrillo *et al.* (2018). 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, $\bar{r}^*$ 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.

4.4 Summary of Quantitative Methods for $\bar{r}^*$ and Outlook

Table 4 summarizes the results of the methodologies we use to compute plausible values for $\bar{r}^*$. The range for this rate, calculated from the average of the minimum and maximum levels obtained with each method, suggests that $\bar{r}^*$ could be located between 1.7% to 3.3% in real terms, and from 4.7% to 6.3% in nominal terms, with midpoints at 2.5% and 5.5%, respectively. To compute the latter, we simply added Banco de México's 3% inflation target.

Table 4

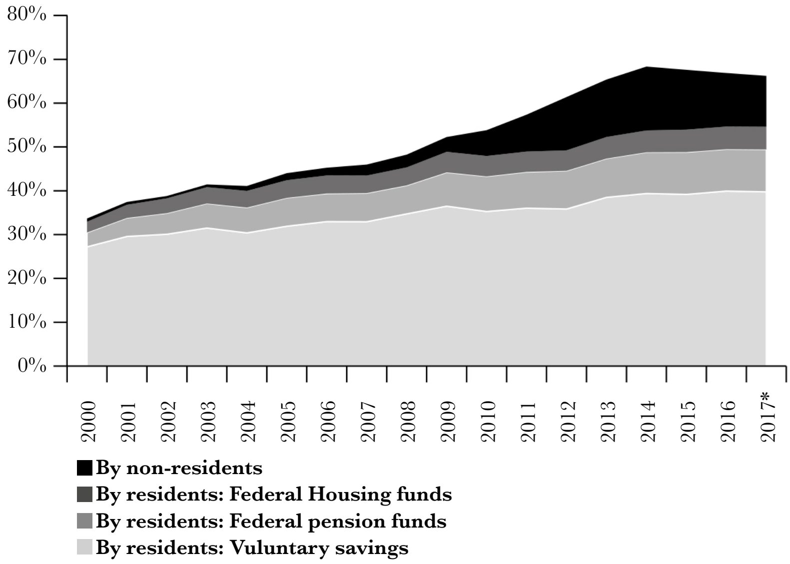
SUMMARY OF QUANTITATIVE RESULTS FOR THE LONG RUN

	Real Neutral $\bar{r}_t^*$		Nominal Neutral $\bar{r}_t^* + \pi_t^e$	
Methods	*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%.

Figure 6

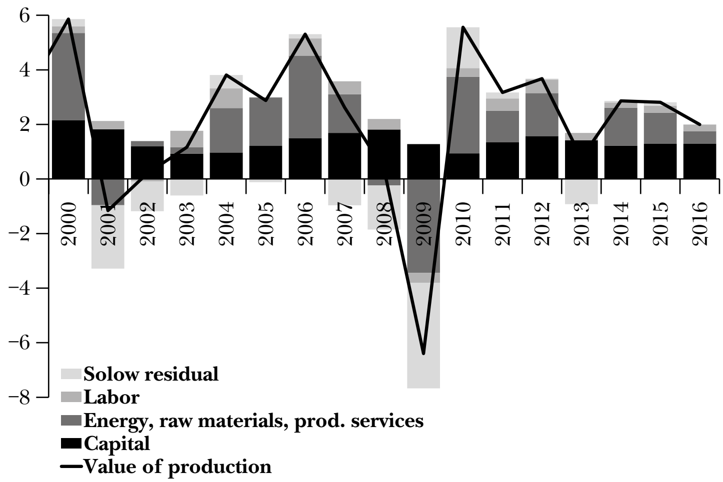
HOLDINGS OF DOMESTIC ASSETS



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

Figure 7

GROWTH ACCOUNTING



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

4.5 Heuristic Analysis of Structural Factors in Mexico

The outlook of $\overline{r}^*$ 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 early 2000s in Mexico. Voluntary savings by residents, distributed among public and private instruments, amounted to 40% of GDP in November 2017 as compared to 27% in 2000. In addition, federal pension and housing funds, a compulsory type of savings, composed 15% of GDP in November 2017 relative to 5.7% in 2000. In addition, domestic asset holdings of non-residents became significant only after 2008. Overall, the trends signal that the supply of loanable funds in the economy will continue to grow, which will exert a downward pressure on $\overline{r}^*$ in the future.

Population. Demographics have also played a role in the determination of $\overline{r}^*$ 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, 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 optimal ability to save in comparison with 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 it predicts that it might reach 0.6% by the end of the 2020s. If capital and labor are 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 relatively slowly. Demographics have, thus, posed downside risks to the long-run convergence level of r^* 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 year of available data. 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 $\bar{r}^*$. However, the latter might revert if the structural reforms recently implemented in Mexico boost productivity in the coming years. Part of these reforms encourages 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 (this reform is currently under revision by the new administration).

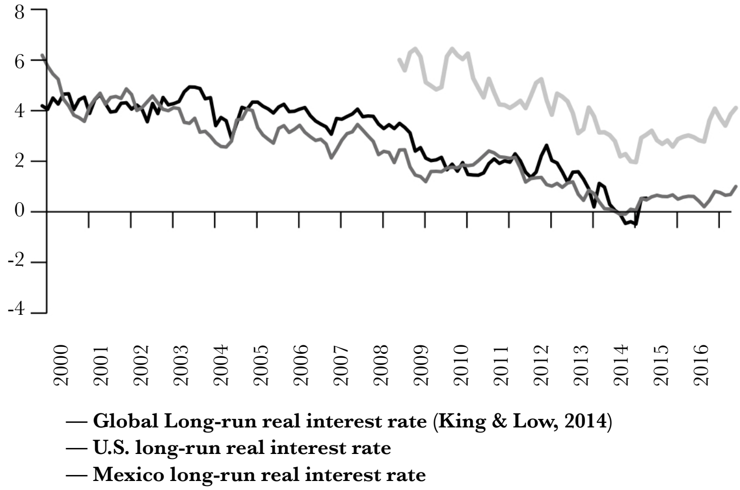
Global cost of money. Figure 8 compares the long-run real interest rates of the world (as computed by King and Low, 2014), the U.S., and Mexico, whose available data begin in 2002. It is noteworthy that the global long-run real interest rate has experienced a clear downward trend for at least 25 years, and Mexico does not seem to be insulated from such a path. Academics and policymakers have hotly debated the drivers behind this trend.[28]

Rachel and Smith (2015) argue 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

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

Figure 8

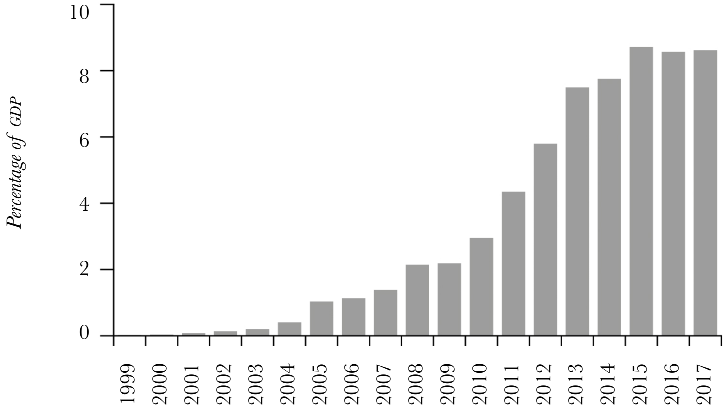
LONG-RUN REAL INTEREST RATES



Source: Own calculations with data from King and Low (2014), the Federal Re- serve, PiP, Valmer and Aguilar-Argaez, Elizondo and Roldán-Peña (2016).

Figure 9

HOLDINGS OF LONG-TERM GOVERMENT BONDS BY NON-RESIDENTS



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

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 negatively affected 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 lower global growth could have contributed to a fall of 100 basis points in the global long-run interest rate.

The global factors just described seem to have affected the Mexican long-run interest rate through international arbitrage. Figure 9 shows that there is a robust increase in the purchases of long-term public debt by foreign investors over the sample period studied. Non-residents holdings of these instruments seem to have accelerated since 2010, growing from 2.2% of GDP in 2009 to about 9% of GDP by 2014. These holdings have stabilized around that level for the last three years of the sample. Part of this acceleration may be due to the inclusion of Mexican peso-denominated debt in the Citigroup's World Government Bond Index (WGBI) in October 2010. This index is used as a benchmark by institutional investors who aim to buy highly-rated long-term debt.

In sum, the data suggest that capital inflows have permanently increased the supply of loanable funds in Mexico in recent years.

5. CONCLUDING REMARKS

In this paper, we argue that foreign and domestic structural factors, such as increasing domestic savings, demographic shifts, and a decreasing global long-run real interest rate, appear to explain the apparent fall in the Mexican neutral rate, from 3% in the period 2001-2008 to 2.5% after 2009. Going forward, 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 larger working-age of the 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.

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