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## **U.S. Monetary Policy Normalization and Global Interest Rates**

by Carlos Caceres, Yan Carrière-Swallow, Ishak Demir, and Bertrand Gruss

**I N T E R N A T I O N A L M O N E T A R Y F U N D**

## IMF Working Paper

Western Hemisphere Department

### U.S. Monetary Policy Normalization and Global Interest Rates

Prepared by Carlos Caceres, Yan Carrière-Swallow, Ishak Demir,<sup>1</sup> and Bertrand Gruss

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#### Abstract

As the Federal Reserve continues to normalize its monetary policy, this paper studies the impact of U.S. interest rates on rates in other countries. We find a modest but nontrivial pass-through from U.S. to domestic short-term interest rates on average. We show that, to a large extent, this co-movement reflects synchronized business cycles. However, there is important heterogeneity across countries, and we find evidence of limited monetary autonomy in some cases. The co-movement of longer term interest rates is larger and more pervasive. We distinguish between U.S. interest rate movements that surprise markets versus those that are anticipated, and find that most countries receive greater spillovers from the former. We also distinguish between movements in the U.S. term premium and the expected path of risk-free rates, concluding that countries respond differently to these shocks. Finally, we explore the determinants of monetary autonomy and find strong evidence for the role of exchange rate flexibility, capital account openness, but also for other factors, such as dollarization of financial system liabilities, and the credibility of fiscal and monetary policy.

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## I. INTRODUCTION

As the U.S. economic outlook strengthens, the Federal Reserve will continue to gradually normalize its monetary stance. After several years of policy rates at the zero lower bound, the deployment of unconventional asset purchase policies, and long-term rates and term premiums at historically low levels, many market analysts and policymakers were anxious about the global implications of the first Fed hike in more than 9 years. In the event, the first decision to increase the federal funds rate on December 16<sup>th</sup>, 2015, turned out to be relatively benign. Central banks with pegs to the U.S. dollar followed the Fed's increase in lock-step, and several central banks in Latin America have also raised their policy rates substantially in the following months. However, there is little evidence of generalized policy rate adjustments elsewhere (Figure 1, panel A). Meanwhile, longer-term interest rates have remained low, with the U.S. term premium compressing further from already low levels (Figure 1, panel B).

The full cycle of U.S. monetary policy tightening will take place over an extended period, progressing from the gradual increase of policy rates to the eventual unwinding of the Federal Reserve's balance sheet. While it is expected to accompany an ongoing economic recovery, the process risks triggering disruptions in global financial markets. First, Federal Reserve rate decisions at the early stages of the tightening cycle may lead agents to revise their expectations about the future path of U.S. short-term rates—which currently anticipate a very gradual pace of adjustment—leading to spikes in longer-term yields. Second, the unprecedented nature of the process could generate uncertainty about the future rate path of interest rates and increased risk aversion, both drivers of global term premiums and capital flows.

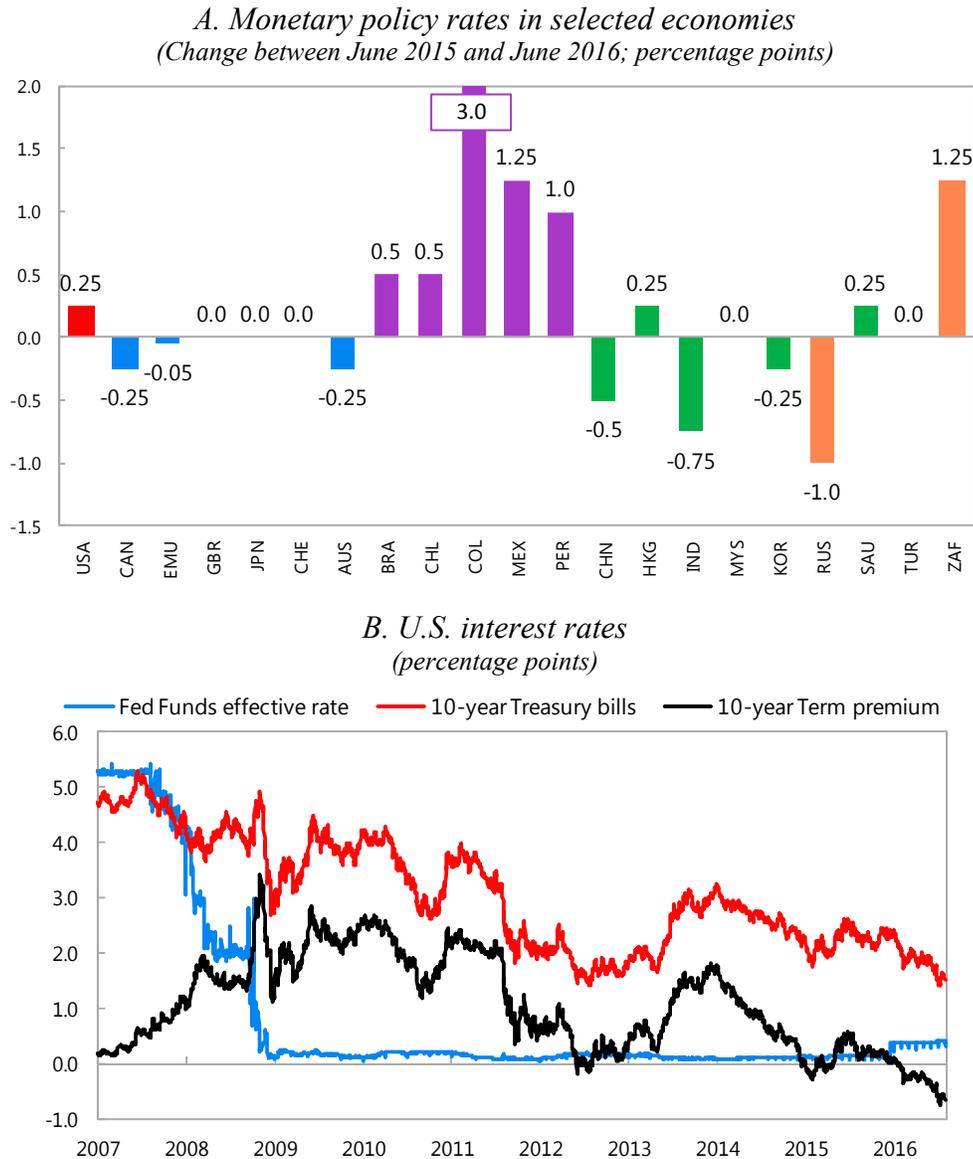
What may be the consequences for interest rates around the world as U.S. monetary policy continues to normalize? Does monetary tightening in the United States have a smaller impact on interest rates in other countries when it reflects good news about the U.S. economy? What are the global effects of shifts in the U.S. term premium? How do the implications from changes in U.S. rates compare with those originating in other major economies, such as the euro area? What policies can grant small open economies control over their monetary stance in the face of tightening global financial conditions?

This paper attempts to address these questions by exploring international financial linkages in a broad sample of 43 emerging and advanced economies since the early 2000s.<sup>1</sup> We begin by analyzing the pass-through of international interest rates for short and long-dated bonds. On average, we find a modest response of short term domestic rates to a change in the federal funds rate—with a 100-basis-point increase in U.S. rates leading to a response of about 20 basis points abroad—but with substantial heterogeneity across countries. The degree of interest rate interdependence is much higher and more homogenous across countries in the case of long-term rates. When 10-year U.S. bond yields rise by 100 basis points, domestic long-term rates increase by between 50 and 80 basis points in two-thirds of the countries in our sample.

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<sup>1</sup> This paper broadens the analysis presented in IMF (2015), which focused on emerging economies in Latin America and the Caribbean.

Figure 1. Recent interest rate developments



Sources: Haver Analytics; Bloomberg. For Saudi Arabia (SAU), the reverse repo rate is shown, provided by the Saudi Arabian Monetary Agency. Estimates of the term premium are provided by the Federal Reserve Bank of New York using the methodology in Adrian, Crump, and Moench (2013).

We then investigate whether the underlying drivers of interest rate movements in major economies determine their impact on interest rates in other countries. To do so, we use a decomposition of movements in U.S. and euro area 10-year bond yields that distinguishes the components that respond to economic news, those that don't, and those movements that are due to changes in risk aversion, and analyze how each are transmitted across borders. We find that monetary policy shocks in large economies that do not reflect a change in their economic outlook have historically played a key role among external drivers of domestic long-term interest rates in small open economies.

A distinctive feature at the current juncture is that the U.S. term premium has been at historically low levels for some time (see Figure 1, panel B). To shed light on the possible consequences of a sudden decompression, we consider a decomposition of movements in 10-year U.S. bond yields into shifts in the expected path of future short-term policy rates and in the term premium, and explore the differential impact these have on long-term interest rates elsewhere. We find that shocks to the U.S. term premium explain the bulk of the variability in long-term interest rates in small open economies that is attributable to changes in U.S. interest rates.

The normalization of U.S. monetary policy finds many economies operating below potential amid a persistent deceleration in economic activity. In many cases, central banks should in principle be able to attenuate the domestic impact of tightening global financial conditions by maintaining short-term interest rates accommodative. Whether they can do so in practice is subject to significant disagreement in policy circles and in the literature.<sup>2</sup> There is a widespread perception that the normalization of the U.S. monetary stance will be followed by interest rate increases in many countries, irrespective of their cyclical position. Indeed, international interest rates co-move strongly with external financial conditions, and with U.S. policy rates in particular.

In a highly integrated global financial system, will monetary authorities around the world be able to avoid a tightening of financial conditions that is not warranted by their domestic cycle? Or will tightening alongside the Federal Reserve become a necessity? What policies can help monetary authorities regain more autonomy vis-à-vis global financial shocks?

Caceres, Carrière-Swallow, and Gruss (2016) argue that inference about monetary autonomy based on estimates of interest rate spillovers can lead us to overstate the limits to monetary autonomy when business cycles are correlated across countries. The starting point of their analysis is the idea that countries facing similar shocks should be expected to respond with similar monetary policies, even if these are completely inward-looking. They propose a two-stage procedure to estimate *autonomy-impairing spillovers* from global financial conditions, which are defined as the subset of responses of domestic rates to foreign shocks that are above and beyond what can be explained by the central bank's pursuit of domestic monetary objectives.

We follow this approach to infer limits to monetary autonomy in a sample of more than 40 advanced and emerging economies since the early 2000s. In order to partial out the policy response to changes in domestic macroeconomic conditions, we start by modeling inward-looking policy reaction functions for each country. In a second stage, we estimate whether the changes in domestic interest rates that are not endogenous reactions to changes in local macroeconomic conditions may have been driven by U.S. interest rates. We find that a majority of countries in our sample have managed to tailor their monetary stance to domestic

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<sup>2</sup> The debate on the ability of open economies to implement autonomous monetary policies in the context of a highly integrated global financial system has intensified recently. See, for instance, Rey (2015), Obstfeld (2015), and Caceres, Carrière-Swallow, and Gruss (2016).

conditions. However, in some cases, monetary conditions did tend to deviate from domestic objectives following movements in U.S. interest rates.

We then analyze the determinants of autonomy-impairing spillover estimates across countries, focusing on the policies that can be used to enhance monetary autonomy. In line with the classical monetary trilemma, we find that exchange rate flexibility plays a key role in ensuring that central banks in open economies can gear monetary policy towards stabilizing the domestic economy. But other policies can also affect monetary autonomy. We find that stronger monetary and fiscal credibility, an active use of reserve requirements, and lower financial dollarization, are also associated with increased monetary autonomy.

Throughout our analysis, we focus on the effects of U.S. monetary policy normalization on global interest rates. Of course, many other dimensions of its potential repercussions are worth exploring. Recent studies have argued that a global financial cycle—largely but not exclusively driven by U.S. monetary policy—affects asset prices, capital flows, and credit growth across countries (e.g., Rey, 2015). Where monetary transmission is incomplete, this could well be the case even where central banks have full autonomy to influence short-term rates. For instance, overall credit growth may be significantly influenced by global financial conditions if domestic firms finance their activities by issuing bonds abroad. These are all interesting dimensions of monetary policy transmission in open economies, but they go beyond the scope of this paper.

The structure of the paper is as follows: Section II presents the data used in the analysis. Section III describes the empirical methodology and presents estimates of pass-through from U.S. to domestic interest rates. Section IV digs deeper into what can be expected from the normalization of U.S. monetary policy by exploring how the pass-through depends on the nature of the factors driving U.S. interest rate movements. Section V moves from interest rate pass-through to spillovers, and discusses the monetary autonomy that countries enjoy to decouple their policies from rising U.S. rates. Section VI explores which factors may improve the degree of monetary policy autonomy in small open economies, and Section VII concludes.

## II. DATA

Financial conditions are determined by the price and quantities of a broad set of financial instruments, including debt and equity. We limit our analysis to interest rates on nominal public debt instruments denominated in local currency that are traded on secondary markets. At the short end of the yield curve, we focus on instruments with remaining maturity between three and six months, and at the long end of the curve, with remaining maturity of 10 years. When government bond yields are not available, quasi-sovereign bonds issued by the central bank for monetary policy operations are used.

Our choice of bond yields rather than monetary policy or money market rates is motivated by two considerations. First, long time series of policy rates are often discontinuous, as the choice of policy instrument changes over time.<sup>3</sup> Second, while money market rates are

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<sup>3</sup> For instance, Chile shifted its policy rate from a real (inflation-indexed) rate to a nominal rate in August 2001.

widely available and are typically more comparable across countries, they are subject to time-varying volatility that is unrelated to monetary policy, which is particularly pronounced among emerging economies with less developed financial systems. For instance, sudden liquidity shortages can lead to very large spikes in nominal rates despite monetary policy having remained unchanged.

One limitation of using yields on public debt is that the data are often subject to gaps. To overcome this limitation and in the hope of using a panel that is as balanced as possible, we interpolate the series using the variability in yields for instruments of similar maturities. For short-term rates, we use instruments with remaining maturities between one month and two years, and for long-term rates, we use remaining maturities between five and 20 years. Our primary data sources are generic bond estimates provided by Bloomberg and the series for treasury bills and government bond yields provided by the IMF's *International Financial Statistics*. While sources vary by country, instrument and time period, we supplement these primary sources with data from the IMF's monetary surveys, Haver Analytics, Global Financial Data, and national authorities.<sup>4</sup> The result is a comprehensive database of interest rates for a set of 43 emerging and advanced countries, at monthly frequency from January 2000 through October 2015.

### III. PASS-THROUGH FROM U.S. TO DOMESTIC INTEREST RATES

How are changes in global or U.S. financial conditions transmitted to interest rates in other economies? We estimate a set of country-specific vector autoregression (VAR) models using monthly data since the early 2000s to quantify the reaction of domestic interest rates to changes in U.S. interest rates. These estimates should be interpreted as correlations between interest rates, or *pass-through*, and are intended to provide a sense of how rates have moved together in the past without implying causality or limits to monetary policy autonomy.

Our analysis is largely focused on the effects of changes in U.S. interest rates, as these are a key driver of global financial conditions (see, for instance, Rey, 2015 and Ricci and Shi, 2016). In section IV, we consider alternative model specifications that include financial conditions in the euro area.

Our workhorse model is a VAR that includes a small open economy block exogeneity assumption, such that lags of domestic conditions do not affect the external variables. The model can be written as:

$$\begin{bmatrix} \mathbf{z}^* \\ \Delta i \end{bmatrix}_t = \mathbf{B}_0 + \sum_{j=1}^2 \mathbf{B}_j \begin{bmatrix} \mathbf{z}^* \\ \Delta i \end{bmatrix}_{t-j} + \begin{bmatrix} \mathbf{v}^* \\ v \end{bmatrix}_t, \quad (1)$$

where  $i$  denotes domestic interest rates in the small open economy (short or long-term rates, depending on the question) and the vector  $\mathbf{z}^*$  includes changes in U.S. interest rates ( $\Delta i^*$ , either the federal funds rate or the 10-year Treasury bond yield, depending on the question)

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<sup>4</sup> See Annex B for further details of index construction, and country-level data sources.

as well as global risk sentiment proxied by the VIX index.<sup>5</sup> The reduced-form error terms  $v$  and  $v^*$  are assumed to be independent and identically distributed with bounded second moments. The matrices  $B_j$  are restricted to ensure the block exogeneity of  $z_t^*$ , such that global variables are not affected by lags of domestic variables.

We focus our analysis on the cumulative Cholesky-orthogonalized impulse response of domestic interest rates ( $\Delta i$ ) to a shock in U.S. rates ( $\Delta i^*$ ).<sup>6</sup> This timing restriction assigns exogeneity to movements in U.S. monetary policy—an assumption that is likely invalid in a context of common global shocks that may have been driving both U.S. and domestic rates simultaneously.

### ***Short-term interest rates***

Table 1 (first column) reports pass-through estimates for our full sample of 43 countries. It shows the cumulative response of the domestic short-term rate 12 months after a shock that raises the Federal Funds rate by 100-basis-points over the same period. The median response for the set of emerging markets is 14 basis points, while the median response for advanced economies is 23 basis points. These statistics mask a good deal of heterogeneity: domestic short-term rates react quite differently across countries to movements in the federal funds rate.

Only 11 countries in our sample—just over one quarter—exhibit a cumulative response that is statistically significant after 12 months at the 10 percent confidence level. Interestingly, these countries are also rather heterogeneous, including both developing and advanced economies, as well as countries with both fixed and flexible exchange rates with respect to the dollar. The magnitude of their responses also varies greatly across economies. For instance, a 100-basis-point hike in the federal funds rate is met by a relatively large increase in short-term interest rates in Mexico and Hong Kong SAR (89 basis points and 65 basis points, respectively). The response to the same shock in Switzerland and Thailand is about 30 basis points, and only five basis points in the case of Japan.

### ***Long-term interest rates***

Movements in short-term rates are only part of the story, since many economic decisions are driven by longer-term rates. Moreover, since the policy rate in the United States hit the zero lower bound during the global financial crisis, the Federal Reserve has conducted monetary policy by influencing the longer end of the yield curve through quantitative easing and forward guidance. Furthermore, the degree of control that central banks exert over short-term interest rates in their respective countries may be substantially greater than their control over longer-term rates.

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<sup>5</sup> Chen, Mancini-Griffoli, and Sahay (2014) find that the VIX may amplify or dampen the effects of U.S. monetary policy. But they also argue that it contains additional information that may affect global asset prices, such as investor sentiment and risk appetite. Based on this, we include the VIX in the exogenous block of the model.

<sup>6</sup> Throughout the paper we focus on models' cumulative impulse response functions after 12 months to allow transmission to be fully realized.

Table 1. Cumulative impulse response of domestic rates after 12 months

Country	Short-term interest rates		Long-term interest rates		
	Pass-through	Spillover	Pass-through	Spillover	
ARG	Argentina	0.42	0.28		
BOL	Bolivia	0.14	0.20		
BRA	Brazil	-0.67	-0.57	1.28 *	1.15 *
CHL	Chile	-0.08	0.02	0.57 *	0.51 *
COL	Colombia	-0.17	-0.32	1.23 *	1.11 *
CRI	Costa Rica	0.02	0.13	0.39	0.56
MEX	Mexico	0.89 *	0.66 *	0.56 *	0.63 *
PER	Peru	0.26	0.39	0.85 *	0.78 *
URY	Uruguay	-0.20	0.08		
ARM	Armenia	1.10	-0.06	-0.37	-0.32
AUS	Australia	0.12	0.07	0.79 *	0.50 *
CAN	Canada	0.60 *	0.37 *	0.72 *	0.64 *
CHN	China	0.16	0.14	0.29	0.14
HRV	Croatia	0.04	0.18	-0.11	-0.09
CZE	Czech Republic	-0.13	-0.03	0.66 *	0.45 *
DNK	Denmark	0.20	0.15	0.68 *	0.53 *
EGY	Egypt	0.26	0.30		
HKG	Hong Kong SAR	0.65 *	0.68 *	1.05 *	0.73 *
HUN	Hungary	-0.06	-0.19	-0.04	-0.03
IND	India	0.51 *	0.46 *	0.45 *	0.34 *
IDN	Indonesia	-0.26	-0.35	0.82 *	0.70 *
ISR	Israel	0.64 *	0.39 *	0.58 *	0.45 *
JPN	Japan	0.05 *	0.02	0.30 *	0.24 *
LVA	Latvia	-0.12	0.05	-0.25	-0.27
MYS	Malaysia	0.06	0.01	0.45 *	0.33 *
NZL	New Zealand	0.23	0.12	0.79 *	0.62 *
NOR	Norway	0.29	0.11	0.67 *	0.50 *
PAK	Pakistan	0.45	0.43 *	1.02 *	0.57 *
PHL	Philippines	0.34	0.34	0.74 *	0.53 *
POL	Poland	0.28	0.19	0.86 *	0.47 *
ROM	Romania	-1.14	0.29	0.76	0.83 *
RUS	Russia	-0.74	-0.41	-0.05	0.08
SAU	Saudi Arabia	0.64 *	0.35 *	0.45	
SGP	Singapore	0.46 *	0.33 *	0.74 *	0.50 *
ZAF	South Africa	0.02	-0.04	0.73 *	0.64 *
KOR	South Korea	0.23	0.12	0.47 *	0.29 *
SWE	Sweden	0.24	0.04	0.67 *	0.48 *
CHE	Switzerland	0.30 *	0.20 *	0.35 *	0.29 *
TWN	Taiwan	0.32 *	0.16 *	0.37 *	0.25 *
THA	Thailand	0.31 *	0.14 *	0.77 *	0.54 *
TUR	Turkey	0.24	0.19	1.76 *	2.31 *
GBR	United Kingdom	0.22	0.04	0.68 *	0.54 *
VNM	Vietnam	-0.04	-0.04		
<b>Median</b>					
	Sample	0.23	0.14	0.67	0.50
	Advanced	0.23	0.12	0.67	0.49
	Emerging	0.14	0.14	0.65	0.54

Note: The table reports the cumulative impulse response of domestic rates after one year to a shock to the federal funds rate that leaves it 100 basis points higher. \* denotes statistical significance at the 10 percent level.

Overall, we find that movements in 10-year U.S. bond yields typically have a greater impact on corresponding domestic rates compared to changes in the federal funds rate (Table 1, third column). Indeed, the response of domestic long-term interest rates appears to be statistically significant at the 10 percent confidence level for 29 out of 38 countries in our sample—roughly three quarters. The responses are also more similar across countries than in the case of short-term rates: following a 100-basis-point increase in the U.S. Treasury yield, roughly two-thirds of these countries exhibit a cumulative response in the range of 50 to 80 basis points. Notable exceptions where pass-through exceeds one-to-one include Turkey, with a response of 176 basis points, followed by Brazil and Colombia with roughly 128 basis points. Overall, the median pass-through to long-term rates for advanced economies is 67 basis points, and 65 basis points for emerging economies.

#### **IV. WHAT CAN BE EXPECTED FROM U.S. MONETARY POLICY NORMALIZATION?**

The analysis presented so far reflects overall historical responses of domestic interest rates to movements in U.S. interest rates. Some of these movements in U.S. interest rates reflected the Federal Reserve's usual response to changes in economic conditions. For instance, rates tend to be increased when aggregate demand is seen to be putting upward pressure on prices. In contrast, others might not reflect a change in the economic outlook, or could deviate from the central bank's historical reaction function. These movements are more likely to surprise financial markets, and may also trigger a change in uncertainty regarding the future path of monetary policy or price inflation.

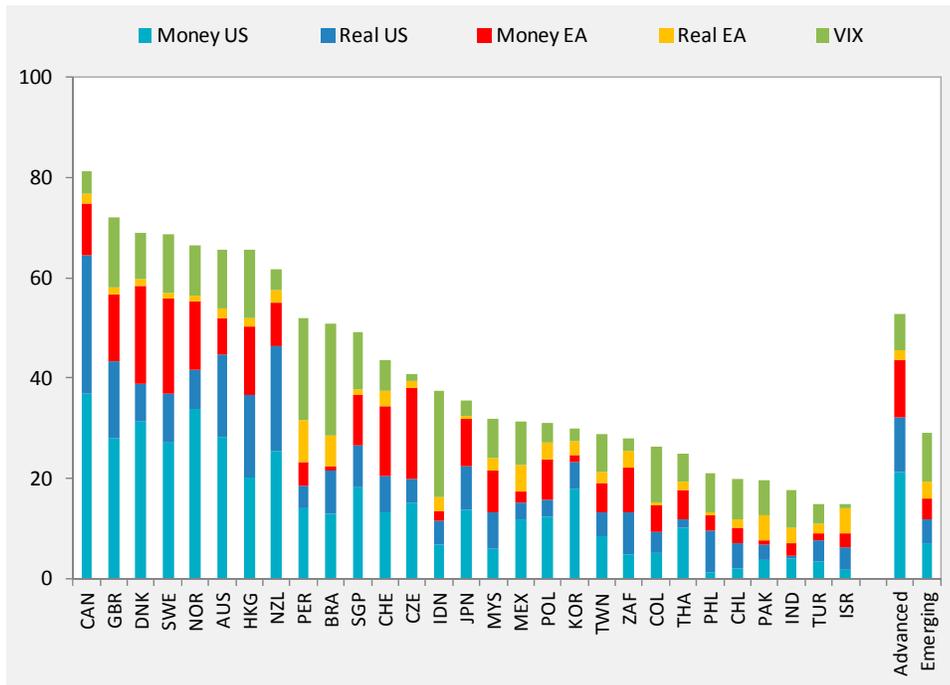
The effects on domestic interest rates are likely to depend on the underlying reason for the change in U.S. rates. Noting that the normalization of U.S. monetary policy reflects the ongoing recovery of the U.S. economy following the global financial crisis, some observers have predicted that its impacts will be small. As alluded to previously, other aspects of the current environment raise important questions. For instance, will normalization be attenuated by accommodative monetary policy in other major advanced economies? Will the term premium remain compressed? In this section, we attempt to shed light on these issues.

##### ***Do the nature and source of global financial shocks matter?***

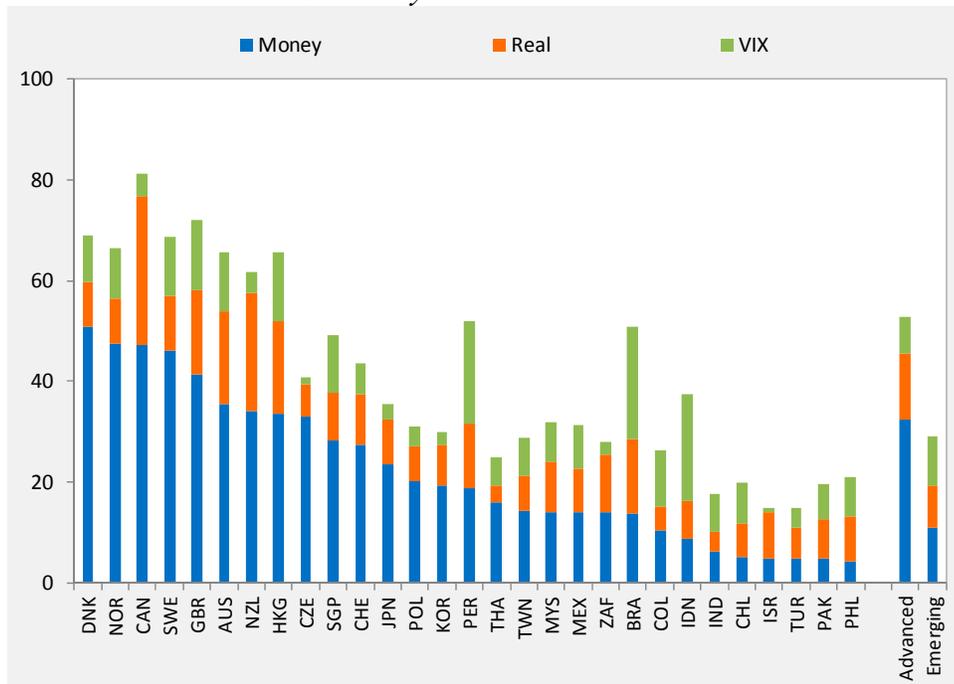
As Romer and Romer (2004) demonstrate, the impact of U.S. interest rate movements on the real economy depends on the underlying developments behind these decisions. Bluedorn and Bowdler (2010) show that the degree to which FOMC announcements are anticipated by markets has a bearing on the transmission of U.S. monetary policy to asset prices in other countries. The impact of a Federal Reserve policy decision will likely differ if it responds to a better economic outlook or reflects tighter monetary conditions alone. One reason is that decisions responding to an improved economic outlook are easier to anticipate, and may already be priced in by financial markets before they occur. An unanticipated rate hike, in turn, is likely to generate sharper adjustments of asset prices than one that has been fully anticipated. Another channel is that the better U.S. outlook will itself have implications for many global variables, including demand for exports and commodity prices, which will affect countries differently.

Figure 2: Drivers of long-term interest rates, 2000-15: Expected and unexpected shocks to 10-year bond yields in the U.S. and the euro area

*A. All shocks*



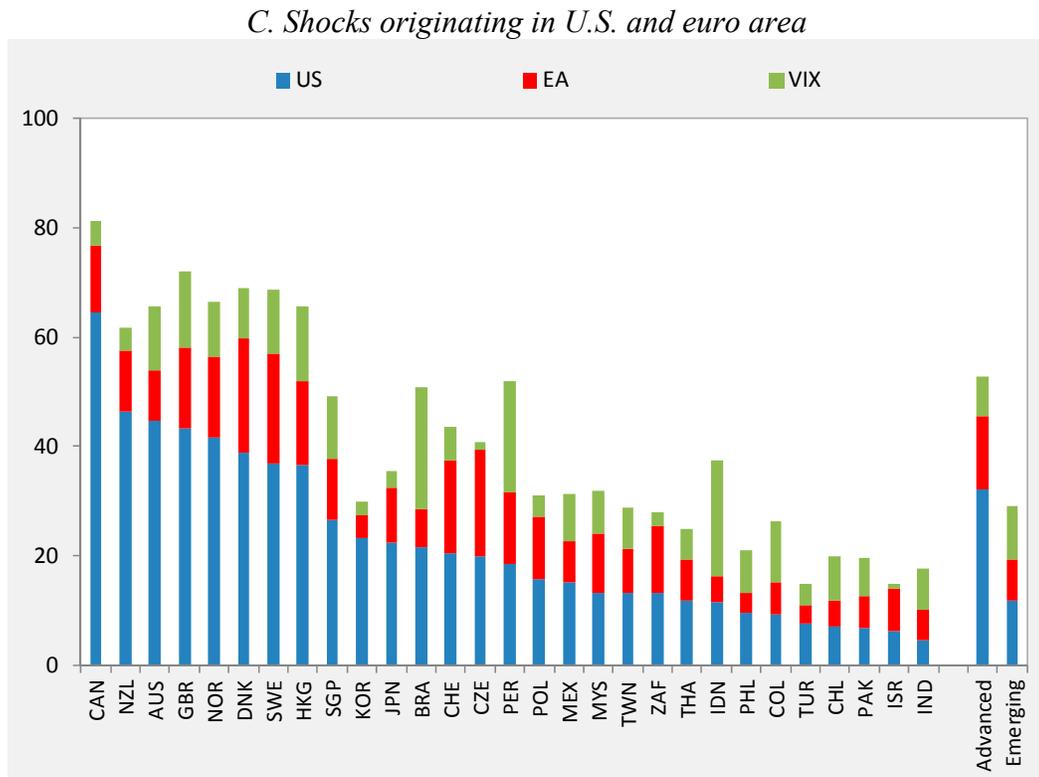
*B. Money versus real shocks*



Source: IMF staff calculations.

Note: The bars denote the variance decomposition (in percent) of domestic long-term interest rates attributable to expected (“Real”) and unexpected (“Money”) changes in 10-year U.S. bond yields (“US”) and in euro area 10-year bond yields (“EA”); and global risk sentiment, as proxied by the VIX index. The shock decomposition and identification are based on Osorio-Buitron and Vesperoni (2015).

Figure 2: Drivers of long-term interest rates, 2000-15: Expected and unexpected shocks to 10-year bond yields in the U.S. and the euro area (continued)



Source: IMF staff calculations.

Note: The bars denote the variance decomposition (in percent) of domestic long-term interest rates attributable to changes in 10-year U.S. bond yields (“US”) and in euro area 10-year bond yields (“EA”); and global risk sentiment, as proxied by the VIX index. The shock decomposition and identification are based on Osorio-Buitron and Vesperoni (2015).

To better understand the impact of anticipated versus unanticipated shocks to interest rates, we consider the decomposition of movements in U.S. and euro area 10-year bond yields constructed by Osorio-Buitron and Vesperoni (2015). Their identification strategy distinguishes between movements in rates that respond to global risk aversion, unexpected monetary tightening, or an improved economic outlook in the two economies. The identification of shocks is based on the sign restriction approach proposed by Matheson and Stavrev (2014) for U.S.-based shocks, and has been further extended by Osorio-Buitron and Vesperoni (2015) in two important ways. First, they orthogonalize their identified shocks to shifts in global risk aversion, proxied by the VIX index. Second, they expand the model and sign restrictions to include euro area yields—constructed as PPP-GDP weighted-average of 10 year bonds issued by France, Germany, Italy, and Spain—and stock returns. The sign restrictions are used to identify “real” shocks, which raise bond yields and depress stock prices, and “money” shocks, which depress both yields and stock prices. Additional sign restrictions distinguish the money and real shocks generated in the United States from those generated in the euro area, by assuming that contemporaneous shocks from the United States

can affect euro area variables but not the other way around. The shocks attributable to these different drivers of long-term yields are orthogonal by construction.<sup>7</sup> This decomposition will enable us to characterize the response of domestic interest rates to four sets of shocks that are in line with economic developments or that deviate from them, and that originate in the United States or in the euro area.

The identification of these shocks is distinct from the strategy employed by the high-frequency event window approach used by Karadi and Gertler (2015) and Gilchrist, López-Salido and Zakrajsek (2015), which measures the surprise element of FOMC decisions using intraday market data surrounding policy meetings. Still, “money” shocks are likely to be less anticipated by markets than their “real” counterparts, to the extent that the central bank and private agents likely share common information about the economic determinants that drive monetary policy decisions.

We conduct the same exercise described in the previous section, but replace the foreign interest rate  $\Delta i^*$  in the vector  $\mathbf{z}^*$  with the five identified drivers of global long-term rates, and include domestic long-term government bond yields in the domestic block. We focus on the 29 countries that showed a significant pass-through to domestic long-term interest rates at the 10 percent confidence level (as reported in Table 1, column 3). Figure 2, Panel A, reports the share of variation in domestic long-term interest rates that can be attributed to each of these drivers; Panel B distinguishes between “money” and “real” shocks; whereas Panel C distinguishes between shocks originating in the United States and the euro area.

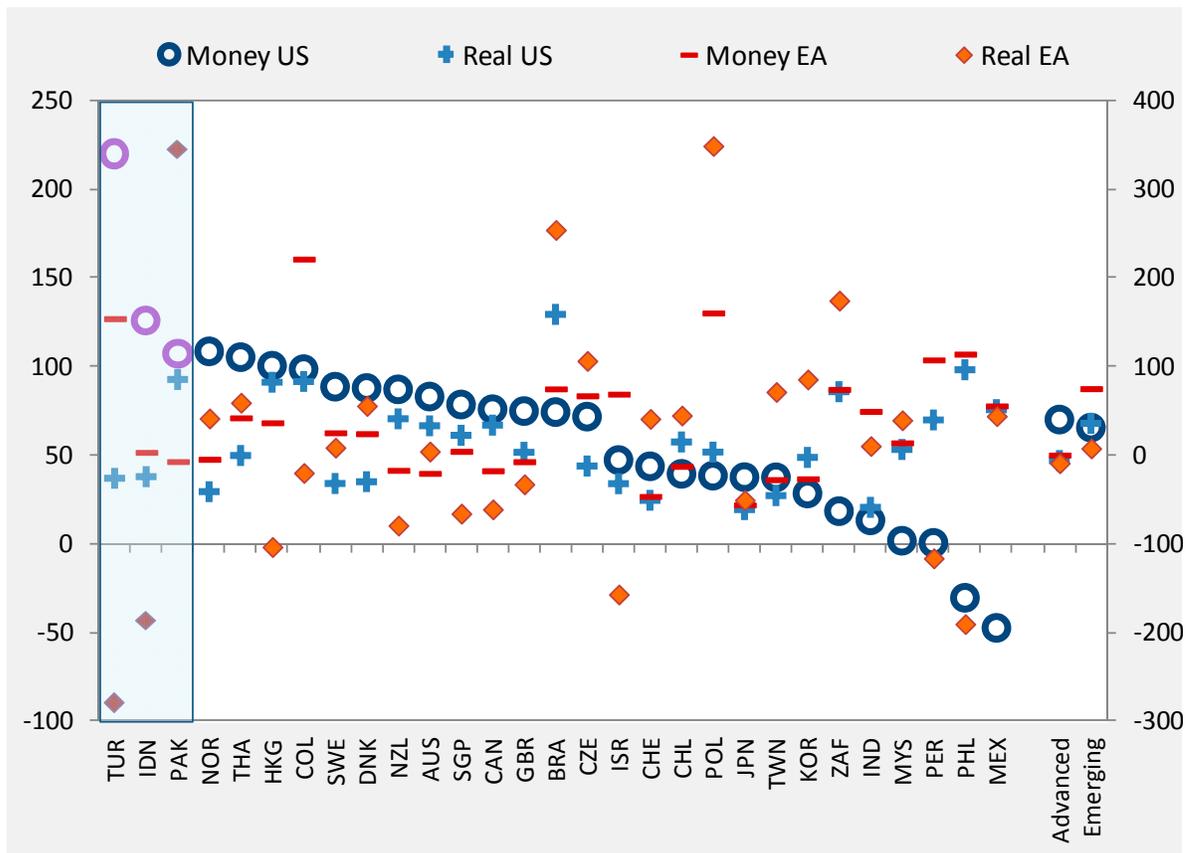
The first result that stands out is that money shocks significantly affect bond yields around the world (Panel B). Indeed, movements in global long-term rates that are driven by an unexpected monetary tightening (“money” shocks) explain a larger fraction of the variability in domestic rates than those driven by an improved economic outlook (“real” shocks). The contribution from monetary surprises among the external factors excluding the VIX is over 70 percent in advanced economies and 60 percent in emerging economies, and is larger than the contribution from “real” shocks in 80 percent of the countries.

Another important feature at the current juncture is that the United States is set to continue normalizing its monetary policy while other major economies, such as the euro area and Japan, maintain a highly accommodative stance. An interesting question in this context is how much of an attenuating effect this asynchronicity of monetary policies will provide for other economies.

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<sup>7</sup> Note however that while the identification strategy cannot distinguish between monetary policy shocks and inflationary surprises, our interest is in distinguishing expected interest rate movements associated with changes in the economic outlook.

Figure 3: Impact on long-term interest rates: Expected and unexpected shocks to 10-year bond yields in the U.S. and the euro area



Source: IMF staff calculations.

Notes: The chart shows the cumulative response of domestic long-term interest rates after 12 months to an identified shock attributable to expected (“Real”) and unexpected (“Money”) changes in 10-year U.S. bond yields (“US”) and in euro area 10-year bond yields (“EA”). The responses to the “Money US” and “Real EA” in the case of Turkey (TUR), Indonesia (IDN), and Pakistan (PAK) are shown on the right scale; all other responses are plotted against the left scale. The model also includes a shock to global risk sentiment, proxied by the VIX index; the corresponding cumulative response is not shown here owing to space considerations.

The results shown in Panel C suggest that the relief most countries may receive from further monetary easing by the ECB will be limited, as movements in U.S. rates are the main source of global financial shocks. Indeed, the share of total variation in domestic bond yields attributable to U.S. shocks is larger than the share corresponding to euro area shocks in most of the countries in our sample, and is particularly large for countries such as Australia, Canada, and South Korea. The few cases where shocks originating in both economies represent roughly the same contribution comprise the Czech Republic, India, Israel, Pakistan, and South Africa. The share of the variance explained by U.S.-originated shocks exceeds the share corresponding to euro area shocks in 93 percent of the countries. On average, they account for around 70 percent of the variability attributable to external factors (excluding the VIX) in advanced economies, and just over 60 percent in emerging economies.

It is worth noting that idiosyncratic factors still explain a large fraction of interest rate movements in emerging economies. For instance, in Canada, Denmark, Norway, Sweden and the United Kingdom the five external factors included in our model explain over two thirds of the overall variance in domestic long-term rates, but these factors explain less than 25 percent of the variance in the case of Chile, India, Pakistan, Philippines, Thailand, and Turkey.

This analysis of the relative importance of each underlying shock to foreign long-term rates is based on a variance decomposition exercise. The advantage of such an approach is that it takes into account not only the estimated impact of each shock when it occurs, but also how often and large those shocks have been in the past. A complementary exercise is to estimate the impact of an identified shock of a given magnitude, irrespective of how likely it is to occur: the impulse responses of domestic long-term interest rates to each shock are reported in Figure 3. A few interesting observations can be gleaned from this exercise. The responses of domestic interest rates to unanticipated U.S. monetary shocks tend to be large and positive for most countries, and tend to outweigh those that accompany good economic news. These are particularly large for Indonesia, Pakistan, and Turkey. In the case of the response to euro area shocks, Pakistan and Poland tend to exhibit the largest responses.

#### *Assessing the effects of shocks to the term premium*

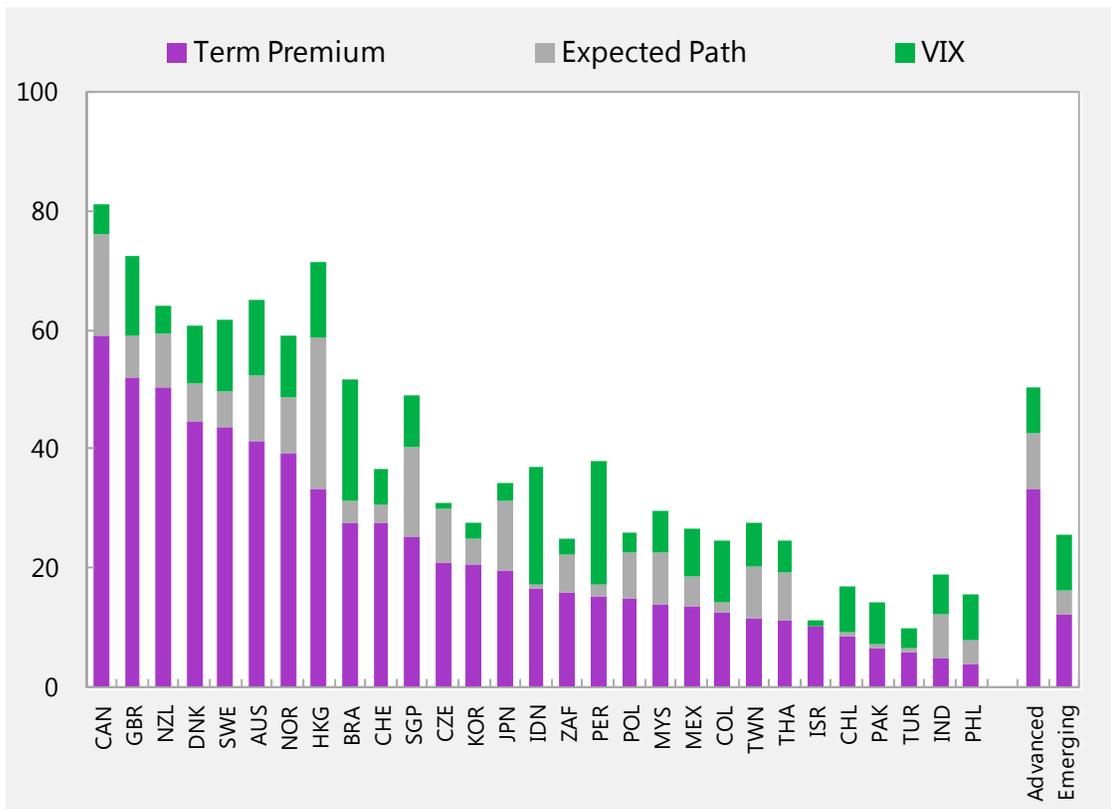
Another potential source of risk surrounding the normalization of U.S. monetary policy is a sudden decompression of the term premium—that is, the difference between the 10-year yield and the average of expected future short-term rates over the same horizon—which is currently at historically low levels.<sup>8</sup>

To assess the potential impact of a rise in the term premium, we include the decomposition of the 10-year U.S. Treasury bond yield into the expected path of short-term interest rates and the term premium as exogenous variables in our workhorse country-specific VAR models. The estimates of the U.S. term premium and the expected path of short-term interest rates are produced by Adrian, Crump, and Moench (2013) and maintained by the Federal Reserve Bank of New York, and have been reported in Figure 1, panel B. Results from this exercise are presented in Figures 4 and 5.

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<sup>8</sup> The term premium can be thought of as the extra return investors require to hold a longer-dated bond instead of investing in a series of short-term securities, and is thought to reflect their uncertainty about the future path of interest rates as well as their degree of risk aversion. As such, movements in the term premium tend to be closely correlated with risk premiums on other assets in global financial markets.

Figure 4: Drivers of long-term interest rates, 2000-15: U.S. term premium, the expected path of U.S. monetary policy, and global risk sentiment

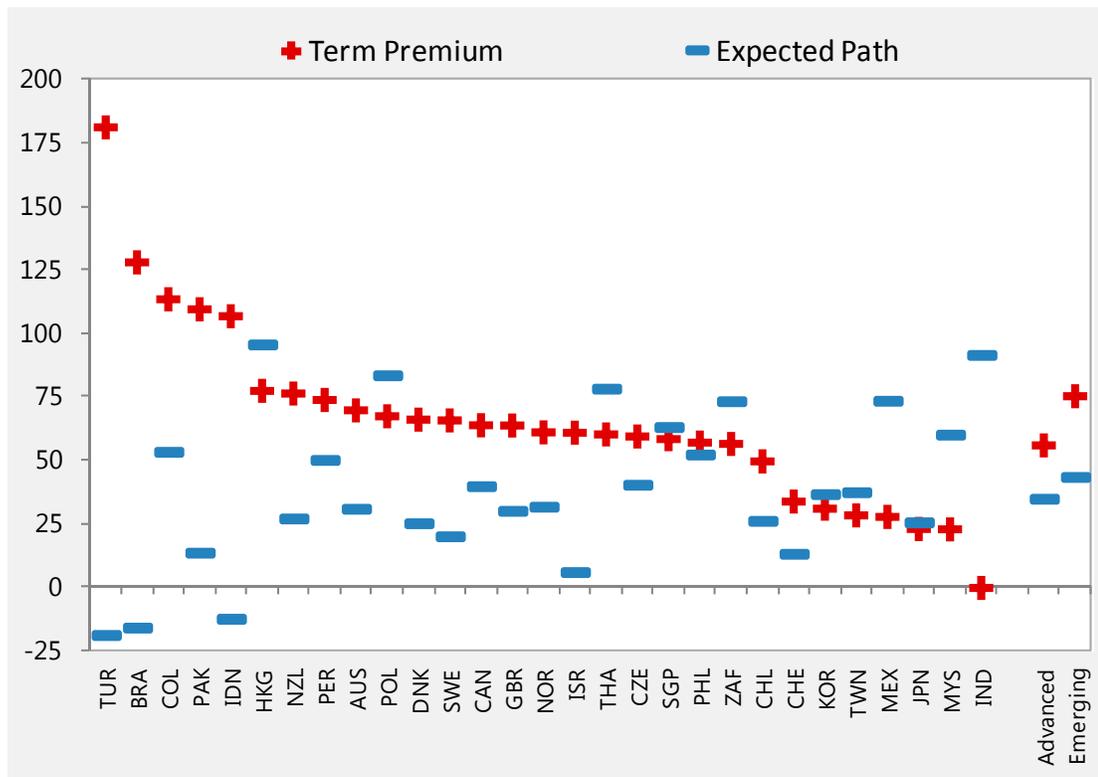


Source: IMF staff calculations.

Note: The bars denote the variance decomposition (in percent) of domestic long-term interest rates attributable to changes in the expected path of U.S. short-term interest rates (“Expected Path”), term premium, and global risk sentiment, as proxied by the VIX index. The Term Premium estimates are produced by Adrian, Crump and Moench (2013), and maintained by the Federal Reserve Bank of New York.

Our results confirm that movements in the term premium are a major source of spillovers from U.S. long-term interest rates. On average, we find that shocks to the U.S. term premium account for about three-quarters of the variance in domestic long-term rates attributable to shocks to U.S. rates (Figure 4), both in advanced and emerging economies. This share is as high as 90 percent for Chile, Indonesia, Israel, Pakistan, and Peru (although the contribution to the overall variance stemming from idiosyncratic factors tends to be large in these countries). In Canada, the same share is about 77 percent, but the U.S. term premium alone explains almost 60 percent of the entire variance of long-term Canadian bond yields. Finally, although these external factors explain a relatively small share of the variance in domestic interest rates in Turkey, the cumulative response of domestic interest rates to a shock to the U.S. term premium tends to be fairly large in that country (Figure 5).

Figure 5: Impact on long-term interest rates: U.S. term premium, the expected path of U.S. monetary policy, and global risk sentiment



Source: IMF staff calculations.

Note: The chart shows the cumulative response of domestic long-term interest rates after 12 months to an identified shock attributable to changes in the expected path of U.S. short-term interest rates (“Expected Path”), and in the term premium. The model also includes a shock to global risk sentiment, proxied by the VIX index; the corresponding cumulative response is not shown here owing to space considerations. The term premium estimates are produced by Adrian, Crump and Moench (2013), and maintained by the Federal Reserve Bank of New York.

## V. THE FED’S NORMALIZATION AND MONETARY AUTONOMY ELSEWHERE

While global financial conditions are expected to tighten as the Federal Reserve continues normalizing its monetary stance, many economies around the world are facing a deceleration in economic activity. Where weak demand is generating negative output gaps and inflation expectations remain well-anchored, supportive monetary conditions are likely to be appropriate. How much monetary leeway do these countries have to avoid an unwanted tightening of financial conditions? While interest rate pass-through at the long end of the yield curve appears widespread, can central banks tailor their short-term policy rates to domestic conditions? Or do they need to follow the interest rate path set by the Federal Reserve, even if that implies implementing pro-cyclical policy? Section III showed that there is a substantial pass-through of U.S. to domestic short-term interest rates in many countries. But is this co-movement of interest rates reflecting a lack of monetary autonomy? Or is it simply a natural consequence of central banks reacting to highly synchronized macroeconomic conditions?

### *From interest rate pass-through to autonomy-impairing spillovers*

To explore how the degree of synchronicity of interest rates across countries compares with that of macroeconomic conditions, we use a principal component analysis of output growth, price inflation, and interest rates for a large set of countries. For each of these series, we explore the correlation of individual country series with the global factor. The global factor (or component) of short-term and long-term interest rates, real output growth and CPI inflation corresponds to the principal component of the time-series of each variable across countries.<sup>9</sup>

Short-term interest rates from both advanced and emerging market economies exhibit a positive correlation with the global component in most countries (Figure 6). The co-movement over the past decade has been particularly strong among advanced economies, with an average correlation of about 0.9. Yet, a relatively high degree of co-movement with the global component is also observed for interest rates in emerging markets, with an average correlation of 0.7 for the financially integrated economies in Asia, Eastern Europe and Latin America.

This synchronicity of interest rates may simply reflect a high degree of co-movement in business cycles across countries. Indeed, all countries in our sample exhibit a positive correlation of real GDP growth with the corresponding global component (Figure 6). On average, countries that exhibit a high degree of synchronicity with the global factor in terms of interest rates also tend to show a high degree of co-movement in terms of output growth and price inflation.

It is often argued that the degree of co-movement in asset prices is increasing over time, driven by deeper integration of financial markets.<sup>10</sup> Indeed, the degree of co-movement of interest rates with respect to the corresponding global factor varies over time, and reached particularly high levels over recent years (Figure 7). However, these fluctuations tend to mimic the variations in synchronization of business cycles across countries.

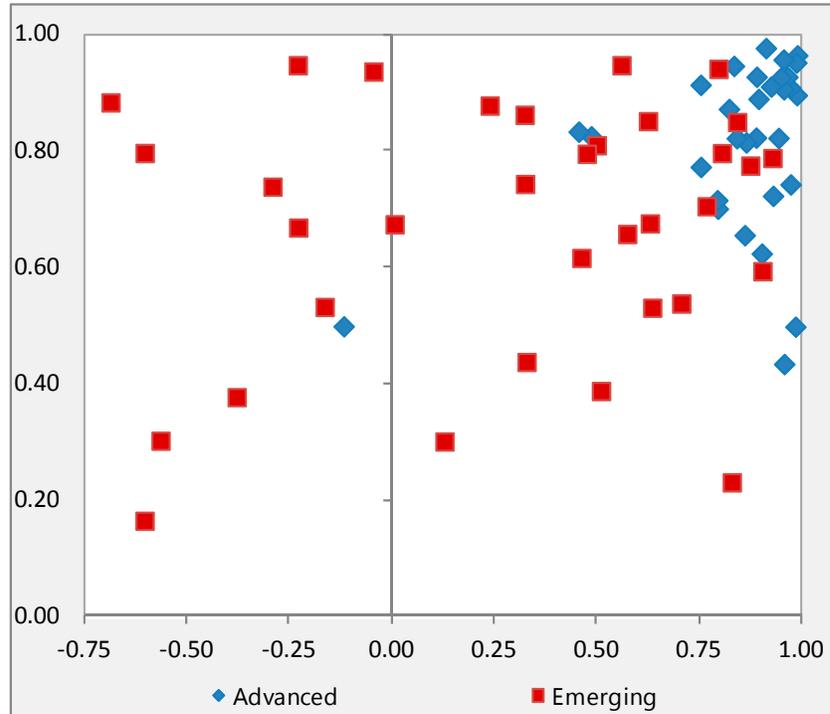
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<sup>9</sup> The principal component is the linear combination of those series that captures the maximum variance in the available data.

<sup>10</sup> See, for instance, Obstfeld, Shambaugh, and Taylor (2010) and Rey (2015).

Figure 6: Synchronicity of global output and interest rate cycles across countries.

*Correlation of real GDP growth (vertical axis) and short-term interest rates (horizontal axis), each against its corresponding "global component"*

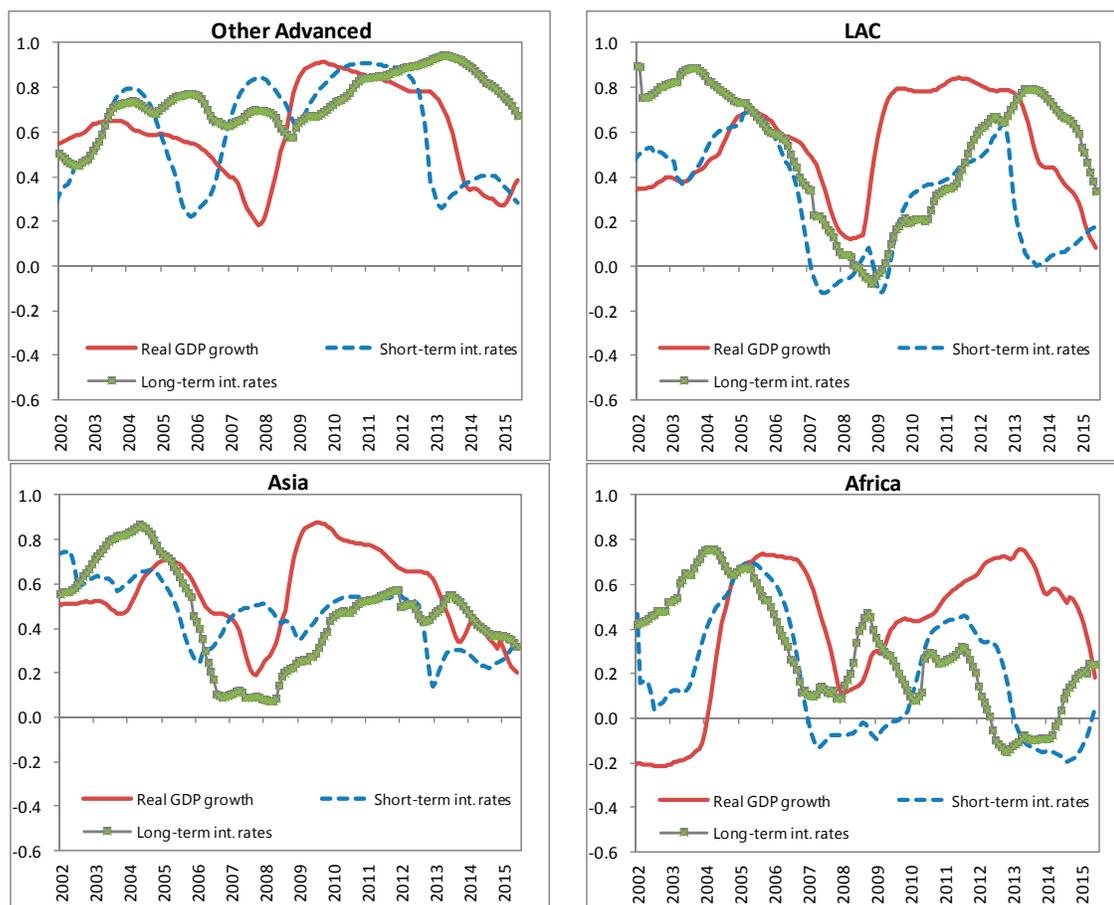


Sources: IMF staff calculations based on data from *International Financial Statistics*.

Note: Advanced economies include Australia, Austria, Belgium, Canada, Czech Republic, Denmark, Finland, France, Germany, Greece, Hong Kong SAR, Iceland, Ireland, Israel, Italy, Japan, Latvia, Malta, Netherlands, New Zealand, Norway, Portugal, Singapore, Slovenia, South Korea, Spain, Sweden, Switzerland, Taiwan POC, U.K., and U.S., Emerging market economies include Albania, Argentina, Armenia, Bangladesh, Bolivia, Brazil, Bulgaria, Chile, China, Colombia, Costa Rica, Croatia, Egypt, El Salvador, Honduras, Hungary, India, Indonesia, Kenya, Malaysia, Mexico, Pakistan, Paraguay, Peru, Philippines, Poland, Romania, Russia, Saudi Arabia, South Africa, Thailand, Turkey, Uruguay, and Vietnam. For each variable, the “global component”, derived from principal component analysis, is computed as the first principal component for all the economies in our sample.

There may be nothing inherently undesirable about domestic financial conditions being synchronized with those of international financial markets. For instance, countries with strong trade and financial linkages to the United States—such as Canada and Mexico—will tend to have an economic cycle that is highly synchronized with the U.S. cycle. In such cases, changes in domestic financial conditions may be broadly aligned with U.S. financial conditions, without posing challenges to achieving price and output stabilization objectives. A tension could emerge, however, in a case where domestic financial conditions are driven by foreign conditions that are out of sync with the domestic business cycle.

Figure 7: Evolution of correlation with global component



Sources: IMF staff calculations based on data from *International Financial Statistics*.

Note: “LAC” includes: Argentina, Bolivia, Brazil, Chile, Colombia, Costa Rica, Mexico, Peru, and Uruguay. “Asia” includes: Bangladesh, China, Hong Kong SAR, India, Indonesia, Malaysia, Philippines, Singapore, South Korea, Taiwan POC, Thailand, and Vietnam. “Africa”: Kenya and South Africa. “Other advanced” includes Australia, Canada, Denmark, Iceland, Israel, Japan, New Zealand, Norway, Sweden, Switzerland, United Kingdom, and United States. For each variable, the global component is computed as the first principal component for all the economies in our sample.

To distinguish between these cases, we follow the procedure proposed in Caceres, Carrière-Swallow, and Gruss (2016) for a sample of 43 countries since the early 2000s. *Autonomy-impairing spillovers* from U.S. interest rates are those movements in domestic short-term interest rates that are triggered by movements in U.S. rates but are unaligned with domestic monetary objectives. A significant spillover estimate can thus be interpreted as evidence that monetary policy in that country is constrained to some extent by foreign developments, and monetary autonomy limited.

In order to partial out the systematic policy response to changes in domestic macroeconomic conditions, a two-stage estimation of VAR models is used (see Annex A for more details). In the first stage, the procedure imposes the benchmark null hypothesis that the central bank

exclusively pursues the objectives of stabilizing domestic output and prices. More precisely, a Taylor-type rule for the dynamic relationship between domestic interest rates and domestic macro conditions is estimated for each country. Each country-specific VAR model includes changes in the nominal domestic short-term interest rate and a vector of variables capturing changes in domestic macroeconomic conditions.<sup>11</sup>

We include forward-looking variables rather than actual outcomes of output and inflation in the vector of domestic macroeconomic conditions. Many studies have used current and lagged values of consumer price inflation and output growth to control for domestic conditions (e.g. Edwards, 2015; Klein and Shambaugh, 2015; and Obstfeld, 2015). While these are natural choices to control for the domestic business cycle, monetary authorities operating under inflation targeting usually justify their monetary policy decisions based on changes in the economic outlook.<sup>12,13</sup>

Ideally, we would use the internal forecasts used by the central bank to inform the policy decision. However, these are only publicly available for a handful of countries and with a significant delay. Instead, we use changes in professional forecasts of economic activity and price inflation as reported by Consensus Economics (12-months-ahead forecasts of inflation and output growth).<sup>14</sup>

The residuals or unexplained components from this estimation can be interpreted as deviations from the historical policy reaction function that characterizes the central bank's efforts to achieve its domestic output and inflation stabilization objectives. These unexplained interest rate movements could reflect other central bank objectives beyond preserving price stability, including financial stability concerns, and thus could well be welfare-enhancing.<sup>15</sup> Nonetheless, they entail changes in domestic monetary conditions

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<sup>11</sup> We use interest rates on short-term government bonds (with maturities of about three months). Although these interest rates are not the monetary policy instrument, they should be closely linked to changes in the monetary policy stance. In fact, if changes in the policy instrument did not heavily influence these short-term interest rates in local currency, it would be hard to argue that the central bank can affect domestic monetary conditions at all

<sup>12</sup> Svensson (1997, 1999) argues that inflation targeting implies inflation *forecast* targeting, where the central bank's inflation forecast is an ideal intermediate target, even in the presence of output and/or interest rate stabilization concerns, and model uncertainty. There is also empirical evidence that central banks do react to changes in *expected* macro conditions rather than actual or lagged changes. For instance, Clarida, Gali, and Gertler (1998) show that the central banks of Germany, Japan and the United States adjust monetary policy rates in response to *anticipated* inflation, as opposed lagged inflation.

<sup>13</sup> It could also be argued that using actual or lagged variables can introduce additional biases in spillover estimates. For example, suppose a given external development is expected to affect aggregate demand both in the United States and in a small open economy sometime in the near future, but has not affected measured activity yet, and both economies adjust their monetary policy stand accordingly in order to achieve their objectives set exclusively in terms of domestic variables. In this context, using actual macro variables would lead to wrongly consider the change in interest rate in the small open economy as a monetary spillover from the United States when, in fact, the domestic authority is acting fully consistently with its policy objective.

<sup>14</sup> Besides being forward looking indicators, using expectations about GDP growth allows controlling for domestic conditions at a monthly frequency, which is not possible using GDP data.

<sup>15</sup> Consider the case of a central bank that decides to increase interest rates in the face of a shock that would otherwise lead to exchange rate depreciation. Our procedure identifies the part of the rate increase that can be

(continued...)

beyond what can be attributed to the central bank's usual response to inflation and output developments.

In a second stage, we quantify to what extent these residual movements in domestic interest rates can be explained by movements in U.S. interest rates. We estimate another country-specific VAR that includes the residual domestic interest rate movements obtained from the first stage, a vector of global variables including changes in the U.S. federal funds rate, and the VIX index to capture global risk sentiment.

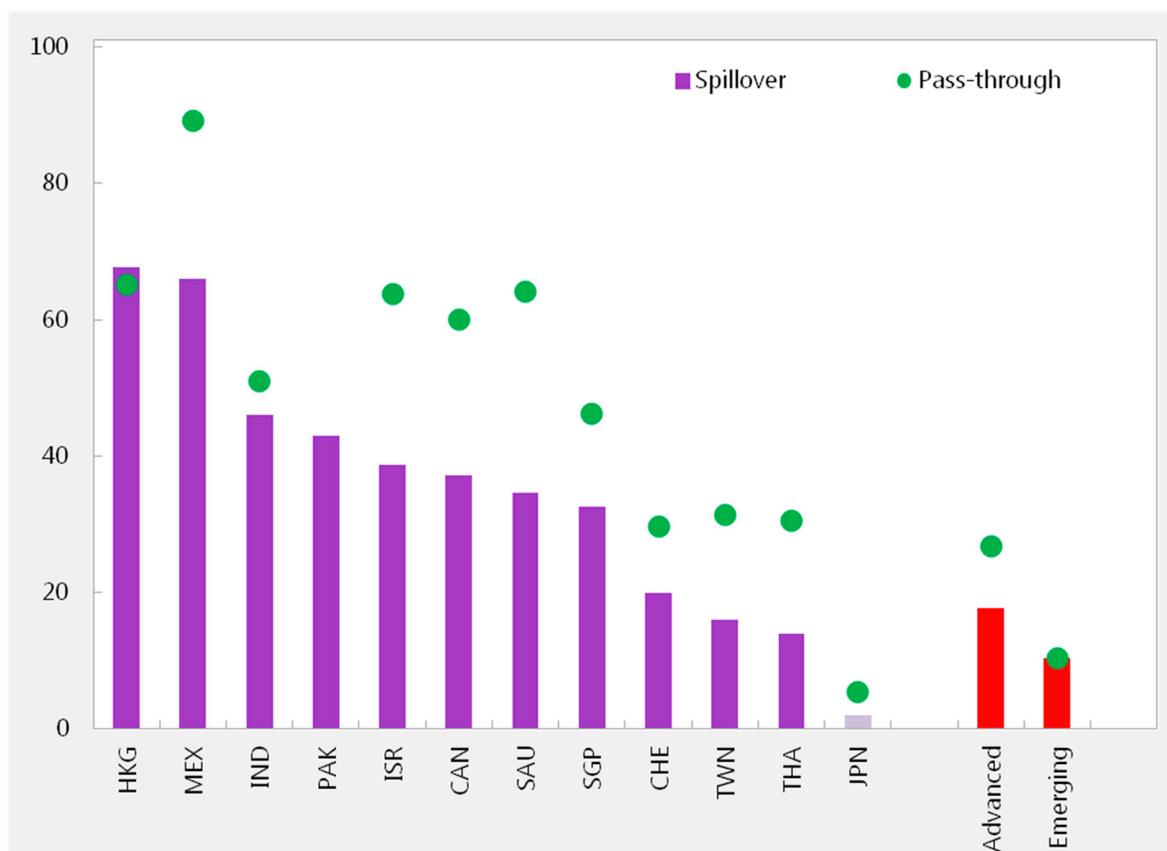
We focus our analysis on the Cholesky-orthogonalized impulse response of the (residual) domestic interest rates to a shock from the U.S. federal funds rate. We interpret these estimates as autonomy-impairing spillovers from U.S. rates, and expect them to be low where monetary autonomy is high. It should be noted that we are not attempting to fully characterize why interest rates deviated from the inward-looking policy rule in the first stage. While a host of factors could be driving these deviations from the policy rule, we focus only on how much of these residuals can be associated with movements in U.S. rates. In general, the autonomy-impairing spillover response of domestic short-term rates following a 100-basis-point increase in the federal funds rate (depicted with bars in Figure 8) is smaller than the overall response reported earlier (20 basis points lower on average). That is to say, an important portion of the co-movement in interest rates is simply a reflection of synchronized business cycles, and thus cannot be construed as inconsistent with full monetary autonomy.

Nonetheless, estimated spillovers to domestic short-term rates are statistically significant at the 10 percent confidence level in 11 out of the 43 advanced and emerging market economies included in our sample, where they average a non-trivial 38 basis points. Interestingly, these economies include countries with fully flexible exchange rates and well-established central banks, such as Canada and Israel. In economies such as Hong Kong SAR and Mexico, autonomy-impairing spillovers from U.S. to domestic short-term interest rates are large, exceeding two-for-three. This is not surprising given their highly open financial systems, compounded with a hard peg to the U.S. dollar in the former and tight financial linkages with the United States in the latter.

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explained by its concern for the second-round effects on inflation, as captured by its historical behavior. The remainder is considered unexplained, even though it could correspond to an explicit intent to contain vulnerabilities from balance sheet mismatches in order to preserve financial stability.

Figure 8: Response of short-term rates to an increase in the federal funds rate



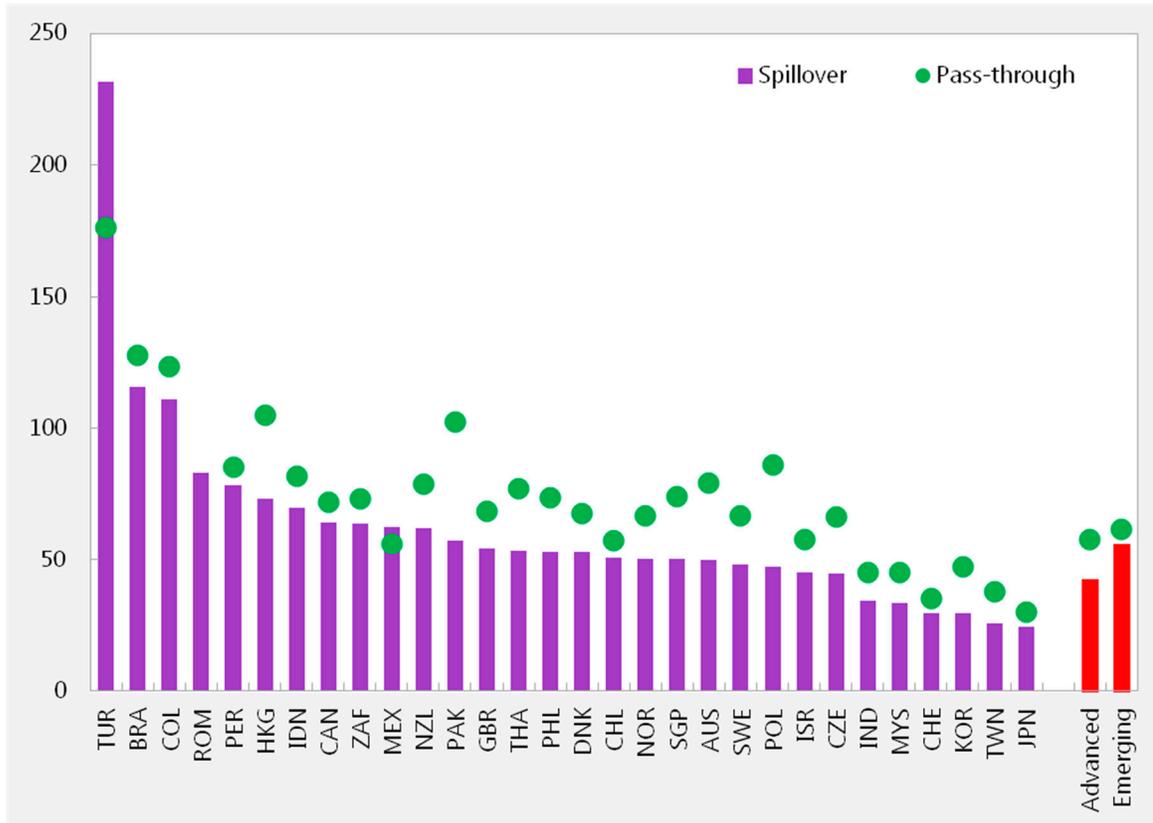
Source: IMF staff calculations.

Note: The chart shows the cumulative response of domestic short-term interest rates (green dots) after 12 months to a shock that increases the federal funds rate by 100 basis points after 12 months. “Spillover” (bars) denotes the estimates of autonomy-impairing spillovers. The dark bars and circles denote the responses that are statistically significant at the 10 percent confidence level.

For completeness, we also apply the two-stage estimation procedure to long-term rates. We find that the difference with respect to simple pass-through estimates is much less relevant at the longer-end of the yield curve. Indeed, the response of long-term interest rates from the two-stage procedure (bars in Figure 9) is essentially the same as the overall pass-through response (dots).

In sum, we find that a large portion of the pass-through to short-term interest rates from movements in U.S. rates can be attributed to the synchronicity of business cycles across countries. However, we also find that movements in U.S. rates generate significant spillovers to domestic short-term rates in several countries—including both advanced and emerging economies—above and beyond what can be explained by business cycle co-movement. Based on historical evidence, these countries seem to have limited monetary autonomy to cope with rising policy rates in the United States.

Figure 9: Response of long-term rates to an increase in the 10-year U.S. Treasury bond yield



Source: IMF staff calculations.

Note: The chart shows the cumulative response of domestic long-term interest rates (green dots) after 12 months to a shock that increases the 10-year U.S. Treasury bond yield by 100 basis points after 12 months. “Spillover” (bars) denotes the estimates of autonomy-impairing spillovers. The chart only shows the responses that are statistically significant at the 10 percent confidence level.

### ***Robustness exercise***

Our approach is subject to common empirical limitations. To estimate the policy response function, we should employ the internal forecasts used by the central bank to inform the policy decision, but these are only publicly available for a handful of countries and with a significant delay. The market forecasts that we use instead are subject to two limitations. First, there is a timing problem because they are not collected on the day as monetary policy decisions.<sup>16</sup> This could potentially bias our autonomy-impairing spillover estimates. For instance, an event that occurs between the forecast date and the policy decision and which affects rates in both countries could be (wrongly) considered a spillover response. However, we find that using alternative timings (that is, forecasts from the same month as the decision

<sup>16</sup> We use lagged market forecasts to ensure that they are predetermined with respect to policy decisions, but this reduces their information content.

or from the following month) does not significantly alter our results. For instance, in the case of Mexico the estimated spillover remains significant and in the order of 60 basis points in all cases. Second, even if timing were not an issue, market forecasts may incorporate expected policy responses.<sup>17</sup> In practice, however, monetary policy only affects economic conditions with a significant delay. Accordingly, movements in 12-months-ahead market forecasts should be highly correlated with movements in the central bank's internal forecasts.

Another potential problem with our estimates is that U.S. policy rates have remained unchanged at the zero lower bound since end-2008, which is a substantial part of our sample period. However, most other economies exhibited positive interest rates during this period, which enabled their central banks to increase or decrease their domestic policy rates while the federal funds rate remained constant. To assess whether this is affecting our spillover estimates, we repeat our procedure using a sample ending in June 2009. The results are broadly unchanged (Table B1 in Annex B), with the main difference being that the number of countries exhibiting a statistically significant spillover (at the 10 percent level) falls from 11 in our baseline estimations to only seven.

Finally, the multi-stage procedure used to quantify spillovers in a way that can be meaningful to infer limits to monetary autonomy relies on an estimate of the historical policy reaction function characterizing the central bank's pursuit of price and output stabilization. Of course, that function could vary over time as monetary policy frameworks change, thus affecting the spillover estimates. As a robustness exercise, we re-estimate dynamic policy rules using a 48-month rolling window in the first stage of our procedure. The results are reported in Table B1 in Annex B. Although the results remain broadly unchanged for the vast majority of countries, the estimated autonomy-impairing spillovers differ substantially in a few cases. For instance, countries such as Bolivia, Malaysia, and Romania now exhibit much larger (and statistically significant) spillover responses to a shock in the federal funds rate compared to our baseline estimations. Conversely, the spillovers in the case of Mexico appear to be much lower (and no longer statistically significant) when the parameters of its Taylor rule are allowed to vary over time. In the case of Canada, the spillover estimate drops by almost half but remains significant at the 10 percent confidence level.

## VI. EXPLORING THE DETERMINANTS OF MONETARY AUTONOMY

What determines the differences in autonomy-impairing spillovers across countries? The traditional trilemma framework points to the degree of exchange rate flexibility and capital account openness as the main determinants of monetary policy autonomy. More recently, Rey (2015) has questioned the dimensions of the trade-off, arguing that autonomy can only be achieved by restricting the capital account—although the arguments and evidence refer to the effect of global financial conditions on longer-term interest rates or credit aggregates, rather than on the central bank's ability to affect short-term rates.

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<sup>17</sup>Under this argument and if the central bank is fully credible, market forecasts might not move at all in response to a shock that would otherwise affect output growth and inflation because agents anticipate that the central bank will do whatever is necessary to neutralize the shock.

Table 2. Country sample for the interacted-panel VAR estimation

<i>Advanced economies</i>			
Australia	Israel	Norway	Switzerland
Canada	Japan	Singapore	United Kingdom
Czech Republic	Latvia	South Korea	
Denmark	New Zealand	Sweden	
<i>Emerging and developing economies</i>			
Bolivia	Costa Rica	Malaysia	Poland
Brazil	India	Nigeria	South Africa
Chile	Indonesia	Peru	Thailand
Colombia	Mexico	Philippines	Turkey

Caceres, Carrière-Swallow, and Gruss (2016) argue that inference about monetary autonomy based on spillover estimates can be biased when economic cycles are synchronized, possibly affecting conclusions about the role of exchange rate flexibility and other factors as determinants of autonomy. They estimate spillovers while modeling the domestic monetary policy reaction function for a panel of advanced and emerging economies, splitting the sample into countries with flexible exchange rates and countries with pegs or soft pegs. They find that spillovers are significantly larger for the latter. Moreover, in the case of countries with flexible exchange rates the response of domestic rates that cannot be accounted for by domestic developments is indistinguishable from zero.

In this section we explore the drivers of monetary autonomy beyond the role played by exchange rate flexibility. We start by assessing the role played by the pillars of the traditional trilemma, including the exchange rate regime and the degree of capital account openness. We then move beyond the trilemma to explore whether, for a given combination of exchange rate flexibility and financial openness, other country-specific factors also matter.

We propose a monthly panel-data approach similar to Hausman and Wongswan (2011) and Bowman, Londono, and Sapriza (2015). Differently from those studies, we interact country characteristics with realized changes in U.S. rates, rather than monetary surprises.

As in Caceres, Carrière-Swallow, and Gruss (2016), we employ the dynamic multi-stage approach described in section V and Annex A. In a first stage, we estimate country-specific policy rules for the dynamic relationship between domestic interest rates and domestic macro conditions, obtaining residual interest rate movements ( $\hat{u}^i$ ). In the second stage, we estimate a recursive Interacted-Panel VAR (IPVAR) model as described in Towbin and Weber (2013) to assess to what extent these residuals are driven by U.S. interest rates. The framework can be considered as a generalized panel VAR regression in which each right hand side variable can vary deterministically with country-specific characteristics. The model is given by:

$$\begin{bmatrix} 1 & 0 & 0 \\ b_0^{21} & 1 & 0 \\ \alpha_{0,ct}^{31} & \alpha_{0,ct}^{32} & 1 \end{bmatrix} \begin{bmatrix} VIX \\ \Delta i^* \\ \hat{u}^i \end{bmatrix}_{c,t} = \gamma' X_{c,t} + \sum_{j=1}^2 \begin{bmatrix} b_j^{11} & b_j^{12} & 0 \\ b_j^{21} & b_j^{22} & 0 \\ \alpha_{j,ct}^{31} & \alpha_{j,ct}^{32} & \alpha_{j,ct}^{33} \end{bmatrix} \begin{bmatrix} VIX \\ \Delta i^* \\ \hat{u}^i \end{bmatrix}_{c,t-j} + \epsilon_{c,t}, \quad (2)$$

where  $c = 1, \dots, N$  denotes countries;  $\mathbf{X}_{c,t}$  is a vector of controls that includes country-specific intercepts;  $\hat{u}^i$  are the residuals from the first-stage estimation; and  $\boldsymbol{\epsilon}_{c,t}$  is a vector of uncorrelated *iid* shocks. We identify external shocks with a small open economy assumption: the external variables do not depend on domestic variables ( $b_j^{13} = b_j^{23} = 0$ , for all  $j$ ), implying exogeneity.

In order to analyze how monetary spillovers vary with country characteristics, we allow for interactions terms. More precisely, the coefficients  $\alpha_{j,ct}^{pq}$  in (2) are given by:

$$\alpha_{j,ct}^{pq} = b_j^{pq} + \boldsymbol{\delta}_j^{pq'} \mathbf{F}_{c,t}, \quad (3)$$

where  $\mathbf{F}_{c,t}$  is a vector of country-specific fundamentals at time  $t$ .<sup>18,19</sup> This implies that domestic interest rates are modeled not only as a function of their own lags and the contemporaneous and lagged U.S. rate and VIX, but also of interactions between these terms and country fundamentals.

We use monthly data from January 2000 to October 2015 for a sample of 30 advanced and emerging market economies. See Table 2 for a list of the countries included in the IPVAR model. We then estimate model (2) and evaluate the coefficients in equation (3) at different values for the country characteristics to compute the corresponding impulse response functions.

### ***The trilemma's pillars: exchange rate flexibility and financial openness***

Before looking beyond the trilemma, we first assess the role played by its two pillars in our sample. More precisely, we first estimate the model with only two fundamentals in  $\mathbf{F}_{c,t}$  capturing the exchange rate regime and the degree of financial openness—the two pillars of the traditional monetary trilemma. The exchange rate regime corresponds to the coarse classification in Reinhart and Rogoff (2004) that has been updated by Ilzetzki, Reinhart, and Rogoff (2009). The financial openness index is from Chinn and Ito (2006) and Aizenman, Chinn, and Ito (2010).

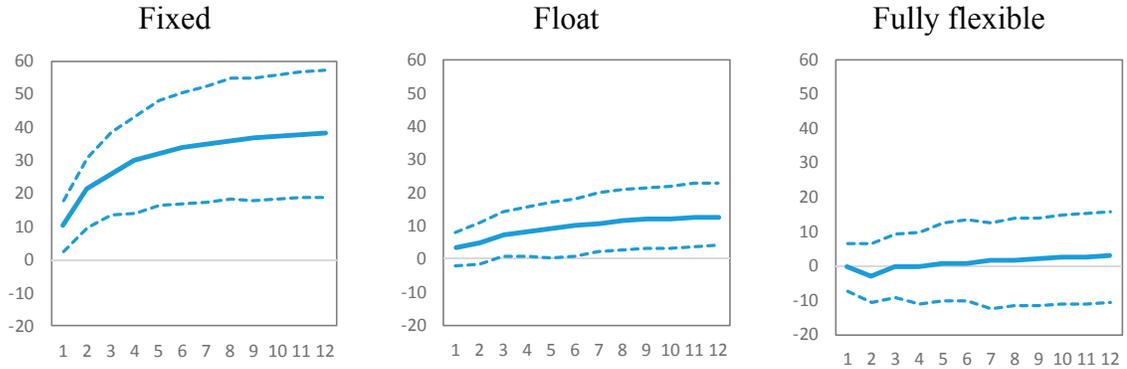
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<sup>18</sup> Note that while the coefficients for the domestic interest rate are allowed to vary with country characteristics, we restrict the dynamics of external variables to be independent of country characteristics (e.g.,  $\alpha_{1,ct}^{11} = b_1^{11}$ ).

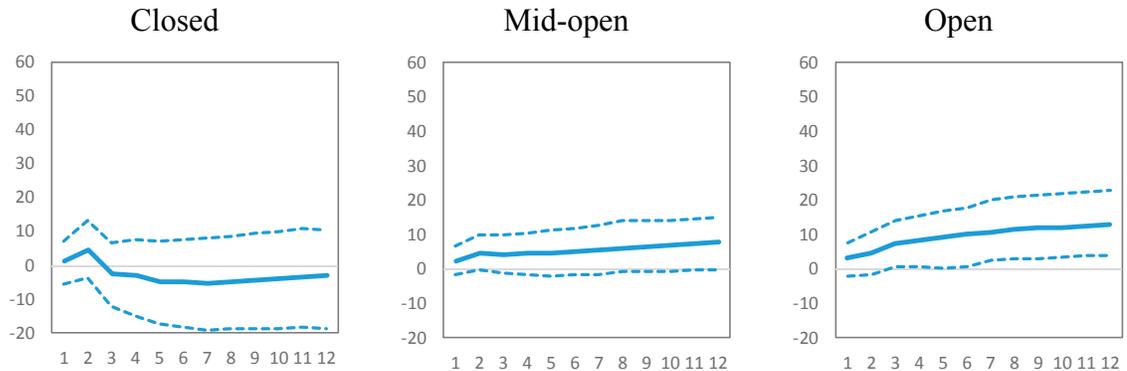
<sup>19</sup> Some of the fundamentals we use are available at annual frequency. In those cases we use linear interpolation to convert the data to monthly frequency.

Figure 10. Determinants of spillovers – Testing the trilemma’s hypothesis

*A. Autonomy-impairing spillovers under different exchange rate regimes, conditional on having high financial openness (basis points vs. months)*



*B. Autonomy-impairing spillovers under different degrees of financial openness, conditional on having a floating exchange rate regime (basis points vs. months)*



Source: IMF staff calculations.

Note: The charts show the cumulative autonomy-impairing spillover (as defined in the text) to a 100-basis cumulative increase in the U.S. federal funds rate. Panel A shows the response under a fixed exchange rate (1<sup>st</sup> decile of the distribution in our sample, corresponding to an index value of 1 in the Reinhart and Rogoff, 2004, course exchange rate classification), floating exchange rate (median in our sample, or index value 3) and fully flexible exchange rate (corresponding to the 9<sup>th</sup> decile in our sample and an index value of 4), while conditioning on high financial openness (the 9<sup>th</sup> decile of Aizenman, Chinn, Ito, 2010, index of financial openness in our sample). Panel B shows the response under a closed, mid-open, and open financial openness, corresponding to the 1<sup>st</sup> decile, the median, and the 9<sup>th</sup> decile of the Aizenman, Chinn, Ito (2010) index in our sample. The solid line reports the median response, conditional on the fundamental values. The dotted lines show a 90 percent confidence interval, calculated based on bootstrap techniques as described in Towbin and Weber (2013).

We find evidence that supports the predictions of the trilemma. Figure 10, panel A, shows monetary policy spillovers under different degrees of exchange rate flexibility, while conditioning on high financial openness. For open economies, maintaining a flexible exchange rate sharply reduces the degree of autonomy-impairing spillovers from U.S. to domestic interest rates.<sup>20</sup> The cumulative response after one year declines from almost 40 basis points under a fixed exchange rate to about 13 basis points under a floating exchange rate and disappears under a fully flexible regime (the median response is three basis points but is indistinguishable from zero at the 10 percent confidence level).

In turn, opening the capital account increases the degree of autonomy-impairing spillovers. Figure 10, panel B, shows the results for different degrees of financial openness while conditioning on having a floating exchange rate regime.<sup>21</sup> The cumulative response declines from 13 to 8 basis points when the degree of financial openness moves from the ninth decile in our sample (corresponding to fully open) to the median, and to -3 (but indistinguishable from zero) when it moves to the first decile.

Given that policy rates in the United States have been at the zero lower bound since end-2008, we conducted the same robustness exercise than in section V and re-estimated the model with data up to June 2009. The results in terms of the role played by the exchange rate and the degree of financial openness remain broadly unchanged (see Figure B1).

### ***Looking beyond the trilemma: do other factors matter?***

We then extend the model to account for a third fundamentals in  $F_{c,t}$  and explore how the response varies when we condition the third fundamental at different values, corresponding to the third and seventh decile of their empirical distribution in our sample, while conditioning on a floating exchange rate regime and high financial openness.<sup>22</sup>

We first explore how the strength of the monetary and fiscal frameworks may affect the extent of autonomy-impairing spillovers. Following a rise in U.S. rates, a less credible central bank may need to deliver a larger interest rate movement to convince agents that the exchange rate depreciation following an opening interest rate differential will not lead to significant second round effects in inflation. Meanwhile, countries with perceived fiscal vulnerabilities may be more susceptible to capital outflows after an increase in U.S. rates, prompting a larger increase in domestic rates.

To explore this, we construct an index of anchored inflation expectations based on the degree of disagreement among professional forecasters of inflation at a 12-month fixed horizon. While disagreement is a function of the variability of supply and demand shocks affecting the economy, Capistrán & Ramos-Francia (2010) argue that it also reflects the predictability and credibility of monetary policy—the more predictable a central bank’s reaction function, the

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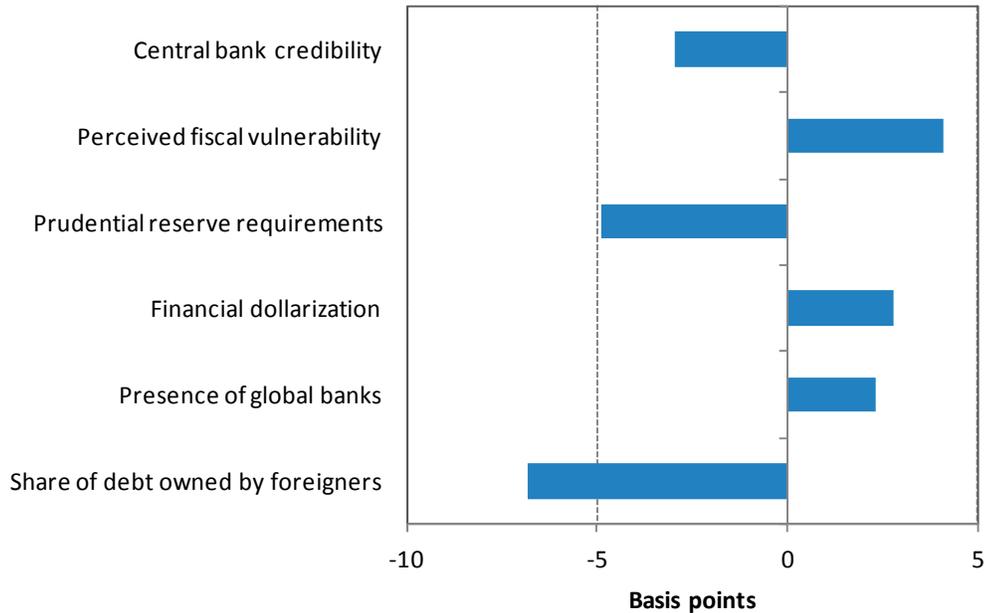
<sup>20</sup> This finding is in line with Obstfeld (2015), Klein and Shambaugh (2015), and Caceres, Carrière-Swallow, and Gruss (2016).

<sup>21</sup> This corresponds to having a value of three under Reinhart and Rogoff’s (2004) coarse classification index.

<sup>22</sup> The results in terms of exchange rate flexibility and financial openness remain valid when we add a third fundamental to the model.

Figure 11. Determinants of spillovers – Looking beyond the trilemma

*Sensitivity of monetary policy spillovers to selected fundamentals, conditional on floating exchange rate and open capital account.*



Source: IMF staff calculations.

Note: Each bar denotes the difference in monetary policy spillovers from an IPVAR model with three fundamentals when that fundamental moves from the 3<sup>rd</sup> to the 7<sup>th</sup> decile of its empirical distribution within our sample, conditioning on the exchange rate regime being floating (an index value of three in the Reinhart and Rogoff, 2004, course exchange rate classification) and the capital account fully open (the highest value in Aizenman, Chinn, Ito, 2010, index of financial openness). See the list of countries included in Table 2.

less likely are forecasters to disagree about the future path of inflation. Relatedly, inflation forecast disagreement has been shown to be related to *de jure* measures of central bank independence in G7 economies (Dovern, Fritsche, and Slacalek, 2012) and to the monetary policy regime in developing economies (Capistrán and Ramos-Francia, 2010).

The index of anchored inflation expectations  $AIE_{i,t}$  for country  $i$  at time  $t$  is constructed as an ordinal ranking of the inverse disagreement among forecasters (measured as the 4-year moving average of the standard deviation across inflation forecasts reported by Consensus Economics,  $MA48(\sigma_{i,t})$ ):

$$AIE_{i,t} = \frac{1}{N} \text{Rank} \left[ \frac{1}{MA48(\sigma_{i,t})} \right].$$

We use sovereign CDS spreads to capture perceived fiscal risks. We find that, conditional on the exchange rate regime and the degree of financial openness, moving the proxies for the

strength of the monetary and fiscal frameworks from the third to the seventh decile of their distribution in our sample leads, in each case, to a decrease in autonomy-impairing spillovers of close to five basis points (see Figure 11).<sup>23</sup>

The use of macroprudential policies has been often posited as an alternative tool that the central bank may use in contexts where adjusting policy rates to offset the effect of global financial conditions on capital flows is at odds with the output and price stability objectives. An interesting question then is whether the active use of macroprudential policies has helped countries attain increased monetary autonomy. One empirical challenge lies in finding measures of macroprudential policies that capture not only their usage, but also the intensity with which they are deployed. Here we use the index constructed by Cordella and others (2014) that is based on the frequency with which reserve requirements are adjusted. We find that, indeed, moving from the third to the seventh decile in terms of the intensity with which reserve requirements are used is associated with a reduction in autonomy-impairing spillovers of about five basis points.

Based on financial stability concerns related to currency mismatches in balance sheets, central banks in countries with a higher degree of financial dollarization may be more concerned about letting the exchange rate react to rising U.S. rates. We explore this by using an updated version of the financial dollarization index proposed in Levy-Yeyati (2006), which is based on the share of bank deposits denominated in foreign currency. We find that, conditional on the exchange rate regime and the degree of financial openness, increasing financial dollarization from the third to the seventh decile of its distribution in our sample is associated with an increase of about three basis points in spillovers associated with limited autonomy.<sup>24</sup>

The structure of the domestic financial system and, in particular, the presence of global banks may affect the way monetary policy responds to changes in global financial conditions. Goldberg (2013) finds some evidence that the presence of global banks may affect monetary autonomy, although the effects are heterogeneous—probably reflecting different business models of global banks—and relatively minor compared to those of the exchange rate regime. Here we use an analogous metric that captures the role of global banks in the provision of domestic credit, and find that a stronger presence of global banks (that is, moving from the third to the seventh decile of its distribution in our sample) is associated with a slightly larger spillover of about two basis points.

The share of sovereign debt in domestic currency that is held by foreigners has been increasing substantially over the last few years, especially in emerging market economies. In this context, portfolio rebalancing by international investors following a rise in U.S. rates can potentially have a larger impact on the capital account. Central banks may then need to raise

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<sup>23</sup> The result for CDS spreads is consistent with the findings in Bowman, Londono, and Sapriza (2015) regarding the response of long-term domestic interest rates to unconventional monetary shocks in the United States.

<sup>24</sup> While this difference may seem small, it should be noted that it corresponds to a rather limited reduction in the degree of dollarization, from 17 percent to six percent. Some countries in our sample have a much larger degree of dollarization (e.g., is about 60 percent in Peru).

policy rates in an attempt to attenuate outflow pressures, irrespective of domestic macro conditions. With this in mind, we assess whether the degree of autonomy-impairing spillovers varies depending on the share of sovereign debt held by foreigners, using the data constructed by Ebeke and Kyobe (2015). Surprisingly, we find the opposite effect in this sample: increasing foreigners' participation from the 3<sup>rd</sup> to the 7<sup>th</sup> deciles is associated with a reduction of autonomy-impairing spillovers of almost seven basis points.

Of course, many of these fundamentals are slow-moving variables, and changing them would require persistent policy action, along with a broader assessment of their welfare implications. All in all, the results from this section confirm that exchange rate flexibility plays the key role in allowing the central bank to gear monetary policy towards stabilizing the domestic economy. But the results suggest that the strength of policy frameworks is also important to attain larger monetary autonomy, although its gains are more modest compared to the exchange rate regime.

## VII. CONCLUSION AND POLICY IMPLICATIONS

Asset prices, and interest rates in particular, exhibit a large degree of co-movement across countries, and these are closely linked to changes in U.S. interest rates. On the basis of this observation, domestic interest rates around the world might be expected to rise as expansionary U.S. monetary policy is normalized. But our analysis suggests that this response will depend on several factors.

First, the nature of the factors driving U.S. interest rates during monetary policy normalization will determine the extent of its implications elsewhere. In general terms, global spillovers are expected to be larger if the pace of interest rate hikes is not commensurate with the recovery of economic indicators. Longer-term interest rates are expected to rise more sharply in other countries if the U.S. term premium were to decompress to more normal levels. Supportive monetary conditions in the euro area will provide some alleviation for certain countries, but interest rates around the world are typically more affected by U.S. rates.

Second, the response of global interest rates will depend on how much autonomy monetary authorities in other countries have to decouple their policies from U.S. monetary policy normalization. Based on historical evidence, we conclude that most countries have been able to tailor their monetary stance to domestic conditions, regardless of the path of U.S. policy rates. But this is not a generalized finding. In the past, spillovers from U.S. monetary policy have been large in a handful of countries.

Finally, our analysis suggests that the extent of monetary autonomy in small open economies vis-à-vis U.S. monetary policy depends crucially on the economic policy framework that is in place. Our results confirm that exchange rate flexibility plays a key role in ensuring greater monetary autonomy when the capital account is unrestricted. This suggests that countries with more flexible exchange rates will be better prepared to cope with the challenges posed by the Fed's normalization going forward.

Our results also suggest that other policies can also help preserve monetary autonomy from global financial conditions. For a given policy choice along the capital account openness and exchange rate flexibility dimensions, improving the credibility of policy frameworks, reducing the extent of financial dollarization, and using macroprudential reserve requirements may help achieve a higher degree of monetary autonomy. However, while some of these other factors are highly relevant for certain countries, they seem to deliver more modest returns compared to the exchange rate regime.

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## Annex A. Estimation of autonomy-impairing spillovers

The estimation of autonomy-impairing spillovers from U.S. interest rates follows the two-stage VAR approach in Caceres, Carrière-Swallow, and Gruss (2016). In the first stage we estimate a country-specific VAR(p) model including domestic variables only:

$$\begin{bmatrix} \Delta \mathbf{X} \\ \Delta i \end{bmatrix}_t = \mathbf{A}_0 + \sum_{j=1}^p \mathbf{A}_j \begin{bmatrix} \Delta \mathbf{X} \\ \Delta i \end{bmatrix}_{t-j} + \begin{bmatrix} \mathbf{e}^X \\ \mathbf{e}^i \end{bmatrix}_t \quad (\text{A1})$$

where  $i$  denotes the nominal domestic interest rate and  $\mathbf{X} = \{\pi, y\}$  are domestic macroeconomic conditions in the small economy.

The reduced-form innovations  $\hat{\mathbf{e}}^X$  and  $\hat{\mathbf{e}}^i$  are orthogonal to lagged values of  $\Delta \mathbf{X}$  and  $\Delta i$ , but they are likely to display substantial contemporaneous correlation. We then regress the innovations  $\hat{\mathbf{e}}^i$  on the residuals from the other equation,  $\hat{\mathbf{e}}^X$ :

$$\hat{\mathbf{e}}^i = \alpha + \beta' \hat{\mathbf{e}}^X + u_t^i. \quad (\text{A2})$$

The residuals  $\hat{u}_t^i$  from this regression are orthogonal to the reduced-form innovations to domestic economic conditions  $\hat{\mathbf{e}}^X$ , corresponding to a timing restriction whereby expectations about the domestic outlook are predetermined with respect to monetary policy.<sup>25</sup> These residuals can then be interpreted as deviations from the central bank's historical policy reaction function characterizing its pursuit of price and output stabilization.

In the second stage, we seek to quantify to what extent these residual movements in domestic interest rates can be explained by movements in U.S. interest rates. To do so, we estimate the following country-specific VAR(p) model:

$$\begin{bmatrix} \mathbf{z}^* \\ \hat{u}^i \end{bmatrix}_t = \mathbf{B}_0 + \sum_{j=1}^p \mathbf{B}_j \begin{bmatrix} \mathbf{z}^* \\ \hat{u}^i \end{bmatrix}_{t-j} + \begin{bmatrix} \mathbf{v}^* \\ \mathbf{v}^i \end{bmatrix}_t, \quad (\text{A3})$$

where vector  $\mathbf{z}_t^*$  is a vector of global variables, including changes in U.S. interest rates ( $\Delta i^*$ ). The matrices  $\mathbf{B}_j$  are restricted to ensure the block exogeneity of  $\mathbf{z}_t^*$ . This restriction assumes that global variables are not affected by lagged domestic variables.

Autonomy-impairing spillovers from U.S. interest rates are defined as the response of  $\hat{u}_t^i$  from a shock to  $\Delta i^*$ , with identification coming from a timing restriction imposed through Cholesky decomposition.

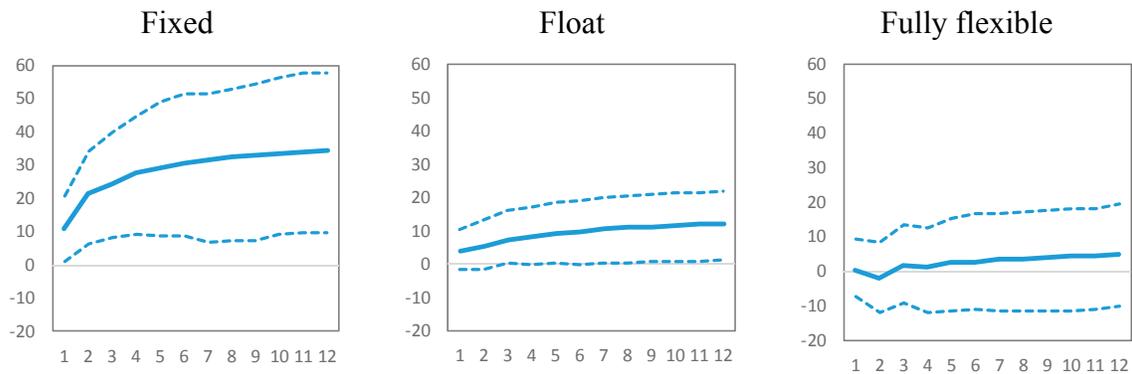
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<sup>25</sup> The timing restriction is the same that would be imposed through a Cholesky decomposition to obtain structural impulse response functions from monetary policy shocks.

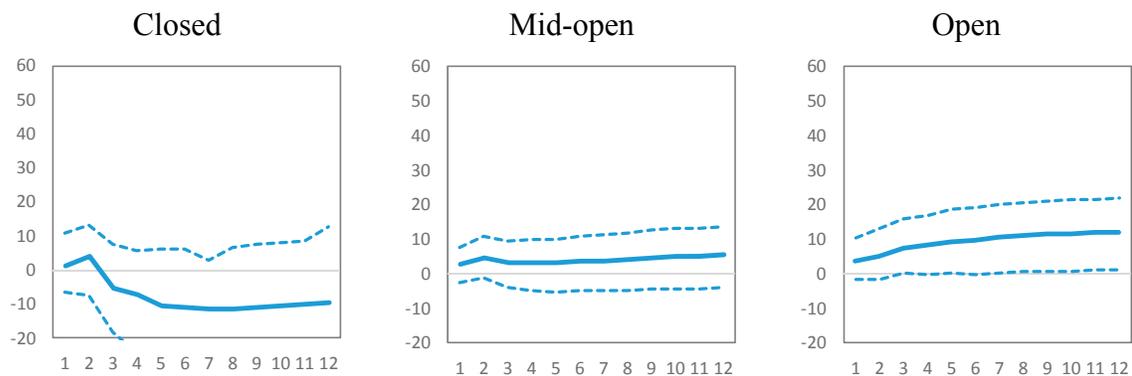
## Annex B. Robustness exercises and data description

Figure B1. Determinants of spillovers – Testing the trilemma’s hypothesis  
Pre zero lower bound sample (January 2000 to June 2009)

*A. Monetary policy spillovers under different exchange rate regimes, conditional on having high financial openness (basis points vs. months following shock)*



*B. Monetary policy spillovers under different degrees of financial openness, conditional having a floating exchange rate regime (basis points vs. months following shock)*



Source: IMF staff calculations.

Note: The charts show the cumulative monetary policy spillover (as defined in the text) to a 100-basis cumulative increase in the U.S. federal funds rate. Panel A shows the response under a fixed exchange rate (1<sup>st</sup> decile of the distribution in our sample, corresponding to an index value of 1 in the Reinhart and Rogoff, 2004, course exchange rate classification), floating exchange rate (median in our sample, or index value 3) and fully flexible exchange rate (corresponding to the 9<sup>th</sup> decile in our sample and an index value of 4), while conditioning on high financial openness (the 9<sup>th</sup> decile of Aizenman, Chinn, Ito 2010 index of financial openness in our sample). Panel B shows the response under a closed, mid-open, and open financial openness, corresponding to the 1<sup>st</sup> decile, the median, and the 9<sup>th</sup> decile of Aizenman, Chinn, Ito (2010) index in our sample. The solid line reports the median response, conditional on the fundamental values. The dotted lines denote the 90 percent confidence interval, calculated using the bootstrap technique described in Towbin and Weber (2013).

Table B1. Cumulative impulse response of domestic rates after 12 months; Robustness

Country	Spillover to short-term interest rates		
	Baseline	Rolling Taylor Rules	Pre-2009M6
ARG Argentina	0.28	-0.16	
BOL Bolivia	0.20	0.53 *	0.18
BRA Brazil	-0.57	-0.56	-0.65
CHL Chile	0.02	0.14	-0.21
COL Colombia	-0.32	-0.30	-0.39
CRI Costa Rica	0.13	-0.12	0.17
MEX Mexico	0.66 *	0.22	0.48
PER Peru	0.39	0.13	0.34
URY Uruguay	0.08	0.36	
ARM Armenia	-0.06	-0.14	
AUS Australia	0.07	0.05	0.05
CAN Canada	0.37 *	0.19 *	0.32 *
CHN China	0.14	0.13	0.06
HRV Croatia	0.18	-0.01	0.26
CZE Czech Republic	-0.03	-0.01	-0.01
DNK Denmark	0.15	0.16 *	0.10
EGY Egypt	0.30	0.32	0.35
HKG Hong Kong SAR	0.68 *	0.53 *	0.67 *
HUN Hungary	-0.19	-0.36	-0.40
IND India	0.46 *	0.45 *	0.32
IDN Indonesia	-0.35	-0.12	-0.35
ISR Israel	0.39 *	0.32 *	0.39 *
JPN Japan	0.02	0.01	0.01
LVA Latvia	0.05	-0.02	-0.11
MYS Malaysia	0.01	0.14 *	-0.05
NZL New Zealand	0.12	0.10	0.10
NOR Norway	0.11	0.08	0.04
PAK Pakistan	0.43 *	0.30	0.35
PHL Philippines	0.34	0.27	0.43
POL Poland	0.19	0.01	0.08
ROM Romania	0.29	1.23 *	0.33
RUS Russia	-0.41	-0.69	-0.32
SAU Saudi Arabia	0.35 *	0.30 *	0.31 *
SGP Singapore	0.33 *	0.23 *	0.31 *
ZAF South Africa	-0.04	-0.01	-0.05
KOR South Korea	0.12	0.01	0.07
SWE Sweden	0.04	0.06	-0.01
CHE Switzerland	0.20 *	0.14 *	0.14 *
TWN Taiwan	0.16 *	0.14 *	0.16 *
THA Thailand	0.14 *	0.08	0.12
TUR Turkey	0.19	0.43	
GBR United Kingdom	0.04	0.04	0.01
VNM Vietnam	-0.04	-0.02	0.32
<b>Median</b>			
Sample	0.14	0.10	0.10
Advanced	0.12	0.09	0.08
Emerging	0.14	0.13	0.17

Note: The table reports the cumulative impulse response of domestic rates after one year to a shock to the federal funds rate that leaves it 100 basis points higher. \* denotes statistical significance at the 10 percent level, and \* denotes statistical significance at the 10 percent level.

<b>Country</b>	<b>Short-term government bond yield in local currency</b>	<b>Long-term government bond yield in local currency</b>
Argentina	March 2002 – June 2015: BCRA 6-month Treasury auction yields in new pesos (GFD). Interpolated using BCRA 1-year treasury auction yields in new pesos (GFD) and 2-year treasury auction yields in new pesos (GFD).	February 2012 – July 2015: 25-year government bond yield in new pesos (GFD).
Armenia	January 2000 – October 2015: 1-year treasury bill rate (IMF MBRF2 line 91160C..XI...). Interpolated using 182-day treasury bill yield (Central Bank of Armenia via Haver Analytics).	March 2000 – October 2015: Yield on Long-term government bonds (Central Bank of Armenia via Haver Analytics). Interpolated using information from 5-year government bond yield in dram (GFD).
Australia	January 2000 – November 2015: 13-week treasury bills (IFS line 19360C..ZF...). Interpolated and spliced using 3-month generic government bond yield (Bloomberg ticker GACGB3M).	January 2000 – November 2015: 10-year generic government bond yield (Bloomberg ticker GACGB10).
Bolivia	January 2000 – November 2015: Treasury bill rate (IFS line 21860C..ZF...).	
Brazil	January 2000 – November 2015: Treasury bill rate (IFS line 22360C..ZF...). Interpolated using Anbima 6-month government bond fixed (Bloomberg ticker BZAD6M) and 6-month generic government bond yield (Bloomberg ticker GEBR06M). Spliced using Anbima 3-month government bond fixed (Bloomberg ticker BZAF3M).	January 2007 – November 2015: 10-year generic government bond yield (Bloomberg ticker GEBR10Y). Interpolated using information from 5-year note yield in real (GFD) and from 5-year generic government bond yield (Bloomberg ticker GEBR5Y).
Bulgaria	January 2000 – November 2015: 3-month treasury bill yield (GFD). Spliced using base interest rate (Bulgarian National Bank via Haver Analytics).	January 2000 – July 2015: 10-year government bond yield (Haver). Spliced using information from 10-year government bond in new lev (GFD).
Canada	January 2000 – November 2015: Treasury bill rate (IFS line 15660C..ZF...). Interpolated using 6-month government bond yield (Bloomberg GCAN6M). Spliced using 3-month government bond yield (Bloomberg GCAN3M).	January 2000 – November 2015: Canadian government bonds 10-year note (Bloomberg ticker GCAN10YR).
Chile	January 2000 – November 2015: 3-month interest rate (OECD MEI series 228.IR3TIB01.ST). Interpolated using 1-year government bond yield in pesos (GFD) and 1-year generic government bond yield (Bloomberg ticker CLGB1Y).	July 2004 – November 2015: Yield on 10-year government bonds (OECD Main Economic Indicators series 228.IRLTLT01.ST). Interpolated using information from General Government 10-year Bond (Bloomberg ticker CLGB10Y) and spliced using information from international rate on 5-year Central Bank of Chile BCP paper (Banco Central de Chile via Haver Analytics).

China	January 2000 – October 2015: 3-month treasury bond trading rate proxy (OECD MEI series 924.IR3TIB01.ST). Spliced using 3-month repo on treasury bills in renminbi (GFD) and prime lending rate (People’s Bank of China via Haver Analytics).	January 2007 – July 2015: Yield on 10-year Government Bond in renminbi (GFD). Missing months have been completed using linear interpolation.
Colombia	January 2000 – November 2015: 1-year treasury notes (IMF MBRF2 23360C.ZB...). Interpolated using 3-month treasury bill yield in pesos (GFD). Spliced using 1-year generic government bond yield (Bloomberg ticker COGR1Y).	March 2001 – November 2015: 10-year government benchmark bonds (OECD Main Economic Indicators series 233.IRLTLT01.ST). Interpolated using information from 10-year generic government bonds (Bloomberg ticker COGR10Y) and 10-year government bond yield (Reuters via Haver Analytics). Spliced using information from 15-year generic government bonds (Bloomberg ticker COGR15Y).
Costa Rica	January 2000 – October 2015: 6-month treasury bill yield in colones (GFD). Interpolated using 12-month treasury bill yield in colones (GFD). Spliced using 1-3 year government bond yields in colones (GFD).	July 2003 – January 2015: 3-7 year government bond yield (GFD). Missing months have been interpolated for up to 3 consecutive months. Gap remains from February to October 2004 and from August 2008 to February 2009.
Croatia	January 2000 – October 2015: 3-month treasury bill yield in kuna (GFD). Interpolated using information from 6-month treasury bill yield in kuna (GFD) and from 1-year government bond yield in kuna (GFD). Spliced using central bank discount rate and Lombard rate (Croatian National Bank via Haver Analytics).	January 2000 – July 2015: 5-year government bond yield in Kuna (GFD). Interpolated using 10-year government bond yield in Kuna (GFD).
Czech Republic	January 2000 – November 2015: Treasury bill rate (IFS line 93560C.ZF...). Interpolated using 1-year government bond yield (Bloomberg CZGB1YR), 1-year government bond yield in koruna (GFD), and 3-year government bond yield (CZGB3YR).	January 2000 – November 2015: Government bond yield (IFS line 93561...ZF...). Spliced using information from 10-year generic government bond (Bloomberg ticker CZGB10YR), 10-year government bond yield (Reuters via Haver Analytics), and 5-year generic government bond (Bloomberg ticker CZGB5YR).
Denmark	January 2000 – November 2015: 3-month treasury bill yield (Bloomberg ticker GDGT3M). Interpolated using 6-month treasury bill yield (Bloomberg ticker GDGT6M) and 2 year government bond yields (Bloomberg ticker GDGB2YR).	January 2000 – November 2015: 10-year government bond yield on secondary market (IMF MBRF2 line 12861...XI...). Interpolated using 10-year government bond yield (Bloomberg ticker GDGB10YR) and 10-year central government bond yield (Haver Analytics).
Egypt	January 2000 – October 2015: Treasury bill rate (IMF MBRF2 line 46960C.ZI...). Interpolated using 3-month treasury bill yield in pounds (GFD).	April 2012 – July 2015: 10-year government bond yield in Pounds (GFD). Interpolated using information from 7-year and 5-year government bond yields in Pounds (GFD).
Guatemala	January 2005 – October 2015: Central bank policy rate (IMF MBRF2).	

Hong Kong SAR	January 2000 – November 2015: Treasury bill rate (IFS MBTS line 53260C..ZI...). Interpolated using 6-month generic bond yield (Bloomberg ticker HKGG6M).	January 2000 – November 2015: 10-year exchange fund notes (Hong Kong Monetary Authority via Haver Analytics). Interpolated using information from 10-year generic bond yields (Bloomberg ticker HKGG10Y).
Hungary	January 2000 – October 2015: 3-month treasury bill yield in forint (GFD). Interpolated using 1-year government bond yield in forint (GFD). Spliced using the base interest rate (National Bank of Hungary via Haver Analytics).	January 2000 – October 2015: Yield on 10-year government debt securities (Haver Analytics).
India	January 2000 – November 2015: 3-month treasury bill yield (Bloomberg ticker IYTB3M). Spliced using 3-month treasury bill yield in rupee (GFD).	January 2000 – November 2015: 10-year government bond yield (Reserve Bank of India via Haver Analytics). Spliced using information from generic 10-year government bond yield (Bloomberg ticker GIND10YR).
Indonesia	January 2000 – November 2011: 6-month sovereign zero-coupon bond yield (Bloomberg ticker I26606M). Spliced using treasury bill yield in rupiah (GFD).	January 2003 – November 2015: 10-year generic government bond (Bloomberg ticker GIDN10YR). Spliced using information from 8-year generic government bond (Bloomberg ticker GIDN8YR).
Israel	January 2000 – November 2015: Treasury bill yield (IFS line 43660C..ZF...). Spliced using 2-year generic government bond yield (Bloomberg ticker GISR2YR).	January 2000 – November 2015: 10-year government bonds (OECD Main Economic Indicators series 436.IRLTLT01.ST). Interpolated using information from 10-year generic government bond yield (Bloomberg ticker GISR10YR) and 10-year government bond yield (Reuters via Haver Analytics).
Japan	January 2000 – November 2015: 6-month treasury discount bill yield (Bloomberg ticker GJTB6MO). Interpolated using 3-month treasury discount bill yield (Bloomberg ticker GJTB3MO). Spliced using Financing bill rate (IFS line 15860C..ZF...).	January 2000 – November 2015: 10-year generic government bond yield (Bloomberg ticker GJGB10).
Latvia	January 2000 – October 2015: Treasury bill rate (IFS line 94160C..ZF...). Spliced using 6-month treasury bill yield in euro (GFD) and central bank policy rate (IFS).	January 2000 – September 2015: Government bond yield (IFS line 94161...ZF...). Sliced using information from 10-year government benchmark bonds (OECD MEI series 941.IRLTLT01.ST) and 10-year government bond yield in Euro (GFD). Missing from March to October 2000.
Malaysia	January 2000 – November 2015: 3-month treasury bill yield (IFS line 54860C..ZF...). Spliced using 1-year Bank Negara Malaysia Treasury bill yield (Bloomberg ticker MGIYBD10).	January 2000 – November 2015: 10-year government bond (IMF MBRF2 lines 54861E..ZB...). Interpolated using information from 10-year Bank Negara Malaysia generic bond (Bloomberg ticker MGIY10Y).

Mexico	January 2000 – November 2015: CETES 90-day yield (MBRF2 line 27360C..ZI...). Interpolated using 3-month treasury bill yield (Bloomberg ticker MPTBC).	January 2000 – November 2015: Government bond yield (IFS line 27361...ZF...). Interpolated using information from 10-year generic government bond yield (Bloomberg ticker GMXN10YR).
New Zealand	January 2000 – November 2015: 3-month treasury bill new issue rate (IFS line 19660C..ZF...). Interpolated using 6-month treasury bill yield (Bloomberg ticker NDTB6M).	January 2000 – November 2015: 10-year government bond yield (Bloomberg ticker GNZGB10).
Nigeria	January 2000 – November 2015: Treasury bill rate (IFS line 69460C..ZF...). Interpolated using 91 day treasury bill yield (Central Bank of Nigeria via Haver Analytics). Spliced using monetary policy rate (Haver Analytics).	July 2007 – October 2015: 10-year treasury bond yield (Central Bank of Nigeria via Haver Analytics). Interpolated using information from 20-year treasury bond yield (Central Bank of Nigeria via Haver Analytics).
Norway	January 2000 – November 2015: 6-month government treasury bill yield (Bloomberg ticker GNGT6M).	January 2000 – November 2015: 10-year government bond yield (Norges Bank via Haver Analytics). Interpolated using information from 10-year government bond yield (Bloomberg ticker GNOR10YR).
Pakistan	January 2000 – October 2015: 6-month government treasury bill rate (IFS line 56460C..ZF...).	December 2000 – October 2015: 10-year government bond yield (Reuters via Haver Analytics). Spliced using information from 10-year government bond yield in Rupee (GFD), and interpolated using information from 5-year government bond yield in rupee (GFD).
Peru	January 2000 – November 2015: 6-month generic government bond yield (Bloomberg ticker GRPE6M). Interpolated using 3-month zero coupon bond yield (Bloomberg ticker I36103M). Spliced using central bank discount rate in new sol (GFD).	May 2006 – November 2015: 10-year generic government bond yield (Bloomberg ticker GRPE10Y). Spliced using 10-year zero-coupon curve (Bloomberg ticker I36110Y) and 10-year sovereign bond yield (Ministerio de Economía del Perú via Haver Analytics). Interpolated using information from 15-year generic government bond yield (Bloomberg ticker GRPE15Y).
Philippines	January 2000 – October 2015: 91-day treasury bill rate (IFS line 56660C..ZF...). Interpolated using PDEX PDST-F Fixing 3-months (Bloomberg ticker PDSF3MO).	January 2000 – October 2015: PDEX PDST-F Fixing 10-year (Bloomberg ticker PDSF10YR). Spliced using 10-year treasury bond mid-yield (Tullett Prebo via Haver Analytics).
Poland	January 2000 – November 2015: Treasury bill rate (IFS line 96460C..ZF...). Interpolated using 1-year government note yield in new zloty (GFD) and 1-year government note yield (Bloomberg ticker POGB1YR).	January 2000 – November 2015: 10-year government bond yield (Reuters via Haver Analytics). Interpolated using variation from 10-year government note (Bloomberg ticker POGB10YR).

Romania	January 2000 – September 2015: 91-day treasury bill rate (IFS line 96860C..ZF...). Spliced using 3-month treasury bill yield in new leu (GFD).	December 2001 - October 2015: Government bond yield (IFS line 96861...ZF...). Interpolated using information from 10-year bid yield on government securities (National Bank of Romania via Haver Analytics) and spliced using long-term government bond yield in new leu (GFD). Missing from September 2003 to June 2004.
Russia	January 2000 – October 2015: 3-month treasury bill yield in ruble (GFD). Interpolated using 1-year government bond yield in ruble (GFD) and 6-month government bond yield in ruble (GFD). Spliced using 1-week repo OMO auction rate (Haver Analytics).	January 2000 – July 2015: 10-year bond yield in ruble (GFD). Missing from April 2011 to February 2012.
Saudi Arabia	January 2000 – October 2015: 13-week treasury bill rate (Saudi Arabian Monetary Agency via Haver Analytics). Spliced using 12-month treasury bill yield in riyal (GFD), 26-week treasury bill rate (IMF MBTS line 45660CC.ZN...), and reverse repo rate (IMF MBTS).	January 2000 – October 2008: 5-year government note yield (GFD). Spliced using information from 10-year government bond yield (GFD). Missing from August to December 2006.
Singapore	January 2000 – November 2015: 3-month treasury bill yield (IFS line 57660C..ZF...). Interpolated using Monetary Authority of Singapore paper 3-month yield (Bloomberg ticker MASB3M) and 6-month yield (Bloomberg ticker MASB6M).	January 2000 – November 2015: Monetary Authority of Singapore 10-year bond yield (Bloomberg ticker MASB10Y).
Slovenia	January 2000 – July 2015: 3-month treasury bill yield rate (Ministry of Finance via Haver Analytics). Spliced using 3-month treasury bill yield in tolar and euro (GFD) and 2-year government bond yield in euro (GFD).	March 2002 – July 2015: 10-year government bond yield in tolar and euro (GFD). Missing from December 2002 to September 2003.
South Africa	January 2000 – November 2015: 91-day treasury bill tender rate (South African Reserve Bank via Haver Analytics). Interpolated using same concept from alternative sources (MBRF2 line 19960C..ZI... and OECD MEI series 199.IR3TIB01.ST). Spliced using 2-year government bond yield (Bloomberg ticker GSAB2YR).	January 2000 – November 2015: 5 to 10-year government bond yield (South African Reserve Bank via Haver Analytics). Interpolated using 10-year government bond yield (Bloomberg ticker GSAB10YR).

South Korea	January 2000 – November 2015: KCMP treasury bond yield (Bloomberg ticker GVSK3MON). Spliced using 3-year government bond yield in won (GFD).	January 2000 – November 2015: 10-year KCMP treasury bond yield (Bloomberg ticker GVSK10YR). Spliced using information from 10-year government bond yields (OECD Main Economic Indicators 542.IRLTLT01.ST), 5-year government bond yield in Won (GFD), and yield on national housing bonds 1 & 2 (IFS line 54261...ZF...).
Sweden	January 2000 – November 2015: 3-month treasury bill yield (Sveriges Riksbank via Haver Analytics). Interpolated using 6-month treasury bill yield (Bloomberg ticker GSGT6M) and 3-month treasury bill yield (Bloomberg ticker GSGT3M).	January 2000 – November 2015: 10-year government bond yield (Sveriges Riksbank via Haver Analytics). Interpolated and spliced using information from 10-year government bond yield (Bloomberg ticker GSGB10YR).
Switzerland	January 2000 – November 2015: Treasury bill rate (IFS line 14660C..ZF...). Interpolated using 1-year government bond yield (Bloomberg ticker GSWISS03) and 1-year government bond yield (Bloomberg ticker GSWISS01).	January 2000 – November 2015: Yield on 10-year confederation bonds (OECD MEI series 146.IRLTLT01.ST). Interpolated and spliced using 10-year government bond yield (Bloomberg ticker GSWISS10).
Taiwan, POC	January 2000 – November 2015: 3-month treasury bill yield in new dollars (GFD). Interpolated using 6-month treasury bill secondary market rate in new dollars (GFD) and 2-year government generic bid yield (Bloomberg ticker GVTW2YR).	January 2000 – November 2015: 10-year government bond yield (Central Bank of Taiwan via Haver Analytics). Interpolated and spliced using bid yield on generic government 10-year note (Bloomberg ticker GVTWTL10).
Thailand	January 2000 – November 2015: Government bill yields (IFS line 57860C..ZF...). Interpolated using Thai bond dealing center 1-month rate (Bloomberg ticker TBDC1M) 6-month rate (Bloomberg ticker TBDC6M).	January 2000 – November 2015: Government bond yield (IFS line 57861...ZF...). Interpolated and spliced using information from 10-year government bond yield (Bloomberg ticker GVTL10YR).
Turkey	January 2000 – November 2015: 3-month treasury bill rate (MBRF2 line 18660C..YI...). Interpolated using 6-month government bond yield (Bloomberg tickers IESM6M and IECM6M).	February 2005 – November 2015: 10-year composite government bond yield (Bloomberg ticker IECM10Y). Interpolated and spliced using information from 5-year composite government bond yield (Bloomberg ticker IECM5Y) and 5-year government bond yield in new lira (GFD).
United Kingdom	January 2000 – November 2011: Treasury bill rate (IFS line 11260C..ZF...). Spliced using 3-month generic government bond (Bloomberg ticker GUKG3M).	January 2000 – November 2015: Generic government 10-year yield (Bloomberg ticker GUKG10).
Uruguay	May 2002 – October 2015: Letras de tesorería en moneda nacional (MBRF2 line 29860C..ZI...). Spliced using monetary policy rate (MBRF2).	