Spillovers of U.S. Monetary Policy to Emerging Market Sovereign Yields

M. Pavel Solís M. *

This draft: August 11, 2020

Abstract

This paper studies whether and how the sovereign bond yields of emerging markets respond to U.S. monetary policy. To better characterize those responses, I propose a novel decomposition of the yields into an expected future short-term interest rate, a term premium and a credit risk premium using a new dataset of nominal and synthetic local currency yields along with survey forecasts for 15 emerging markets from 2000 to 2019. This decomposition actually provides several new insights about the dynamics of bond yields in emerging markets. I find that U.S. monetary policy influences each of the yield components and thus gives rise to a reassessment of policy rate expectations and a repricing of interest and credit risks in those countries, consistent with a portfolio rebalancing channel. U.S. monetary policy has therefore not only monetary but also fiscal implications for emerging markets. Furthermore, even though the yields of emerging markets are less globally connected than those of advanced economies, their response to U.S. monetary policy lasts longer.

Keywords: Monetary policy spillovers, emerging markets, synthetic yields, term premium, credit risk, affine term structure models.

JEL Classification: E43, F34, G12, G15, H63.

*Address: Wyman Park Building 544E, 3400 N. Charles Street, Baltimore, MD 21218, United States. Email: msolism1@jhu.edu.

1 Introduction

[Go2ToC]

Although it is well-known that U.S. monetary policy has effects beyond its borders, the extent to which it influences financial conditions in emerging markets is not yet well understood. Given the increasingly important role these countries play in the global economy, it is pressing to understand how their domestic interest rates respond to U.S. monetary policy.

The analysis of U.S. monetary policy spillovers on interest rates has so far mainly focused on advanced economies, which partly reflects data constraints in emerging markets. Not so long ago, they used to issue bonds at short maturities and in foreign currency, but now it is more common for them to issue bonds with longer maturities and in local currency (LC). LC bonds have in fact become an important source of funds for emerging markets over the last two decades (Du and Schreger, 2016b; Ottonello and Perez, 2019; Galli, 2020).

This paper studies the response of the LC bond yields of emerging markets to U.S. monetary policy. Since bond yields compensate investors for bearing risks, yield decompositions provide valuable information. In fact, to better understand how the yields of emerging markets respond to U.S. monetary policy, this paper decomposes them into an expected future short-term interest rate, a term premium and a credit risk premium.[1] This last component is characteristic of emerging markets and, to the best of my knowledge, it has not been accounted for in the literature that decomposes the bond yields of emerging markets.

The credit risk premium acknowledges that not all sovereign bonds are equal. Contrary to the debt issued by advanced countries, international investors demand a credit risk premium to hold bonds issued by emerging markets (Du and Schreger, 2016a). Indeed, even though countries have the ability to print their own currency to avoid default-

[1]The term premium compensates investors for interest rate risk, and the credit risk premium compensates them for the risk of not receiving the promised payments. Credit risk here is broadly defined including, for example, (selective) default risk, currency convertibility risk, regulation risk, capital controls, jurisdiction risk and, if any, liquidity risk. Therefore, when investors require compensation for any of these risks, it is considered that they demand a premium for credit risk even if the country does not default per se.

ing on their debt, emerging markets are prone to default (Reinhart and Rogoff, 2011; Erce and Mallucci, 2018).[2] To account for credit risk, I use synthetic LC yield curves, which essentially swap the U.S. yield curve into a LC, something akin to the U.S. issuing bonds in that currency.[3] These synthetic yields can be seen as being free of credit risk and can thus be decomposed into a future expected short-term interest rate and a term premium using affine term structure models.[4] The difference between the nominal (or actual) yields and the synthetic ones captures the credit risk premium in the LC debt of emerging markets (Du and Schreger, 2016a).

The proposed decomposition of yields can actually help to refine the analysis of the transmission of monetary policy in emerging markets as well as risk management and asset allocation strategies. Indeed, I show that it provides several new insights about the dynamics of emerging market yields. For instance, there is evidence of a trade-off between explicit and implicit defaults and of a downward trend in the term premia of emerging markets as is the case for advanced countries (Wright, 2011). Similarly, the levels of the long-term expected real interest rates in emerging markets are also in line with the evidence for advanced countries (Holston, Laubach, and Williams, 2017). Furthermore, long-term policy rate expectations in emerging markets have been stable over time, consistent with the fact that all but one of the countries in the sample have adopted (before or during the sample period) an inflation targeting regime.[5]

To analyze the spillover effects of U.S. monetary policy, I consider three types shocks. They capture unanticipated changes to the current policy rate, to its future path and to the large-scale asset purchase (LSAP) programs that were implemented by the U.S. Federal Reserve (Fed) as part of its unconventional monetary policy response to the global financial crisis (GFC). The shocks are identified using high-frequency data around the

[2]Examples of actual defaults in LC debt include El Salvador (2017), Ecuador (2008), Argentina (2001), Russia (1998); and in 1999 after an earthquake, Turkey retroactively taxed its debt. Du and Schreger (2016b) and Galli (2020) provide theoretical explanations for these episodes.

[3]This implicitly assumes that the U.S. yield curve and the financial instruments used to swap it are free of credit risk. I argue that these are reasonable assumptions in section 2.

[4]I use data on survey forecasts to obtain robust decompositions of the synthetic yield curves (see Guimarães, 2014). Surveys are especially relevant for emerging markets because they help to overcome both small sample sizes and regime changes, which are common in those countries.

[5]Although Malaysia does not follow an inflation targeting regime, it has different characteristics that are aligned with such a regime, see Pennings et al. (2015).

Fed's monetary policy announcements, which is by now a well-established strategy to overcome endogeneity concerns because it isolates the surprise component of monetary policy decisions (Gürkaynak and Wright, 2013; Nakamura and Steinsson, 2018).

The main finding of this paper is that U.S. monetary policy influences each of the components of emerging market yields. That is, surprises in Fed's policy decisions give rise to a reassessment of policy rate expectations in emerging markets and a repricing in their interest and credit risks, consistent with a portfolio rebalancing channel. In particular, since U.S. monetary policy influences the repricing of credit risks in emerging markets, it has therefore not only monetary but also fiscal implications for those countries. Furthermore, although the yields of emerging markets are less globally connected than those of advanced economies, the response of emerging market yields to U.S. monetary policy lasts longer than those of advanced countries. Specifically, the path and LSAP shocks—more prevalent since the GFC—have a larger and longer effect on EM yields.

This paper contributes to different branches of the literature. It pioneers the application of term structure models on synthetic yields, which have been widely used recently to study deviations from covered interest parity (CIP).[6] Instead of concentrating on the CIP deviations, this paper focuses on the synthetic yields themselves to decompose the nominal bond yields of emerging markets more finely. It also contributes to the literature on the international comparison of sovereign yields (Dahlquist and Hasseltoft, 2016)—so far mainly focused on advanced economies—and extends the results in Wright (2011) by considering emerging markets in the international comparison of term premia.

On the spillover effects of U.S. monetary policy, this paper extends the work of Gilchrist, Yue, and Zakrajšek (2019) by studying its effects not only on sovereign yields but on its components, as in Curcuru, Kamin, Li, and Rodriguez (2018) and Adrian, Crump, Durham, and Moench (2019), but focusing on the yields of emerging markets.[7]

[6] Du, Tepper, and Verdelhan (2018c) show that there are persistent and systematic deviations from CIP reflecting a higher regulatory burden for financial intermediaries. Du, Im, and Schreger (2018b) show that CIP deviations reflect differences in the convenience yield of advanced countries relative to the U.S. Du and Schreger (2016a) meanwhile show that CIP deviations capture a credit risk premium for emerging markets.

[7] Hofmann, Ilhyock, and Shin (2019) study the link between the U.S. monetary policy, the exchange rate and the credit risk premium in emerging markets.

The rest of the paper proceeds as follows. Section 2 explains how to construct the local currency yield curves. Section 3 presents the affine term structure model. Section 4 analyzes the decompositions of yields. Section 5 studies the U.S. monetary policy spillovers to the yields of emerging markets. The last section concludes.

2 Local Currency Yield Curves

[Go2ToC]

This section explains how to construct the nominal and synthetic local currency (LC) yield curves of emerging markets, and explains that the difference between the two captures a credit risk premium. In the next section, the synthetic yield curve will be decomposed into an expected future short-term interest rate and a term premium.

2.1 Construction of Synthetic Yield Curves

The main idea to construct the synthetic LC yield curves is to swap the U.S. yield curve into LC using the forward premium at each maturity (see Du, Im, and Schreger, 2018a). The forward premium compensates investors for the expected depreciation of a currency, an increase in the exchange rate when it is expressed in LC per U.S. dollar (USD), as is used throughout this paper. The key assumption behind this approach is that the U.S. yield curve is free of default risk; as such, it serves as the benchmark for all the other countries in the sample.

The calculation of the forward premium depends on the maturity. For maturities of less than one year, the forward premium is calculated as the annualized difference between the forward and the spot exchange rates. For maturities equal or larger than one year, the forward premium is calculated using cross-currency swaps (XCS) since outright forwards are less liquid. Although, the fixed-for-fixed XCS rates are rarely observed in the market directly, they can be constructed using cross-currency basis swaps and interest rate swaps. The idea is to start by swapping fixed payments in LC into floating-rate cash flows in USD using cross-currency basis swaps (referenced to the Libor—London interbank offered rate—in USD), and swap them into fixed-rate cash flows also in USD using interest rate swaps. Both types of swaps are liquid and collateralized instruments and so the bilateral counterparty risk in XCS is small.

Let $y_{t,n}^{US}$ denote the zero-coupon yield for an n-period U.S. Treasury security at time t, and $\rho_{t,n}$ the n-period forward premium from USD to LC at time t. The zero-coupon synthetic LC yield for the n-period bond at time t, $\widetilde{y}_{t,n}^{LC}$, is calculated as

$$\widetilde{y}_{t,n}^{LC} = y_{t,n}^{US} + \rho_{t,n}. \tag{1}$$

In contrast, the actual or nominal zero-coupon yield, $y_{t,n}^{LC}$, can be obtained directly from the quotes of the bonds traded in the market. Notice that the construction of the synthetic yield curve $\widetilde{y}_{t,n}^{LC}$ does not require information about the nominal yield curve $y_{t,n}^{LC}$,[8] as it relies on the U.S. yield curve and XCS rates.

According to the CIP condition, the nominal (direct) and the synthetic (indirect) LC interest rates should be equal. Thus, CIP implies that an issuer should be able to borrow directly or indirectly (synthetically) in LC at the same yield. Du, Tepper, and Verdelhan (2018c) show, however, that there are persistent and systematic deviations from CIP. The spread between the nominal and synthetic yields ($y_{t,n}^{LC} - \widetilde{y}_{t,n}^{LC}$) therefore measures CIP deviations in sovereign yields. In the case of advanced countries, Du, Im, and Schreger (2018b) argue that the CIP deviations reflect differences in convenience yields relative to the U.S.

CIP deviations have a different interpretation for emerging markets. Whereas the nominal yields of advanced countries are usually considered free of credit risk, it is reasonable for the nominal yields of emerging markets to include a credit risk premium since emerging markets are prone to default (Reinhart and Rogoff, 2011; Erce and Mallucci, 2018). The credit risk in the components of the synthetic yields in equation (1) is small, a synthetic yield can therefore be seen as the borrowing rate paid by a hypothetical issuer in LC with no credit risk. Du and Schreger (2016a) indeed show that the nominal-synthetic spread captures a credit risk premium in the sovereign yields of emerging markets. In particular, the spread is highly correlated with the rates of credit default swaps (CDS)

[8] For the U.S., $\widetilde{y}_{t,n}^{US} = y_{t,n}^{US}$ since there is no forward premium for the USD relative to the USD.

(CDS)—financial derivatives aimed to protect investors against default by a bond issuer.

Although CDS capture credit risk in the medium to long term (Palladini and Portes, 2011), they are not used in this paper to account for credit risk for several reasons. First, credit risk in CDS themselves is not eliminated, it simply shifts from the bond issuer to the CDS seller, i.e. there is counterparty credit risk. Second, a credit event is not clearly defined and, thus, borrowers can intentionally circumvent the CDS payout. Third, since investors do not need to hold the underlying bond to buy a CDS, there is a chance for market manipulation.

2.2 Construction of Nominal Yield Curves [Go2ToC]

I use the Bloomberg Fair Value (BFV) curves to estimate the nominal yield curve $y_{t,n}^{LC}$. Since these curves report coupon-equivalent par yields, I convert them into continuously-compounded yields (see Gürkaynak, Sack, and Wright, 2007) to obtain the implied zero-coupon curves.[9] This is done for all but two countries for which there are no BFV curves. For Brazil and Israel, Bloomberg provides zero-coupon yields with coupon-equivalent compounding. I also convert these yields, known as IYC curves, into continuously-compounded yields.[10]

The resulting continuously-compounded zero-coupon curve for each country is what this paper refers to as the nominal yield curve $y_{t,n}^{LC}$.

2.3 Yield Curve Data [Go2ToC]

Nominal and synthetic yield curves are constructed for 15 emerging markets (EMs) and, to compare the results, 10 advanced economies (AEs).[11] The countries are selected

[9]As a robustness check, I estimate the nominal yield curves from actual prices for some of the countries in the sample following Nelson and Siegel (1987). These estimated yield curves follow those reported by Bloomberg closely.

[10]For some emerging markets, Bloomberg reports both BFV and IYC curves. BFV curves are preferred for several reasons: they have a longer history than IYC curves, IYC curves are not available for advanced countries—the benchmark for some of the results reported later—and, compared to the BFV curves, the short end of the IYC curves seems disconnected from the rest of the curve for some countries and dates.

[11]For each of the following countries, the currency identifier is shown in parenthesis. The 15 EMs

based on data availability. AEs are sometimes split into two groups to assess whether the type of country plays a role. The first group (non-U.S. G3) is comprised by Germany, Japan and the U.K., and the rest of the countries make up the second group (A-SOE), which are basically advanced small open economies, arguably more directly comparable to emerging markets.

It is worth noting that all of the EMs in the sample, with the exception of Malaysia, have adopted an inflation targeting regime. In fact, the central banks in Hungary, Philippines, Indonesia, Russia and Turkey adopted inflation targeting during the sample period.[12]

The data for nominal and synthetic yields is available daily. The sample starts in January 2000 and ends in January 2019. Even though the sample ends on the same date for all countries, the starting dates vary per country. All the yields for advanced countries start no later than September 2001. The sample sizes for emerging markets, however, are smaller. The nominal yields of 9 and the synthetic yields of 7 emerging markets start before March 2004; both types of yields for the rest of the countries start no later than June 2007. Thus, there are at least 10 years of data for most of the emerging markets in the sample.[13] In principle, this is a reasonable time period for the estimation of the affine term structure model presented in section 3.1, but in practice there may be too few interest rate cycles per country. This small sample problem is addressed using surveys of professional forecasters in section 3.4.

The yields have maturities of 3 and 6 months, and 1 through 10 years, ranging from a minimum of nine to a maximum of twelve maturities per country.[14] The maximum

are: Brazil (BRL), Colombia (COP), Hungary (HUF), Indonesia (IDR), Israel (ILS), Korea (KRW), Malaysia (MYR), Mexico (MXN), Peru (PEN), Philippines (PHP), Poland (PLN), Russia (RUB), South Africa (ZAR), Thailand (THB) and Turkey (TRY). The 10 AEs are: Australia (AUD), Canada (CAD), Denmark (DKK), Germany (EUR), Japan (JPY), Norway (NOK), New Zealand (NZD), Sweden (SEK), Switzerland (CHF) and the U.K. (GBP).

[12] Hungary in June 2001, Philippines in January 2002, Indonesia in July 2005, Turkey in January 2006 and Russia in 2014. In addition, Hungary and Poland were accepted to join the European Union in April 2003.

[13] For Turkey, the nominal yields with a maturity of up to 10 years start on June 2010; nevertheless, its synthetic yields start on May 2005. For Russia, data on both types of yields start in 2007 but there were extreme observations in the beginning of the sample due to low liquidity so its sample starts in July 2009.

[14] All countries have data from 3 months to 5 years and 10 years. All countries except Brazil have data for the 7-year maturity. Data for 6, 8 and 9 years vary per country.

maturity considered is 10 years because bonds and swaps with larger maturities tend to have less history and be less liquid, especially for emerging markets who do not issue longer-term bonds as often as advanced countries.

The construction of the LC synthetic yield curves involves data from the U.S. yield curve and the forward premium for different maturities, as explained in section 2.1. Data for the U.S. zero-coupon yield curve is obtained from two sources. For maturities of one through ten years, the yields come from the dataset constructed by Gürkaynak, Sack, and Wright (2007) (hence GSW).[15] Treasury securities with less than one year to maturity are less actively traded than longer-maturity ones. The estimates from the Center for Research in Security Prices (CRSP) are thought to be more robust at the short end of the curve (Duffee, 2010). Specifically, the 3- and 6-month yields are the annualized 13- and 26-week rates (bid/ask average) in the CRSP Risk-Free Rates Series.[16]

As mentioned before, the calculation of the forward premium varies by maturity. To compute the forward premium for maturities of less than one year, I use data on the spot exchange rate along with 3- and 6-month forwards from Bloomberg for all but three countries; for Korea, Philippines and Thailand the data is obtained from Datastream. To construct the XCS rates, I use data on cross-currency basis swaps and interest rate swaps for each available maturity from one through ten years. For all the countries in the sample, the data for these swap curves comes from Bloomberg.[17]

2.3.1 Timing

The parameters of the affine term structure models are estimated using end-of-month data, as explained in section 3.4. Since the U.S. yield curve is the benchmark for the synthetic yield curves, those dates are the last business days of each month according to the U.S. calendar.

[15] Available at: https://www.federalreserve.gov/pubs/feds/2006/200628/200628abs.html.

[16] The 3- and 6-month yields implied by the GSW fitted model are highly correlated with the CRSP yields (0.9985 and 0.9995, respectively) but the former are on average higher (by 16 and 10 basis points, respectively) using data since 1983.

[17] A spreadsheet with the tickers used in the construction of the forward premiums and in the estimation of the nominal yield curves is available upon request. The file consolidates and expands (with tenors and tickers) similar files kindly posted online in Wenxin Du and Jesse Schreger's websites.

Getting the timing right is key to adequately measure the responses of emerging market yields to the U.S. monetary policy shocks. The analysis of monetary policy spillovers in section 5.2 uses daily changes in nominal and synthetic yields, which need to correctly capture the surprises in Fed announcements. Since the nominal yields reported by Bloomberg are pulled at around 16 hours New York time, they are shifted one day forward for the non-Western Hemisphere countries so that their daily changes adequately capture surprises in Fed announcements. The credit risk premium for those countries is thus calculated using the shifted nominal yields.

3 Methodology

[Go2ToC]

This section describes the affine term structure model used to estimate the dynamics of the yield curve of each country in the sample. It then discusses the difficulty in identifying the parameters of the model and how survey data helps in the identification.

3.1 Affine Term Structure Model

Let $P_{t,n}$ be the price at time t of a zero-coupon risk-free bond with maturity n. The continuously compounded yield on that bond is then $y_{t,n} = -\ln P_{t,n}/n$. In particular, the one-period continuously compounded risk-free rate is $i_t = y_{t,1} = -\ln P_{t,1}$.

If there is no arbitrage, there exists a strictly positive stochastic discount factor (SDF), M_{t+1}, that prices all assets. Accordingly, the bond price today equals the expectation, conditioned on all current information, of tomorrow's discounted price; recursively defined as follows:

$$P_{t,n} = \mathrm{E}_t^{\mathbb{P}} \left[M_{t+1} P_{t+1,n-1} \right], \tag{2}$$

where $\mathrm{E}_t^{\mathbb{P}}[\cdot]$ denotes the conditional expectation at time t of a random variable taken using the actual or physical probability measure, $\mathbb{P}$, that generates the data. The existence of the SDF also implies that there exists a theoretical risk-neutral or risk-adjusted pricing measure, $\mathbb{Q}$—different from the $\mathbb{P}$ measure—that is defined as follows:[18]

$$P_{t,n} = \mathrm{E}_t^{\mathbb{Q}}\left[\exp\left(-i_t\right) P_{t+1,n-1}\right], \tag{3}$$

where $\mathrm{E}_t^{\mathbb{Q}}[\cdot]$ also denotes conditional expectation at time t but taken using the $\mathbb{Q}$ measure.

A discrete-time affine term structure model assumes that the dynamics of a $\mathrm{K} \times 1$ vector of unobserved pricing factors or state variables, X_t, follow a first-order vector autoregression, VAR(1), under the risk-neutral measure $\mathbb{Q}$:

$$X_{t+1} = \mu^{\mathbb{Q}} + \Phi^{\mathbb{Q}} X_t + \Sigma \nu_{t+1}^{\mathbb{Q}}. \tag{4}$$

where $\mu^{\mathbb{Q}}$ is a $\mathrm{K} \times 1$ vector and $\Phi^{\mathbb{Q}}$ is a $\mathrm{K} \times \mathrm{K}$ matrix of parameters, Σ is a $\mathrm{K} \times \mathrm{K}$ lower triangular matrix with positive diagonal elements, and $\nu_{t+1}^{\mathbb{Q}}$ is a $\mathrm{K} \times 1$ independent and identically distributed, normal vector with zero mean and covariance equal to the identity matrix conditional on the pricing factors, which is written as $\nu_{t+1}^{\mathbb{Q}} | X_t \sim \mathcal{N}_{\mathrm{K}}\left(0, I\right)$.

The dynamics of the one-period interest rate are driven by the pricing factors:

$$i_t = \delta_0 + \delta_1' X_t, \tag{5}$$

where δ_0 is a scalar and δ_1 is a $\mathrm{K} \times 1$ vector of parameters.

These assumptions imply that the bond price is an exponentially affine function of the pricing factors:

$$P_{t,n} = \exp\left(A_n + B_n X_t\right),$$

such that the corresponding continuously compounded yield of the bond is an affine function of those factors:

$$y_{t,n}^{\mathbb{Q}} = A_n^{\mathbb{Q}} + B_n^{\mathbb{Q}} X_t, \tag{6}$$

where $A_n^{\mathbb{Q}} = -\frac{1}{n} A_n$, $B_n^{\mathbb{Q}} = -\frac{1}{n} B_n$, where in turn the scalar $A_n = \mathcal{A}(\delta_0, \delta_1, \mu^{\mathbb{Q}}, \Phi^{\mathbb{Q}}, \Sigma, n)$

[18] Since the price at maturity of a zero-coupon bond is 1, recursive substitution of equation (2) implies that today's price equals the conditional expectation of the product of SDFs over the life of the bond, $P_{t,n} = \mathrm{E}_t^{\mathbb{P}}\left[\Pi_{j=1}^{n} M_{t+j}\right]$. Similarly, recursive substitution of equation (3) implies that ...

and the $1 \times \mathrm{K}$ vector $B_n = \mathcal{B}(\delta_1, \Phi^{\mathbb{Q}}, n)$ are loadings that satisfy the recursive equations:[19]

$$A_{n+1} = -\delta_0 + A_n + B_n'\mu^{\mathbb{Q}} + \frac{1}{2}B_n'\Sigma\Sigma'B_n, \quad A_0 = 0, \tag{7}$$

$$B_{n+1} = -\delta_1 + \Phi^{\mathbb{Q}\prime}B_n, \quad B_0 = 0. \tag{8}$$

The yields $y_{t,n}^{\mathbb{Q}}$ are the model's fitted yields, which means that the risk-neutral measure $\mathbb{Q}$ is sufficient for pricing bonds. However, to be able to decompose the yields into a future expected short-term interest rate and a term premium, the model also specifies the dynamics for the market prices of risk, which control the transformation between the $\mathbb{Q}$ and $\mathbb{P}$ measures. Following Duffee (2002), the $\mathrm{K} \times 1$ vector of market prices of risk, λ_t, is also an affine function of the pricing factors:

$$\lambda_t = \lambda_0 + \lambda_1 X_t, \tag{9}$$

where λ_0 is a $\mathrm{K} \times 1$ vector and λ_1 is $\mathrm{K} \times \mathrm{K}$ matrix of parameters.

Given this structure for the market prices of risk, the dynamics of the pricing factors can also be described by a VAR(1) under the physical measure $\mathbb{P}$ as follows:

$$X_{t+1} = \mu^{\mathbb{P}} + \Phi^{\mathbb{P}}X_t + \Sigma\nu_{t+1}^{\mathbb{P}}. \tag{10}$$

where $\mu^{\mathbb{Q}} = \mu^{\mathbb{P}} - \Sigma\lambda_0$, $\Phi^{\mathbb{Q}} = \Phi^{\mathbb{P}} - \Sigma\lambda_1$, $\nu_{t+1}^{\mathbb{P}}|X_t \sim \mathcal{N}_{\mathrm{K}}(0, I)$. Note that the covariance matrix of the shocks is the same under both measures; that is, it is measure independent. Finally, the SDF is conditionally lognormal:[20]

$$M_{t+1} = \exp\left(-i_t - \frac{1}{2}\lambda_t'\lambda_t - \lambda_t'\nu_{t+1}^{\mathbb{P}}\right). \tag{11}$$

The yields consistent with the expectations hypothesis of the yield curve—as if investors were actually risk-neutral ($\lambda_0 = \lambda_1 = 0$)—are obtained as:

$$y_{t,n}^{\mathbb{P}} = A_n^{\mathbb{P}} + B_n^{\mathbb{P}}X_t,$$

[19] The price coefficients are obtained recursively after combining the no-arbitrage condition and the functional form for bond prices.

[20] The law of motion of the vector of pricing factors in equation (10) and the SDF in equation (11) can be formalized separately or jointly, see Gürkaynak and Wright (2012). For instance, in a utility maximization framework, the SDF is usually interpreted as the intertemporal marginal rate of substitution.

where $A_n^{\mathbb{P}} = -\frac{1}{n}A_n$, $B_n^{\mathbb{P}} = -\frac{1}{n}B_n$, and the loadings $A_n = \mathcal{A}(\delta_0, \delta_1, \mu^{\mathbb{P}}, \Phi^{\mathbb{P}}, \Sigma, n)$ and $B_n = \mathcal{B}(\delta_1, \Phi^{\mathbb{P}}, n)$ satisfy the same recursions as those above but using the parameters of the law of motion of the pricing factors under the $\mathbb{P}$ rather than the $\mathbb{Q}$ measure.[21]

The expected yield over the life of an n-period bond equals the conditional expectation under $\mathbb{P}$ of the average short rate over the period, thus:

$$y_{t,n}^e = A_n^e + B_n^e X_t,$$

where $A_n^e = -\frac{1}{n}A_n$, $B_n^e = -\frac{1}{n}B_n$, where in turn $A_n = \mathcal{A}(\delta_0, \delta_1, \mu^{\mathbb{P}}, \Phi^{\mathbb{P}}, 0, n)$ and $B_n = \mathcal{B}(\delta_1, \Phi^{\mathbb{P}}, n)$; that is, A_n^e and B_n^e also satisfy the recursions under the $\mathbb{P}$ measure but setting $\Sigma = 0$ (see Appendix C of Guimarães (2014)).[22] In this sense, information about $y_{t,n}^e$ could help in identifying the parameters under the $\mathbb{P}$ measure, $\{\mu^{\mathbb{P}}, \Phi^{\mathbb{P}}\}$.

Since the long-term forecasts are the expectation between 6 and 10 years ahead, it is matched to the 5-year forward rate starting 5 years hence. The forward rate from n to m periods hence given by $y_{t,n|m} = (m y_{t,m} - n y_{t,n}) / (m - n)$ becomes

$$y_{t,n|m}^e = A_{t,n|m}^e + B_{t,n|m}^e X_t.$$

where $A_{t,n|m}^e = (m A_m^e - n A_n^e) / (m - n)$ and $A_{t,n|m}^e = (m B_m^e - n B_n^e) / (m - n)$.

Finally, the term premium for maturity n at time t, $\tau_{t,n}$, can then be estimated as the difference between the yields obtained under the $\mathbb{Q}$ and $\mathbb{P}$ measures:[23]

$$\tau_{t,n} = y_{t,n}^{\mathbb{Q}} - y_{t,n}^{\mathbb{P}}. \tag{12}$$

A key assumption behind this model is that the yield $y_{t,n}$ is free of credit risk, a reasonable assumption for advanced but not for emerging countries since investors demand a credit risk premium from them (Du and Schreger, 2016a). This implies that while the nominal yield curve $y_{t,n}^{LC}$ is the relevant one for advanced countries, it is not so for emerging markets because it is not free of credit risk. In that case, the synthetic yield curve $\widetilde{y}_{t,n}^{LC}$ better

[21]See Lloyd (2018) for a derivation of the loadings under both measures.

[22]The difference between $y_{t,n}^{\mathbb{P}}$ and $y_{t,n}^e$ is a convexity term due to Jensen's inequality, which increases with maturity. In practice, however, this term usually becomes relevant for maturities beyond ten years. Further, the term is constant across maturities in homoskedastic models like the ones used in this paper.

[23]Note that $\tau_{t,n}$ can also be written as $(A_n^{\mathbb{Q}} - A_n^{\mathbb{P}}) + (B_n^{\mathbb{Q}} - B_n^{\mathbb{P}})X_t$; that is, the term premium is also an affine function of the pricing factors.

aligns with the risk-free assumption and, in turn, with the affine model.

3.2 Identification Problem [Go2ToC]

In principle, the estimation of the parameters in the affine term structure model only requires zero-coupon bond yields as an input. This is enough to estimate the pricing coefficients under the $\mathbb{Q}$ measure, $\{\mu^{\mathbb{Q}}, \Phi^{\mathbb{Q}}\}$, but not to identify the parameters under the $\mathbb{P}$ measure, $\{\mu^{\mathbb{P}}, \Phi^{\mathbb{P}}\}$, which are necessary to estimate the term premium in equation (12).

This identification problem owes to the high persistence of bond yields, which results in small sample bias (Kim and Orphanides, 2012). In fact, there might be too few interest rate cycles per country. Accordingly, the dynamics of the pricing factors will tend to mean-revert too quickly, overestimating the stability of the expected path of the short rate and attributing much of the variability in yields to fluctuations in the term premium.

The instability in the decomposition of the yield curves of advanced countries is a well-known phenomenon. From the different solutions proposed in the literature,[24] Guimarães (2014) shows that incorporating survey data on interest rate forecasts in the estimation provides robust decompositions of the U.S. and U.K. yield curves.[25]

The importance of survey data is particularly relevant to get reliable decompositions for the bond yields of emerging markets since they help to address two common concerns when working with emerging market data, namely small sample sizes and regime changes. First, the sample size for these countries is relatively shorter, compared with data from advanced countries. Second, since the early 2000s there has been a considerable reduction in the inflation rates of many emerging markets, a phenomenon that can be considered a regime change. Indeed, structural breaks is one of the reasons U.S.-focused studies use different sample periods.

[24] They include restrictions on parameters (Duffee, 2010) (Bauer 2018), bias-corrected estimators (BRW 2012) and complementing bond yield data with survey forecasts of future interest rates (Kim and Wright, 2005; Kim and Orphanides, 2012).

[25] For instance, he finds that the term premium estimated with the aid of surveys remains essentially the same after varying the number of pricing factors (from 3 to 5) and the sample periods, even with a sample starting in 1972, which includes the U.S. Great Inflation period.

In addition to supplementing the information from bond yields in the estimation of affine models, surveys can also be used to obtain a model-free estimate of the term premium, calculated as the difference between the long-term interest rate and the expected future short rate from surveys (over the same horizon). This model-free estimate serves as a robustness check for the term premium obtained from the affine model.

Although surveys have good forecasting properties, it is important to acknowledge that they are not a panacea. For instance, surveys might not represent the market expectations or the expectations of the marginal investor.

3.3 Survey Data [Go2ToC]

Long-term forecasts are particularly helpful for the model to pin down the parameters under the $\mathbb{P}$ measure. Twice a year Consensus Economics provides 5-year ahead and long-term (between 6 and 10 years ahead) forecasts for consumer inflation and real GDP growth for most of the emerging countries in the sample. For this paper, the data is available from March 2001 to October 2017.[26] Figures 1 and 2 respectively plot the long-horizon forecasts for inflation and real GDP growth. With the exception of Brazil and Turkey, inflation expectations in emerging markets have been stable or even declining, and are generally within the upper and lower bounds for the domestic inflation target.[27] Notice that the 5-year ahead and long-term forecasts are highly correlated for both inflation and GDP growth.

Although there is no source for long-term forecasts for the short rate of emerging markets, they can be inferred from existing data. Since emerging markets are small open economies, I assume that the real interest rate is determined abroad. The implied forecast for the domestic nominal short-term rate is then equal to the forecast for the international real short-term rate plus the domestic inflation forecast for the same matu-

[26] Data availability varies by country; for example, data for the Philippines starts in 2009, whereas it ends in October 2013 for Latin American countries.

[27] For Israel and South Africa, figure 1 shows the inflation trend, see below for more information. The upper and lower bounds are the most recent ones for each country. Russia has updated its target range almost every year since early 2000s; the plotted band corresponds to the highest and lowest bounds since 2009.

rity. This approach requires to define the international real interest rate and long-term forecasts for it. As with synthetic yields, the U.S. is taken as the benchmark for emerging markets. Forecasts for the U.S. real interest rate are obtained as the difference between the forecast for the three-month Treasury bill (T-bill) rate and the inflation forecast for the same maturity. The required data is available quarterly from the Survey of Professional Forecasters. Specifically, I use the 5- and 10-Year CPI inflation forecasts and, for the T-bill rate, the 10-year forecast and the second longest available one since there is no 5-year forecast for the T-bill rate.[28] Figure 3 shows the implied long-term forecasts for the domestic short-term rates. The estimates are sensible, their level is in line with the synthetic 10-year yield in each country.

An alternative way to infer the embedded expectations for the policy rate is to use Taylor rule-type regressions, and assume that the estimated parameters for inflation and real GDP growth apply at each of the survey maturities.[29] The two approaches yield similar values for the implied forecasts for the domestic short-term rates. The correlation between the two approaches is 0.75 and 0.83 for the 5- and 10-year tenors, respectively.

An advantage of the small open economy approach is that it only requires forecasts for inflation or a proxy in the case of countries with no long-term forecasts available, namely Israel and South Africa. Inflation expectations are hoped to match measures of inflation that exclude unexpected shocks and better reflect the inflation environment. Different measures of core inflation exist. I use the inflation trend obtained by applying the Hodrick-Prescott filter to the series of realized inflation of each country. Of course, the filter is sensitive to the sample period used. The resulting trend can also be outside of

[28]The specific series are CPI5YR, CPI10, BILL10 and TBILLD. The BILL10 series is released in the first-quarter surveys only, so I use linear interpolation for the second to fourth quarters in the respective year. Consensus Economics forecasts are considered at the end of the month in which they are published, at that time the most recent value for the U.S. real interest rate forecast is used to calculate the implied forecast for the domestic short rate.

[29]I regress the policy rate on its lag, the year-on-year consumer price inflation and the year-on-year real GDP growth for all the countries except Israel and South Africa. The coefficient for the lag of the policy rate is a smoothing parameter that improves the fit of the model to the data. A potential drawback of this approach is precisely that it requires to know the expectation of the policy rate for the previous forecast horizon. Nevertheless, it is reasonable to assume stationarity for the long-term forecasts (5 and 10 years), in which case only survey data for inflation and GDP growth are needed after dividing their coefficients by 1 minus the coefficient for the lag of the policy rate (due to stationarity). Data from the policy rate statistics of the Bank for International Settlements is used for the dependent variable.

the target inflation band due to the innate dynamics of the series, which would be at odds with survey data (see figure 1). Fortunately, unlike other countries, there is no marked upward or downward trend in the inflation of the two countries during the sample period. For each country, trend inflation is calculated for the whole period but only considered within the time range for which survey data is available for the rest of the countries, and as long as the trend is within the inflation target band. Figure 4 shows the realized and trend inflation for Israel and South Africa, and compares them with those of Malaysia and Thailand, two countries with a similar pattern for inflation (i.e. no marked trend) and for which survey data is available. Trend inflation seems to be a good proxy for the long-term inflation forecasts of Israel and South Africa. Finally, since the 5-year and long-term forecasts closely follow each other (see figure 1), I use trend inflation for both tenors.

3.4 Estimation [Go2ToC]

Affine term structure models can be estimated by maximum likelihood, but the convergence to the global optimum has been traditionally subject to computational challenges and multiple local optima. Joslin, Singleton, and Zhu (2011) (hence JSZ) propose a normalization of the affine model that improves the convergence to the global optimum of the likelihood function.[30]

The JSZ normalization allows for the near separation of the model's likelihood function into the product of the $\mathbb{P}$ and $\mathbb{Q}$ likelihood functions, and reduces the dimension of the parameter space from $(\delta_0, \delta_1, \mu^{\mathbb{Q}}, \Phi^{\mathbb{Q}}, \Sigma)$ to $(i_\infty^{\mathbb{Q}}, \lambda^{\mathbb{Q}}, \Sigma)$, where $i_\infty^{\mathbb{Q}}$ is the short rate under $\mathbb{Q}$ in the long-run and $\lambda^{\mathbb{Q}}$ is a $\mathrm{K} \times 1$ vector of ordered eigenvalues of $\Phi^{\mathbb{Q}}$.

The affine model can be estimated using the JSZ normalization following a two-step procedure. First, the $\mathbb{P}$ parameters are estimated by OLS of the VAR in equation (10) using the K principal components of the yield curve. This provides initial values for the maximum likelihood estimation of the matrix Σ. Then, taking $\hat{\mu}^{\mathbb{P}}$ and $\hat{\Phi}^{\mathbb{P}}$ as given, the $\mathbb{Q}$ parameters can be estimated by maximum likelihood.

[30] Adrian et al. (2013) estimate the model using linear regressions. This approach leads to overidentification. Goliński and Spencer (2019) correct for it based on JSZ.

As mentioned before, to estimate the affine model, the relevant (risk-free) yield curve for advanced countries is the nominal one $y^{LC}_{t,n}$, whereas for emerging markets it is the synthetic one $\widetilde{y}^{LC}_{t,n}$. In addition, the model for emerging markets is augmented with survey forecasts of the short rate.[31]

The parameters of the model are estimated using end-of-month data. For advanced countries, the model is estimated only with yield data. For emerging markets, the model is augmented with survey data on the last day of the month for which the data was published.[32] Since yield curve data is monthly and survey data is available twice a year, the latter can be considered as missing in non-release dates.

The survey-augmented model is estimated by maximum likelihood using the Kalman filter, which is well-suited to handle missing data. The transition equation is the law of motion of the pricing factors under the $\mathbb{P}$ measure given in equation (10). The dimension of the observation equation varies depending on the availability of survey data.

On months in which there is no data on survey expectations, the observation equation adds measurement error to the fitted yields in equation (6) for each of the N maturities:

$$\mathbf{y}_t = \mathbf{A} + \mathbf{B}X_t + \mathbf{\Sigma}_Y \mathbf{u}_t, \tag{13}$$

where $\mathbf{y}_t$ is an $\mathrm{N} \times 1$ vector of observed bond yields, $\mathbf{A}$ is an $\mathrm{N} \times 1$ vector with elements $A^{\mathbb{Q}}_n$, $\mathbf{B}$ is an $\mathrm{N} \times \mathrm{K}$ matrix with rows equal to $B^{\mathbb{Q}}_n$ for $n = 1, \ldots, \mathrm{N}$, $\mathbf{u}_t \sim \mathcal{N}_{\mathrm{N}}(0, I)$ and $\mathbf{\Sigma}_Y$ is a lower triangular $\mathrm{N} \times \mathrm{N}$ matrix with positive elements on the diagonal.

On months when survey data is available, the observation equation increases by the number of survey forecasts S as follows:

$$\begin{bmatrix} \mathbf{y}_t \\ \mathbf{y}^S_t \end{bmatrix} = \begin{bmatrix} \mathbf{A} \\ \mathbf{A}^S \end{bmatrix} + \begin{bmatrix} \mathbf{B} \\ \mathbf{B}^S \end{bmatrix} X_t + \begin{bmatrix} \mathbf{\Sigma}_Y \mathbf{u}_t & \mathbf{0} \\ \mathbf{0} & \mathbf{\Sigma}_S \mathbf{u}^S_t \end{bmatrix} \tag{14}$$

[31]The model for advanced countries in the sample is not augmented with survey data because I do not have access to it. They are, however, not the main focus of this paper, the affine model is estimated for them just for comparison purposes. Moreover, the results reported later for them are more comparable with other studies that do not use survey data. Finally, there are less concerns about small sample sizes for advanced countries and, for some of them, even for a regime change during the sample period.

[32]From 2001 to 2014, data for countries covered in the Eastern European release is available in March and September; starting in October 2014, it is released on April and October. For the rest of emerging markets the forecasts have always been released on April and October.

where $\mathbf{y}_t^S$ is a S×1 vector of survey forecasts with elements $i_{t,n}^{survey}$, $\mathbf{A}^S$ is a S×1 vector with elements A_n^e, $\mathbf{B}^S$ is a S × K matrix with rows equal to B_n^e for $n = 1, \ldots, \mathrm{S}$, $\mathbf{u}_t^S \sim \mathcal{N}_{\mathrm{S}}(0, I)$ and $\mathbf{\Sigma}_S$ is a lower triangular S × S matrix with positive elements on the diagonal.

To estimate the survey-augmented model, I follow Guimarães (2014) and Lloyd (2018) in two aspects. First, I use the estimated parameters from the JSZ normalization as initial values for the Kalman filter. Second, I assume homoskedasticity in yields and survey errors, so that $\mathbf{\Sigma}_Y = \sigma_y I_{\mathrm{N}}$ and $\mathbf{\Sigma}_S = \sigma_s I_{\mathrm{S}}$, where I_{N} and I_{S} are N × N and S × S identity matrices, respectively. This reduces the number of parameters to be estimated.

Finally, although survey data helps to discipline the model, relying too much on it can be counterproductive due to, for instance, measurement error[33] and overfitting. To strike a balance between these two considerations, I follow Kim and Orphanides (2012) in fixing σ_s at a conservative level of 75 basis points. In essence, survey data is seen as a noisy measure of expectations for the short rate.[34]

3.4.1 Estimation with Daily Data [Go2ToC]

The parameters of the model are estimated using end-of-month data because at the daily frequency there is noise that can undermine the estimation. However, once the parameters are estimated with monthly data, they can be used to estimate the model at the daily frequency (Adrian et al., 2013).

The estimation procedure explained above gives estimates for both the parameters and the pricing factors. Principal component analysis on the bond yields generates a matrix of weights or loadings for the different maturities to construct each principal component, which could be used to estimate the daily pricing factors. However, the pricing factors for the survey-augmented model are not necessarily the same as the principal components, so I obtain the matrix of loadings implied by those pricing factors using OLS. I regress the monthly pricing factors on the end-of-month observed yields. The weights so obtained are then applied to the daily yields to estimate the daily pricing factors. Finally, the

[33] Recall that Consensus Economics does not report long-term forecasts for the short rate of emerging markets.

[34] When σ_s is allowed to be estimated, its average value among all emerging markets is 41 basis points.In this case, the estimated term premia remains largely the same for most of the countries in the sample.

parameters estimated with monthly data along with the daily pricing factors are used to fit—and decompose—the yields at the daily frequency.

4 Decomposing the Yields of Emerging Markets [Go2ToC]

This section shows that the yield decompositions provide several new insights about the dynamics of emerging market yields. This highlights the benefits of using synthetic curves and survey data when analyzing the yields of emerging markets.

4.1 Model Fit [Go2ToC]

As is common in the literature (Joslin et al., 2011), the pricing factors for all countries are the first three principal components (PCs) of the respective yield curves, they are commonly known as level, slope and curvature (Litterman and Scheinkman, 1991). On average, they explain more than 99.5% of the variation in the synthetic yields of emerging markets and in the nominal yields of advanced countries. Interestingly, the level factor plays a relatively bigger role in advanced countries than in emerging markets (96% vs 88%), so that the slope (4 vs 10%) and curvature (0 vs 2%) factors are relatively more important in EMs.

Figure 5 illustrates the fit of the model for the 10-year synthetic yields of emerging markets. The model fits the data reasonably well for most of the countries. The squared root of the average (across months and maturities) squared difference between the actual and the fitted yields is commonly used to summarize the fitting errors. For the advanced countries in the sample, those fitting errors are small at around 5 basis points in line with previous studies (Wright, 2011; Adrian et al., 2019). The dynamics of emerging market yields, however, are relatively harder to capture, reflected in an average fitting error of 16 basis points.[35] In general, emerging market yield curves are less smooth than those of advanced countries, likely due to a shallower investor base.

[35]For instance, the fitting errors for the short end of the yield curves of Indonesia and Philippines are on average larger than for the rest of the countries.

4.2 Assessing the Decompositions [Go2ToC]

Once the affine model is estimated, the nominal yields of emerging markets can be decomposed into three parts. This section assesses the sensibility of those decompositions. Unless otherwise stated, the analysis focuses on the 10-year maturity for the sake of brevity.

Figure 6 decomposes the nominal yields of each country. Several patterns emerge from the figure. First, the main component of the nominal yields of most countries is the expectation of the future short rate. Moreover, a downward trend over the sample period can be seen for most of the countries, consistent with evidence for advanced countries (Adrian et al., 2019).

Second, the credit risk premium and the term premium are both time-varying and both play an important role in the dynamics of emerging market yields. Although the term premium plays a relatively bigger role in explaining yield variation for most countries, the relative importance of the two premia varies by country and can even change over time as can be seen, for instance, for Hungary and the Philippines.

Third, the literature for advanced countries documents a downward trend in their term premia, which is consistent with a decline in inflation uncertainty (Wright, 2011). Adrian et al. (2019) find a similar decline for emerging markets but without taking into account credit risk. Figure 6 shows that their term premia indeed decreased over the sample period even after correcting for credit risk, although not equally widespread since a few exceptions can be seen; for instance, the term premium for Malaysia, Turkey and South Africa did decline earlier in the sample but increased again afterwards.

4.2.1 Credit Risk Premium

The dynamics of the credit risk premium are in line with the results reported by Du and Schreger (2016a) who, in particular, show that it is highly correlated with the CDS of the respective country. Unlike the term premium, no clear trend is visible for the credit risk premium.

The role of the credit risk premium in explaining yield variation is non-negligible in general. On average, it is positive for all the countries. In fact, it is not realistic for it to be negative. Instances in which the premium actually turns negative are brief and can reflect financial market frictions (Du and Schreger, 2016a), including market segmentation between foreign and local investors and short selling constraints. Although far from perfect, the premium is a valid measure of credit risk, and definitely better than ignoring it. Otherwise, estimates of the term premium would be biased. Therefore, it does matter which curve is used (nominal or synthetic) when decomposing the yields of emerging markets.

4.2.2 Expected Future Short Rate

Incorporating survey data in the estimation helps the model pinning down the parameters under the $\mathbb{P}$ measure, and in addressing the small sample problem characteristic in emerging market data. Figure 7 shows that the model-implied expectation of the short rate for the 10-year maturity tracks the interest rate forecasts reasonably well, even though the model does not rely too much on them given the conservative value of σ_s. An alternative model-free measure of the expectation of the future short-term interest rate is the 2-year yield, the correlation between the two is 93%.

The long-term expectation for the short rate implied by the model can also be assessed in terms of real interest rates. The implied long-term forecast for the short rates of emerging markets are based on the long-term U.S. real interest rate (see section 3.3). Therefore, the expected long-term real interest rate—the difference between the model-implied long-term expectation for the short rate and the long-term inflation forecast from Consensus Economics—should be similar. Figure 8 verifies that this is the case. They all fluctuate in a range around zero. Estimates for advanced countries show that real interest rates have converged and declined toward zero over the past years (Holston, Laubach, and Williams, 2017), a phenomenon that is partly explained by the increase in savings from East Asian countries following the regional crises of the late-1990s (Obstfeld, 2020). After correcting for credit risk and inflation, the real interest rates of emerging markets behave similarly to those of advanced countries.

4.2.3 Term Premium

While the (bond) risk premium is usually associated with the term premium in AEs, for emerging markets they are two different concepts. By estimating the model using synthetic yields and surveys, I am able to estimate a genuine term premium.

There are two model-free measures commonly used to assess the model-implied term premium. One of them is the survey-based term premium obtained as the difference between the synthetic yield and the survey-expectation of the short rate for the same maturity. Since the model-implied expectations track the interest rate forecasts very closely, as mentioned above, the model-implied and the survey-based term premia are highly correlated as well. The other alternative measure is the residual from the regression of the 10-year yield on the 3-month yield (both synthetic). The model-implied term premium is also closely related to this residual, the average correlation between the two measures among all EMs is close to 70%.

The term premia in EMs is generally higher than the term premia in AEs. Nevertheless, in addition to the declining trend in the term premia of both AEs and EMs since the start of the QE program, figure 6 also shows that for some EMs,[36] a declining trend in their term premia can be perceived even before the GFC; this is consistent with the evidence for AEs documented by Wright (2011) with data going back to 1990. He argues that this downward trend reflects a reduction in inflation uncertainty. The largest declines can be seen in Asian and European countries.[37] Interestingly, Mexico, Russia and Turkey experienced reversals in their term premia, they increased after declining to historic lows.

It is noteworthy that more than half of the EMs in the sample experienced negative term premia—which implies that investors considered EM bond yields as hedges and were therefore willing to give up some return in their EM bond investments. For comparison, the U.S. term premium estimated by Kim and Wright (2005) has been negative for most of the time since mid-2011, fluctuating between -1 and 0%. In addition, Wright (2011) reports that the 5-to-10-year forward term premium for six of the AEs considered here

[36] HUF, IDR, KRW, MXN, PHP, PLN

[37] HUF, IDR, KRW, PHP, PLN, THB.

turned negative even before the GFC. [38] Moreover, for the EMs with a negative term premia, the phenomenon can be seen during and after the GFC, in some cases it coincided with the QE announcements. Interestingly, for Brazil and Turkey the decline in their term premia between 2008 and 2013 happened even when their inflation expectations were increasing (see figure 1).

Finally, one can also compare the term premia across maturities per country, known as the term structure of term premia. In general, the term premium increases with maturity. As one would expect, when long-term bonds are seen as riskier than short-term bonds, investors would require a higher compensation for holding long-term bonds.

4.2.4 Risk Premia in EM Yields [Go2ToC]

Given that the two premia play an important role in the dynamics of EM yields, a natural question is whether and how they are related.

For all EMs their estimated term premium correlates negatively with their credit risk premium.[39] When EMs face difficulties servicing their debt, they can either default or inflate away their debt, which can be referred to as explicit and implicit default. The first option would increase the credit risk premium, whereas the second essentially generates inflation risk that would be reflected in a higher term premium. Choosing one of the two options reduces the need for the other, which would explain the negative correlation. The evidence for EMs thus shows a trade-off between explicit and implicit defaults.

Term premia captures the uncertainty risk of investing in bond yields. It is expected then for the term premium to be related to risk and uncertainty measures. Baker, Bloom, and Davis (2016) construct an economic policy uncertainty (EPU) index based on the frequency of articles in local newspapers containing key words such as 'economy', 'uncertainty' and 'central bank'. That index is, however, only available for 5 of the EMs in the sample, Brazil, Colombia, Mexico, Russia and South Korea. In principle, the EPU index

[38] The estimates of the term premia for AEs obtained here, however, do not turn negative. In his estimation, Wright (2011) augments the affine model with data from macroeconomic variables. This might support the case of supplementing the estimation of the affine model of AEs either with macro or survey data.

[39] The correlation is not statistically significant only in the case of Indonesia.

can be correlated with any of the two components of bond risk premia in EMs. Indeed, it is positively correlated with the term premia in Colombia and Mexico, and with the credit risk premium in Russia and South Korea, but it is negatively correlated with the term premia in South Korea and with the credit risk premium in Mexico. The EPU index is not related with any of the two components in the case of Brazil.

Two variables commonly used in the global financial cycle literature (see Rey, 2013) are the monetary policy stance in the U.S. and the Cboe's volatility index (Vix), the latter is usually considered as a measure of risk aversion and economic uncertainty. A natural variable in the former case, is the U.S. term premia as estimated by Kim and Wright (2005). The correlation of the term premia in EMs with the U.S. term premium is 62%, compared to 77% in the case of the term premia in AEs. At the same time, increases in the Vix are also positively associated with increases in the term premia of both EMs (32%) and AEs (37%).

Finally, the term premia in EMs is positively correlated with the local inflation rate.

4.3 Connectedness [Go2ToC]

Although a global factor plays a role in the variation of yields in both AEs and EMs, the bond yields in EMs are not as tightly connected as they are in AEs. Pooling all yields together, there is a global-level factor—associated with the first PC—that explains around 50% of the variation in EMs but as high as 90% in AEs. When yields are separated by maturity, a similar pattern emerges for the 1-year yields, whereas for the 10-year yields, a global factor explains around 95% in AEs and 40% in EMs. These percentages are largely the same after the GFC.

A similar conclusion can be drawn for the yield components. A global factor explains around 40% of the variation of all the three components of yields in EMs.[40] For AEs, in contrast, a global factor explains 90% of the variation in the expected short rate and 80% in their term premia. After the GFC, these percentages increased for the term premia in

[40] Du and Schreger (2016a) also show that the credit risk premium has a low reaction to global variables.

AEs and for the two components of the synthetic yields in EMs.

The role of a global factor seems to be different throughout the yield curve. Obstfeld (2015) shows that the effects of the global financial cycle are different at the short than at the long end of the yield curve. In particular, he argues that long-term bonds are more globally connected. Kalemli-Özcan (2019) meanwhile argues the opposite, that short-term yields suffer more the effects of global influences. When yields are separated by maturity, a global factor explains around 50% of the variation in the 1-year yields of EMs, compared to 90% for AEs. For the 10-year yields, a global factor explains around 40% of the variability in EMs, relative to 95% in AEs. However, the role of the global factor is larger for the two components of the synthetic yield curve in the long end than for the short end. In sum, the role of a global component is larger for short-term yields than for the long-term yields, but it is more relevant for the components of the longer-term yields.

Kalemli-Özcan (2019) further argues that U.S. spillovers work through changes in risk premia, especially for EMs. For instance, the expected short rate in EMs reacts to changes in the U.S. term premium. In this case, the effect is similar among all countries. On average, the 10-year U.S. term premium explains around 40% of the variation in the expected short rates in both EMs and AEs. Moreover, for EMs, the 10-year U.S. term premium explains more of the variation of the expected short rate than the U.S. expected rate itself. After the GFC, the 10-year U.S. term premium explained more than the U.S. expected rate in both AEs and EMs.

Is There A Non-U.S. Common Factor? To assess whether the global factor is associated with U.S. variables, I regress the components of the yield curves of AEs and EMs on the respective components of the U.S. yield curve based on the methodology of Kim and Wright (2005). I then perform a PC analysis on the residuals (i.e. the part of a country's term premium orthogonal to the U.S. term premium) to see whether there is a common non-U.S. factor.

The expectation of the short rate and the term premium in the U.S. explain around 70% of the variation—measured by the R^2 statistic—in the respective components of the yield curves in AEs, and around 30% and 40%, respectively, for the components of the yields in EMs. After the GFC, these numbers largely remained for the term premia but declined for expected part.

There seems to be a non-U.S. common factor. Again, its relevance is lower for EMs than for AEs. It explains up to 85% of the variation in the yields of AEs and up to 50% in the yields of EMs. The importance of a non-US factor slightly increased after the GFC.

These results show that while the yields of EMs are indeed globally connected, they are less so than the yields of AEs; that is, they are still far from being fully integrated in the global financial markets, at least relative to the levels seen for AEs. Several factors can explain these patterns, including segmented markets, capital controls, home bias and regional differences. Local factors therefore seem to play a bigger role in the yields of EMs, and its components, than they do for AEs.

5 U.S. Monetary Policy Spillovers to Emerging Market Bond Yields

[Go2ToC]

This section analyzes how the sovereign yields of emerging markets respond to U.S. monetary policy shocks. It shows that each component of the yields reacts to the news but that the response depends on the type of information conveyed.

5.1 U.S. Monetary Policy Shocks

Monetary policy has more than one dimension since asset prices respond to different types of news about monetary policy (Gürkaynak et al., 2005; Swanson, 2018). This paper considers three types of U.S. monetary policy shocks, namely unexpected changes to the current policy rate (Kuttner, 2001), surprise changes to the future path of the policy rate (Gürkaynak et al., 2005), and unanticipated announcements about the Fed's LSAP programs (Swanson, 2018). They will be referred to hereinafter as target, path and LSAP shocks, respectively.

The shocks are identified using high-frequency data around Fed's monetary policy announcements. The shocks are essentially surprises in monetary policy decisions that represent a change in the information set of market participants (Gürkaynak and Wright, 2013; Nakamura and Steinsson, 2018). The construction of the shocks therefore requires to calculate the change in asset prices that capture market expectations of monetary policy decisions. These changes are measured around windows containing monetary policy decisions. The windows start 15 minutes before and end 1 hour and 45 minutes after each Fed's Federal Open Market Committee (FOMC) meeting since 2000.[41]

The assets used to measure the monetary policy shocks are the current federal funds rate future (FF0) contract, the 8-quarters ahead eurodollar future (ED8) contract and the 10-year U.S. Treasury bond. The target shock uses rescaled changes in the price of the FF0 contract; I follow Gürkaynak et al. (2005) who implement an intraday version of the daily measure proposed by Kuttner (2001). The price change in the ED8 contract is a measure of surprises in the 2-year U.S. Treasury yield. The path shock is the residual of a regression of the intraday change in the ED8 contract on the target shock; this shock is highly correlated with the path factor in Gürkaynak et al. (2005).[42] Finally, I follow Swanson (2018) for measuring the last type. The LSAP shock is the residual of a regression of the intraday change in the 10-year yield on the target and path shocks. A positive shock represents a tightening of the monetary policy stance, whereas a negative shock indicates an easing.

The relevance of the shocks has varied over time. After 2008 the target shocks are essentially zero, since then there were no surprise changes in the current policy rate. By contrast, the meaning of LSAP shocks is unclear before 2008. Path shocks, nevertheless, have been relevant before and after the GFC. As a result, target shocks are considered from 2000 to 2008, LSAP shocks are considered starting in 2009, and path shocks span the whole sample period. Figure 9 shows all three monetary policy shocks.

[41] It is common to exclude from the analysis the meeting of September 2001 that followed the terrorist attacks earlier in the month, see Gürkaynak et al. (2005) and Nakamura and Steinsson (2018).

[42] The change in the 4-quarters ahead eurodollar future (ED4) contract could also be used to measure the path shocks. However, after 2011 intraday changes in ED4 became essentially zero since market participants expected the policy rate to remain at zero for at least a year.

5.2 Panel Local Projections [Go2ToC]

The impact of the U.S. monetary policy shocks on emerging market yields is estimated using panel local projections. These regressions are useful to understand not only how the yields respond to the policy surprise at the time of the shock but over time. To have a better understanding of the monetary policy spillovers, the regressions are performed for the daily changes in yields and their components.

The panel local projections follow Jordà (2005):

$$y_{i,t+h} - y_{i,t-1} = \alpha_i + \beta_h \epsilon_t + \gamma_h y_{i,t-2} + \phi_h f_{i,t-1} + u_{i,t+h}, \tag{15}$$

where i and t are respectively the country and time indexes, h indicates the horizon (in days) with $h = 0, 1, \ldots, 90$, α_i is a country fixed effect, ϵ_t represents a type of monetary policy shock, $f_{i,t-1}$ is a one-day lag of the exchange rate (LC per USD) and $u_{i,t+h}$ is the error term. The regressions are run for the 2- and 10-year yields and each of their components (the expected short rate, the term premium and the credit risk premium) for every type of monetary policy shock.[43]

The parameters of interest are β_h. They measure the average response of the yield or one of its components to a monetary policy shock at horizon h. These contemporaneous effects of the shocks are obtained when $h = 0$ in equation (15).

The Pesaran test of cross-sectional independence is rejected at any significance level. Therefore, standard errors are computed following Driscoll and Kraay (1994), so that they are robust to heteroskedasticity, autocorrelation and cross-sectional dependence.

5.3 Spillovers

The response of U.S. yields to U.S. monetary policy shocks is used as a benchmark to assess the responses of the yields of emerging markets. Figures 11 and 12 show the responses of the 2- and 10-year U.S. yields, respectively, to all three shocks. All shocks

[43] By definition, the shocks are unanticipated by the market and thus there is no need to control for past or future shocks. Also, by construction, the shocks are not correlated among themselves so, there is no need to control for the other types of shocks.

have a similar on-impact effect on the 2-year yield. Nevertheless, target shocks did not influence the 10-year yield. Path and LSAP shocks do indeed influence the 10-year yield. In fact, one of the goals of those policies was to lower long-term bond yields and, in particular, the term premium.[44] LSAP shocks, however, have larger effects than path shocks. Moreover, the effects of these shocks last for about a month.

The main finding of this paper is reported in figures 13 and 14. They show the responses of the 2- and 10-year EM yields, respectively, to all three shocks. U.S. monetary policy influences each of the components of emerging market yields. This means that surprises in Fed's policy decisions give rise to a reassessment of policy rate expectations in emerging markets and a repricing in their interest and credit risks. This evidence is consistent with a portfolio rebalancing channel.

Interestingly, the shocks have an opposite effect on the synthetic yields and on the credit risk premium. Whereas a tightening shock in the U.S. increases synthetic yields in emerging markets, it decreases the credit risk premia. These results have both monetary and fiscal implications for those countries. Affects financing costs for the government.

Figures 13 and 14 also show that although the yields of emerging markets are less globally connected than those of advanced economies (see section 4.3), the response of emerging market yields to U.S. monetary policy lasts longer than those of advanced countries. In particular, the path and LSAP shocks have a larger and longer effect on EM yields.

6 Concluding Remarks [Go2ToC]

This paper studies how emerging market sovereign bond yields respond to U.S. monetary policy shocks. To better characterize those responses, this paper proposes a novel decomposition of the sovereign yields of emerging markets into an expected future short-term interest rate, a credit risk premium and a (pure) term premium. This decomposition provides several new insights about the dynamics of bond yields in emerging markets.

[44] Kuttner (2018) discusses the effects of the QE announcements on U.S. bond yields.

I show that U.S. monetary policy not only influences each of the components of emerging market yields but that the effect lasts longer than for the yields of advanced countries. U.S. monetary policy thus gives rise to a reassessment of policy rate expectations and a repricing of interest and credit risks in emerging markets. This evidence is consistent with a portfolio rebalancing channel.

The results in this paper have several important implications for policymakers in emerging markets. For instance, since U.S. monetary policy drives a repricing of interest and credit risks, spillovers have not only monetary but also fiscal implications for those countries.

Figure 1. Long-Horizon Forecasts of Inflation

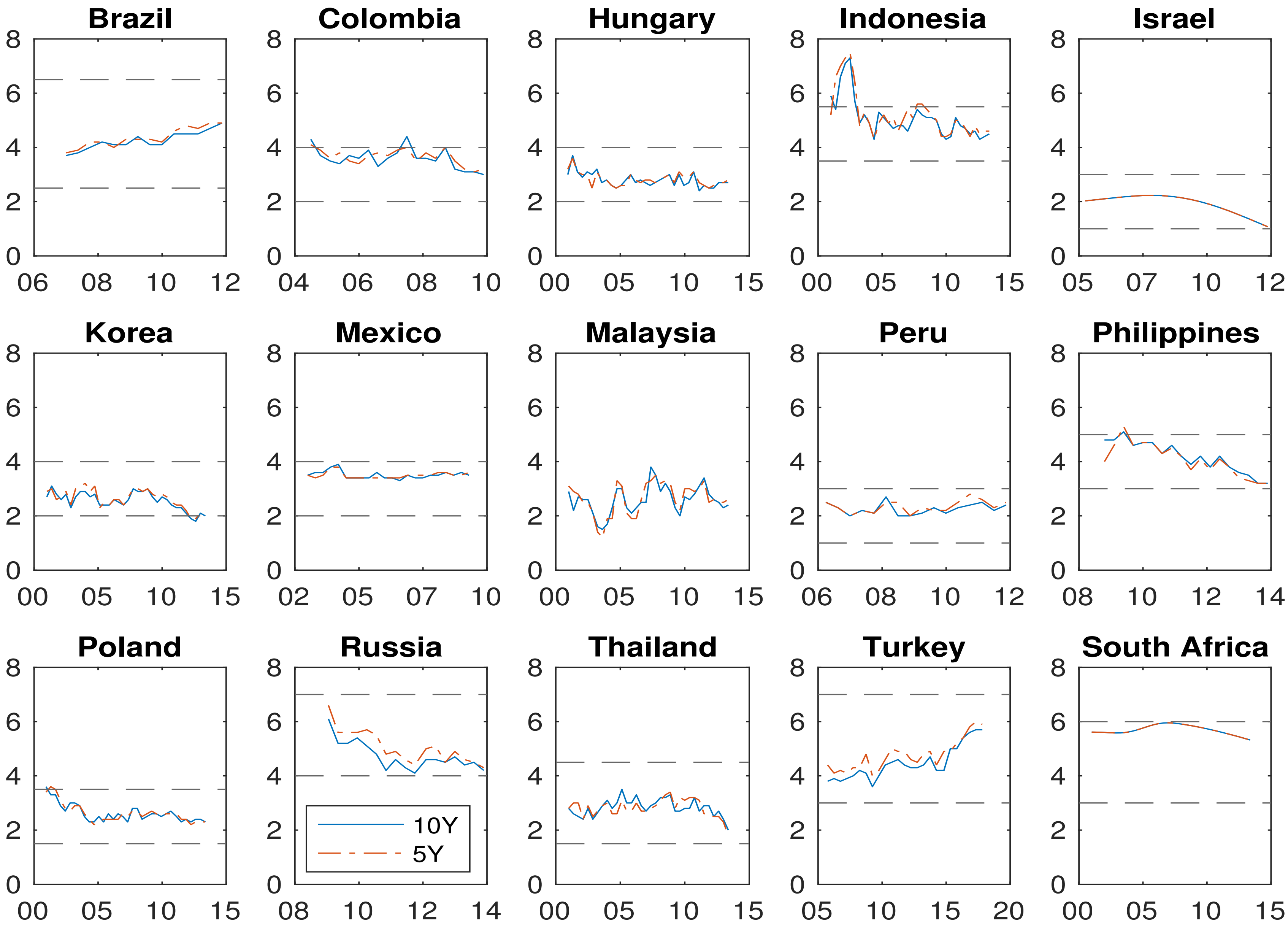


Figure 2. Long-Horizon Forecasts of Real GDP Growth

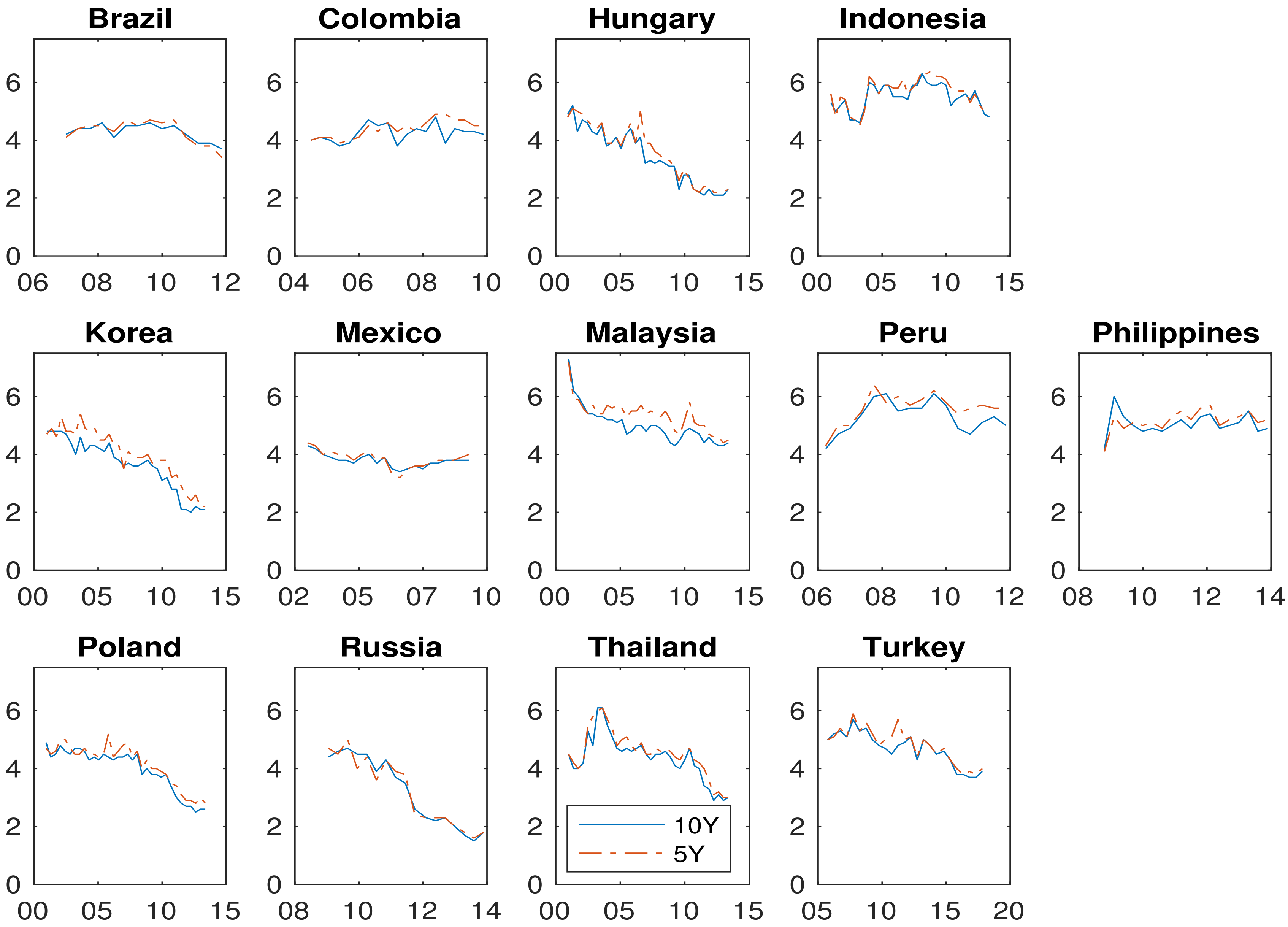


Figure 3. 10-Year Synthetic Yield and Long-Horizon Implied Forecasts for the Short Rate

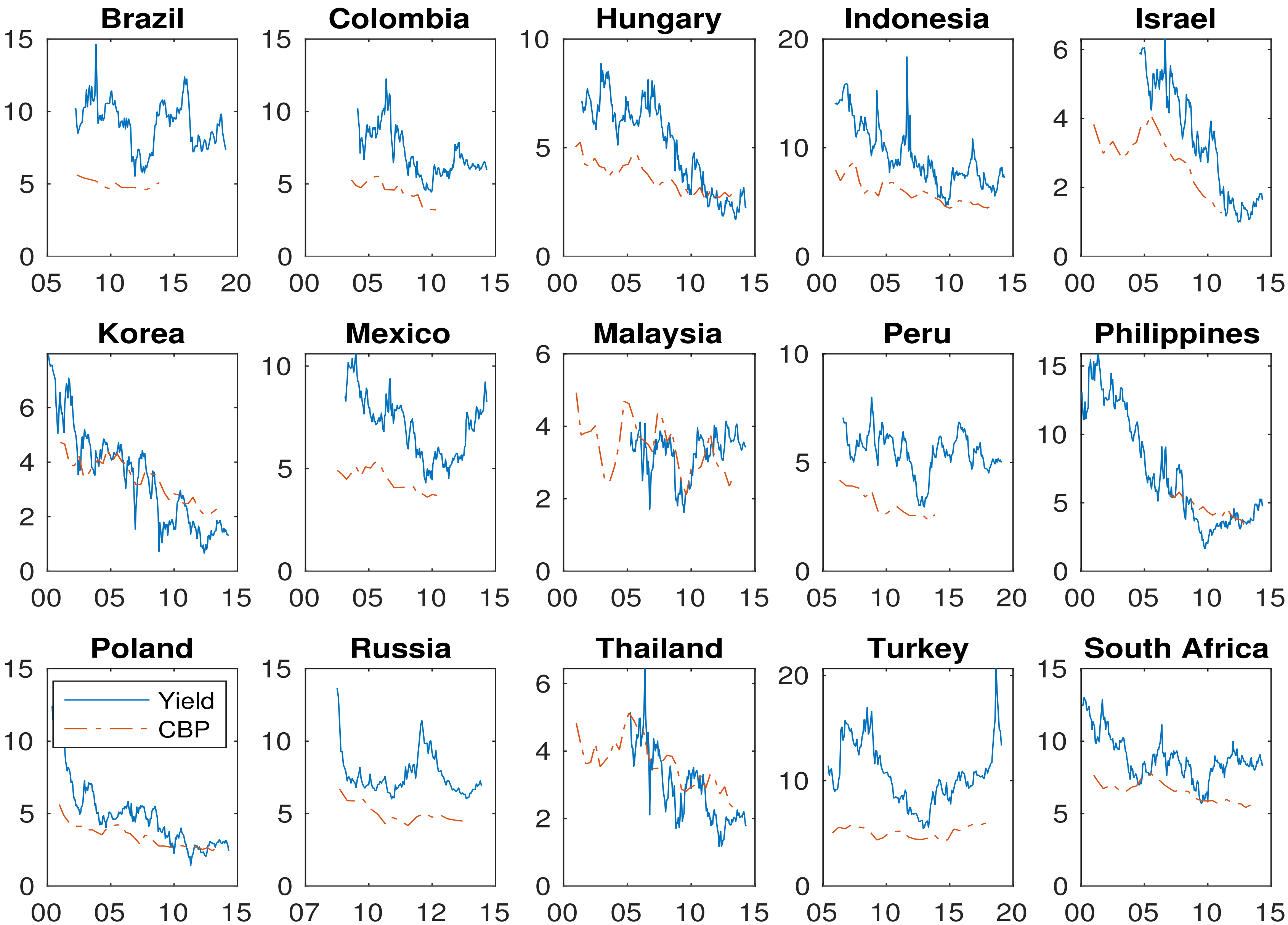


Figure 4. Trend versus Long-Horizon Forecasts of Inflation

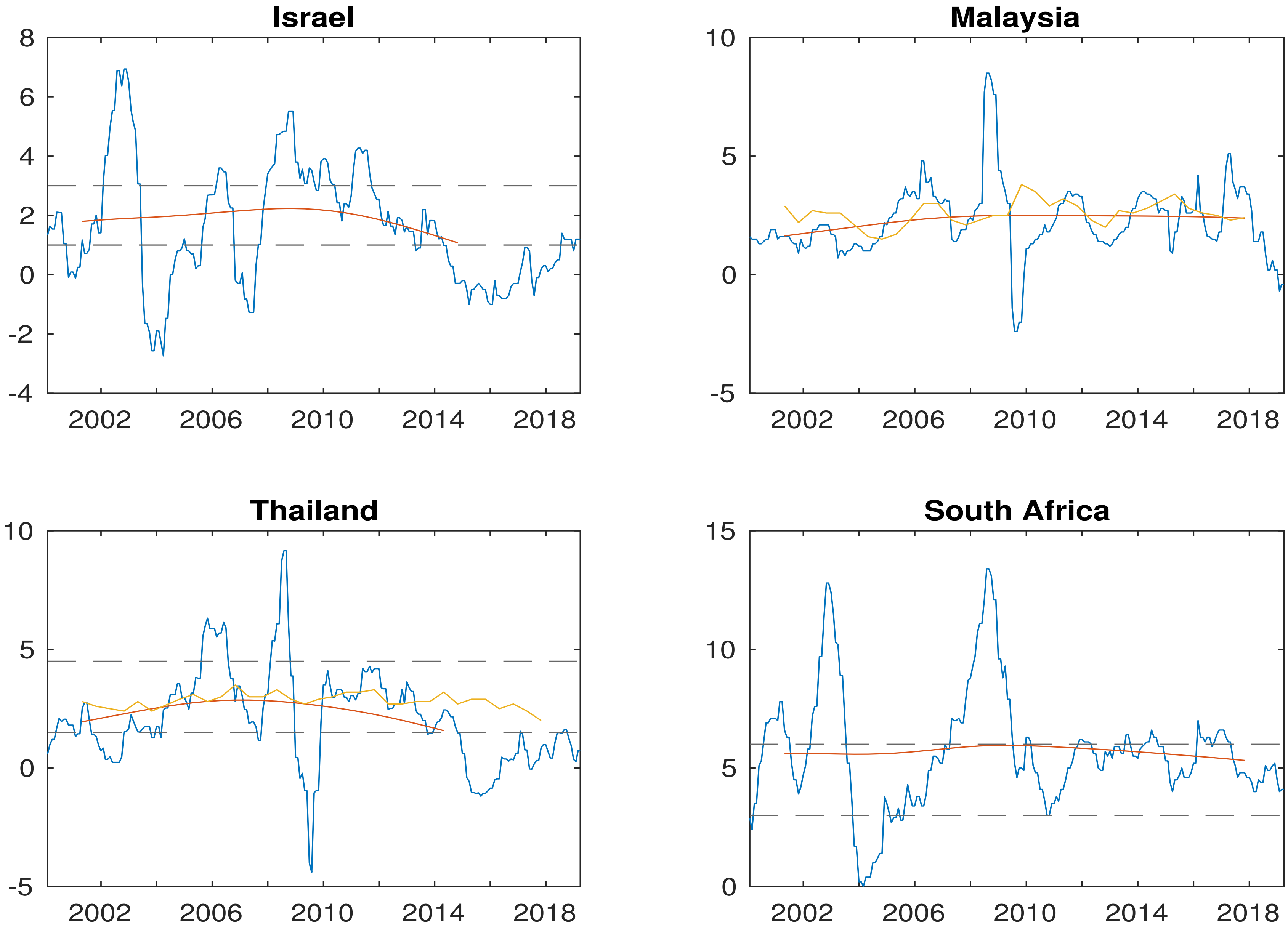


Figure 5. Model Fit for Emerging Markets: 10-Year Synthetic Yields

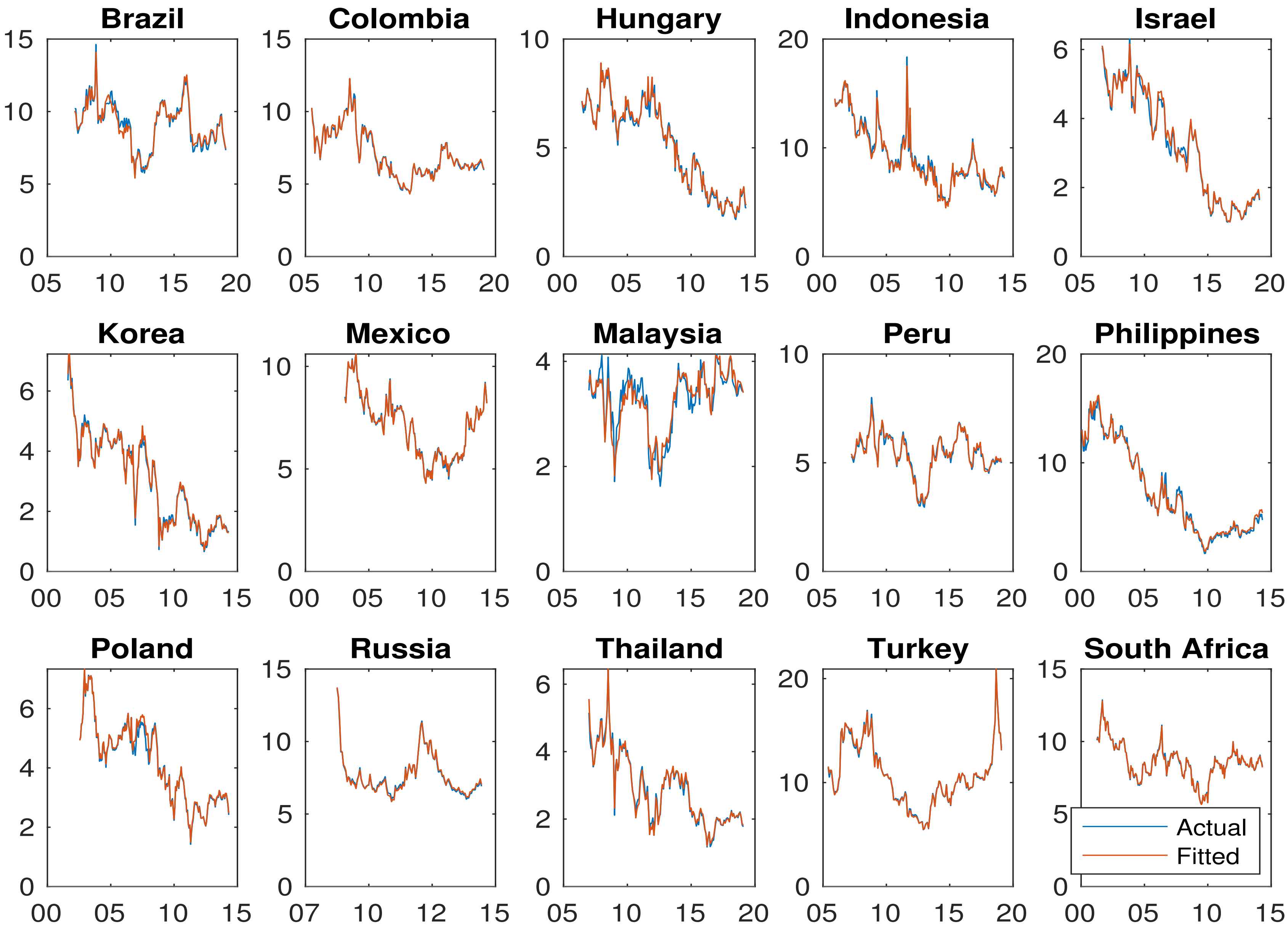


Figure 6. Decomposition of EM Nominal 10-Year Yields

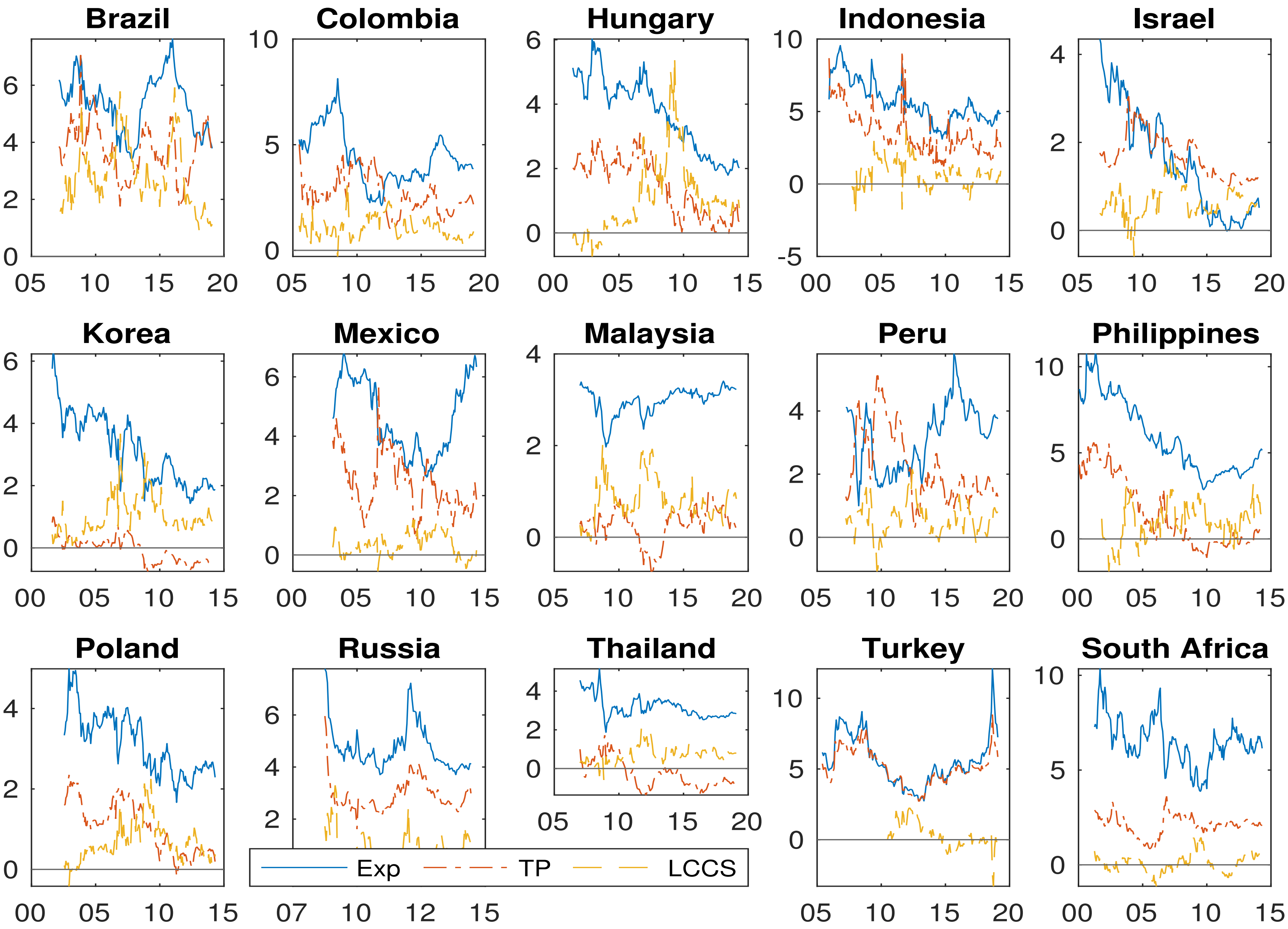


Figure 7. Interest Rate Forecasts vs Model-Implied Expectation: 10-Year Yields

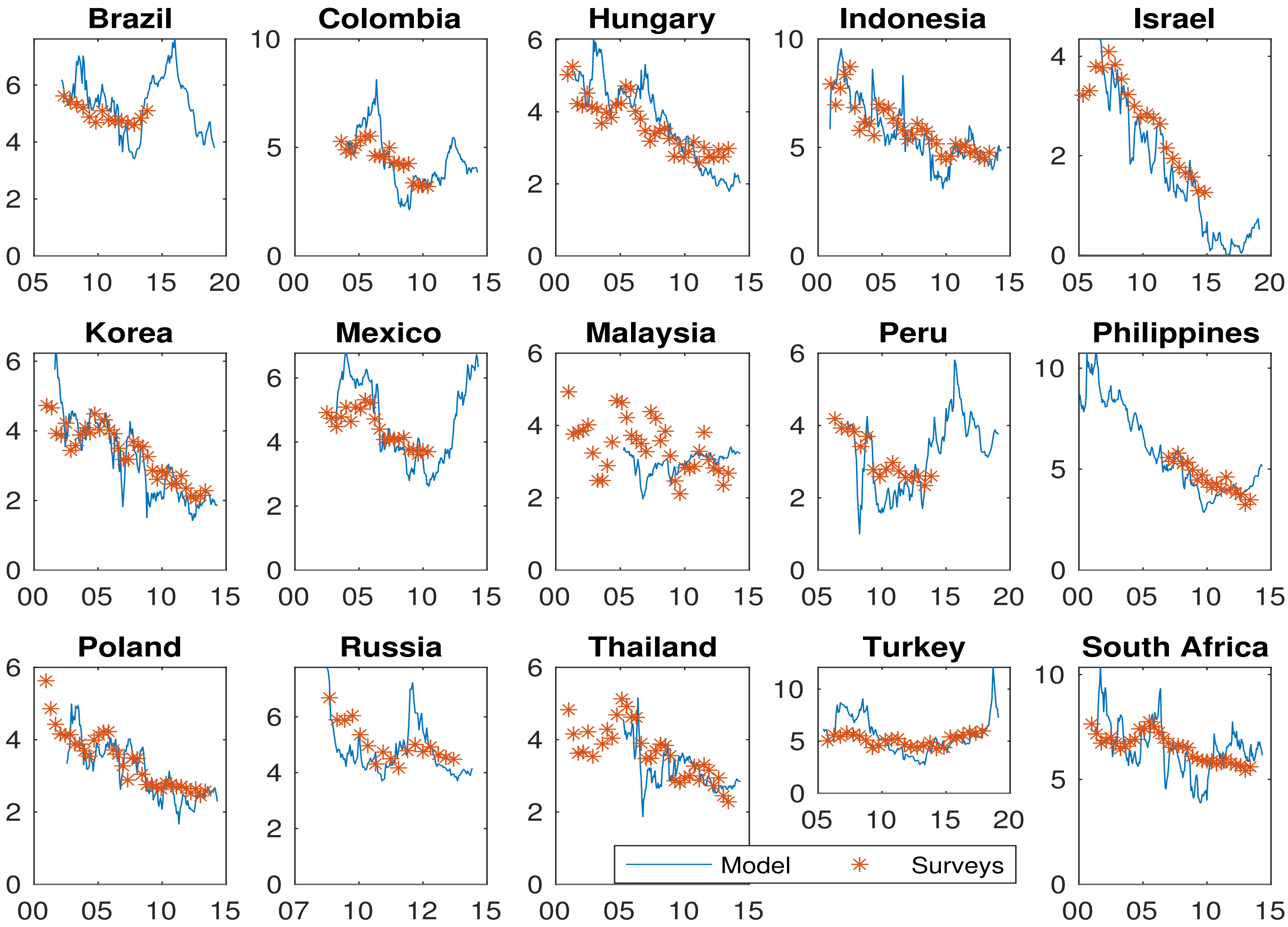


Figure 8. 10-Year Expected Real Interest Rate

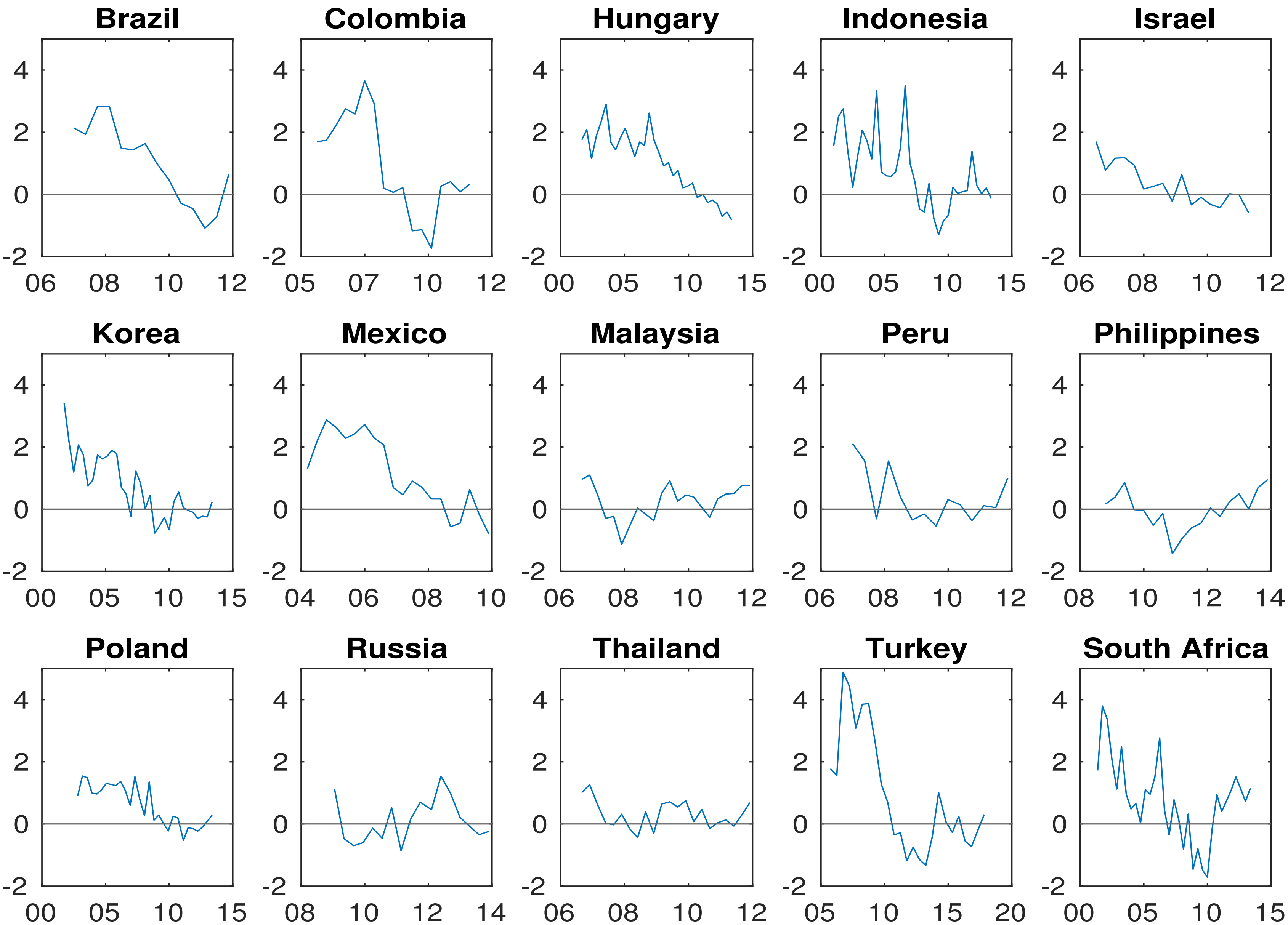
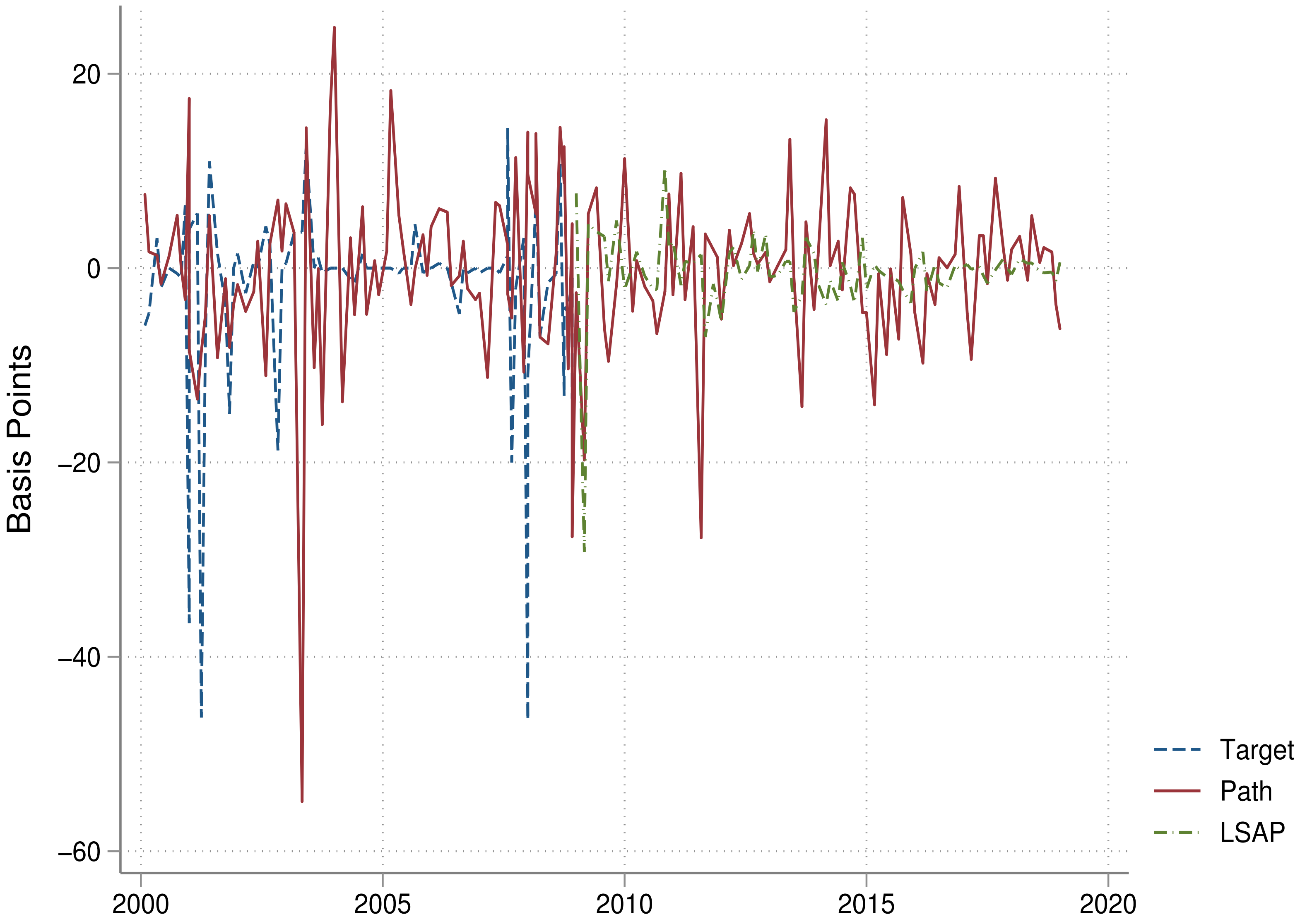


Figure 9. U.S. Monetary Policy Shocks



References

T. Adrian, R. K. Crump, and E. Moench. Pricing the Term Structure with Linear Regressions. *Journal of Financial Economics*, 110(1):110–138, 2013.

T. Adrian, R. K. Crump, J. B. Durham, and E. Moench. Sovereign Yield Comovement. *Working Paper*, 2019.

S. R. Baker, N. Bloom, and S. J. Davis. Measuring Economic Policy Uncertainty. *The Quarterly Journal of Economics*, 131(4):1593–1636, 2016.

S. E. Curcuru, S. B. Kamin, C. Li, and M. Rodriguez. International Spillovers of Monetary Policy: Conventional Policy vs. Quantitative Easing. *International Finance Discussion Paper*, 2018(1234), 2018.

M. Dahlquist and H. Hasseltoft. International Bond Risk Premia. In *Handbook of Fixed-Income Securities*, pages 169–190. 2016.

W. Du and J. Schreger. Local Currency Sovereign Risk. *Journal of Finance*, 71(3): 1027–1070, 2016a.

W. Du and J. Schreger. Sovereign Risk, Currency Risk, and Corporate Balance Sheets. *HBS Working Paper*, No. 17-024, 2016b.

W. Du, J. Im, and J. Schreger. A Dataset for Covered Interest Rate Parity Deviations Between Government Bond Yields. In *NBER IFM Data Sources Project*, 2018a.

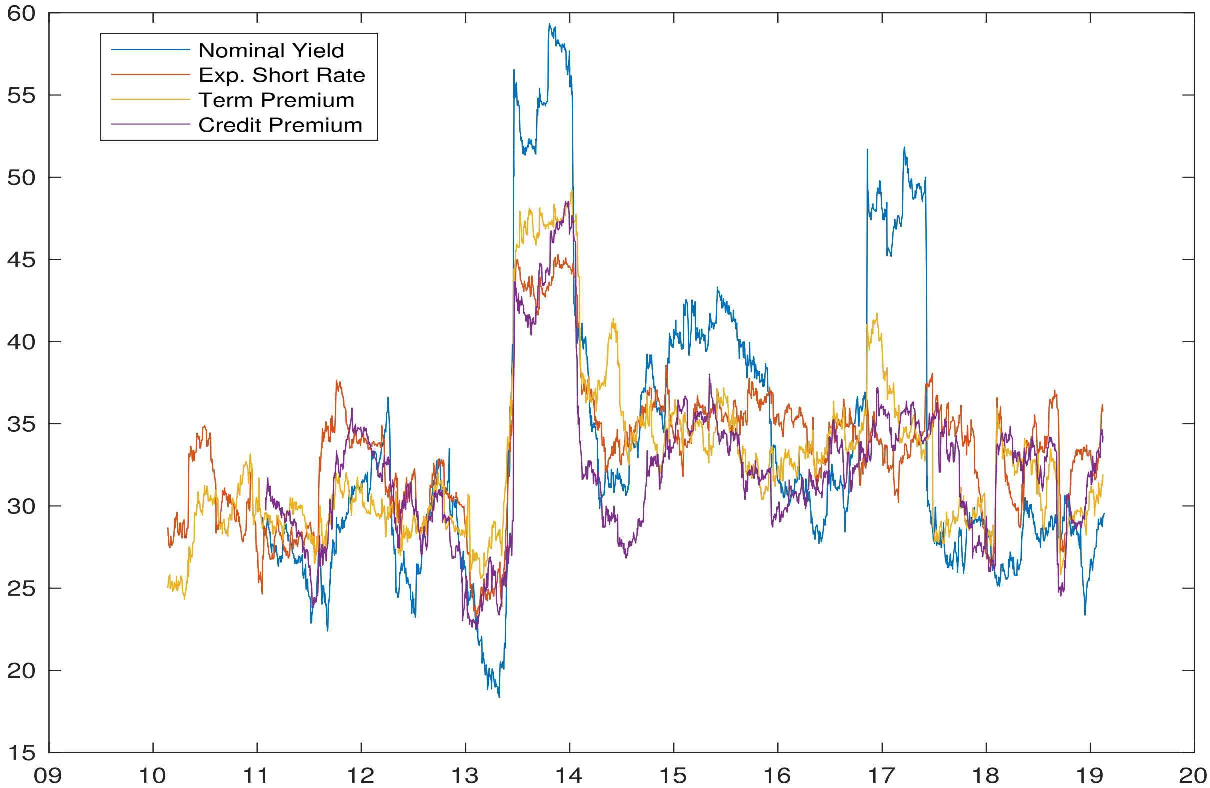
W. Du, J. Im, and J. Schreger. The U.S. Treasury Premium. *Journal of International Economics*, 112:167–181, 2018b.

W. Du, A. Tepper, and A. Verdelhan. Deviations from Covered Interest Rate Parity. *Journal of Finance*, 73(3):915–957, 2018c.

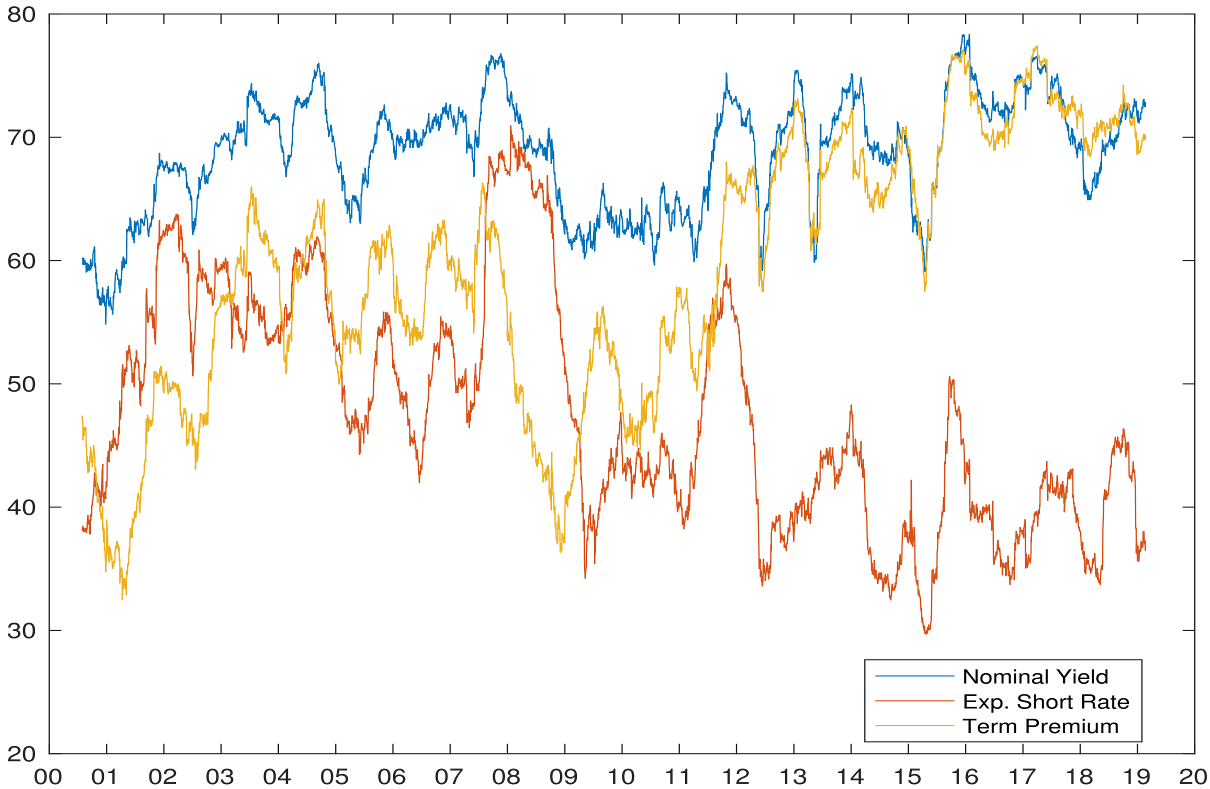
G. R. Duffee. Term Premia and Interest Rate Forecasts in Affine Models. *Journal of Finance*, 57(1):405–443, 2002.

G. R. Duffee. Sharpe Ratios in Term Structure Models. *Working Paper*, 2010.

Figure 10. Connectedness of Sovereign 10-Year Yields

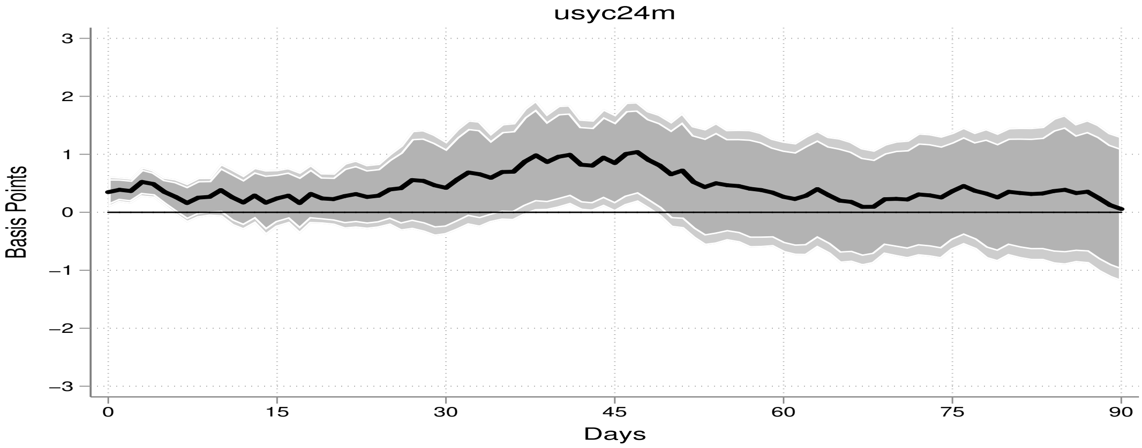


(a) Emerging Markets

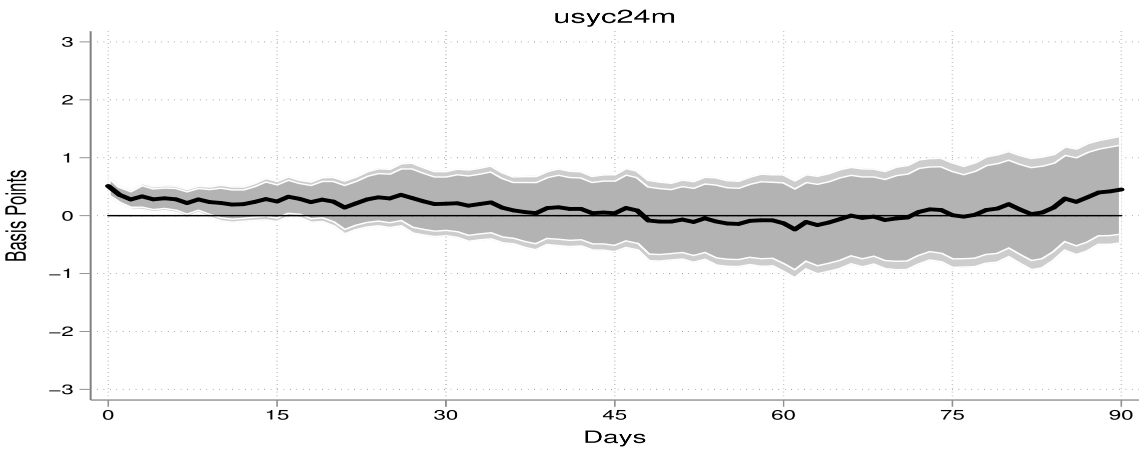


(b) Advanced Countries

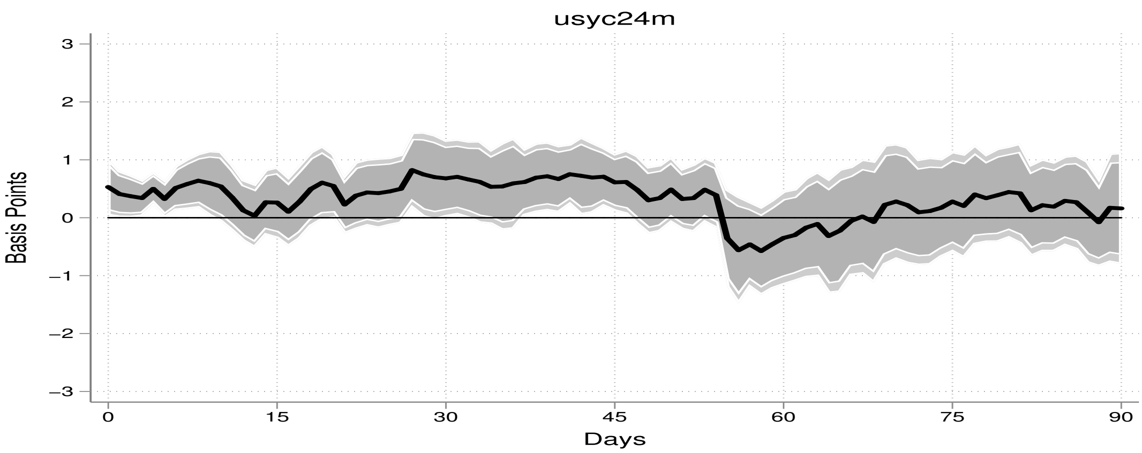
Figure 11. Response of 2Y U.S. Yield to U.S. Monetary Policy Shocks



(a) Target Shock: 2000-2008

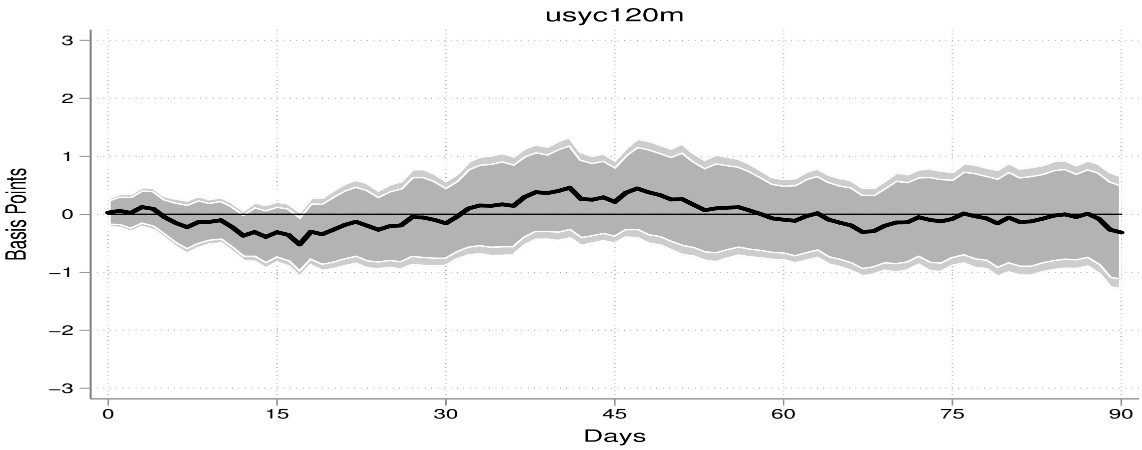


(b) Path Shock: 2000-2019

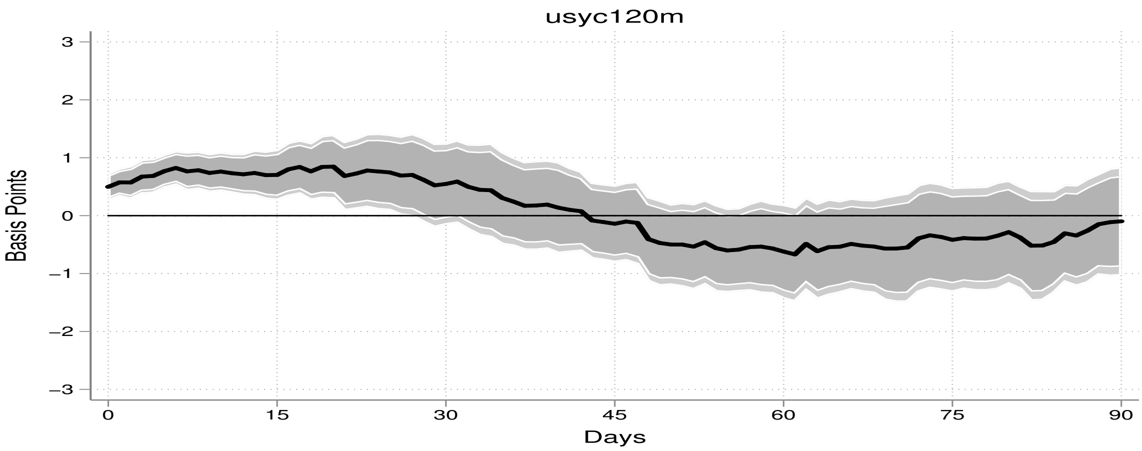


(c) LSAP Shock: 2009-2019

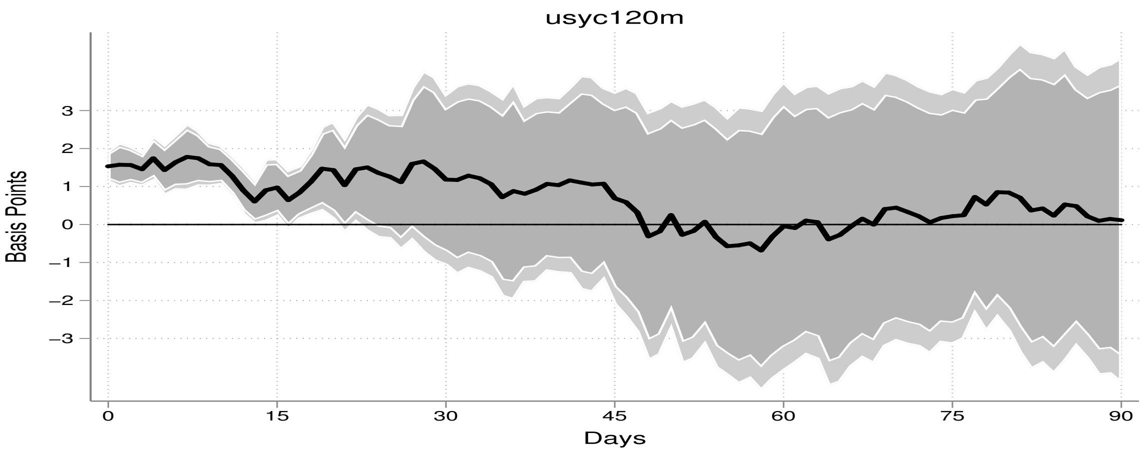
Figure 12. Response of 10Y U.S. Yield to U.S. Monetary Policy Shocks



(a) Target Shock: 2000-2008

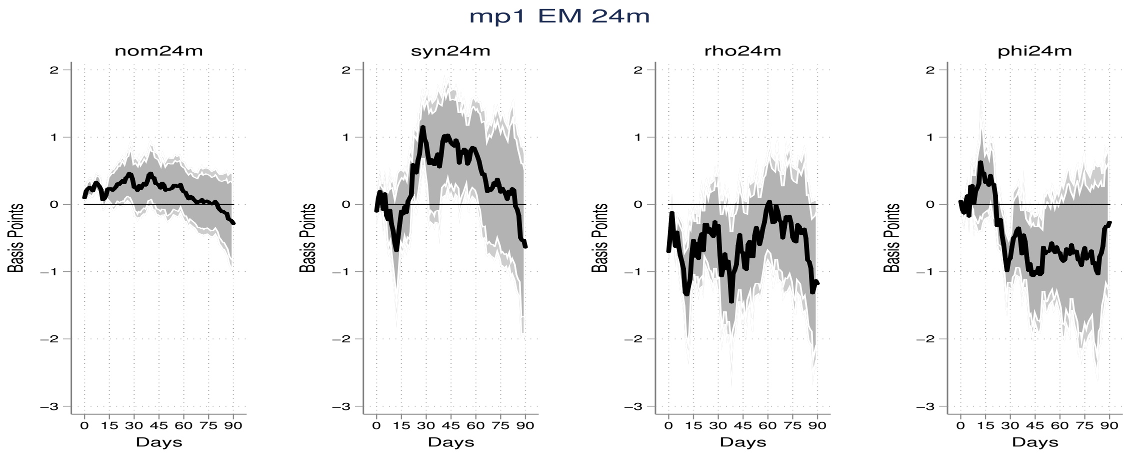


(b) Path Shock: 2000-2019

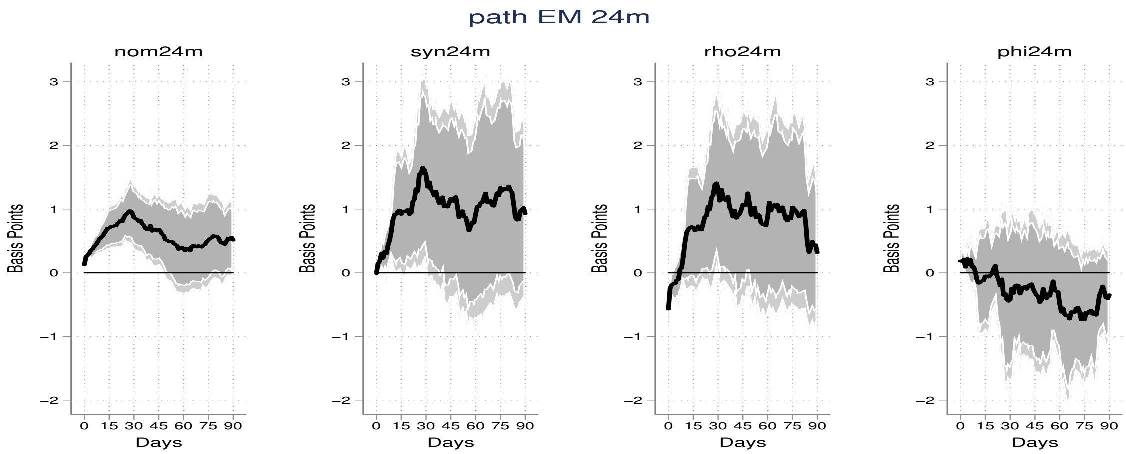


(c) LSAP Shock: 2009-2019

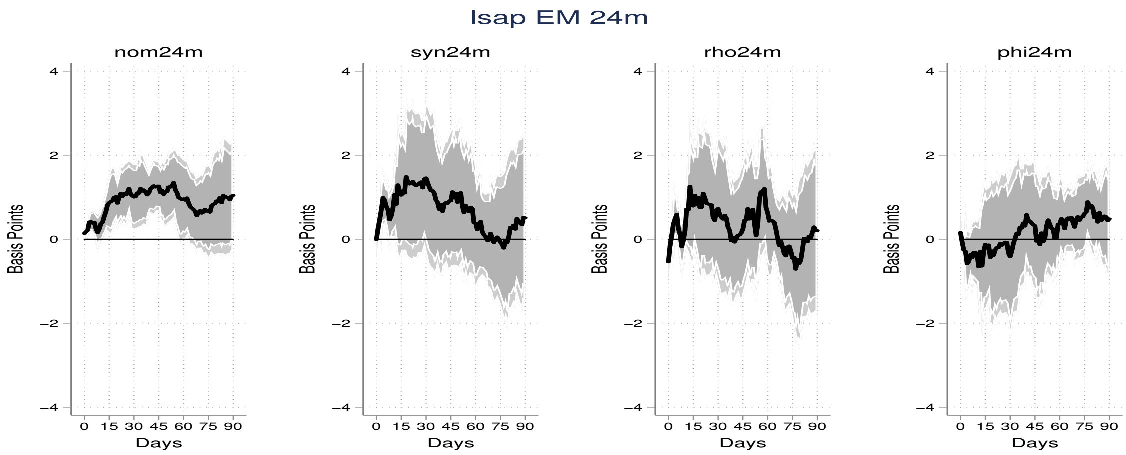
Figure 13. Response of 2Y EM Yields to U.S. Monetary Policy Shocks



(a) Target Shock: 2000-2008

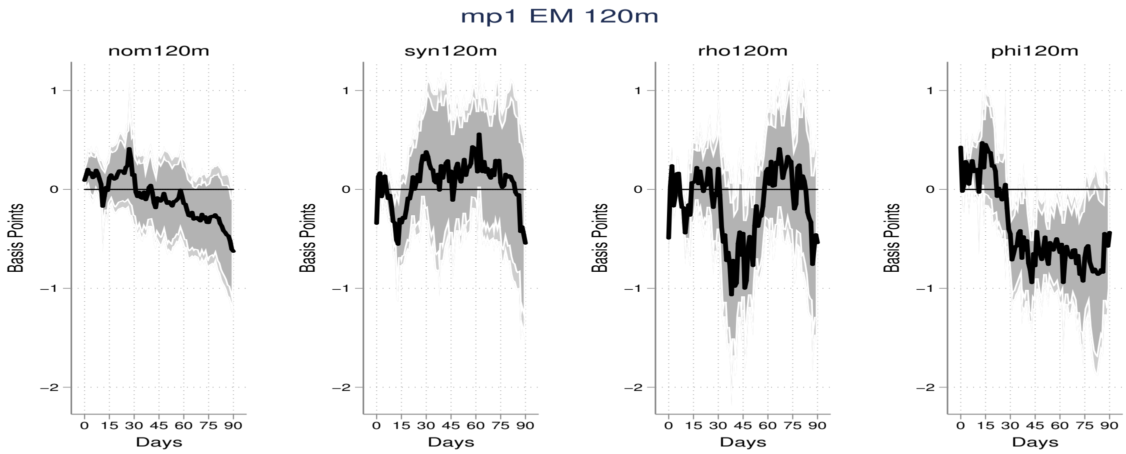


(b) Path Shock: 2000-2019

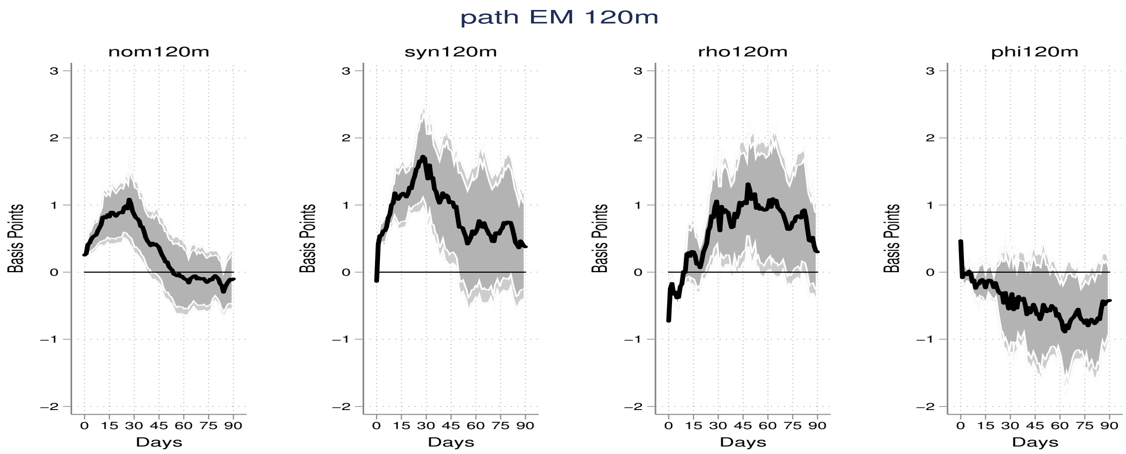


(c) LSAP Shock: 2009-2019

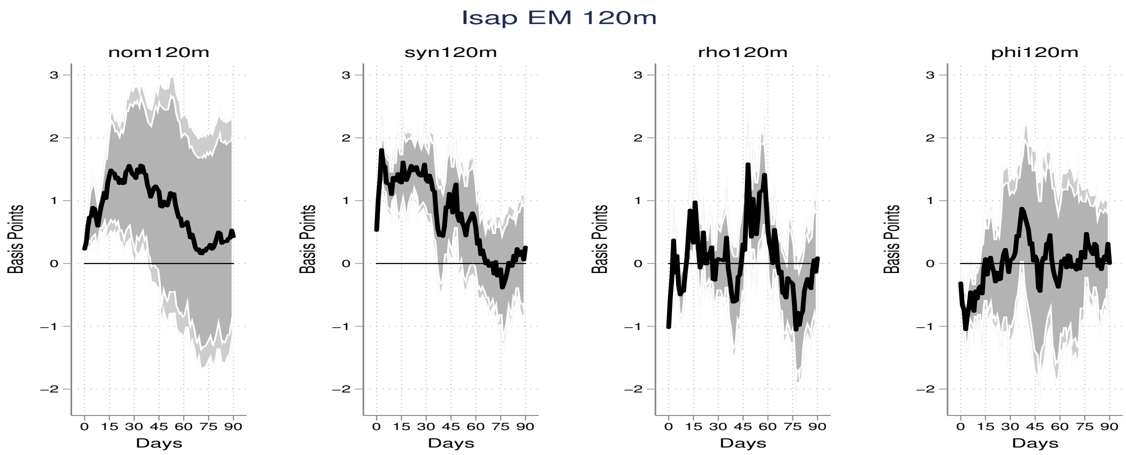
Figure 14. Response of 10Y EM Yields to U.S. Monetary Policy Shocks



(a) Target Shock: 2000-2008



(b) Path Shock: 2000-2019



(c) LSAP Shock: 2009-2019

A. Erce and E. Mallucci. Selective Sovereign Defaults. *Board of Governors of the Federal Reserve System Discussion Paper*, 1239, 2018.

C. Galli. Inflation, Default Risk and Nominal Debt. *Working Paper*, 2020.

S. Gilchrist, V. Yue, and E. Zakrajšek. U.S. Monetary Policy and International Bond Markets. *Journal of Money, Credit and Banking*, 51(S1):127–161, 2019.

A. Goliński and P. Spencer. Estimating the Term Structure with Linear Regressions: Getting to the Roots of the Problem. *Journal of Financial Econometrics*, pages 1–25, 2019.

R. Guimarães. Expectations, Risk Premia and Information Spanning in Dynamic Term Structure Model Estimation. *Bank of England Working Paper*, 489, 2014.

R. S. Gürkaynak and J. H. Wright. Macroeconomics and the Term Structure. *Journal of Economic Literature*, 50(2):331–367, 2012.

R. S. Gürkaynak and J. H. Wright. Identification and Inference Using Event Studies. *The Manchester School*, 81(S1):48–65, 2013.

R. S. Gürkaynak, B. P. Sack, and E. T. Swanson. Do Actions Speak Louder Than Words? The Response of Asset Prices to Monetary Policy Actions and Statements. *International Journal of Central Banking*, 1(1):55–93, 2005.

R. S. Gürkaynak, B. P. Sack, and J. H. Wright. The U.S. Treasury Yield Curve: 1961 to the Present. *Journal of Monetary Economics*, 54(8):2291–2304, 2007.

B. Hofmann, S. Ilhyock, and H. S. Shin. Bond Risk Premia and the Exchange Rate. *BIS Working Paper*, 775, 2019.

K. Holston, T. Laubach, and J. C. Williams. Measuring the Natural Rate of Interest: International Trends and Determinants. *Journal of International Economics*, 108:S59–S75, may 2017.

Ò. Jordà. Estimation and Inference of Impulse Responses by Local Projections. *American Economic Review*, 95(1):161–182, mar 2005.

S. Joslin, K. J. Singleton, and H. Zhu. A New Perspective on Gaussian Dynamic Term Structure Models. *Review of Financial Studies*, 24(3):926–970, 2011.

S. Kalemli-Özcan. U.S. Monetary Policy and International Risk Spillovers. *Working Paper*, 2019.

D. H. Kim and A. Orphanides. Term Structure Estimation with Survey Data on Interest Rate Forecasts. *Journal of Financial and Quantitative Analysis*, 47(1):241–272, 2012.

D. H. Kim and J. H. Wright. An Arbitrage-Free Three-Factor Term Structure Model and the Recent Behavior of Long-Term Yields and Distant-Horizon Forward Rates. *Board of Governors of the Federal Reserve System Discussion Paper*, 33, 2005.

K. N. Kuttner. Monetary Policy Surprises and Interest Rates: Evidence from the Fed Funds Futures Market. *Journal of Monetary Economics*, 47(3):523–544, 2001.

K. N. Kuttner. Outside the Box: Unconventional Monetary Policy in the Great Recession and Beyond. *Journal of Economic Perspectives*, 32(4):121–146, 2018.

R. Litterman and J. Scheinkman. Common Factors Affecting Bond Returns. *Journal of Fixed Income*, 1(1):54–61, 1991.

S. P. Lloyd. Estimating Nominal Interest Rate Expectations: Overnight Indexed Swaps and the Term Structure. *Bank of England Working Paper*, 763, 2018.

E. Nakamura and J. Steinsson. Identification in Macroeconomics. *Journal of Economic Perspectives*, 32(3):59–86, 2018.

C. R. Nelson and A. F. Siegel. Parsimonious Modeling of Yield Curves. *Journal of Business*, 60(4):473–489, 1987.

M. Obstfeld. Trilemmas and Trade-Offs: Living with Financial Globalisation. *BIS Working Paper*, 2015.

M. Obstfeld. Global Dimensions of U.S. Monetary Policy. *International Journal of Central Banking*, 16(1):73–132, 2020.

P. Ottonello and D. J. Perez. The Currency Composition of Sovereign Debt. *American Economic Journal: Macroeconomics*, 11(3):174–208, 2019.

G. Palladini and R. Portes. Sovereign CDS and Bond Pricing Dynamics in the Euro-Area. *NBER Working Paper*, 17586, 2011.

S. Pennings, A. Ramayandi, and H. C. Tang. The Impact of Monetary Policy on Financial Markets in Small Open Economies: More or Less Effective During the Global Financial Crisis? *Journal of Macroeconomics*, 44:60–70, 2015.

C. M. Reinhart and K. S. Rogoff. The Forgotten History of Domestic Debt. *Economic Journal*, 121(552):319–350, 2011.

H. Rey. Dilemma not Trilemma: The Global Financial Cycle and Monetary Policy Independence. *Proceedings of the 2013 Federal Reserve Bank of Kansas City Economic Symposium at Jackson Hole*, pages 285–333, 2013.

E. T. Swanson. Measuring the Effects of Federal Reserve Forward Guidance and Asset Purchases on Financial Markets. *Working Paper*, 2018.

J. H. Wright. Term Premia and Inflation Uncertainty: Empirical Evidence from an International Panel Dataset. *American Economic Review*, 101(4):1514–1534, 2011.

Contents

Todo list