Inflation Expectations and Risk Premia in Emerging Bond Markets: Evidence from Mexico

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Abstract

To study inflation expectations and associated risk premia in emerging bond markets, we provide estimates for Mexico based on an arbitrage-free dynamic term structure model of nominal and real bond prices that accounts for their liquidity risk. Beyond documenting the existence of large and weakly correlated liquidity premia in nominal and real bond prices, our results indicate that long-term inflation expectations in Mexico are well anchored close to the Bank of Mexico's inflation target. Furthermore, Mexican inflation risk premia are larger and more volatile than those in Canada and the United States.

JEL Classification: D84, E31, E43, E44, E47, E52, E58, G12

Keywords: term structure modeling, liquidity risk, financial market frictions, central bank credibility

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We thank participants at the Southern Finance Association 2021 Annual Meeting, the 34th Australasian Finance and Banking Conference, the Financial Markets and Corporate Governance Conference 2022, and the IBEFA 2022 Summer Meeting for helpful comments, including our discussants I-Hsuan Ethan Chiang, Lorenzo Menna, Hee Su Roh, and Weidong Tian. We also thank seminar participants at the Bank of Mexico, the Federal Reserve Bank of New York, and the Central Bank of Chile for helpful comments and suggestions. Furthermore, we thank Bernardo Keiserman for useful comments and suggestions. Finally, we thank Gus Kmetz for excellent research assistance. The views in this paper are solely the responsibility of the authors and should not be interpreted as reflecting the views of the Federal Reserve Bank of New York, the Federal Reserve Bank of San Francisco, or the Federal Reserve System.

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This version: November 1, 2022.

1 Introduction

Breakeven inflation (BEI)—the difference between yields on comparable-maturity nominal and real debt—is a widely used indicator of inflation expectations. In particular, long-term BEI is frequently used to measure the credibility of the central bank's inflation objective.[1] However, BEI is a noisy measure of expected inflation because it contains both an inflation risk premium and differential liquidity premia. As a consequence, better measures of investors' underlying inflation expectations could be obtained by subtracting both inflation risk premia and the differential liquidity premia in nominal and real yields from BEI rates.

The literature on inflation expectations and associated inflation risk premia extracted from nominal and real yields in advanced economies is burgeoning.[2] However, these topics have received far less attention for emerging economies, although they arguably matter more there for policymakers and bond investors due to higher macroeconomic uncertainty, in general, and larger inflation variability, specifically. The main contribution of this paper is to fill this gap by building on recent advances in fixed-income analysis.

The challenge in accounting for the differential liquidity premia in nominal and real bond prices is to distinguish them from more fundamental factors such as inflation risk premia that would affect asset prices even in a world without any frictions to trading. To achieve this separation, we augment a flexible dynamic term structure model of nominal and real bond prices studied in Carriero et al. (2018) with separate liquidity risk factors for nominal and real bonds using the approach described in Andreasen et al. (2021, henceforth ACR). For each class of bonds, the identification of the liquidity risk factor comes from its unique loading, which mimics the idea that, over time, an increasing amount of the outstanding notional of individual securities gets locked up in buy-and-hold investors' portfolios. This increases their sensitivity to variation in the market-wide liquidity risk captured by the corresponding liquidity risk factor. By observing prices for panels of nominal and real bonds, their respective liquidity risk factors can be separately identified. This separation is particularly salient in emerging bond markets as they tend to be much less deep and liquid than the well-established major international bond markets in the United States and elsewhere.

To better understand the properties and dynamics of inflation expectations and associated risk premia in emerging bond markets, we choose to focus on a country where they are likely to play a first-order role, namely Mexico. Several motivations underlie this choice. First, Mexico has a long history of high and fairly volatile inflation.[3] Second and equally important

[1]Provided the objective is credible, it should be reflected in inflation expectations for the distant future as any current inflation shocks should be considered temporary and not affect long-run inflation expectations.

[2]For Canada, see Christensen et al. (2021); for the euro area, see Hördahl and Tristani (2012, 2014); for the United Kingdom, see Joyce et al. (2010) and Carriero et al. (2018); for the United States, see Christensen et al. (2010), Abrahams et al. (2016), D'Amico et al. (2018), among many others.

[3]For the 2010-2019 period, the year-over-year inflation in Mexico as measured by the consumer price index averaged 3.96 percent with a standard deviation of 1.02 percent. For comparison, the corresponding statistics in the United States were 1.75 percent and 0.85 percent, respectively, while the matching statistics for Canada were 1.74 percent and 0.65 percent, respectively.

for our analysis, it has well-functioning markets for both standard nominal fixed-coupon government bonds, so-called bonos, and real inflation-indexed government bonds, known as udibonos. Finally, we consider Mexico and its government bond market to be representative of the wider set of large emerging bond markets; for example, it typically makes up between 10 percent and 13 percent of most emerging market bond indices.

In terms of our empirical findings, we make a number of observations. First, our results show that the average liquidity premia embedded in both nominal and real Mexican bond yields exhibit notable time variation. For nominal yields, our sample covers the period from January 2007 through December 2019, and their estimated liquidity premia average 45 basis points with a standard deviation of 22 basis points. For real yields, our sample contains data from May 2009 through December 2019, and their estimated liquidity premia average 48 basis points with a standard deviation of 34 basis points. Thus, the liquidity premia of Mexican inflation-indexed government bonds are larger and more variable than those of standard Mexican nominal government bonds. These results are consistent with the findings of ACR, who report that the average liquidity premium in U.S. Treasury Inflation-Protected Securities (TIPS) is estimated at 34 basis points for the 1997-2013 period, which is well above measures of liquidity premia in regular U.S. Treasury bonds. The difference in liquidity premium levels across the U.S. TIPS and the Mexican udibonos markets is likely to be due to the much greater relative liquidity of U.S. Treasury securities. Importantly, the nominal and real bond liquidity risk premia we estimate are practically uncorrelated in levels and their monthly changes are only mildly positively correlated. These results suggest that inflation-indexed Mexican bonds are less liquid and less desirable than nominal Mexican bonds.

Second, we find that the model delivers estimates of investors' inflation expectations that are robust to a range of different model implementations. In addition, we compare the calendar-year ahead model-implied inflation forecasts to those reported in the monthly Consensus Forecasts and find our model to be reliable, but slightly less accurate than the survey forecasts. Furthermore, we examine the dynamic properties of the model-implied inflation expectations using both a statistical measure from the literature and model-based projections of the outlook for long-term inflation expectations in Mexico. The results suggest that bond investors' long-term inflation expectations are likely to remain close to the 3 percent inflation target of the Bank of Mexico in the foreseeable future, which we take as a sign that long-term inflation expectations are well anchored in Mexico. Crucially, model updates through September 2022 continue to support this conclusion.

Third, the stable anchoring of investors' long-term inflation expectations in Mexico implies that most of the variation in the liquidity-adjusted BEI rates is driven by fluctuations in the inflation risk premium. As for the determinants of Mexican inflation risk premia, we perform regression analysis with a large battery of explanatory variables. The regressions have large explanatory power with adjusted R^2s of 0.75 and above. Using informed priors allows us to identify four variables of particular importance, namely the foreign share of the bonos market, oil prices, the VIX, and the ten-year U.S. Treasury yield. While the foreign share of the bonos market has a negative effect on Mexican inflation risk premia, the other three variables have positive coefficients meaning that increases in oil prices, global risk aversion, and U.S. long-term interest rates tend to boost long-term inflation risk premia in Mexico. Importantly, model updates during the COVID-19 pandemic and post-pandemic period show a large increase in the inflation risk premium in response to these shocks consistent with the predictions from our preferred regression model.

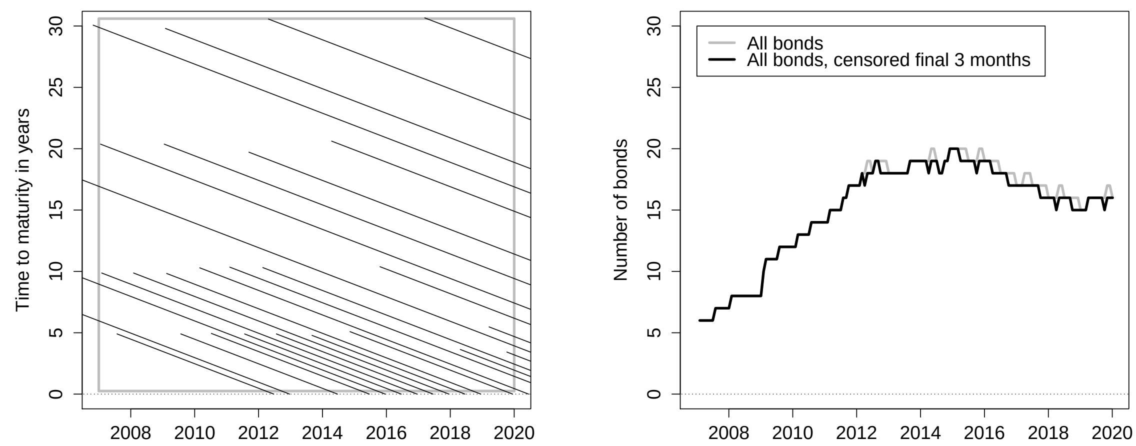
Finally, we compare our estimated inflation risk premium series to estimates from Canada and the United States and find them to be positively correlated, although the estimated Mexican inflation risk premium is clearly larger and more volatile, as anticipated. Still, for extended periods, inflation risk in Mexico only commands a premium slightly above those seen in Canada and the United States, which is somewhat surprising. These findings support the claim that long-term inflation expectations in Mexico are well anchored, as also noted by De Pooter et al. (2014), and they underscore that the inflation target of the Bank of Mexico is viewed as being credible by financial market participants.[4]

The remainder of the paper is structured as follows. Section 2 describes the data, while Section 4 details the model and empirical results. Section 5 examines the BEI decomposition and scrutinizes the estimated inflation risk premia. Finally, Section 6 concludes. An online appendix contains details of the Mexican government bond data along with additional analysis, estimation results, and robustness exercises.

2 Mexican Government Bond Data

This section first describes the Mexican government bond data we use in the model estimation before we proceed to a discussion of the credit risk, bond holdings, and bid-ask spreads in the markets for these bonds.

[4] The ability of the Bank of Mexico to affect asset prices, the exchange rate, and portfolio flows through its conduct of monetary policy is documented in Solís (2020).



(a) Distribution of bonos

(b) Number of bonos

Figure 1: **Overview of the Mexican Government Bonos Data**
Panel (a) shows the maturity distribution of the Mexican government fixed-coupon bonos considered in the paper. The solid gray rectangle indicates the sample used in the empirical analysis, where the sample is restricted to start on January 31, 2007, and limited to bonos prices with more than three months to maturity after issuance. Panel (b) reports the number of outstanding bonos at a given point in time.

2.1 Bonos

The available universe of individual Mexican government fixed-coupon bonds, known as bonos, is illustrated in Figure 1(a).[5] Each bond is represented by a solid black line that starts at its date of issuance with a value equal to its original maturity and ends at zero on its maturity date. These bonds are all marketable non-callable bonds denominated in Mexican pesos that pay a fixed rate of interest semiannually. We note that we track the entire universe of bonos issued since January 2007. In addition, we include a set of bonos outstanding at the start of our sample period identical to those analyzed by Christensen et al. (2021, henceforth CFS). As a consequence, our sample entirely encompasses theirs. In general, the Mexican government has been issuing five-, ten-, twenty- and thirty-year bonos on a fairly regular basis during the period shown. As a result, there is a wide variety of bonds with different maturities and coupon rates in the data throughout our sample. It is this variation that provides the foundation for the econometric identification of the factors in the yield curve models we use.

Figure 1(b) shows the distribution across time of the number of bonds included in the sample. We note a gradual increase from six bonds at the start of the sample to 16 at its end. Combined with the cross sectional dispersion in the maturity dimension observed in Figure 1(a), this implies that we have a very well-balanced panel of Mexican bonos prices.

Figure 2 shows the time series of the yields to maturity implied by the observed Mexican

[5]The contractual characteristics of all 28 bonos securities in our sample are reported in the online appendix.

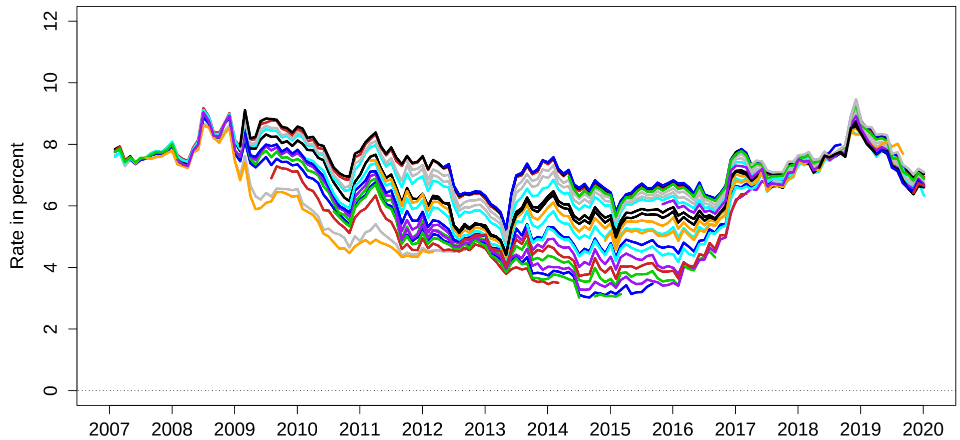


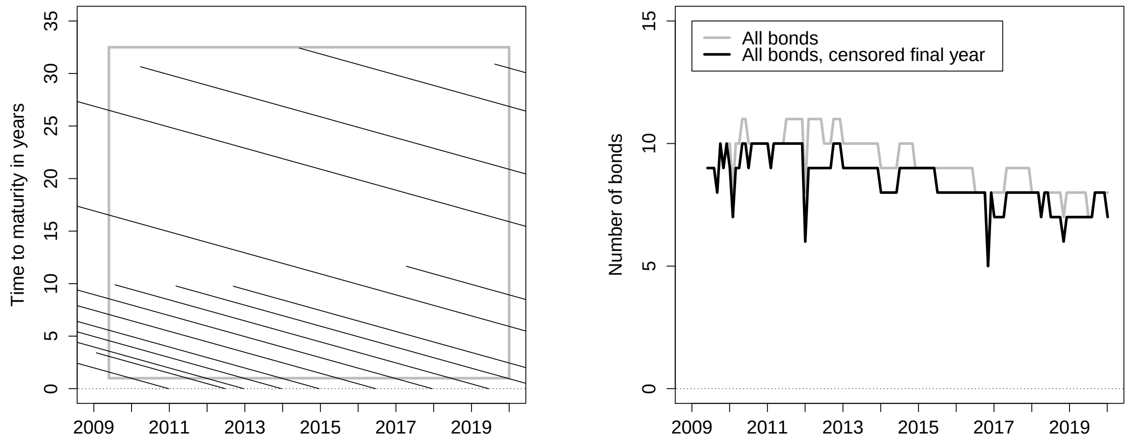
Figure 2: **Yield to Maturity of Mexican Government Bonos**
Illustration of the yields to maturity implied by the Mexican government fixed-coupon bonos prices downloaded from Bloomberg. The data is monthly covering the period from January 31, 2007, to December 31, 2019, and censors the last three months for each maturing bond.

bonos prices downloaded from Bloomberg. First, we note that the general yield level in Mexico has been fairly stable since 2007, unlike government bond yields in advanced economies, which have declined significantly during this period, see Holston et al. (2017) and Christensen and Rudebusch (2019), among others. Second, as in U.S. Treasury yield data, there is notable variation in the shape of the yield curve. At times like in mid-2018, yields across maturities are relatively compressed. At other times, the yield curve is steep with long-term bonos trading at yields that are 300-400 basis points above those of shorter-term securities like in 2015.

Finally, regarding the important question of a lower bound, the Bank of Mexico has never been forced to lower its conventional policy rate even close to zero, and the bond yields in the data have remained well above zero throughout the sample period. Thus, there is no need to account for any lower bounds to model these fixed-coupon bond prices, which motivates our focus on Gaussian models.

2.2 Udibonos

The Mexican government also issues inflation-indexed bonds, known as udibonos. These bonds pay semiannual interest based upon a real interest rate. Unlike standard fixed-coupon marketable bonds, interest payments on udibonos are adjusted for changes in the general price level. Technically, their payoff is measured in a unit called Unidad de Inversión (UDI), which is calculated and published daily by the Bank of Mexico. UDI changes with the biweekly release of the National Consumer Price Index, abbreviated INPC in Spanish, according to



(a) Distribution of udibonos

(b) Number of udibonos

Figure 3: **Overview of the Mexican Government Udibonos Data**
Panel (a) shows the maturity distribution of the Mexican government inflation-indexed udibonos considered in the paper. The solid gray rectangle indicates the sample used in the empirical analysis, where the sample is restricted to start on May 29, 2009, and limited to udibonos prices with more than one year to maturity after issuance. Panel (b) reports the number of outstanding bonos at a given point in time.

the procedure determined by the Bank of Mexico as originally laid out in Mexico's Federal Official Gazette on April 4, 1995. UDI represents the accumulated inflation in Mexico since April 1, 1995, denominated in Mexican pesos, and it is the factor used to convert the real return of udibonos into the corresponding value measured in current pesos at any given point in time.

The Mexican government launched its udibonos program in 1996. However, due to the quality of the available data from Bloomberg, we are limited to starting our sample in May 2009. The available universe of udibonos and their maturity distribution across time is shown in Figure 3(a).[6] It includes the entire universe of udibonos issued since May 2009 combined with the outstanding stock of udibonos at the start of our sample. We note that the issuance is concentrated in ten-year udibonos with occasional issuance of twenty- and thirty-year udibonos.

The total number of udibonos in our sample across time is shown in Figure 3(b). As with the nominal bonos, we stress that the sample of udibonos we use is very well-balanced across maturities at all times, which underpins the econometric identification of the state variables in the term structure model we use.

Figure 4 shows the yields to maturity implied by the udibonos prices. Similar to what we observe for the nominal bonos yields, the yields of udibonos have fluctuated around a fairly

[6] The contractual details of all 16 udibonos in our sample are reported in the online appendix.

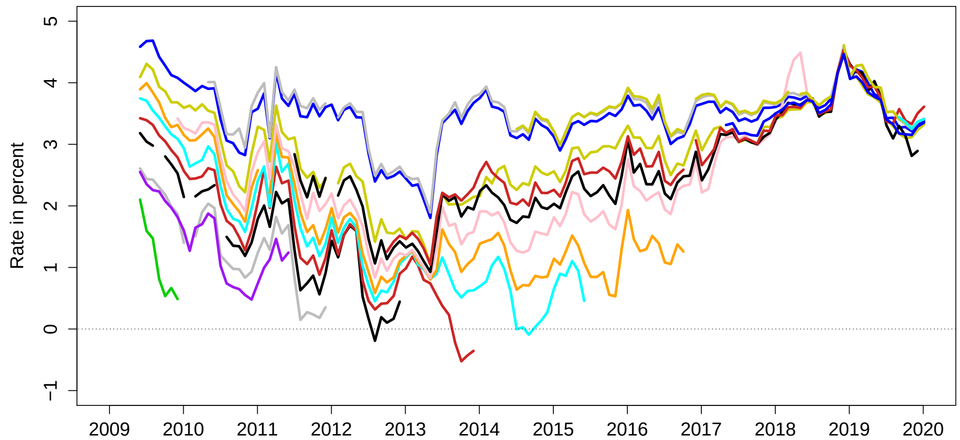


Figure 4: **Yield to Maturity of Mexican Government Udibonos**
Illustration of the yield to maturity implied by the Mexican government inflation-indexed udibonos prices considered in this paper, which are subject to two sample choices: (1) sample limited to the period from May 29, 2009, to December 31, 2019; (2) censoring of a bond's price when it has less than one year to maturity.

stable level during the shown period, but with some variation in the steepness of the udibonos yield curve. Our model exploits this variation to deliver estimates of their liquidity premia, as explained in Section 4. Also, the greater dispersion in the udibonos yields across maturities in the early part of the sample is a tangible sign that bond-specific premia in this market are likely to play a particularly large role during this period.

2.3 The Credit Risk of Mexican Government Bonds

To gauge whether there are any material credit risk issues to consider in modeling Mexican government bond prices, we use rates on credit default swap (CDS) contracts. They reflect the annual rate investors are willing to pay to buy protection against default-related losses on these bonds over a fixed period of time stipulated in the contract. Such contracts have been used to price the credit risk of many countries, including Mexico, since the early 2000s.

In Figure 5, we plot the series for the one- and five-year Mexican CDS rate since 2007 with solid gray and black lines, respectively. Also shown with a solid red line is the spread between these two CDS rates. We note that the five-year CDS rate has fluctuated in a fairly narrow range between 100 and 200 basis points, except for a brief period during the global financial crisis when Mexican CDS rates temporarily spiked above 300 basis points. This is a level of credit risk on par with most investment-grade firms in the United States, and its variation is mostly very gradual. This suggests that credit risk-related components are unlikely to be the driver of the results we present later on. Furthermore, we note that

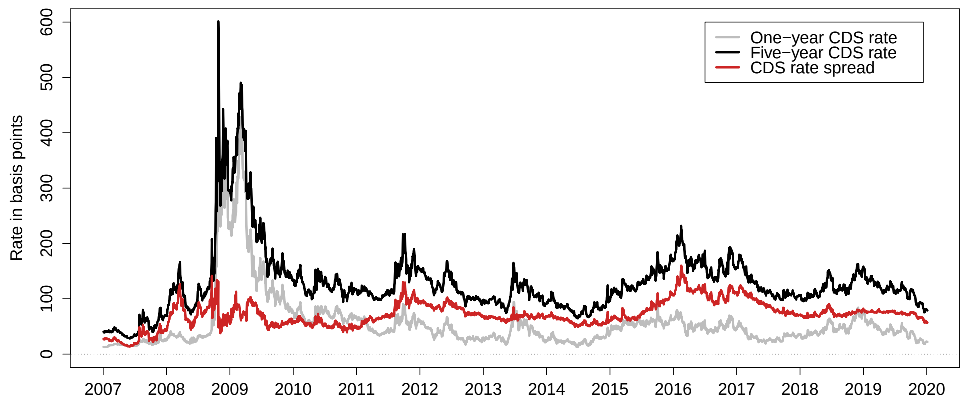
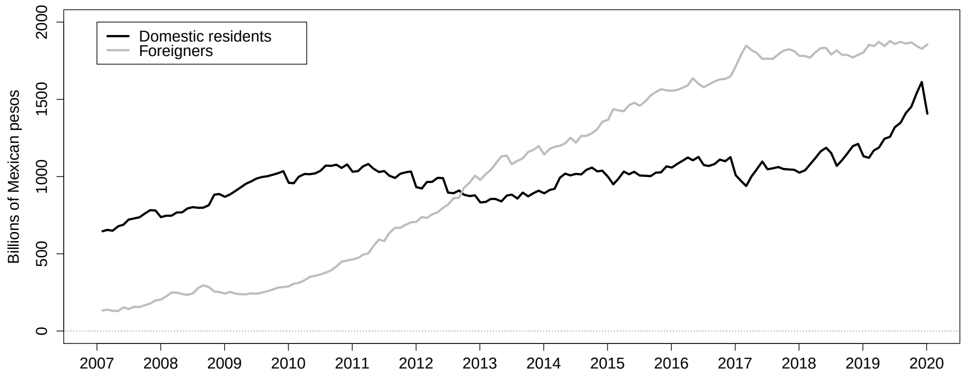


Figure 5: **Mexican CDS Rates**

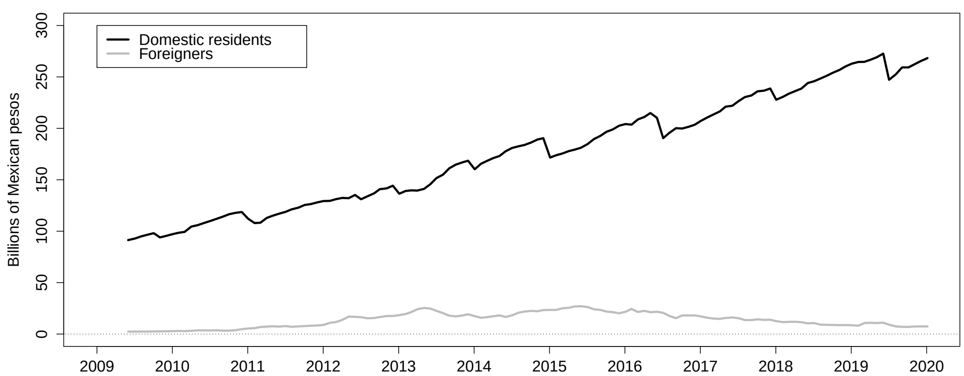
our measure of the Mexican government debt relative to GDP never goes above 52 percent, which is not a high value by international standards. Finally, the slope of the CDS rate curve measured as the difference between the five-year and one-year CDS rates is always positive, is fairly stable, fluctuates in a narrow range, and is mostly uncorrelated (11%) with the one-year CDS rate. Thus, the steepness of the CDS rate curve for Mexican government debt has little connection to the near-term level of the priced credit risk of the Mexican government. We take this as a sign that the bulk of the variation in Mexican CDS rates reflect investor sentiment and risk aversion rather than actual credit risk.[7] This view finds further support in Gamboa-Estrada and Romero (2022), who analyze CDS rates across Latin-American countries, including Mexico, and find that their levels are mainly driven by a common component and global financial conditions leaving little room for country-specific factors in their determination. Overall, we take this evidence to imply that credit risk is not likely to materially affect our results, and we are therefore comfortable not accounting for credit risk premia in our analysis.

More importantly, on a practical note, there are no differences in the credit risk of bonos and udibonos in the sense that they will receive the same treatment in case the Mexican government stops servicing its debt. Thus, using arguments similar to those made by Fleckenstein et al. (2014) for U.S. Treasuries and TIPS, there is no reason to believe that there are any differentials in the pricing of bonos and udibonos tied to credit risk. By implication, our measures and decompositions of Mexican BEI are unaffected by variation in the credit risk premia of Mexican government debt.

[7]This is a phenomenon also seen in the pricing of corporate bonds and frequently referred to as the credit spread puzzle; see Christensen (2008) and references therein.



(a) Bonos



(b) Udibonos

Figure 6: **Holdings of Mexican Bonos and Udibonos**

2.4 Domestic and Foreign Mexican Government Bond Holdings

In addition to the bond price data described above, our regression analysis later on utilizes data on domestic and foreign holdings of Mexican government debt securities that the Bank of Mexico requires financial intermediaries to report as a way to track market activity in the Mexican sovereign bond markets. These data have been collected since 1978 and are available at a daily frequency up to the present. A key strength of the data set is that it covers any change in Mexican government debt holdings by either domestic or foreign investors. For each transaction, the reporting forms also identify the type of Mexican government security. Therefore, we are able to exploit the data reported for holdings of both bonos and udibonos. Although the data are available at a daily frequency, we use only the observations at the end of each month to align them with our bond price data.

Figure 6(a) shows the monthly level of bonos holdings by domestic residents and foreigners over the period from January 2007 through December 2019. We note that foreigners overtook domestic residents in total holdings by late 2012 and have continued to increase their share

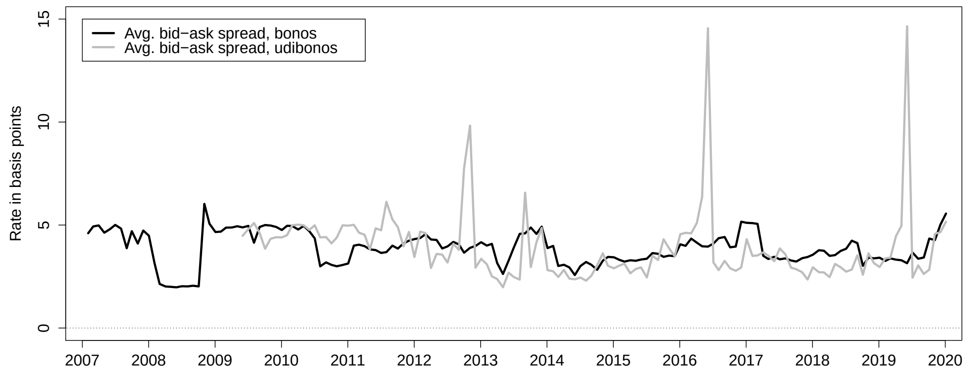


Figure 7: **Bid-Ask Spreads of Mexican Bonos and Udibonos**

quite notably such that they now exceed those of domestic residents by a wide margin. In contrast, for the udibonos holdings shown in Figure 6(b) covering the period from May 2009 through December 2019, we note that only a very small share of this market is held by foreigners. We note that the difference in the foreign participation across the two markets could lead to significant differences in their trading dynamics and perceived liquidity risks.

To provide a sense of the relative size of the market for udibonos, we note that, as of December 31, 2019, the total outstanding amount of bonos was 3,261 billion pesos, or about USD160 billion, while the total amount of udibonos outstanding was 276 billion pesos, or only about USD14 billion.[8] Hence, udibonos represent about 8.5 percent of the government's long-term debt denominated in Mexican pesos. Finally, we add that the total outstanding domestic debt of the Mexican federal government was 7,586 billion pesos, or almost USD380 billion, while its gross long-term foreign debt is reported as USD102 billion as of September 30, 2019.[9] Thus, the vast majority of the government's debt is issued in local currency, which underscores the importance of the domestic government bond markets analyzed here.

2.5 Bid-Ask Spreads of Mexican Government Bonds

In this section, to shed light on the trading frictions in the markets for bonos and udibonos, we compare the average bid-ask spread of the udibonos in our sample to the average bid-ask spread of the bonos in our sample. These series are four-week moving averages and shown in Figure 7. Note that the two series tend to be close to each other. Thus, most of the time, there is no discernible difference in this measure of current liquidity across the two markets. Still, there are occasional large spikes in the average bid-ask spread of udibonos, which are driven

[8] See data from the Bank of Mexico at: https://www.banxico.org.mx/SieInternet/consultarDirectorioInternetAction.do?sector=7&accion=consultarDirectorioCuadros&locale=en

[9] See CEIC data at: https://www.ceicdata.com/en/mexico/gross-external-debt/gross-external-debt-federal-government-long-term

by erratic pricing of individual udibonos as they approach maturity. This phenomenon is much less pronounced in the bonos market and further mitigated by the much larger number of bonos in our sample.

The key takeaway from this evidence is twofold. First, with a level of bid-ask spreads around 5 basis points, the trading of these securities is indeed associated with some amount of liquidity risk. Second, the occasional large spikes in the bid-ask spreads of udibonos would be another sign that the liquidity risk of these securities is somewhat greater than that of bonos.

Motivated by the evidence in this section, our model assumes that both nominal and real bond prices contain liquidity premia that investors demand to assume their liquidity risk. The purpose of the remainder of the paper is to quantify the relative importance of these bond risk premia in the pricing of bonos and udibonos and what adjustments for them may imply about bond investors' underlying inflation expectations and associated inflation risk premia.

3 Analysis of Nominal and Real Bond-Specific Factors

Motivated by the evidence in the previous section, the dynamic term structure model we use in the empirical analysis assumes that both nominal and real bond prices contain liquidity premia that investors demand to assume their liquidity risk. In this section, we aim to build further support for that assumption. In doing so, we follow D'Amico et al. (2018, henceforth DKW), who note that, in a world without any financial market frictions, nominal yields y_t^N must be the sum of the matching real yield y_t^R, expected inflation π_t^e, and the inflation risk premium ϕ_t:

$$y_t^N(\tau) = y_t^R(\tau) + \pi_t^e(\tau) + \phi_t(\tau). \tag{1}$$

Furthermore, the frictionless real yield is the sum of the neutral or natural real rate r_t^* and a real term premium:

$$y_t^R(\tau) = r_t^* + TP_t^R(\tau).$$

In turn, the frictionless BEI that would prevail in a world without any financial frictions is then given by

$$BEI \equiv y_t^N(\tau) - y_t^R(\tau) = \pi_t^e(\tau) + \phi_t(\tau). \tag{2}$$

Comparing equations (1) and (2) motivates DKW to regress BEI on the first three principal components (PC) of nominal yields, which normally explain more than 99 percent of the nominal yield variation. If there are no priced frictions or other deviations from the law of one price in the data, this could be expected to yield high R^2s, in particular if the frictionless real yields y_t^R have stationary dynamics, as suggested by our Mexican data.

As for the observed nominal and real yields, denoted $\overline{y}_t^N$ and $\overline{y}_t^R$, respectively, they may each contain unobserved liquidity premia, denoted η_t^N and η_t^R, respectively. Hence, we have the following relationships to the frictionless yields discussed above:

$$\begin{aligned} \overline{y}_t^N(\tau) &= y_t^N(\tau) + \eta_t^N(\tau), \\ \overline{y}_t^R(\tau) &= y_t^R(\tau) + \eta_t^R(\tau). \end{aligned}$$

This implies that the observed nominal yield can be written as

$$\overline{y}_t^N(\tau) = r_t^* + TP_t^R(\tau) + \pi_t^e(\tau) + \phi_t(\tau) + \eta_t^N(\tau),$$

while the observed BEI becomes

$$\overline{BEI}_t(\tau) \equiv \overline{y}_t^N(\tau) - \overline{y}_t^R(\tau) = \pi_t^e(\tau) + \phi_t(\tau) + \eta_t^N(\tau) - \eta_t^R(\tau).$$

Returning to DKW's proposed regression, it is now clear that, provided $\eta_t^N(\tau) - \eta_t^R(\tau) \approx 0$ and r_t^* and $TP_t^R(\tau)$ are stationary, it will be the case that observed $\overline{BEI}$ regressed on the first three PCs of nominal yields should generate fairly large R^2s. On the other hand, if either of these two assumptions are violated, we are likely to see fairly low R^2, which then suggests that either (a) the frictionless real yields contain some trending component; or (b) nominal and/or real yields contain some sizable persistent priced frictions or liquidity premia that prevent the condition $\eta_t^N(\tau) - \eta_t^R(\tau) \approx 0$ from being satisfied.

DKW study U.S. Treasury and TIPS data, where it is reasonable to assume that, indeed, $\eta_t^N(\tau) \approx 0$. Also, they implicitly assume that there are no trends in TIPS yields, although that may be a questionable assumption given that evidence provided in Laubach and Williams (2016) and Christensen and Rudebusch (2019), among many others, points to a long-term secular decline in the natural real rate in the United States. As a consequence, when they obtain really low R^2s in their regressions, DKW conclude that TIPS yields contain a significant liquidity premium η_t^R.

For our Mexican data, we first estimate arbitrage-free Nelson-Siegel (AFNS) models from Christensen et al. (2011) for the Mexican bonos and udibonos prices separately. This gives us fitted nominal and real yield curves at all relevant maturities, which are then used to calculate the corresponding fitted BEI rates. We then regress those fitted BEI rates on the three filtered state variables from the AFNS model estimation based on our sample of Mexican bonos prices, which serve as our equivalent of the first three principal components of nominal yields in the analysis of DKW:

$$\widehat{BEI}_t(\tau) = \alpha^\tau + \beta_L^\tau \widehat{L}_t + \beta_S^\tau \widehat{S}_t + \beta_C^\tau \widehat{C}_t + \varepsilon_t^\tau.$$

The results at four maturities from two to ten years are reported in Table 1. The level and slope factors in the AFNS model are highly statistically significant across all considered

Coef.	BEI			
	$\tau = 2$	$\tau = 5$	$\tau = 7$	$\tau = 10$
α	0.01	0.01	0.00	0.00
	(0.01)	(0.00)	(0.00)	(0.00)
β_L	0.34**	0.40**	0.43**	0.46**
	(0.06)	(0.06)	(0.06)	(0.06)
β_S	0.13**	0.14**	0.13**	0.13**
	(0.03)	(0.02)	(0.02)	(0.02)
β_C	0.03	0.06**	0.07**	0.07**
	(0.02)	(0.02)	(0.02)	(0.02)
R^2	0.40	0.52	0.58	0.66

Table 1: **Breakeven Inflation Regressions**
The table reports the results of regressions with breakeven inflation as the dependent variable and the estimated level, slope, and curvature factors from an AFNS model of nominal bonos prices. Standard errors computed by the Newey-West estimator (with three lags) are reported in parentheses. Asterisks * and ** indicate significance at the 5 percent and 1 percent levels, respectively.

maturities and have very stable coefficients. As for the curvature factor, its loading is insignificant at shorter maturities, but highly statistically significant at medium- and long-term maturities. This is consistent with its hump-shaped loading structure across maturities.

Despite this very stable and significant pattern in the regression coefficients of the three factors, the obtained R^2 values fall in the range from 0.40 to 0.66 and decline as the maturity shortens. DKW interpret this kind of pattern as evidence of the existence of liquidity premia in the underlying bond yields. In our case, a similar interpretation is likely to apply. First, in our Mexican data, the udibonos real yields do not appear to have a trend. Hence, when we get R^2s ranging from 0.40 to 0.66, it is most likely caused by $\eta_t^N(\tau) - \eta_t^R(\tau) \approx 0$ not being satisfied. Second, the decline in R^2 as maturity is shortened is consistent with a liquidity premium interpretation as well given that yields at shorter maturities in our data primarily reflect the prices of seasoned bonds that have been outstanding for many years. As a consequence, shorter-term BEI rates are likely to be more biased by liquidity premia than the ten-year sector where a majority of the bond issuance has taken place historically. To summarize, building on the findings of DKW for the large U.S. TIPS market, where they find sizable liquidity premia, our regression results imply the existence of large and time-varying liquidity premia in the much smaller market for udibonos.

To provide evidence of the existence of important liquidity premia in Mexican bonos beyond that provided in CFS, we again follow DKW. For U.S. data, they document that TIPS BEI tends to be *below* the inflation forecasts reported in surveys of both consumers and professional forecasters. Furthermore, as demonstrated by DKW, this is due to the existence of large positive liquidity premia in TIPS yields in combination with small and negligible liquidity premia in Treasury yields.

To repeat this exercise in our setting, Figure 8 compares the ten-year fitted BEI considered

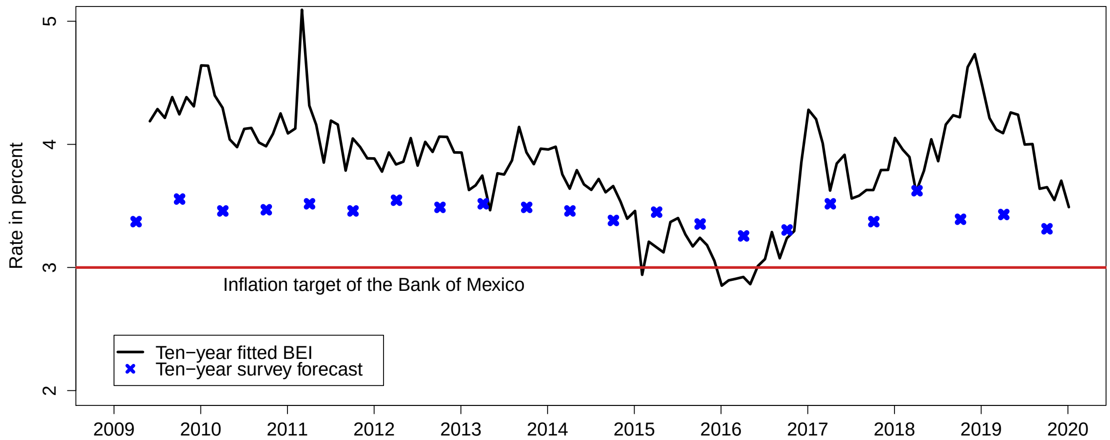


Figure 8: **Survey Inflation Forecasts and Fitted BEI**
Illustration of the ten-year fitted BEI obtained by fitting an AFNS model to Mexican bonos and udibonos prices separately. Also shown are the ten-year inflation forecasts from the semiannual Consensus Forecasts survey of professional forecasters tracking the Mexican economy.

earlier with the ten-year CPI inflation forecasts that can be constructed from the long-term economic forecasts reported semiannually in the Consensus Forecasts surveys of professional forecasters tracking the Mexican economy.

In the Mexican data, we see the opposite pattern of DKW whereby BEI tends to be *above* the survey forecasts of inflation. This leaves the possibility that there could be large liquidity premia in nominal bonos yields that more than offset the negative effects from the liquidity premia in the udibonos prices. Alternatively, this could be a sign that there are large positive and time-varying inflation risk premia in bonos prices.

As in DKW, we explore this further by correlating the difference between the ten-year fitted BEI and the ten-year survey inflation forecasts with measures of the priced frictions in the bonos market. It turns out that the difference is weakly positively correlated with the average bonos bid-ask spread in our sample (17%) and with the mean absolute fitted errors of the bonos prices from the AFNS model estimation used in the construction of the fitted BEI (19%).[10] The small number of observations (21) prevents us from further substantiating this result. However, similar to DKW, we take this as weak evidence of the existence of liquidity premia in the bonos prices, even though we note that this variation could equally well reflect changes in inflation risk premia independent of the bonos and udibonos liquidity premia.[11]

Finally, we stress that it is a key purpose of the empirical analysis in the paper to quantify the relative magnitudes of these three different types of risk premia in the pricing of bonos

[10]This is a noise measure of arbitrage capital frictions similar to the one developed in Hu et al. (2013).
[11]Hördahl and Tristani (2012) report a similar pattern for euro-area BEI rates and tie it to positive inflation risk premia.

and udibonos and what they imply about bond investors' underlying inflation expectations.

4 Model Estimation and Results

In this section, we first describe the dynamic term structure model of nominal and real yields that we use to account for the liquidity bias in their pricing. We then detail how BEI is decomposed within the model before we proceed to a discussion of the stationarity assumption imposed in the model estimation. We end the section with a brief overview of the main estimation results, including the estimated bonos and udibonos liquidity premia.

4.1 An Arbitrage-Free Model of Nominal and Real Yields with Liquidity Risk

In order to precisely measure nominal and real liquidity premia, we need an accurate model of the instantaneous nominal and real rate, r_t^N and r_t^R. With that goal in mind we choose to focus on the tractable affine dynamic term structure model of nominal and real yields briefly summarized below. We emphasize that, even though the model is not formulated using the canonical form of affine term structure models introduced by Dai and Singleton (2000), it can be viewed as a restricted version of the corresponding canonical Gaussian model.

To begin, let $X_t = (L_t^N, S_t^N, C_t^N, X_t^N, L_t^R, S_t^R, X_t^R)$ denote the state vector of our seven-factor model, which we refer to as the $G^{X^N,X^R}(7)$ model extending the terminology of ACR. Here, (L_t^N, S_t^N, C_t^N) represent level, slope, and curvature factors in the nominal yield curve,[12] while (L_t^R, S_t^R) represent separate level and slope factors in the real yield curve.[13] Finally, (X_t^N, X_t^R) represent the added nominal and real liquidity risk factors. Our joint model of nominal and real yields is a liquidity-augmented extension of the five-factor model used by Carriero et al. (2018) to analyze nominal and real U.K. gilt yields.

The instantaneous nominal and real risk-free rates are defined as

$$r_t^N = L_t^N + S_t^N, \tag{3}$$
$$r_t^R = L_t^R + S_t^R. \tag{4}$$

The risk-neutral $\mathbb{Q}$-dynamics of the state variables used for pricing are given by

[12]To motivate this choice, we note that Espada et al. (2008) show that the first three principal components in their sample of Mexican government bond yields have a level, slope, and curvature pattern in the style of Nelson and Siegel (1987) and account for more than 99 percent of the yield variation.

[13]Chernov and Mueller (2012) provide evidence of a hidden factor in the nominal yield curve that is observable from real yields and inflation expectations. Our model accommodates this stylized fact via the (L_t^R, S_t^R) factors.

$$\begin{pmatrix} dL_t^N \\ dS_t^N \\ dC_t^N \\ dX_t^N \\ dL_t^R \\ dS_t^R \\ dX_t^R \end{pmatrix} = \begin{pmatrix} 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & \lambda^N & -\lambda^N & 0 & 0 & 0 & 0 \\ 0 & 0 & \lambda^N & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & \kappa_N^Q & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & \lambda^R & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & \kappa_R^Q \end{pmatrix} \left[\begin{pmatrix} 0 \\ 0 \\ 0 \\ \theta_N^Q \\ 0 \\ 0 \\ \theta_R^Q \end{pmatrix} - \begin{pmatrix} L_t^N \\ S_t^N \\ C_t^N \\ X_t^N \\ L_t^R \\ S_t^R \\ X_t^R \end{pmatrix} \right] dt + \Sigma \begin{pmatrix} dW_t^{L^N,\mathbb{Q}} \\ dW_t^{S^N,\mathbb{Q}} \\ dW_t^{C^N,\mathbb{Q}} \\ dW_t^{X^N,\mathbb{Q}} \\ dW_t^{L^R,\mathbb{Q}} \\ dW_t^{S^R,\mathbb{Q}} \\ dW_t^{X^R,\mathbb{Q}} \end{pmatrix},$$

where Σ is assumed to be a diagonal matrix as per Christensen et al. (2011).

Due to the liquidity risk in the markets for nominal and real bonds, their yields are sensitive to liquidity pressures. As a consequence, the pricing of nominal and real bonds is not performed with the frictionless short rates in equations (3) and (4), but rather with discount functions that account for the liquidity risk as in ACR:

$$\bar{r}_t^{N,i} = r_t^N + \beta^{N,i}(1 - e^{-\delta^{N,i}(t-t_0^i)})X_t^N = L_t^N + S_t^N + \beta^{N,i}(1 - e^{-\delta^{N,i}(t-t_0^i)})X_t^N, \quad (5)$$
$$\bar{r}_t^{R,j} = r_t^R + \beta^{R,j}(1 - e^{-\delta^{R,j}(t-t_0^j)})X_t^R = L_t^R + S_t^R + \beta^{R,j}(1 - e^{-\delta^{R,j}(t-t_0^j)})X_t^R, \quad (6)$$

where t_0^i and t_0^j denote the dates of issuance of the specific nominal and real bonds, respectively, and $\beta^{N,i}$ and $\beta^{R,j}$ are their sensitivities to the variation in their respective liquidity risk factors. Furthermore, the decay parameters $\delta^{N,i}$ and $\delta^{R,j}$ are assumed to vary across securities.

Christensen and Rudebusch (2019) show that the net present value of one unit of currency paid by nominal bond i at time $t + \tau^i$ has the following exponential-affine form[14]

$$\begin{aligned} P_t^N(t_0^i, \tau^i) &= E^{\mathbb{Q}}\Big[e^{-\int_t^{t+\tau^i} \bar{r}^{N,i}(s,t_0^i)ds}\Big] \\ &= \exp\Big(B_1^N(\tau^i)L_t^N + B_2^N(\tau^i)S_t^N + B_3^N(\tau^i)C_t^N + B_4^N(t, t_0^i, \tau^i)X_t^N + A(t, t_0^i, \tau^i)\Big). \end{aligned}$$

By similar arguments, the net present value of one unit of the consumption basket paid by real bond j at time $t + \tau^j$ has the following exponential-affine form

$$\begin{aligned} P_t^R(t_0^j, \tau^j) &= E^{\mathbb{Q}}\Big[e^{-\int_t^{t+\tau^j} \bar{r}^{R,j}(s,t_0^j)ds}\Big] \\ &= \exp\Big(B_1^R(\tau^j)L_t^R + B_2^R(\tau^j)S_t^R + B_3^R(t, t_0^j, \tau^j)X_t^R + A(t, t_0^j, \tau^j)\Big). \end{aligned}$$

Now, consider the whole value of the nominal bond i issued at time t_0^i with maturity at

[14] The calculations leading to our bond pricing results can be found in the online supplementary appendix.

$t+\tau^i$ that pays an annual coupon C^i semiannually. Its price is given by[15]

$$\begin{aligned}\overline{P}_t^{N,i}(t_0^i,\tau^i,C^i) &= C^i(t_1-t)E^{\mathbb{Q}}\left[e^{-\int_t^{t_1}\overline{r}^{N,i}(s,t_0^i)ds}\right]+\sum_{k=2}^{n}\frac{C^i}{2}E^{\mathbb{Q}}\left[e^{-\int_t^{t_k}\overline{r}^{N,i}(s,t_0^i)ds}\right]\\ &\quad +E^{\mathbb{Q}}\left[e^{-\int_t^{t+\tau^i}\overline{r}^{N,i}(s,t_0^i)ds}\right].\end{aligned}$$

Next, consider the whole value of the real bond j issued at time t_0^j with maturity at $t+\tau^j$ that pays an annual coupon C^j semiannually. Its clean price is given by[16]

$$\begin{aligned}\overline{P}_t^{R,j}(t_0^j,\tau^j,C^j) &= C^j(t_1-t)E^{\mathbb{Q}}\left[e^{-\int_t^{t_1}\overline{r}^{R,j}(s,t_0^j)ds}\right]+\sum_{k=2}^{n}\frac{C^j}{2}E^{\mathbb{Q}}\left[e^{-\int_t^{t_k}\overline{r}^{R,j}(s,t_0^j)ds}\right]\\ &\quad +E^{\mathbb{Q}}\left[e^{-\int_t^{t+\tau^j}\overline{r}^{R,j}(s,t_0^j)ds}\right].\end{aligned}$$

The only minor omission in the bond price formula above is that we do not account for the lag in the inflation indexation of the real bond payoff, but the potential error should be modest in most cases; see Grishchenko and Huang (2013) and DKW for evidence in the case of the U.S. TIPS market.

To complete the model description, we need to specify the risk premia that connect the factor dynamics under the $\mathbb{Q}$-measure to the dynamics under the objective $\mathbb{P}$-measure, where we use the essentially affine risk premium specification introduced in Duffee (2002). In the Gaussian framework, this specification implies that the risk premia Γ_t depend on the state variables; that is,

$$\Gamma_t = \gamma^0 + \gamma^1 X_t,$$

where $\gamma^0 \in \mathbf{R}^7$ and $\gamma^1 \in \mathbf{R}^{7\times 7}$ contain unrestricted parameters. Thus, the resulting unrestricted $G^{X^N,X^R}(7)$ model has $\mathbb{P}$-dynamics given by

$$dX_t = K^{\mathbb{P}}(\theta^{\mathbb{P}} - X_t) + \Sigma dW_t^{\mathbb{P}},$$

where $K^{\mathbb{P}}$ is an unrestricted 7×7 mean-reversion matrix, $\theta^{\mathbb{P}}$ is a 7×1 vector of mean levels, and Σ is a 7×7 lower triangular volatility matrix. This is the transition equation in the extended Kalman filter estimation of this model.

[15] This is the clean price that does not account for any accrued interest and maps to our observed bond prices.

[16] Unlike U.S. TIPS, Mexican udibonos have no embedded deflation protection option, which makes their pricing straightforward.

4.2 Decomposing BEI

Christensen et al. (2010) show that the price of a nominal zero-coupon bond with maturity in τ years can be written as

$$P_t^N(\tau) = P_t^R(\tau) \times E_t^{\mathbb{P}}\left[\frac{\Pi_t}{\Pi_{t+\tau}}\right] \times \left(1 + \frac{cov_t^{\mathbb{P}}\left[\frac{M_{t+\tau}^R}{M_t^R}, \frac{\Pi_t}{\Pi_{t+\tau}}\right]}{E_t^{\mathbb{P}}\left[\frac{M_{t+\tau}^R}{M_t^R}\right] \times E_t^{\mathbb{P}}\left[\frac{\Pi_t}{\Pi_{t+\tau}}\right]}\right),$$

where $P_t^R(\tau)$ is the price of a real zero-coupon bond that pays one consumption unit in τ years, M_t^R is the real stochastic discount factor, and Π_t is the price level.[17]

By taking logarithms, this can be converted into

$$y_t^N(\tau) = y_t^R(\tau) + \pi_t^e(\tau) + \phi_t(\tau),$$

where $y_t^N(\tau)$ and $y_t^R(\tau)$ are nominal and real *frictionless* zero-coupon yields as described in the previous section, while the market-implied average rate of inflation expected at time t for the period from t to $t+\tau$ is

$$\pi_t^e(\tau) = -\frac{1}{\tau} \ln E_t^{\mathbb{P}}\left[\frac{\Pi_t}{\Pi_{t+\tau}}\right] = -\frac{1}{\tau} \ln E_t^{\mathbb{P}}\left[e^{-\int_t^{t+\tau}(r_s^N - r_s^R)ds}\right] \tag{7}$$

and the associated inflation risk premium for the same time period is

$$\phi_t(\tau) = -\frac{1}{\tau} \ln \left(1 + \frac{cov_t^{\mathbb{P}}\left[\frac{M_{t+\tau}^R}{M_t^R}, \frac{\Pi_t}{\Pi_{t+\tau}}\right]}{E_t^{\mathbb{P}}\left[\frac{M_{t+\tau}^R}{M_t^R}\right] \times E_t^{\mathbb{P}}\left[\frac{\Pi_t}{\Pi_{t+\tau}}\right]}\right).$$

This last equation demonstrates that the inflation risk premium can be positive or negative. It is positive if and only if

$$cov_t^{\mathbb{P}}\left[\frac{M_{t+\tau}^R}{M_t^R}, \frac{\Pi_t}{\Pi_{t+\tau}}\right] < 0. \tag{8}$$

That is, the riskiness of nominal bonds relative to real bonds depends on the covariance between the real stochastic discount factor and inflation and is ultimately determined by investor preferences, as in, for example, Rudebusch and Swanson (2012).

Now, the BEI rate is defined as

$$BEI_t(\tau) \equiv y_t^N(\tau) - y_t^R(\tau) = \pi_t^e(\tau) + \phi_t(\tau),$$

that is, the difference between nominal and real yields of the same maturity. Note that it can be decomposed into the sum of expected inflation and the inflation risk premium.

[17] The full details of the decomposition can be found in the online supplementary appendix.

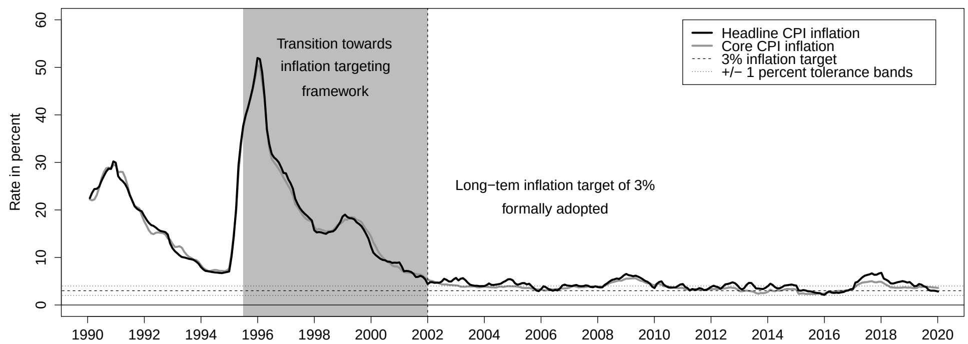


Figure 9: **Mexican CPI Inflation**

4.3 Model Estimation

Due to the nonlinearity of the bond pricing formulas, the model cannot be estimated with the standard Kalman filter. Instead, we use the extended Kalman filter. In addition to the nominal and real bond prices already described, the data used in the estimation includes inflation forecasts from the Consensus Forecasts surveys of professional economists. While the details of the estimation and its execution are provided in the appendix, we will elaborate on the validity of the imposed stationarity assumption in the following.

4.3.1 Stationarity Assumption

To begin the model estimation, we assume that the state variables are stationary and therefore start the Kalman filter at the unconditional mean and covariance matrix. Ultimately, the validity of this assumption hinges on the stationarity of the inflation process and that of real rates. In this section, we therefore examine their statistical behavior and note that it is sufficient to demonstrate that both are stationary.

To support the assumption of stationarity for the inflation process, Figure 9 shows Mexican headline and core CPI inflation measured as 12-month changes back to 1990. The high and volatile inflation of the 1990s is evident. Highlighted in the figure is also the seven-year period it took to transition to the current inflation-targeting framework that officially began in 2002 with the formal adoption of the Bank of Mexico's current 3 percent inflation target. Chiquiar et al. (2010) provide evidence that suggests that Mexican inflation became stationary at some point in the early 2000s, while De Pooter et al. (2014) document that measures of long-term inflation expectations from both surveys and the Mexican government bond market appear to have been anchored close to the 3 percent inflation target of the Bank of Mexico at least since 2003. Overall, we take this evidence to show that Mexican inflation behaved in manner consistent with our stationarity assumption during our 2007-2019 sample period.

Next, we evaluate the statistical behavior of real rates where we choose to focus on the equilibrium real rate of interest r_t^*. Following Christensen and Rudebusch (2019), we define r_t^* as

$$r_t^* = \frac{1}{5}\int_{t+5}^{t+10} E_t^{\mathbb{P}}[r_{t+s}^R]ds, \tag{9}$$

that is, the average expected real short rate over a five-year period starting five years ahead where the expectation is with respect to the objective $\mathbb{P}$-probability measure. As discussed in Christensen and Rudebusch (2019), this 5yr5yr forward average expected real short rate should be little affected by short-term transitory shocks. Alternatively, r_t^* could be defined as the expected real short rate at an infinite horizon. However, this quantity will depend crucially on whether the factor dynamics exhibit a unit root. As is well known, the typical spans of time series data that are available do not distinguish strongly between highly persistent stationary processes and non-stationary ones. Our model follows the finance literature and adopts the former structure, so strictly speaking, our infinite-horizon steady state expected real rate is constant.

First, we compare our finance-based estimate of the Mexican r_t^* with a pure macro-based estimate of r_t^* taken from Carrillo et al. (2018, henceforth CERR). This is a filtered estimate generated by applying a small open economy extension of the Laubach and Williams (2003) model to a combination of Mexican and U.S. macroeconomic series. Both estimates are shown in Figure 10 and share a mild common upward trend during the overlapping period between May 2009 and December 2017. More importantly, the macro-based r_t^* estimate was close to 3 percent in the years prior to the global financial crisis.[18] Thus, we consider the observed increase since 2014 to be a reversal towards the pre-crisis levels.

Figure 10 also shows the U.S. finance-based r_t^* estimate obtained by Christensen and Rudebusch (CR) (2019) using solely the prices of U.S. TIPS as well as the Canadian finance-based r_t^* estimate reported by Christensen et al. (CRS) (2021) using Canadian nominal and real yields. In contrast to the Mexican estimates, these two series are highly positively correlated and with a pronounced downward trend.

The key takeaway is that the natural real rate in Mexico appears to be stationary during our sample and unaffected by the factors that were pushing down long-term real rates in advanced economies such as Canada and the United States. By extension, it seems reasonable to assume that Mexican real rates and bond risk premia are stationary as well, as also suggested by visual inspection of the individual yield series depicted in Figure 4.

Importantly, we stress that the Canadian and U.S. models discussed here are both estimated using the same stationarity assumption that we are imposing in our model estimation. Clearly, this does not prevent these models from producing trending estimates of r_t^* in case the yield data call for it. Hence, we conclude that our stable r_t^* estimate for Mexico is *not* caused by that assumption.

[18]CERR consider several other approaches to estimating r_t^* and find their results to be robust.

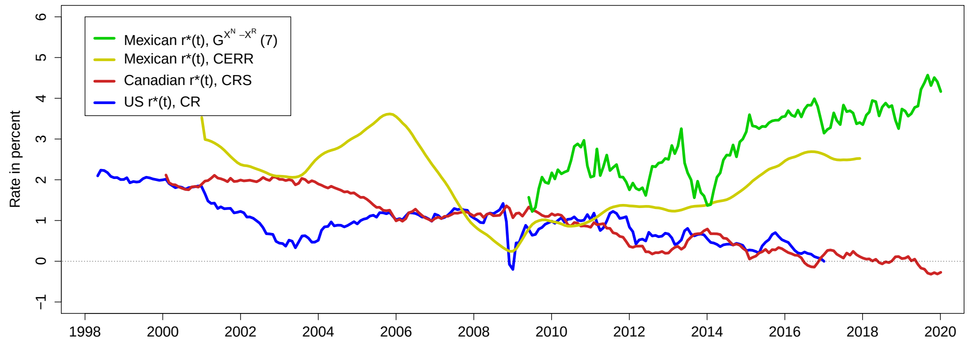


Figure 10: **Estimates of r***

Finally, combining the stationarity results for the inflation process and real rates, we can extend the stationarity assumption to include the Mexican nominal government bond yields, which is again supported by visual inspection of the individual yield series shown in Figure 2.

4.4 Results

In this section, we briefly summarize the main estimation results, while additional details are provided in the online appendix.

To examine the model fit, pricing errors are computed based on the implied yield on each coupon bond to make these errors comparable across securities. That is, for the price on the ith coupon bond $P_t^i(\tau, C^i)$, we find the value of $y_t^{i,c}$ that solves

$$P_t^i(\tau^i, C^i) = C^i(t_1 - t)\exp\left\{-y_t^{i,c}(t_1 - t)\right\} + \sum_{j=2}^{N}\frac{C^i}{2}\exp\left\{-y_t^{i,c}(t_j - t)\right\} + \exp\left\{-y_t^{i,c}(t_N - t)\right\}. \quad (10)$$

For the model-implied estimate of this bond price, denoted $\hat{P}_t^i(\tau, C^i)$, we find the corresponding implied yield $\hat{y}_t^{i,c}$ and report the pricing error as $y_t^{i,c} - \hat{y}_t^{i,c}$. For the nominal bonos the root mean-squared error (RMSE) for all bonds combined is 4.17 basis points, while the corresponding statistics for the real udibonos is 7.56 basis points. Thus, the $G^{X^N,X^R}(7)$ model provides a very good fit to both sets of bond prices.

As for the monthly data on inflation forecasts for the following full calendar year and semiannual data on five-year, ten-year, and so-called 5yr5yr inflation forecasts, the mean errors are 5.14 basis points, -7.54 basis points, -0.42 basis points, and 6.71 basis points, respectively, while the corresponding RMSEs are 45.72 basis points, 26.32 basis points, 28.50 basis points, and 34.24 basis points, respectively, which are all well below the 75 basis points assumed in the model estimation. Thus, the model is also able to simultaneously deliver an

$K^{\mathbb{P}}$	$K^{\mathbb{P}}_{\cdot,1}$	$K^{\mathbb{P}}_{\cdot,2}$	$K^{\mathbb{P}}_{\cdot,3}$	$K^{\mathbb{P}}_{\cdot,4}$	$K^{\mathbb{P}}_{\cdot,5}$	$K^{\mathbb{P}}_{\cdot,6}$	$K^{\mathbb{P}}_{\cdot,7}$	$\theta^{\mathbb{P}}$		Σ
$K^{\mathbb{P}}_{1,\cdot}$	6.5463	2.6817	2.0520	0.2885	-8.9583	-2.1303	-1.1902	0.0912	σ_{11}	0.0120
	(0.8349)	(0.5190)	(0.4329)	(0.3594)	(0.9948)	(0.5940)	(0.4585)	(0.0164)		(0.0007)
$K^{\mathbb{P}}_{2,\cdot}$	8.1454	3.2797	2.7250	0.1269	-11.9725	-2.4465	-1.3970	-0.0286	σ_{22}	0.0185
	(0.9417)	(0.6683)	(0.6186)	(0.4836)	(1.1110)	(0.7666)	(0.5647)	(0.0319)		(0.0029)
$K^{\mathbb{P}}_{3,\cdot}$	-1.9068	1.3745	-1.4645	0.9883	7.5536	-1.3208	-0.3108	-0.0146	σ_{33}	0.0200
	(0.9116)	(0.7087)	(0.5303)	(0.5024)	(1.1164)	(0.7777)	(0.5228)	(0.0392)		(0.0043)
$K^{\mathbb{P}}_{4,\cdot}$	0.9258	3.1202	-1.8217	1.8822	4.4028	-3.3416	-1.1167	-0.0182	σ_{44}	0.0411
	(1.0834)	(0.9275)	(0.8837)	(0.7440)	(1.1512)	(0.9343)	(0.6462)	(0.0708)		(0.0123)
$K^{\mathbb{P}}_{5,\cdot}$	-3.4198	-1.1278	-1.9236	0.2601	6.7996	0.7562	0.5373	0.0287	σ_{55}	0.0087
	(0.7244)	(0.4977)	(0.3787)	(0.2849)	(0.9690)	(0.5646)	(0.3399)	(0.0140)		(0.0010)
$K^{\mathbb{P}}_{6,\cdot}$	-7.3269	-5.2611	-1.9902	-0.4671	11.6805	5.4507	2.6102	-0.1061	σ_{66}	0.0192
	(0.9944)	(0.8186)	(0.7194)	(0.5010)	(1.0896)	(0.8530)	(0.7469)	(0.1218)		(0.0040)
$K^{\mathbb{P}}_{7,\cdot}$	2.2263	-0.1050	1.7390	-0.3899	-5.5887	0.3768	-0.0254	0.2216	σ_{77}	0.0130
	(0.8573)	(0.6112)	(0.5045)	(0.3966)	(0.9669)	(0.7367)	(0.3931)	(0.2903)		(0.0058)

Table 2: **Estimated Dynamic Parameters of the $G^{X^N,X^R}(7)$ Model**
The table shows the estimated parameters of the $K^{\mathbb{P}}$ matrix, $\theta^{\mathbb{P}}$ vector, and diagonal Σ matrix for the $G^{X^N,X^R}(7)$ model. The estimated value of λ^N is 0.2641 (0.0121), while $\lambda^R = 0.4863$ (0.0437), $\kappa^{\mathbb{Q}}_N = 2.2614$ (0.3868), $\theta^{\mathbb{Q}}_N = 0.0106$ (0.0014), $\kappa^{\mathbb{Q}}_R = 0.3333$ (0.0950), and $\theta^{\mathbb{Q}}_R = 0.0136$ (0.0025). The maximum log likelihood value is 19,770.21. The numbers in parentheses are the estimated parameter standard deviations.

accurate fit to the full term structure of available survey inflation forecasts.

The estimated dynamic parameters in the $G^{X^N,X^R}(7)$ model are reported in Table 2. We note that the estimated mean and volatility parameters for the four nominal factors $(L^N_t, S^N_t, C^N_t, X^N_t)$ are very similar to those reported by CFS for their shorter and smaller sample of 21 bonos price series. Thus, the nominal side of our joint model of bonos and udibonos prices fits the bonos data in much the same way as their nominal model.

4.5 The Estimated Bonos and Udibonos Liquidity Premia

We now use the estimated $G^{X^N,X^R}(7)$ model to extract the liquidity premium in the bonos and udibonos prices. To compute these premia we first use the estimated parameters and the filtered states $\{X_{t|t}\}_{t=1}^T$ to calculate the fitted bond prices $\left\{\hat{P}^i_t\right\}_{t=1}^T$ for all outstanding securities in our sample. These bond prices are then converted into yields to maturity $\left\{\hat{y}^{c,i}_t\right\}_{t=1}^T$ by solving the fixed-point problem

$$\begin{aligned} \hat{P}^i_t &= C(t_1 - t)\exp\left\{-(t_1 - t)\hat{y}^{c,i}_t\right\} + \sum_{k=2}^{n} \frac{C}{2}\exp\left\{-(t_k - t)\hat{y}^{c,i}_t\right\} \\ &\quad + \exp\left\{-(T - t)\hat{y}^{c,i}_t\right\}, \end{aligned} \tag{11}$$

for $i = 1, 2, ..., n$, meaning that $\left\{\hat{y}^{c,i}_t\right\}_{t=1}^T$ is approximately the rate of return on the ith bond if held until maturity (see Sack and Elsasser 2004). To obtain the corresponding yields without correcting for liquidity risk, a new set of model-implied bond prices are computed from the

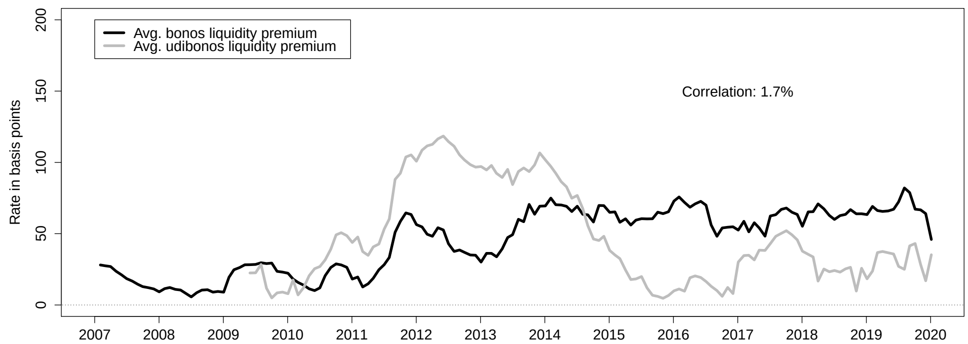


Figure 11: **Average Estimated Liquidity Premia of Mexican Bonos and Udibonos** Illustration of the average estimated liquidity premium of Mexican bonos and udibonos for each observation date implied by the $G^{X^N,X^R}(7)$ model. The liquidity premia are measured as the estimated yield difference between the fitted yield to maturity of individual bonds and the corresponding frictionless yield to maturity with the liquidity risk factor turned off. The bonos data cover the period from January 31, 2007, to December 30, 2019, while the udibonos data cover the period from May 31, 2009, to December 30, 2019.

estimated $G^{X^N,X^R}(7)$ model but using only its frictionless part, i.e., using the constraints that $X^N_{t|t} = 0$ for all t as well as $\sigma_{44} = 0$ and $\theta^Q_N = 0$ for the nominal bonos, and $X^R_{t|t} = 0$ for all t as well as $\sigma_{77} = 0$ and $\theta^Q_R = 0$ for the real udibonos. These prices are denoted $\left\{\tilde{P}^i_t\right\}_{t=1}^T$ and converted into yields to maturity $\tilde{y}^{c,i}_t$ using (11). They represent estimates of the prices that would prevail in a world without any financial frictions. The liquidity premium for the ith bond is then defined as

$$\Psi^i_t \equiv \hat{y}^{c,i}_t - \tilde{y}^{c,i}_t. \tag{12}$$

This can be calculated for bonos and udibonos separately.

Figure 11 shows the average bonos and udibonos liquidity premium series, denoted $\bar{\Psi}^N_t$ and $\bar{\Psi}^R_t$, across the outstanding set of each type of bond at each point in time. The bonos average liquidity premium series averages 45.41 basis points with a standard deviation of 22.05 basis points, while the average udibonos liquidity premium has a slightly higher mean equal to 47.60 basis points with a standard deviation of 34.12 basis points. Furthermore, their correlation in levels for the overlapping period is 2 percent, while it is 17 percent in first differences. Thus, the liquidity risk in the two markets is practically uncorrelated.

5 Empirical BEI Decomposition

In this section, we explore the properties of the BEI decomposition implied by the $G^{X^N,X^R}(7)$ model with a particular emphasis on both the model-implied expected inflation and the associated inflation risk premium that investors in bonos demand to assume their inflation risk.

First, we examine the BEI decomposition and the outlook for long-term inflation expectations in Mexico before we turn to an analysis of the inflation risk premia and their determinants, including an international comparison.

5.1 BEI Decomposition

In this section, we examine the BEI decompostion implied by the estimated $G^{X^N,X^R}(7)$ model. To be consistent with the existing literature, we focus on a horizon long enough into the future that most transitory shocks to the economy can be expected to have vanished. At the same time, the horizon must be practically relevant and covered by the available maturities in the underlying bond data. Balancing these considerations, we limit our analysis to the five-year forward BEI rate that starts five years ahead, denoted 5yr5yr BEI.

The result of decomposing 5yr5yr BEI as described in Section 4.2 is shown in Figure 12. The solid gray line shows the fitted 5yr5yr BEI obtained by estimating a standard three-factor arbitrage-free Nelson-Siegel (AFNS) model to nominal bonos and real udibonos prices separately. This can be compared to the estimated 5yr5yr frictionless BEI implied by the $G^{X^N,X^R}(7)$ model and shown with a solid black line in the figure. The difference between these two measures of 5yr5yr BEI represents the net liquidity premium or distortion of the observed BEI series due to bond-specific liquidity risk premia in both bonos and udibonos prices. Whenever the frictionless BEI is above the fitted BEI it means that the distortions due to liquidity risk are larger in the real yields compared to those in the nominal yields.

Due to its theoretical consistency, the $G^{X^N,X^R}(7)$ model allows us to break down the 5yr5yr frictionless BEI into an expected inflation component, shown with a solid red line in Figure 12, and the residual inflation risk premium, shown with a solid green line. Also shown in the figure with a solid black horizontal line is the 3 percent inflation target of the Bank of Mexico formally adopted in 2002. For comparison, the figure also shows the 5yr5yr expected CPI inflation in Mexico reported semiannually in the Consensus Forecasts surveys. Although these survey inflation forecasts are included in the model estimation, the model-implied expected inflation does deviate quite notably from them for extended periods thanks to the assumed standard deviation of 75 basis points for the associated measurement errors. Still, the closeness of the model's expected inflation to all the considered survey forecasts reported earlier underscores its ability to appropriately capture the term structure of inflation expectations among investors in the Mexican bonos and udibonos market. Finally, Figure 12 also shows the year-over-year change in the Mexican CPI with a solid cyan line to provide a measure of the actual inflation outcomes during this ten-year period.

Note that annual CPI inflation has averaged 4.02 percent during the shown period, somewhat above the Bank of Mexico's target, but mostly within the acceptable $\pm$ 1 percentage point tolerance band around the target. As a consequence, it seems reasonable that both the survey inflation forecasts and the model-implied expected inflation are generally somewhat

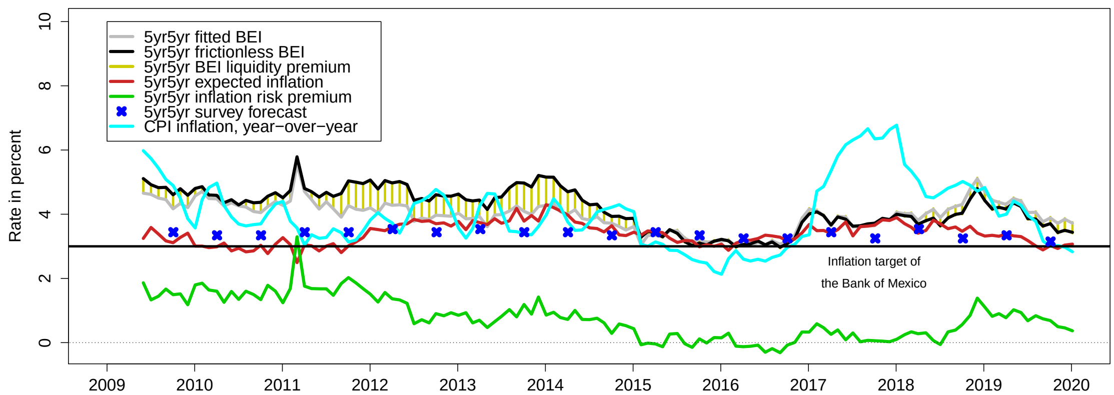


Figure 12: **Decomposition of 5yr5yr BEI**
Illustration of the fitted 5yr5yr BEI obtained by fitting an AFNS model to Mexican bonos and udibonos prices separately and its decomposition based on the $G^{X^N,X^R}(7)$ model estimated with an unrestricted specification of $K^{\mathbb{P}}$ and a diagonal specification of Σ into: (1) the estimated frictionless BEI, (2) expected inflation, and (3) the residual inflation risk premium. The difference between the fitted and frictionless 5yr5yr BEI is highlighted in yellow and represents the net liquidity premium of the observed 5yr5yr BEI. Also shown is the 12-month change in the CPI. The data cover the period from May 31, 2009, to December 30, 2019.

above the announced inflation target. However, given that the Bank of Mexico implements monetary policy with a $\pm$ 1 percentage point tolerance band around its 3 percent target, both the survey inflation forecasts and the model-implied inflation expectations can be viewed as anchored at a level consistent with the central bank's inflation target. Furthermore, the 5yr5yr expected inflation from the model is positively correlated with the year-over-year change in the CPI, as one could expect, but only weakly so with a correlation of 40 percent.

Turning to the estimated inflation risk premia, we first note that, historically, they have been thought of as positive in terms of their sign. This is also the key motivation behind the issuance of inflation-indexed bonds, namely that governments can save paying this premium by switching issuance away from standard nominal bonds and towards inflation-indexed debt, see Price (1997) for a discussion. Our estimated 5yr5yr inflation risk premium is mostly positive, consistent with this prior. However, it did switch sign and turned negative during the 2015-2016 period. In our setting, this happens when the conditional covariance between the real stochastic discount factor and inflation in equation (8) becomes positive. In practice, this means that investors are concerned about a potential economic downturn happening in the context of low inflation instead of the conventional fear of inflation spinning out of control and causing a recession through the reaction of the monetary authority. In the former case, a nominal bond would serve as a good hedge to help smooth consumption through a recession. This explains why, under those circumstances, its price can switch from being penalized with a positive inflation risk premium discount to instead getting a boost from a negative inflation risk premium.

The decline in the Mexican inflation risk premia in 2015 coincided with fears about a softening in the outlook for economic growth in China. In response, oil and other commodity prices fell sharply throughout 2015. For example, the WTI oil price was above 100 US dollars in the summer of 2014 and traded below 35 US dollars by early 2016. As an oil producing country, this weakened the economic outlook for the Mexican economy and simultaneously put downward pressure on inflation in Mexico. As a result, the switch in the sign of Mexican inflation risk premia during this period is consistent with economic theory.

5.1.1 Comparison with Another BEI Decomposition

To further validate the BEI decomposition implied by the $G^{X^N,X^R}(7)$ model, we compare it to the BEI decomposition from an existing model of Mexican nominal and real government bond yields described in Aguilar-Argaez et al. (2016, henceforth AER).[19] The AER model is a three-factor affine model that uses seven nominal bond yields with maturities ranging from one month to ten-year years in combination with the ten-year real yield and the year-over-year change in the CPI to produce estimates of both investors' inflation expectations and the premium they demand to be exposed to inflation risk. A key limitation of this model is that it makes no adjustments for the liquidity risk in either nominal or real yields. Furthermore, it is very parsimonious relying only on three factors for the joint modeling of the two yield curves and the included inflation series.

Figure 13 compares the estimated 5yr5yr expected inflation from the two models. We note that the long-term inflation expectations implied by the AER model are very stable. Blake et al. (2015) report long-term nominal short rate expectations for the Mexican bonos market that are similarly stable. We interpret this as evidence that these models suffer from finite-sample bias in the estimated factor dynamics, as discussed in Bauer et al. (2012). This means that the state variables are expected to revert back to steady state much faster than actually anticipated by investors. In contrast, for the $G^{X^N,X^R}(7)$ model, this problem is significantly mitigated by including medium- and long-term inflation forecasts from surveys in the information set used for the model estimation, as recommended by Kim and Orphanides (2012). As a result, the $G^{X^N,X^R}(7)$ model-implied long-term inflation expectations exhibit cyclical variation, which is positively correlated with the cyclical variation in the realized year-over-year inflation, as also noted in Section 5.1.

Figure 14 compares the estimated ten-year inflation risk premium from the two models. The two series align very closely with each other and have a correlation of 76 percent for the overlapping period. This closeness gives us extra confidence in the $G^{X^N,X^R}(7)$ model's

[19] The updated data for this analysis is taken from: https://www.banxico.org.mx/publicaciones-y-prensa/informes-trimestrales/%7B67E312ED-E93D-EA9C-2A3F-8C20FEE6C215%7D.pdf. This also determines the January 2010 start date for the comparison.

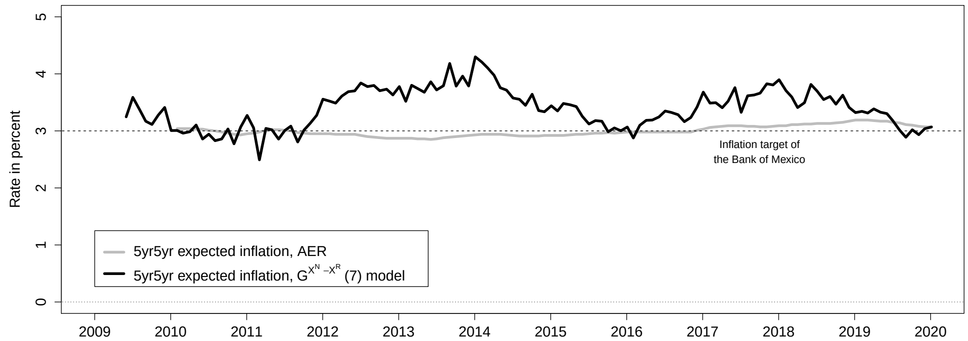


Figure 13: **Comparison of Market-Based 5yr5yr Expected Inflation**

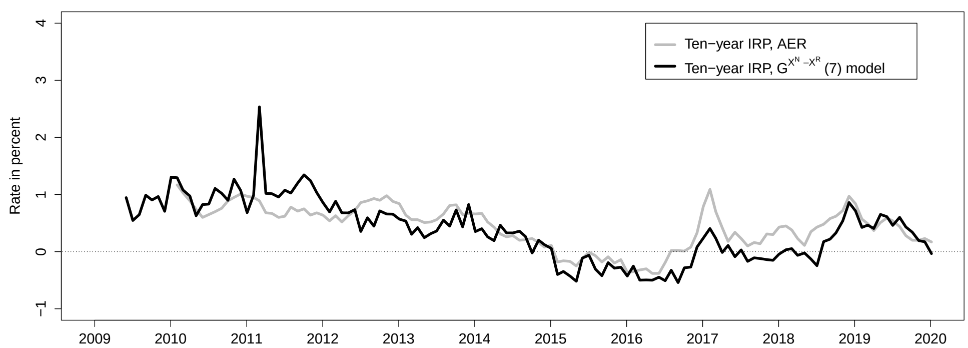


Figure 14: **Comparison of Ten-Year Inflation Risk Premium**

estimated inflation risk premia.

5.2 Analysis of the Model-Implied Inflation Expectations

In this section, we examine the properties of the inflation expectations implied by the $G^{X^N,X^R}(7)$ model in greater detail. First, we evaluate its ability to forecast inflation for the coming calendar year by comparing its performance to that of the Consensus Forecast survey. Second, we assess how anchored inflation and inflation expectations appear to be in Mexico using a statistical measure from the literature before we end the section by exploiting the estimated model dynamics to study the outlook for the 5yr5yr expected inflation over a three-year horizon.

5.2.1 Performance Comparison with Consensus Forecasts

In this section, we explore whether the desirable properties of the $G^{X^N,X^R}(7)$ model-implied long-term inflation expectations documented so far allow it to also generate realistic short-

Model	Mean	RMSE	MAE
Random walk	-14.86	168.14	120.34
BEI	20.74	134.79	92.95
Consensus Forecasts	19.23	118.98	75.67
$G^{X^N,X^R}(7)$ model	21.83	126.12	82.12

Table 3: **Summary Statistics of CPI Inflation Forecast Errors**
This table reports the mean forecasting errors (Mean), the root mean squared forecasting errors (RMSE), and the mean absolute forecasting errors (MAE). The $G^{X^N,X^R}(7)$ model forecasts are computed from the full sample estimation results. The forecast errors are reported as the true value minus the model-implied prediction, and all numbers are reported in annual basis points.

term inflation dynamics.

We structure the forecast exercise to match the monthly Consensus Forecasts survey. At the start of each month, the professional forecasters are asked about their expectations for the change in the CPI for the coming calendar year in addition to their expectations about the change for the current calendar year. To have a series of pure forecasts not distorted by incoming information on realized inflation outcomes, we focus on the monthly survey forecasts of CPI inflation over the coming calendar year. We then use the estimated $G^{X^N,X^R}(7)$ model to generate the matching model-implied CPI inflation forecasts. This has the advantage that the model-implied forecasts reflect information available at the end of each month and therefore lag the official survey dates by between one and two weeks. Thus, this exercise is by design conservative, although we stress the model forecasts are based on the full-sample estimates unlike the survey forecasts, which are real-time forecasts by construction. Finally, to align the exercise with the available observed udibonos prices, we start it in May 2009 and end it in November 2019, a total of 127 forecasts.

As benchmarks, we include two additional forecasting methods. The first is the classic random walk assumption of no change for which the one-year inflation forecast each month equals the past twelve-month change in the Mexican CPI. Hence, the fact that the forecast does not start until the beginning of the next calendar year is without importance. The second is constructed from the observed BEI rates and equals the one-year forward BEI rate that starts at the beginning of the coming calendar year and hence align exactly with the forecast horizon in the Consensus Forecasts surveys.[20]

The summary statistics of the 127 monthly forecast errors from the four forecast methods are reported in Table 3. First, we note that observed BEI rates outperform the random walk assumption. This suggests that the bond yield data are informative about inflation dynamics. However, as an inflation forecasting tool, observed BEI rates are inferior to both the survey forecasts and the $G^{X^N,X^R}(7)$ model-implied forecasts because of the noise added by both the

[20]Similar to Figure 12, the BEI rates are obtained by estimating a standard three-factor arbitrage-free Nelson-Siegel (AFNS) model to nominal bonos and real udibonos prices separately.

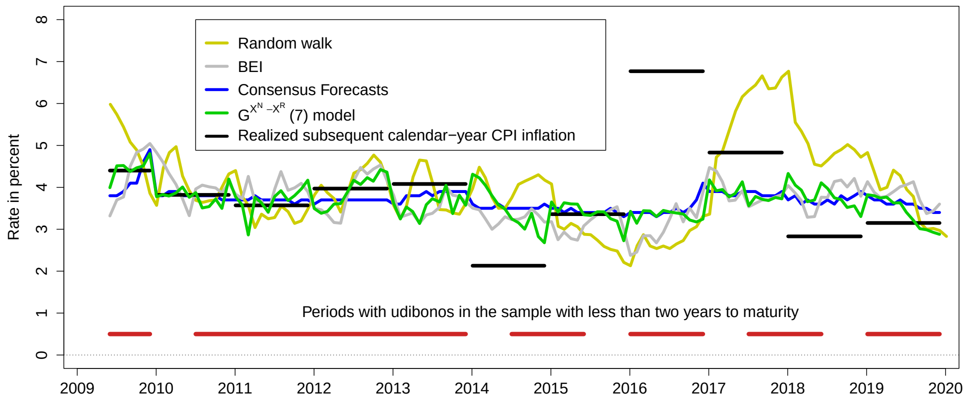


Figure 15: **CPI Inflation Forecasts and Realizations**

inflation risk premium and differential liquidity premia. This explains its higher root mean square forecast error and mean absolute forecast error. Importantly, this also underscores the value of adjusting for these risk premia within the $G^{X^N,X^R}(7)$ model. Finally, in terms of the direct comparison to the survey forecasts, we note that the $G^{X^N,X^R}(7)$ model produces slightly higher forecast errors as measured by all three reported statistics. Given the flexible structure of the $G^{X^N,X^R}(7)$ model and its high number of parameters and state variables, we consider this an encouraging outcome.

In comparing the forecast series, Figure 15 shows that the survey forecasts are very stable, even at the short calendar-year-ahead horizon examined here, another sign that inflation expectations in Mexico appear to be well anchored. In contrast, BEI rates and the $G^{X^N,X^R}(7)$ model-implied forecasts exhibit a greater level of variation that is closer to that reflected in the subsequent CPI inflation realizations also shown in Figure 15 with solid black lines. Lastly, the random walk forecasts are the most volatile as they span the full swings in realized one-year inflation by construction.

To better understand the periodic deviations between the survey and $G^{X^N,X^R}(7)$ model-implied forecasts, we note that the deviations are positively correlated with the periods during which there are udibonos with less than two years to maturity in our sample, highlighted with solid red lines in the figure. Given that the latter are periods when the udibonos data may be particularly informative about investors' near-term inflation expectations, it seems reasonable that these would also be times when the model-implied inflation expectations are more likely to differ from those reported in the surveys.

Overall, these results and findings lead us to conclude that the $G^{X^N,X^R}(7)$ model is able to generate realistic inflation dynamics with properties that match those of the actual CPI series, even though we stress that no inflation data is included in the model estimation.

5.2.2 A Statistical Measure of Inflation Anchoring

For an inflation-targeting central bank like the Bank of Mexico, an important policy question is to what extent inflation expectations in Mexico appear to be anchored at a level consistent with the announced inflation target. In this section, to focus more squarely on that question, we consider a statistical measure of inflation anchoring inspired by Grishchenko et al. (2019).

This measure is centered around the conditional probability of our chosen anchoring measure—the 5yr5r expected inflation, here denoted $\pi^e_{t+\tau}(5yr5yr)$—being within the (2%, 4%) tolerance band used by the Bank of Mexico. That is, we are interested in the following conditional probability

$$P\Big(\pi^e_{t+\tau}(5yr5yr) \leq 0.04\Big|X_t\Big) - P\Big(\pi^e_{t+\tau}(5yr5yr) \leq 0.02\Big|X_t\Big),$$

where τ is the considered horizon. Hence, this measure emphasizes whether the crucial 5yr5yr expected inflation among bond investors and other financial market participants is likely to remain within the tolerance band.

Since $\pi^e_t(5yr5yr)$ is affine in the state variables

$$\pi^e_t(5yr5yr) = A^\pi + \Big(B^\pi\Big)' X_t,$$

it follows from the Gaussian dynamics of our model that

$$\pi^e_{t+\tau}(5yr5yr) \sim N\Big(A^\pi + \Big(B^\pi\Big)' E^{\mathbb{P}}_t[X_{t+\tau}], \Big(B^\pi\Big)' V^{\mathbb{P}}_t[X_{t+\tau}]B^\pi\Big).$$

Thus, the involved probabilities are easily calculated given that the first and second moments of X_t within the $G^{X^N,X^R}(7)$ model follow well-known formulas.

Figure 16 shows these probabilities based on our estimated model at the one- and three-year horizon. As noted in Figure 12, the estimated value of $\pi^e_{t+\tau}(5yr5yr)$ has tended to be within the tolerance band. As a consequence, it is not surprising that the probability of it remaining within the band one year ahead has fluctuated close to 60 percent during our sample period. However, as we increase the considered horizon to three years, the probability drops uniformly to a level close to 40 percent. This is thanks to the increase in the uncertainty in the underlying projections as we lengthen the forecast horizon.

Mapping to the results reported for the U.S. and the euro area in Grishchenko et al. (2019), we note that the probabilities we obtain for Mexico are notably lower. This underscores the point that inflation overall is more volatile in an emerging market economy like Mexico compared to advanced economies. For that reason expected inflation in Mexico is also more volatile. As a consequence, measures of anchoring in Mexico such as the one considered here will—all else equal—show a lower probability of remaining within a certain fixed band than a similar measure applied to advanced economies.

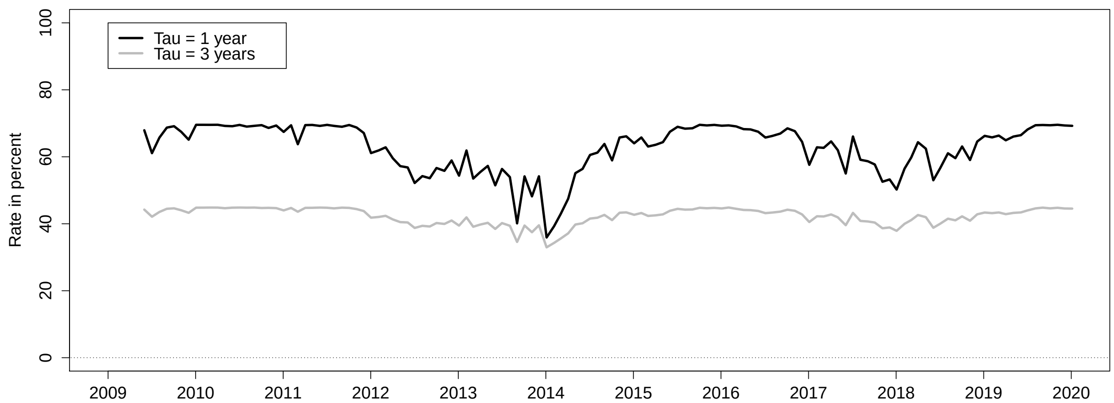


Figure 16: **Probability of 5yr5yr Expected Inflation Remaining Anchored**

5.2.3 Outlook for Long-Term Inflation Expectations

To assess the outlook for long-term inflation expectations based on the $G^{X^N,X^R}(7)$ model while taking the full distribution of potential outcomes into account, we follow the approach of Christensen et al. (2015) and simulate 10,000 factor paths over a three-year horizon, conditioned on the shapes of the nominal and real yield curves and investors' embedded forward-looking expectations as of the end of December 2019 (i.e., using the estimated state variables and factor dynamics as of December 30, 2019). The simulated factor paths are then converted into forecasts of 5yr5yr expected inflation. Figure 17 shows the median projection and the 5th and 95th percentile values for the simulated 5yr5yr expected inflation over the three-year forecast horizon.[21]

The model projections indicate that the long-term inflation expectations are likely to very gradually trend higher from their December 2019 estimate of 3.05 percent. Thus, long-term inflation expectations in Mexico appear to be well anchored at a level close to the inflation target of the Bank of Mexico, although it is important to stress the sizable uncertainty surrounding estimates of long-term inflation expectations, as reflected in the wide 90% confidence band and consistent with the probabilities reported in the previous section.

As an out-of-sample robustness exercise and to assess the reasonableness of the projections, we examine the results from real-time updates of our model during the COVID-19 pandemic and its immediate aftermath. Given that the extraordinary and unusual economic shocks caused by the pandemic were entirely unexpected and exogenous to economic developments in Mexico at the time, this period represents a near-ideal natural experiment to both stress test the model and assess the robustness of our conclusions and findings. To achieve

[21]Note that the lines do not represent paths from a single simulation run over the forecast horizon; instead, they delineate the distribution of all simulation outcomes at a given point in time.

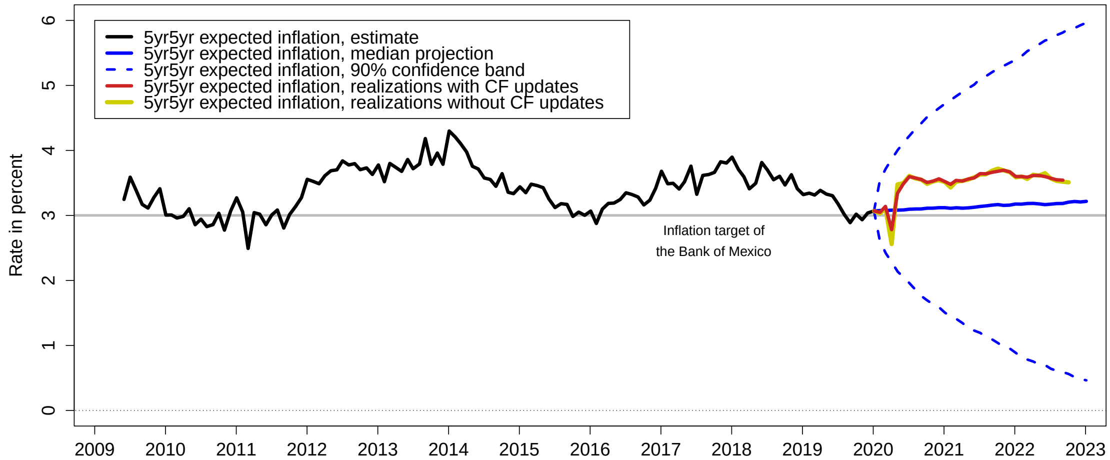


Figure 17: **Three-Year Projections of 5yr5yr Expected Inflation**

this, we update our data each month from January 2020 to September 2022 and re-estimate the $G^{X^N,X^R}(7)$ model in two ways. In one, both yield and survey forecasts are updated, in the other only the yield data are updated. This allows us to examine to what extent the updated results are influenced by the survey information. These updated estimates of the 5yr5yr expected inflation are shown in Figure 17 with solid red and yellow lines, respectively. Despite the unprecedented nature of the involved economic shocks, including Russia's invasion of Ukraine in February 2022, it is comforting to see that the subsequent realizations of the estimated 5yr5yr expected inflation have been very close to, although slight above, our median projection. The fact that the 90% confidence band is much wider than the subsequent realizations suggests that the model is able to fully account for the uncertainty in investors' inflation expectations.

Finally, we add that, while inflation expectations in Mexico remain anchored close to the 3 percent inflation target, the policy path needed to arrive there likely entailed trade-offs. The Bank of Mexico responded early and forcefully to the on-going inflation pressures starting in June 2021. By June 2022, it had raised its benchmark overnight reference rate by a cumulative 375 basis points to 7.75 percent—the sharpest 12-month increase in more than 20 years—and indicated further tightening of the policy rate would likely be appropriate to bring inflation back down to the target over the medium term. This is likely to lower economic growth in Mexico in both 2022 and 2023.

5.2.4 Summary

In this section, we have performed a careful examination of the $G^{X^N,X^R}(7)$ model-implied inflation dynamics. First and most importantly, we find that long-term inflation expectations in Mexico appear to be stable at a level slightly above 3 percent. This makes us draw the same conclusion as the one reached by De Pooter et al. (2014), namely that inflation expectations in Mexico appear to be well-anchored close to the 3 percent inflation target of the Bank of Mexico. Crucially, we stress that model updates during the COVID-19 pandemic and post-pandemic periods do not give us reason to alter this view. If anything, the consistent results during this unique and extraordinary period give us even greater confidence in drawing this inference. Furthermore, this mitigates concerns one could have based on the lower probabilities of statistical measures of inflation anchoring that we find compared to existing results reported for advanced economies. Lastly, the documented reasonableness of the model's estimated inflation dynamics also gives us confidence in its estimated inflation risk premia, which we analyze next.

5.3 Analysis of Inflation Risk Premia

In this section, we first explore what determines the size of, and variation in, Mexican inflation risk premia using regression analysis. This is followed by an international comparison to Canadian and U.S. inflation risk premia.

5.3.1 Determinants of Inflation Risk Premia

While the long-term inflation expectations in Mexico are largely determined by the inflation target of the Bank of Mexico, it is less clear what factors would matter for the size of Mexican long-term inflation risk premia. To explain the variation of the 5yr5yr Mexican inflation risk premium series, we therefore run a battery of standard regressions with it as the dependent variable and a wide set of explanatory variables that are thought to play a role for inflation risk premia as explained in the following.

To begin, we are interested in the role of factors that are believed to matter for bonos and udibonos market liquidity specifically or bond market liquidity more broadly as they could matter for the estimated inflation risk premia, even though we have explicitly accounted for bonos and udibonos liquidity premia in the model estimation. Building on the findings of CFS, we include the foreign-held share of the bonos market as a key explanatory variable. Second, we use the average bid-ask spread in the bonos market shown in Figure 7. Third, we add the average bonos age and the one-month realized volatility of the ten-year bonos yield as additional proxies for bond liquidity following the work of Houweling et al. (2005). Inspired by the analysis of Hu et al. (2013), we also include a noise measure of bonos prices to control for variation in the amount of arbitrage capital available in this market. Combining these five explanatory variables tied to bonos market liquidity and functioning produces the results reported in regression (1) in Table 4. We note a high adjusted R^2 of 0.63. The foreign share has a significant negative coefficient. This implies that an increased presence of foreigners in the bonos market is associated with lower inflation risk premia in addition to its positive effects on bonos liquidity premia documented in CFS.

Next, we repeat the above regression exercise, but now focus on the corresponding explanatory variables derived from the udibonos market, i.e., we include the foreign-held share of the udibonos market, the average bid-ask spread of the udibonos in our sample (also shown in Figure 7), the average age of the udibonos in our sample, the one-month realized volatility of daily changes in the fitted ten-year udibonos yield, and the noise measure constructed from fitted errors of our sample of daily udibonos prices. These five explanatory variables tied to the udibonos market produce the results in regression (2) in Table 4. They generate a slightly higher adjusted R^2 of 0.64, but we again have the foreign share as an important variable with the same sign of its regression coefficient as before.

In a third step, we combine all ten explanatory variables tied to bonos and udibonos market functioning and liquidity. This produces the results reported for regression (3) in Table 4. This yields an even higher adjusted R^2 of 0.72. However, with the exception of the foreign held share of udibonos, no other single variable really stands out as notably more important than any of the others.

After having explored the role of liquidity factors, we examine the effects of factors reflecting risk sentiment domestically and globally on the inflation risk premia. This set of variables includes the VIX, which represents near-term uncertainty about the general stock market as reflected in options on the Standard & Poor's 500 stock price index and is widely used as a gauge of investor fear and risk aversion. The set also contains the yield difference between seasoned (off-the-run) U.S. Treasury securities and the most recently issued (on-the-run) U.S. Treasury security of the same ten-year maturity mentioned earlier. This on-the-run (OTR) premium is a frequently used measure of financial frictions in the U.S. Treasury market. To control for factors that affect emerging market sovereign bonds more broadly, we include the J.P. Morgan Emerging Market Bond Index (EMBI). The fourth variable is the U.S. TED spread, which is calculated as the difference between the three-month U.S. LIBOR and the three-month U.S. T-bill interest rate. This spread represents a measure of the perceived general credit risk in global financial markets. As an additional indicator of credit risk and credit risk sentiment, we use the five-year credit default swap (CDS) rate for Mexico shown in Figure 5. The final variable in the set is the ten-year U.S. Treasury yield from the Federal Reserve's H.15 database, which is included to control for reach-for-yield effects in advanced economies. This may be particularly relevant for our sample during the period between December 2008 and December 2015 when U.S. short-term interest rates were constrained by the zero lower bound.

The results of the regression with these six explanatory variables is reported in regression (4) in Table 4. We note a relatively modest adjusted R^2 of 0.57. Furthermore, it is only the VIX and the ten-year U.S. Treasury yield, which are significant and with the expected positive sign.

Explanatory variables	(1)	(2)	(3)	(4)	(5)	(6)
Foreign share of bonos	-3.72***		-0.18			0.20
	(0.71)		(1.30)			(1.08)
Bonos bid-as-spread	-4.06		-10.48			1.97
	(7.60)		(6.46)			(4.97)
Avg. bonos age	-1.97		-49.27***			21.03
	(5.91)		(12.32)			(11.96)
One-month bonos yield vol.	1.06		0.33			-0.17
	(0.62)		(0.62)			(0.45)
Bonos noise measure	-0.10		-4.18			-2.29
	(3.23)		(2.56)			(1.95)
Foreign share of udibonos		-8.67***	-9.88***			-6.75***
		(1.51)	(2.75)			(2.11)
Udibonos bid-ask spread		1.49	-0.28			0.43
		(2.81)	(2.10)			(0.99)
Avg. udibonos age		-25.55***	25.34			20.62
		(4.58)	(15.89)			(11.49)
One-month udibonos yield vol.		1.21	0.96			0.56
		(0.70)	(0.79)			(0.66)
Udibonos noise measure		2.81**	2.84			2.10
		(1.39)	(1.57)			(1.64)
VIX				6.86***		0.38
				(1.24)		(0.85)
OTR premium				-0.63		0.40
				(1.05)		(1.08)
EMBI				-0.43***		0.03
				(0.13)		(0.14)
TED spread				-0.69		-0.13
				(0.77)		(0.33)
CDS rate				0.14		0.10
				(0.25)		(0.22)
10yr US Treasury yield				33.57**		14.02
				(13.52)		(8.90)
Peso/USD exchange rate					1.41	1.17
					(6.39)	(6.28)
WTI					0.07	0.12
					(0.30)	(0.43)
CPI Inflation					-12.80***	-2.76
					(3.60)	(4.70)
Debt-to-GDP ratio					-11.94***	-16.98***
					(2.18)	(4.83)
MSCI one-month return					-0.49	-0.20
					(0.48)	(0.46)
One-month cetes rate					17.37***	-9.47
					(4.93)	(9.50)
Intercept	272.63***	243.89***	282.69***	34.58	534.75***	577.65***
	(48.23)	(34.72)	(46.78)	(49.57)	(62.55)	(158.97)
Adjusted R^2	0.63	0.64	0.72	0.57	0.76	0.84

Table 4: **Regression Results for the 5yr5yr Inflation Risk Premium**
The table reports the results of regressions with the estimated 5yr5yr inflation risk premium as the dependent variable and 22 explanatory variables. Standard errors computed by the Newey-West estimator (with three lags) are reported in parentheses. Asterisks *, ** and *** indicate significance at the 10 percent, 5 percent and 1 percent levels, respectively.

In the final exercise, we assess the role played by standard macro variables for Mexican inflation risk premia. In this set of variables, we first include the Mexican peso-U.S. dollar exchange rate. As an open emerging market economy, inflation dynamics in Mexico and the premium investors attach to the associated risk is likely to be sensitive to exchange rate developments. The second variable is the West Texas Intermediate (WTI) Cushing crude oil price. Because Mexico is a major oil producing country, the revenue and bond issuance of the Mexican government are affected by changes in oil prices, and so is Mexican inflation itself. Thus, this variable could matter for Mexican inflation risk premia. To capture inflation risk in a more direct way, we include the year-over-year change in the Mexican consumer price index (CPI). Given our focus on longer-term inflation risk as reflected in the 5yr5yr inflation risk premium, the outlook for public finances and any lingering risk of inflating away an outsized debt burden could matter as well. To capture such effects, we include the public debt-to-GDP ratio as measured by the OECD. Furthermore, we include the monthly return of the MSCI Mexican stock index as a measure of the general economic developments in the Mexican economy of importance to investors.[22] Finally, we include the one-month cetes rate.[23] In addition to capturing the stance of monetary policy, this rate serves as a proxy for the opportunity cost of holding money and the associated liquidity convenience premia of bonos, as explained in Nagel (2016).

The results of the regression with these six standard macroeconomic explanatory variables is reported in regression (5) in Table 4. They produce a sizable adjusted R^2 of 0.76. Among the six variables, the year-over-year change in the CPI, the public debt-to-GDP ratio, and the one-month cetes rate stand out as significant. Unfortunately, the CPI inflation and the debt-to-GDP ratio have counterintuitive *negative* coefficients, which are hard to rationalize.

To assess the robustness of the results from these five regressions, we include all explanatory variables with the results reported in column (6) in the table. Although this joint regression produces a high adjusted R^2 of 0.84, we also see a surprising number of switches in signs of the estimated coefficients. Among the affected variables we find the bonos and udibonos age, EMBI, and the one-month cetes rate.

Given the mixed results from the large regression models, we use informed priors to identify a simple preferred regression model to explain the variation in the 5yr5yr inflation risk premium series. First, it follows from the findings of CFS that the foreign share of the large bonos market—and not the foreign share of the much smaller udibonos market—should be a key variable. Second, the VIX is widely used as a measure of global risk aversion that matters for risk premia in both bond and stock markets. Third, as already explained, oil prices should play a first order role for both economic growth and inflation in Mexico. Finally, as a small open economy bordering the United States, the Mexican government bond market is

[22]The MSCI index is a free-float weighted equity index designed to measure the performance of the large and mid cap segments of the Mexican stock market. The index is reported in U.S. dollars.

[23]Cetes are short-term instruments issued by the Mexican government similar to U.S. Treasury bills.

Explanatory variables	(1)	(2)	(3)	(4)	(5)
Foreign share of bonos	-3.80***				-1.89***
	(0.45)				(0.59)
VIX		6.45***			2.45**
		(0.79)			(1.11)
WTI			1.93***		1.20***
			(0.32)		(0.18)
10yr US Treasury yield				59.37***	17.43*
				(13.07)	(9.49)
Intercept	271.13***	-35.31**	-61.67***	-65.70*	1.32
	(23.56)	(15.97)	(22.22)	(35.05)	(62.28)
Adjusted R^2	0.62	0.32	0.39	0.28	0.75

Table 5: **Preferred Regression Results for the 5yr5yr Inflation Risk Premium**
The table reports the results of regressions with the estimated 5yr5yr inflation risk premium as the dependent variable and the four representative explanatory variables identified in the initial round of regressions. Standard errors computed by the Newey-West estimator (with three lags) are reported in parentheses. Asterisks *, **, and *** indicate significance at the 10 percent, 5 percent and 1 percent levels, respectively.

significantly affected by the interest rate level prevailing in the U.S. Treasury market, which we proxy with the ten-year U.S. Treasury yield. Thus, we run a second set of regressions with these four variables individually and combined. This allows us to identify a final preferred regression model for our Mexican 5yr5yr inflation risk premium series.

The results are reported in Table 5. Regression (5) with all four representative variables combined delivers an adjusted R^2 of 0.75. This is on par with or better than that produced by any of the groups of variables we explored in the initial round of regressions. Hence, this supports our selection of this particular set of representative variables. These results also underscore that our four representative variables are responsible for the vast bulk of the significant explanatory power. Furthermore, all four variables have some statistical significance. Most importantly, their regression coefficients have consistent and sensible signs. As a consequence, we consider regression (5) to be our preferred explanatory regression model for the Mexican 5yr5yr inflation risk premium series.

As for the involved magnitudes, we note that a one percentage point increase in the foreign participation in the bonos market lowers the 5yr5yr inflation risk premium by almost 2 basis points. Thus, the 40 percentage point increase between 2007 and 2017 in foreign bonos holdings documented by CFS helped push down significantly Mexican inflation risk premia during this period. On the other hand, increases in both oil prices and the VIX will tend to put significant upward pressure on Mexican inflation risk premia. Lastly, a one-percentage-point increase in the ten-year U.S. Treasury yield will tend to boost the Mexican 5yr5yr inflation risk premium by about 17 basis points.

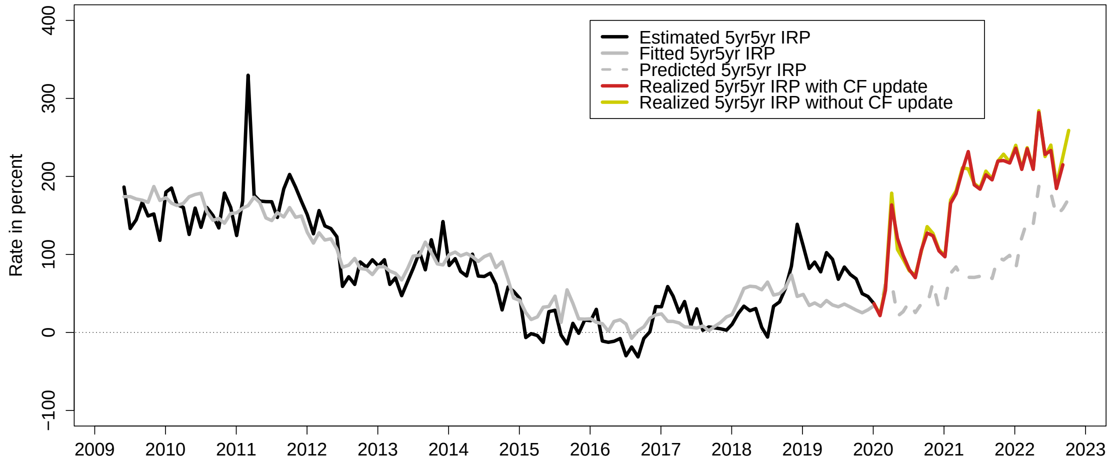


Figure 18: **Fit and Predictability of Preferred Regression Model of 5yr5yr Inflation Risk Premia**
Illustration of the fit of the preferred regression model with the estimated 5yr5yr inflation risk premium as the dependent variable. Also shown is its out-of-sample predictive accuracy measured against the subsequent real-time estimates of the 5yr5yr inflation risk premium. The shown data cover the period from May 31, 2009, to June 30, 2022.

To validate the preferred regression model, we examine both its in-sample fit over the period from May 31, 2009, to December 30, 2019, and the accuracy of its out-of-sample predictability. The latter is performed by comparing the model's predicted values for the period from January 31, 2020, to September 30, 2022, i.e., we fix the estimated regression coefficients at their December 2019 values while updating the explanatory variables, to the subsequent real-time estimated values of the 5yr5yr inflation risk premium. As described in the previous section, these are obtained from real-time estimations of the $G^{X^N,X^R}(7)$ model with expanding samples with and without updates of the survey forecasts. This produces the realized 5yr5yr inflation risk premium series shown in the figure with solid red and yellow lines, respectively. We note that the predicted series shown with a grey dashed line trends up in tandem with the realized series, but at a lower level. This seems reasonable given that the upside inflation surprises since early 2021 have been much larger and persistent than most observers anticipated. In response, investors in the Mexican bond markets have significantly raised the premia they demand for being exposed to inflation risk. Furthermore, we add that the accuracy of the out-of-sample predicted values from the regression model are only marginally lower than the tight fit obtained in sample. These results lead us to conclude that the preferred regression model is robust and accurately captures the key determinants of Mexican inflation risk premia.

Finally, we leave it for future research to explore to what extent these results apply more

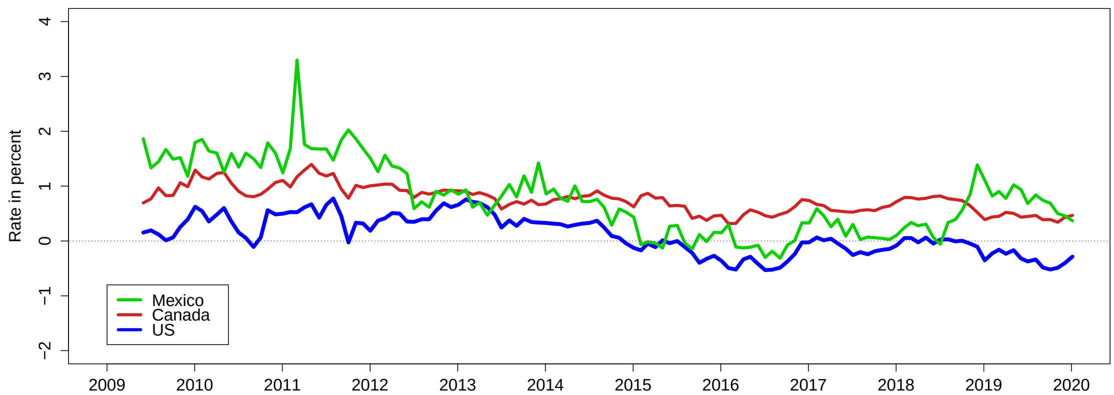


Figure 19: **International Panel of 5yr5yr Inflation Risk Premia**
Illustration of the estimated 5yr5yr inflation risk premium series from Mexican, Canadian, and U.S. nominal and real bond prices as described in the text. The shown data cover the period from May 31, 2009, to December 30, 2019.

broadly to other open emerging market economies with inflation-targeting monetary policy regimes or whether they are unique to Mexico and maybe reflect its very close ties to the U.S. economy in general and U.S. financial markets in particular.

5.3.2 International Comparison of Inflation Risk Premia

To go beyond the regression analysis above, we compare the estimated 5yr5yr inflation risk premium for Mexico with matching estimates from Canadian and U.S. nominal and real yields.[24] Figure 19 shows all three series for the available overlapping sample period.

The Canadian and U.S. inflation risk premia are highly positively correlated (85%). The Mexican inflation risk premium series is also positively correlated with each, but less so, 66 percent and 62 percent, respectively. Thus, both in terms of size and time variation, Mexican inflation risk premia share similarities with those observed in Canadian and U.S. bond markets. However, as expected, Mexican inflation risk premia are more volatile with a standard deviation of 66 basis points compared with 24 basis points and 35 basis points for the Canadian and U.S. series, respectively.

The mostly positive and small inflation risk premia in Canada and the United States are consistent with the findings from simple macro-finance representative agent models; see Hördahl and Tristani (2012). For the United States, DKW also report empirical estimates of inflation risk premia, which are mostly positive and relatively small. In turn, to observe larger and more volatile inflation risk premia in an emerging market economy, such as Mexico,

[24]The Canadian estimate is taken from Christensen et al. (2021), while the U.S. estimate comes from an update of the model described in ACR using all available TIPS.

would seem like a reasonable result given the higher and more volatile CPI inflation in Mexico compared with Canada and the United States.

6 Conclusion

In this paper, we introduce a flexible joint model of nominal and real yields that accounts for liquidity risk premia in both nominal and real bond prices. We estimate the model on a representative sample of nominal and real bond prices from Mexico. This allows us to be the first to provide estimates of the liquidity-adjusted frictionless BEI in a major open emerging market economy, along with its decomposition into investors' underlying inflation expectations and associated inflation risk premia.

Our results indicate that long-term inflation expectations in Mexico appear to have remained well anchored during our sample period at a level close to the 3 percent inflation target of the Bank of Mexico. Still, it is important to note that inflation in Mexico was notably more volatile during our sample period compared to its neighboring advanced economies to the North. As a consequence, inflation uncertainty represents a real risk for investors in the Mexican bonos market, which likely explains why we find that Mexican inflation risk premia are larger and more volatile than corresponding estimates from Canada and the United States.

A comprehensive analysis of the determinants of long-term inflation risk premia in Mexico identifies four variables of particular importance, namely the foreign share of the bonos market, oil prices, the VIX, and the ten-year U.S. Treasury yield. While the foreign share of the bonos market has the expected negative effect following the analysis by CFS, the other three variables have positive coefficients meaning that increases in oil prices, global risk aversion, and U.S. long-term interest rates tend to boost long-term inflation risk premia in Mexico. Thus, to maintain the credibility of its monetary policy target, the Bank of Mexico will have to carefully navigate these global influences on its domestic bond markets. We leave it for future research to explore whether this holds for other open emerging market economies with inflation-targeting central banks.

Finally, we feel compelled to stress that our model framework can be applied to other emerging market economies with established nominal and real bond markets such as Brazil, Chile, and Colombia, among many others. However, we also leave those applications for future research.

Appendix: Model Estimation and Econometric Identification

In this appendix, we detail the estimation of our model and the restrictions needed for it to be identified econometrically.

Due to the nonlinearity of the bond pricing formulas, our model cannot be estimated with the standard Kalman filter. Instead, we use the extended Kalman filter as in Kim and Singleton (2012); see Christensen and Rudebusch (2019) for details. To make the fitted errors comparable across bonds of various maturities, we follow ACR and scale each bond price by its duration. Thus, the measurement equation for the nominal bond prices takes the following form:

$$\frac{\overline{P}_t^N(\tau^i)}{D_t^N(\tau^i)} = \frac{\widehat{P}_t^N(\tau^i)}{D_t^N(\tau^i)} + \varepsilon_t^{N,i},$$

where $\widehat{P}_t^N(\tau^i)$ is the model-implied price of nominal bond i and $D_t^N(\tau^i)$ is its duration, which is fixed and calculated before estimation. Similarly, the measurement equation for the real bond prices takes the following form:

$$\frac{\overline{P}_t^R(t_0^j, \tau^j)}{D_t^R(\tau^j)} = \frac{\widehat{P}_t^R(t_0^j, \tau^j)}{D_t^R(\tau^j)} + \varepsilon_t^{R,j},$$

where $\widehat{P}_t^R(\tau^j)$ is the model-implied price of real bond j and $D_t^R(\tau^j)$ is its duration, which is again fixed and calculated before estimation. See Andreasen et al. (2019) for evidence supporting this formulation of the measurement equations.

Since the liquidity factors are latent factors that we do not observe, their levels are not identified without additional restrictions. As a consequence, when we include the nominal liquidity factor X_t^N, we let the first thirty-year bonos issued after the start of our sample window have a unit loading on the liquidity factor, that is, bonos number (10) in our sample issued on January 29, 2009, with maturity on November 18, 2038, and a coupon rate of 8.5 percent has $\beta^{N,i} = 1$. When we include the real liquidity factor X_t^R, we let the first thirty-year udibonos in our sample have a unit loading on this factor, that is, udibonos number (5) issued on January 5, 2006, with maturity on November 22, 2035, and a coupon rate of 4.5 percent has $\beta^{R,j} = 1$.

Furthermore, we note that the liquidity decay parameters, $\delta^{N,i}$ and $\delta^{R,j}$, can be hard to identify if their values are too large or too small. As a consequence, we impose the restriction that they fall within the range from 0.0001 to 10, which is without practical consequences based on the evidence presented in CFS. Also, for numerical stability during the model optimization, we impose the restrictions that the liquidity sensitivity parameters, $\beta^{N,i}$ and $\beta^{R,j}$, fall within the range from 0 to 250, which turns out not to be a binding constraint at the optimum.

In addition, we assume that all nominal bond price measurement equations have *i.i.d.* fitted errors with zero mean and standard deviation σ_ε^N. Similarly, all real bond price measurement equations have fitted errors that are assumed to be *i.i.d.* with zero mean and standard deviation σ_ε^R.

We also incorporate long-term forecasts of inflation from the Consensus Forecasts survey for Latin America in our model estimation. These include monthly data on inflation forecasts for the following full calendar year and semiannual data on five-year, ten-year, and so-called 5yr5yr inflation forecasts, which represent long-term inflation forecasts covering a five-year period starting five years ahead.[25] As demonstrated by Kim and Orphanides (2012), the inclusion of long-term survey forecasts can help the model better capture the appropriate persistence of the factors under the objective $\mathbb{P}$-dynamics, which can otherwise suffer from significant finite-sample bias.[26]

The measurement equation for the survey expectations incorporating these long-term forecasts takes the

[25] Similar to Christensen et al. (2010) and Abrahams et al. (2016), we do not include inflation data in the model estimation. This omission is expected to, at most, have a small impact on our results due to the relatively long maturities of most of our real yield observations, see DKW for evidence.

[26] Also, see Bauer et al. (2012).

form

$$\pi_t^{CF}(\tau) = \pi_t^e(\tau) + \varepsilon_t^{CF},$$

where $\pi_t^e(\tau)$ is the model-implied τ-year expected inflation calculated using equation (7), which is affine in the state variables, while the measurement error is $\varepsilon_t^{CF} \sim \mathcal{NID}\left(0, (\sigma_\varepsilon^{CF})^2\right)$ and identical for all survey forecasts independent of their horizon as we consider all survey inflation forecasts to be equally informative. As for the value of σ_ε^{CF}, we follow DKW and fix it at 75 basis points in order to not overly influence the estimation results by including the survey forecasts. Alternatively, this approach can be interpreted as treating the survey forecasts as relatively noisy measures of bond investors' inflation expectations. We perform a comprehensive sensitivity analysis to assess the impact of this assumption on our results in online Appendix B, while we examine the effect on our estimation results of excluding the survey information from the model in online Appendix C.

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