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Abstract

This paper investigates how a rise in US long-term interest rates would have an effect on other international markets. We document that a significant portion of long-term interest rates is due to term premium, which can be interpreted as compensation for inflation risk. Based on an assumed interest rate shock of a 200-basis-point increase in the US term premium, we find that most economies experience a stronger response to an interest rate shock of the same magnitude following the taper tantrum episode in May 2013. A fixed-effect panel data regression analysis suggests that country specific fundamentals such as financial integration and inflation are crucial determinants of their response.

Keywords: Spillovers, term premium, financial integration, vector autoregression, emerging markets, affine term structure model, sovereign bond markets

JEL classification: C32, F34, G15, H63

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1. Introduction

US inflation risk has tilted to the upside recently amid concerns over surging energy inflation and expectations of a large-scale fiscal expansion under the Trump administration at a time when economic slack in the US is already diminishing (see Figure 1). Such a repricing of inflation risk quickly reverberates globally, leading to widespread and notable increases in long-term yields not only in advanced economies (AEs), but also in many emerging market economies (EMEs) as shown in Figure 2. Against this background, this paper assesses the potential spillover impact of higher interest rate expectations in the US on sovereign bond markets in other economies.

In theory, long-term interest rates at any maturity can be decomposed into two key components according to the expectations hypothesis: (1) an expectation of future short-term rates; and (2) term premium. The latter is the additional return that compensates investors for holding a long-term bond as opposed to rolling over a sequence of short-term bonds over the same period. Given that inflation erodes the nominal value of long-term bonds more than their short-term counterparts, a positive term premium can be interpreted as a compensation for inflation risk. As shown in Figure 1, we can see that the both 10-year US Treasury yields and term premium edged up tangibly when the market became concerned about a repricing of inflation risk under the Trump Administration. Thus, instead of assessing sovereign bond yields directly, we examine the issue through assessing term premium that captures the transmission of uncertainty about inflation in this paper.

Term premium can be obtained from either survey or model-based estimates. However, survey estimates are typically infrequent and may not be available for many EMEs. Hence, we follow previous studies and use an affine term structure model (affine model) to extract term premium for each economy.[1] The model assumes that the driving force of the yield curve are the first three principal components of the yield curve and imposes no arbitrage condition

[1] See Piazzesi (2010) for a survey of the affine model.

to derive the expectations components and term premium.[2] Although it is well known that affine models may be problematic under a near-zero interest rate environment and alternative models have been proposed recently (e.g. Monfort et al. (2015), Wu and Xia (2016) and Filipovic et al. (2017)), we opt to use the affine model as it is relatively simple to estimate and can be readily applied to our sample of both AEs and EMEs.[3]

In an influential study, Wright (2011) argues that the fall in long-term interest rates in OECD countries from 1990 to 2007 is mainly due to a fall in term premia. He establishes that the reduction in term premia is largely driven by a better anchoring of inflation expectations due to inflation targeting. By extending both the sample period to include the post-global financial crisis and EME experiences, this paper sheds additional insights on the dynamic interaction between term premia globally. Including yield behaviour after the global financial crisis can capture the effect of flight-to-quality and/or search-for-yield phenomenon affecting government securities that happened not only in many AEs, but also in many EMEs.[4] Moreover, the inclusion of EMEs is meaningful to an understanding of global term premia because EMEs have played a major role in global financing flows after years of Quantitative Easing (QE) adopted by major central banks.[5]

We find that the share of term premium in explaining yield variation is maturity dependent. For short-term bond yields such as the 1-year yield variation, the share of term premium is roughly 30%, as opposed to 80% in 10-year bond yields. Taken at face value, these figures suggest while the co-movement of short-term yields is accounted for by the risk-neutral expectation of future short-term rates, the simultaneous decline in long-term yields is due more to declining long-term term premia.

[2] The use of affine model for extracting term premium is common. For example, see Kim and Wright (2005) and Adrian et al. (2013).

[3] The dynamics of interest rates in the affine model are symmetric due to its normality assumption. However, this assumption will be problematic when interest rates are near zero as there is a limit on how far interest rates can fall below zero. This concern is less relevant for EMEs as their policy rates are still significantly higher than zero.

[4] The flight-to-quality and/or search-for-yield phenomenon for government securities after the global financial crisis could have been attributable to the fact that (i) as policy rates remain low in many economies, long-term sovereign bonds remain a viable and attractive investment; and (ii) banks and insurers are now required to hold more safe assets such as government securities because of regulatory requirements.

[5] There is a voluminous literature on the spillover of QE on EMEs. For recent studies, see Chen et al. (2016).

Many recent studies suggest that US monetary policy is a global factor affecting global financial markets. For instance, Bruno and Shin (2015) find that US monetary policy can influence global liquidity and leverage working through the channel of branch and subsidiaries networks of global banks. Rey (2015) argues that US monetary policy affects the global financial cycle, leverage and returns. Georgiadis (2016) finds that US monetary policy generates substantial output spillovers to other economies. The transmission of monetary policy is conventionally viewed as an effect rippling from the short end of the yield curve, which is managed by the US Federal Reserve, to the long end, over which it has no control. Hence, it is possible that spillovers from US long-term interest rates to other economies could be substantial.

To address the possible impact arising from a change in US term premium, we estimate a vector autoregression (VAR) and conduct an impulse response analysis based on the estimated term premia. We consider an interest rate shock of a 200-basis-point increase in the US term premium, which mimics a rise of the US term premium from the current level of almost zero percent to its long-run pre-crisis mean level of 2% between 1980 and 2008. We focus on two sample periods in our analysis, covering the periods January 2011 to May 2013 and June 2013 to December 2016. We find that the estimated responses of all economies in the post-tapering period are mostly stronger than those in the pre-tapering period. The stronger response in the post-tapering period can also be seen from a 100-week rolling window estimation of the VAR which shows that the estimated responses are tangibly higher after May 2013. To shed light on why some economies are more responsive to the shock, we regress the estimated responses from the rolling window VAR on various country characteristics in a panel regression. We find that trade and financial linkages and inflation are the key determinants.

The rest of this paper is organised as follows. Section 2 describes the data and provides some preliminary analysis. Section 3 describes both the static and rolling window VARs and reports the impulse response analysis. Section 4 reports the panel regression. Section 5 concludes. A brief discussion of the affine model is provided in the Annex.

2. Data

We obtain weekly zero-coupon sovereign bond yields for twenty six economies from Bloomberg.[6] The earliest available data for some economies is from March 1989 but the sample period of yield-curve data for some EMEs is rather short. The sample ends in December 2016.[7] For the cross sectional dimension, we select maturities of 3-month, 6-month, 1-year, 2-year, 3-year, 5-year, 7-year and 10-year. To make the assessment more comprehensive, we include seven AEs that have a sovereign bond market worth more than one trillion USD as at June 2015. The selection criteria of EMEs are based on the availability of long-term yield data and criteria commonly used in previous studies.[8] Besides the US which is the main shock origin, other economies can be classified into the following four groups: (1) AEs excluding US which include Germany, Japan, the UK, Italy, Canada, France and Spain; (2) Latin America which include Brazil, Chile, Colombia, Mexico and Peru; (3) Emerging Asia which include China, India, Indonesia, Hong Kong, Singapore, South Korea, Thailand and the Philippines; and (4) Emerging Europe and Africa (EMEA) which include Czech Republic, Hungary, Poland, South Africa and Turkey. Table 1 provides a summary of the regional classification and the earliest availability of the zero-coupon yield curve for each economy.

3. Empirical analysis

[6] According to the data vendor, the zero-coupon yields are stripped from the most recent auctioned on-the-run sovereign bonds using standard bootstrapping.

[7] It is worth noting that only some economies in the sample have yield curve data starting in 1989. For each economy, we take the longest possible data from Bloomberg as a sufficiently long data is less prone to identification problems inherited in the estimation of affine term structure model. For details, see Bauer et al. (2013)

[8] The selected EME has to satisfy either one of the following three criteria. (1) A member of either the IMF's emerging or developing economies or World Bank's low and middle-income countries; (2) Constituents of Barclays, JP Morgan, Markit or Merrill Lynch emerging-market government bond indices; and (3) Stock of public debt exceeding USD 10 billion or long-term sovereign credit rating above BB/Ba.

In this section, we first document the contribution of term premia in short- and long-term bond yields, which explains a significant portion of long-term interest rates. We then use estimated term premium to assess the impact of a rise in US term premium using a VAR. Finally, we use a panel regression to link up the determinants with the estimated impulse response of the VAR.

3.1 Role of term premia in short- and long-term bond yields

Let $\mathrm{Y^m}, \mathrm{RN^m}$ and $\mathrm{TP^m}$ denote the yield, expectations component and term premium respectively for a m-tenor sovereign bond. By construction, $\mathrm{Y^m} = \mathrm{RN^m} + \mathrm{TP^m}$. Using the concept of no-arbitrage, modern financial theory postulates that RN can be computed using the notion of a risk-neutral measure.[9] Hence, given any observed yield $\mathrm{Y^m}$ and estimated expectations component $\mathrm{RN^m}$, $\mathrm{TP^m}$ can be obtained residually for any maturity of a bond.

Partly due to the availability of yield curve data for EMEs and partly due to the excessive volatility in the aftermath of the global financial crisis, all of the empirical analysis conducted below is based on a balanced panel of 1-year and 10-year term premia for twenty six economies from January 2011 to December 2016. The second column of Table 2 summaries the contribution of term premium to 1-year sovereign bond yields. As RN and TP may not always be positive, the share in Table 2 is approximated by $\frac{\Delta \mathrm{TP}^2}{\Delta \mathrm{TP}^2 + \Delta \mathrm{RN}^2}$, where Δ is the difference operator.[10] As can be seen, term premium explains about one third of the fluctuations in short-term bond yields on average. Alternatively, we can interpret this as saying that two thirds of the fluctuations in short-term bond yields are due to changes in the expectations component, which measures investors' perceptions of the bond yield in the next year. This finding suggests that short-term bond yield movements are more related to expectations of future interest rates, which are in turn related to expectations of future monetary policy actions of the underlying sovereignty.

[9] See Annex 1 for more description on the formal definition of expectation component and term premium.

[10] The denominator is approximately equal to ΔY^2 given that $\Delta TP \Delta RN \approx 0$.

The last column of Table 2 shows the share of the term premium of 10-year sovereign bond yields. Term premium explains a significant amount of the fluctuations in long-term bond yields, with an explanatory power of 82% on average during the sample period. While Wright (2011) only focuses on OECD countries and finds that long-term interest rates are overwhelmingly explained by term premium, we find that this finding is also applicable to many EMEs. For instance, term premium on average explain 86%, 83% and 79% of fluctuations in Latin America, Emerging Asia and EMEA respectively.

3.2 How do economies respond to a rise in the US term premium?

As US monetary policy is often regarded as a global factor in previous studies such as Bruno and Shin (2015) and Rey (2015), we conduct a scenario analysis of a hypothetical one-shot increase in the US term premium. We focus on 10-year bond yields as they are more affected by movements in term premium.

Specifically, we run the following weekly frequency VAR(1)

$$\Delta \mathrm{TP}_t = \mathrm{c} + \Theta \Delta \mathrm{TP}_{t-1} + \alpha \Delta \mathrm{VIX}_t + \beta \Delta \mathrm{DXY}_t + \varepsilon_t \qquad (1)$$

where c and ε_t are the constant and error terms of the regression. To control for the effect of global factors that could affect global financial markets, we include changes in the VIX index and the US dollar index (DXY) as exogenous factors in Eq. (1).

We first estimate Eq. (1) on two adjacent periods, covering the periods from January 2011 to May 2013 (pre-tapering) and from June 2013 to December 2016 (post-tapering). We separate the analysis into two adjacent periods because our previous research found that global sovereign bond markets have become more synchronised following the taper tantrum in May 2013.[11] Based on the estimated coefficients in both periods, we conduct an impulse response analysis to evaluate how term premia in other economies would respond to an interest rate shock of a 200-basis-point increase in the US term premium. This interest rate shock mimics

[11] For details, see Fong et al. (2016).

a rise of the US term premium from the level of 0% at December 2016 to its long run pre-crisis mean level of 2% from 1980 to 2008.

Figure 3 shows the cumulative 10-week impulse responses in term premium to the US shock during the two sample periods, with the economies in each economy group ranked according to the size of their impulse responses in the post-tapering period. Taking Hong Kong as an example, the estimated increase is 89 basis points in the pre-tapering period, compared with the increase of 143 basis points in the post-tapering period. We summarise the key findings as follows:

First of all, the estimated responses of all economies in the post-tapering period are mostly stronger than those in the pre-tapering period, except for South Africa, Chile, and Japan. On average, the estimated change in term premium for all economies is 70 basis points in the pre-tapering period and 117 basis points in the post-tapering period. This result suggests that the differentiation between the valuation of the US and other economies' long-term sovereign bonds has narrowed since the taper tantrum.

Second, by comparing the estimated increases of EMEs in the post-tapering period, economies in EMEA are the most responsive to the US shock on average (146 bps), followed by those in Latin America (123 bps) and Emerging Asia (96 bps). This probably reflects the fact that geo-political instability remains a key risk confronting emerging economies in EMEA, while relatively stronger economic fundamentals eases part of the risk in Emerging Asian economies in the post-tapering period.[12]

Finally, the spillover impact on AEs would be comparable with that on EMEs. On average, the estimated increase in term premium in AEs is 117 basis points in the post-tapering period. The commensurate response may partially stem from heightened economic and political uncertainties in some core European economies with closer trade and financial linkages with the US.

[12] As a reference, the average real GDP growth from June 2013 to December 2016 in Emerging Asia is 4.6%. The corresponding figures for AEs, Latin America and EMEA are 1.4%, 2.2% and 3.1% respectively.

We can also obtain a time-varying measure of the US term premium spillover by estimating Eq. (1) using a rolling window of 100 weeks. That is, for the full sample period from January 2011 to December 2016, we estimate Eq. (1) for every 100 weeks of data and the 100 week window is rolled forward every week. In each of the window, we conduct the same interest rate shock of a 200-basis-point increase in the US term premium and record the 10-week cumulative shock impact for every economy. For ease of illustration, we only report the regional average in Figure 4. Since the taper tantrum in May 2013, the responses of all economy groups is tangibly higher. Meanwhile, consistent with the findings in Figure 1, economies in EMEA are the most responsive to the interest rate shock. It is worth noting that all economies register a temporary spike in their interest rate responses shortly after Trump's victory in the US president election. This probably reflects a repricing of inflation risk that accompanied expectations of expansionary fiscal policies by the Trump administration which will likely raise US long-term interest rates.

3.3 Determinants of the US term premium spillover

To understand why some economies are more responsive to the US term premium shock, we consider the following panel regression:

$$s_{it} = \rho s_{it-1} + \gamma Z_t + \alpha_i + u_t + \varepsilon_t \tag{2}$$

where s_{it} is economy i's 10-week cumulative impulse response to the US shock from the rolling window estimations of Eq. (1), which we define as US spillovers.[13] Z_t is a vector of explanatory variables commonly used in previous studies, outlined in more detail below. α_i and u_t are country fixed effects and time effects to capture the unobserved heterogeneity across the sample and over time. Finally, ε_t is the error term in the regressions. In Eq. (2), we include the lagged term s_{it-1} to control for the persistence of the dependent variable. We estimate Eq. (2) using the panel GMM approach pioneered by Arellano and Bond (1991),

[13] s_{it} is available in weekly frequency from the estimation of Eq. (1), we use its quarter-end figures in Eq. (2).

Arellano and Bover (1995) and Blundell and Bond (1998).[14] Specifically, we use a GMM in first differences in the estimations. The quarterly frequency regression starts from the first quarter of 2011 to the fourth quarter of 2016. Table 3 shows the estimation results.[15] The Hansen J-statistic and the AR(2) serial correlation tests suggest that our model is adequate. Key findings are summarized as follows:

1. Lagged spillover: The lagged spillover is positive and statistically significant, indicating that the effect of US term premium shock is persistent with a significant estimate of the autocorrelation coefficient of 0.78.

2. Trade integration: Trade integration is measured by the sum of a country's export and import to its GDP. Previous studies have mixed findings on whether trade can enlarge or dampen foreign shocks. On the one hand, Clark and Van Wincoop (2001) and Baxter and Kouparitsas (2005) suggest that the trade channel is important in business cycle synchronization and spillovers. On the other hand, trade can dampen the impact of external shocks by making sudden stops and current account reversal less likely (see Rey and Martin (2006), Cavallo and Frankel (2008) and Calvo et al. (2008)). We find that the beneficial role of a trade channel outweighs its possible negative effects on financial stability as the coefficient of trade integration is significantly negative which reduces the pass through of the US term premium shock.

3. Financial integration: Besides trade linkage, financial linkages across countries can give rise to large swings in capital flows and asset prices. Previous studies have suggested that EMEs that are more financially open are prone to abrupt reversals in their current account positions and sudden stops (see Rey and Martin (2006), Edwards (2007) and Milesi-Ferretti and Tille (2011)). We follow previous studies and proxy financial integration by gross foreign assets and liabilities relative to GDP as in

[14] According to Eq. (2), α_i is a part of the process that generates $s_{\mathrm{it}-1}$ such that $E(s_{\mathrm{it}-1}\alpha_i) > 0$. Thus, $s_{\mathrm{it}-1}$ is correlated with the individual fixed effects and is not strictly exogenous, which means standard panel OLS estimation is not applicable for Eq. (2).

[15] The reported standard errors are the first-stage standard errors.

Lane and Milesi-Ferretti (2007). The positive estimate in Table 3 suggests that economies that are more financially integrated with the global economy are more vulnerable to shocks from the US.

4. Sovereign bond market depth: An economy with deeper sovereign bond markets (as measured by its size over GDP) typically has finer pricing and better price discovery. Hence, it should be better able to withstand shocks in global financial markets. However, we find it the coefficient is not significant at conventional significance levels.

5. Volatility of exchange rate: The volatility of exchange rates is measured by the square of the log difference between an economies' foreign exchange rate against the US dollar in the current quarter compared with the previous quarter. The volatility of the exchange rate is a proxy for the exchange rate regime in place in that economy. Holding other factors equal, a more flexible exchange rate regime enables the economy to adjust swiftly to negative shocks, but the coefficient estimate in Table 3 is not significant at conventional significance levels.

6. Inflation: It is well known that a significant portion of long-term interest rates reflects compensation for inflation risks. In responding to bond market shocks, it is conceivable that economies with higher inflation would be subjected to a larger surge in its sovereign bond yields. The positive and significant coefficient for inflation appears to confirm this intuition.

4. Concluding Remarks

Against a background of higher upside risks to US inflation, this paper examines how a rise in US term premium would affect other economies. Our empirical findings suggest that the influence of the US Treasury bond market on other sovereign bond markets has increased since the taper tantrum episodes in 2013, and that higher interest rates and tighter financial conditions in the US would significantly affect many AEs and EMEs. If a repricing of inflation risk leads to a rapid surge in the long end of the US yield curve, the impact on other economies could be significant.[16] In particular, increases in sovereign bond yields would likely lead to higher borrowing costs in the private sector that would have a material impact on EMEs with weaker underlying growth and a heavier sovereign debt financing burden. How the expansionary fiscal policies proposed by the new US administration may impact the US economy and affect the future trajectory of US long-term interest rates should be closely monitored.[17]

[16] It is worth noting that an increase in short-term interest rates due to the Fed tightening may not always lead to an increase in the long-term interest rates. One recent example is the Greenspan conundrum in 2005, during which US long-term interest rates remained flat when the US Fed started the tightening cycle as the term premium actually fell (see Chart B1.1). For details on the Greenspan conundrum and its association with term premium, see Backus and Wright (2007),

[17] If expansionary fiscal policies lift up the US economy significantly leading to substantial inflationary pressures, the US Fed would respond by raising its policy rate which could increase US long-term interest rates.

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Annex 1 The Affine Term Structure Model

In this section, we briefly describe the model specification of the term structure model. We adopt the yield-only Gaussian dynamic term structure model formulated by Joslin et al. (2011). The short-term interest rate (i.e., three-month rate) is linked by observable risk factors X_t (i.e., the state variables) by

$$r_t = \delta_0 + \delta_1^T X_t \tag{A1}$$

The state variables following a first order VAR under real world probability measure *P*

$$X_t = \mu + \phi X_{t-1} + \Sigma \varepsilon_t \tag{A2}$$

where Σ is lower triangular and ε_t is an i.i.d. standard normal random vector. In the absence of arbitrage, it can be shown that there exists a risk-neutral probability measure *Q* such that the law of motion of the risk factors in Eq. (A2) can be rewritten as[18]

$$X_t = \mu^Q + \phi^Q X_{t-1} + \Sigma \varepsilon_t^Q \tag{A3}$$

With the assumptions in Eqs. (A1) to (A3), it can be shown that a stochastic discount factor exists and the price of bond with maturity of *m* period is

$$P_t^m = E_t^Q \left[\exp\left(- \sum_{i=0}^{m-1} r_{t+i} \right) \right] \tag{A4}$$

This expectation can be solved analytically which is exponentially affine in the risk factors

$$P_t^m = \exp(A_n + B_n^T X_t) \tag{A5}$$

where the loadings A_n and B_n are functions of model parameters ($\mu^Q, \phi^Q, \delta_0, \delta_1, \Sigma$) which satisfies the following recursion

$$A_n = -\delta_0 + A_{n-1} + B_{n-1}^T \mu^Q + \frac{1}{2} B_{n-1} \Sigma \Sigma^T B_{n-1}^T \tag{A6}$$

$$B_n = -\delta_1^T + \phi^Q B_{n-1}^T \tag{A7}$$

with starting values $\mathrm{A}_1 = -\delta_0$ and $\mathrm{B}_1^\mathrm{T} = -\delta_1^\mathrm{T}$. Yield is also affine in the risk factors because $y_t^m = -m^{-1} \log(P_t^m) = -m^{-1}(A_n + B_n^T X_t)$. If the risk-neutral parameters μ^Q and ϕ^Q in Eqs. (A6) and (A7) are replaced by the real world parameters μ and ϕ , we can define the risk-neutral yield $\hat{y}_t^m$ analogously. For each maturity *m*, the risk-neutral yield approximately measures the expected interest rates (or policy expectations) over the life of the bond (i.e., $m^{-1} \sum_{h=0}^{m-1} E_t r_{t+h}$). Term premium is defined as the difference between the model yield and the risk neutral yield. (i.e., $\mathrm{ytp}_\mathrm{t}^\mathrm{m} = \mathrm{y}_\mathrm{t}^\mathrm{m} - \hat{\mathrm{y}}_\mathrm{t}^\mathrm{m}$)

Following the previous studies, the state vector X_t consists of the first three principal components of zero-coupon yields constructed from the yield curve of each economy. Litterman and Scheinkman (1991) shows that the three principal are sufficient to capture most of the variation in the yield curve.

In estimating the term-structure model, we follow the estimation strategy developed by Joslin et al. (2011) in separately estimating the risk neutral and real world parameters. Specifically, as the state vector X_t is observed, the real world parameters μ and ϕ can be estimated from fitting a VAR on Eq. (A2) using ordinary least square. Assuming observed bond yields are equal to the model-implied yields plus i.i.d normal measurement errors, the remaining parameters $\mu^*, \phi^*, \delta_0, \delta_1$ and Σ are estimated by maximum likelihood.

[18] The risk-neutral and real world parameter are related such that $\mu^Q = \mu^P - \Sigma\lambda_0$ and $\Phi^Q = \Phi^P - \Sigma\lambda_1$. λ_0 and λ_1 are the loadings of the market price of risk λ to the observed factors X_t (i.e., $\lambda = \lambda_0 + \lambda_1 X_t$)

Table 1: The regional classification and the first date at which zero-coupon yield curve existed for each economy

Country	Date
US	31 March 1989
Advanced Economy	
Germany	30 December 1994
Japan	07 April 1989
UK	30 December 1994
Italy	30 December 1994
Canada	30 December 1994
France	30 December 1994
Spain	30 December 1994
Latin America	
Brazil	30 March 2007
Chile	30 September 2005
Colombia	28 April 2006
Mexico	08 August 2003
Peru	05 May 2006
Emerging Asia	
China	30 April 2004
Hong Kong	30 December 1994
South Korea	29 November 2002
the Philippines	28 June 1996
Singapore	30 December 1994
Thailand	30 December 1994
Indonesia	16 May 2003
India	13 November 1998
Emerging Europe and Africa	
Czech	15 December 2000
Hungary	16 March 2001
Poland	15 December 2000
Turkey	01 April 2005
South Africa	03 February 1995

Source: Bloomberg

Table 2: Share of term premium in short-term and long-term sovereign bonds

Country	one-year	ten -year
Advanced Economy		
US	41%	83%
Germany	29%	83%
Japan	32%	80%
UK	41%	74%
Italy	29%	75%
Canada	33%	90%
France	31%	80%
Spain	34%	78%
Latin America		
Brazil	40%	89%
Chile	51%	95%
Colombia	25%	69%
Mexico	41%	78%
Peru	58%	97%
Emerging Asia		
China	53%	90%
Hong Kong	26%	83%
South Korea	50%	65%
the Philippines	51%	79%
Singapore	46%	91%
Thailand	21%	93%
Indonesia	49%	91%
India	15%	71%
Emerging Europe and Africa		
Czech	19%	76%
Hungary	27%	80%
Poland	7%	88%
Turkey	35%	81%
South Africa	38%	65%
AEs' average	*33%*	*80%*
Latin America's average	*43%*	*86%*
Emerging Asia's average	*39%*	*83%*
Emerging Europe and Africa's average	*25%*	*78%*
All economies' average	**35%**	**82%**

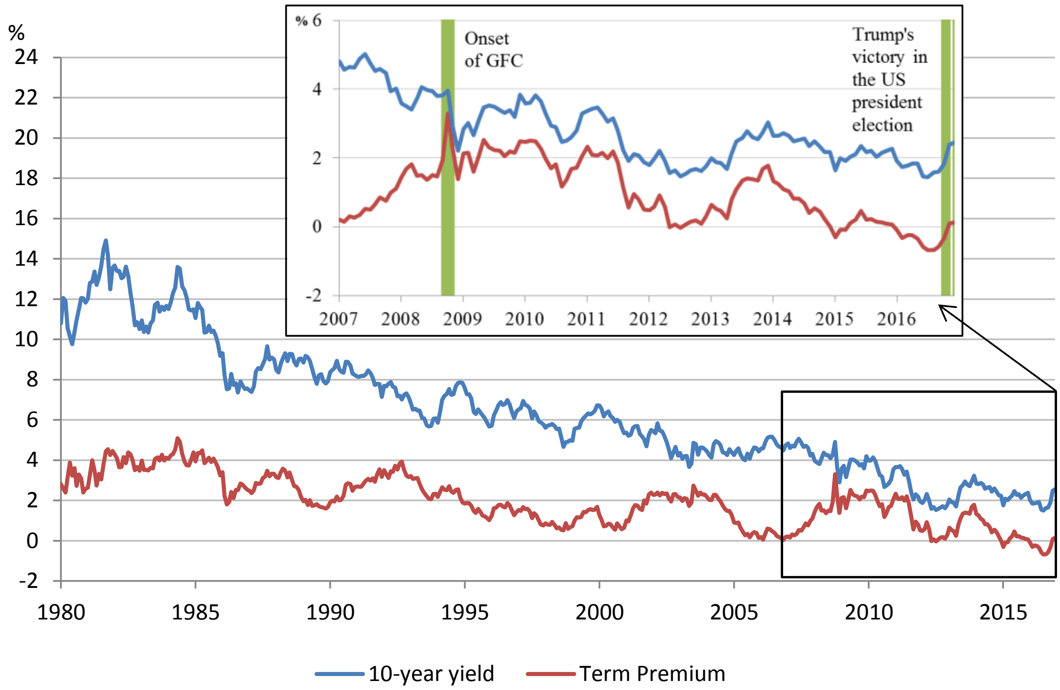
Source: Staff estimates.

Table 3. Estimation result on the determinants of the US spillover in Eq. (2)

	(1)
Lagged spillover	0.777
	(0.075)***
Trade Integration	-2.075
	(0.592)***
Financial integration	0.290
	(0.146)**
Sovereign bond market depth	-4.298
	(4.595)
Volatility of exchange rate	-17.020
	(27.580)
Inflation	5.813
	(2.931)**
Country fixed effect	Y
Time fixed effect	Y
Hansen test (P-value)	0.168
AR(2) test (P-value)	0.398

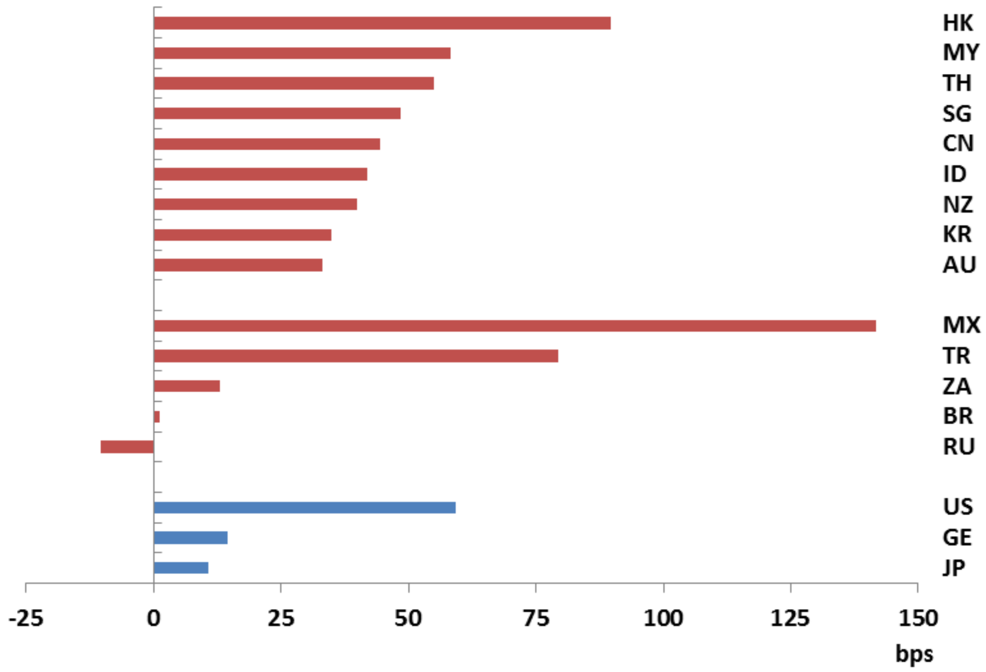
Note: Panel regression of the **impulse response estimated from the rolling VAR of Eq. (1)** to its own lag and country-specific factors. 'Hansen test' reflects the Hansen J-statistic and 'AR(2) test' indicates the second-order serial correlation test of residuals. Robust standard errors are shown in parentheses, and ***, ** and * denote significance at the 1%, 5% and 10% levels, respectively.
Source: Staff estimates.

Figure 1: 10-year US Term Premium and Treasury Yield



Sources: Blomberg and Federal Reserve Bank of New York.

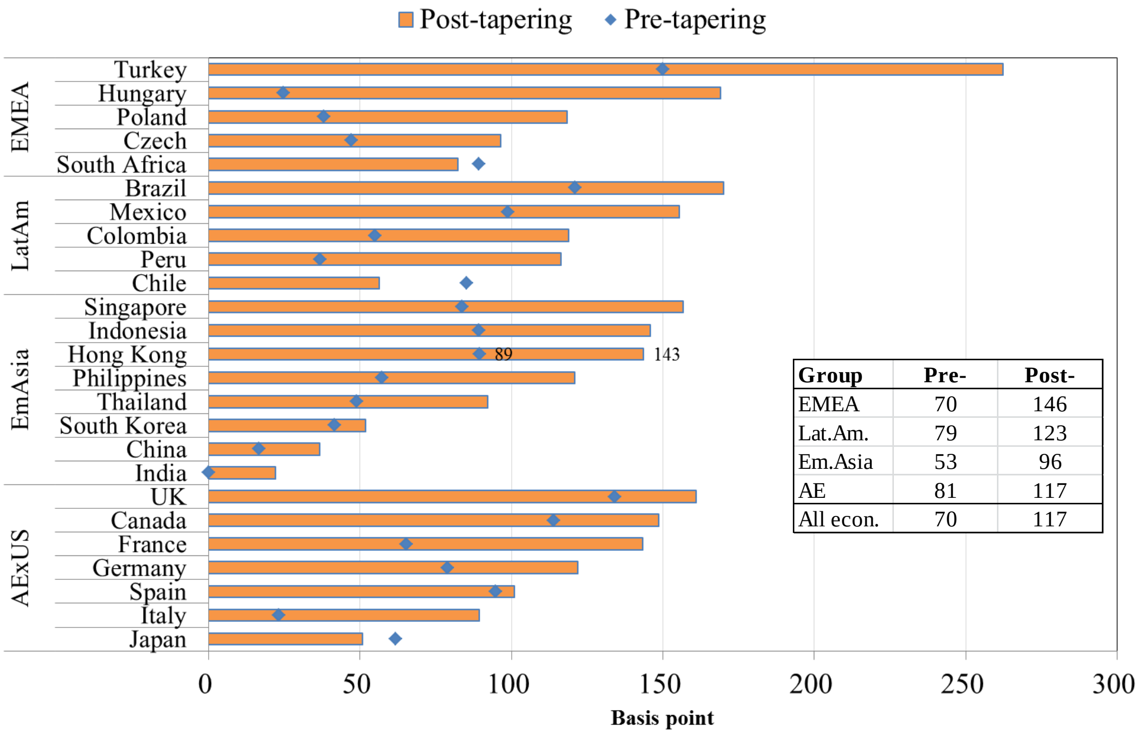
Figure 2: Change in 10-year sovereign bond yields for selected economies since the US election



Note: From 7 Nov 2016 – 6 Jan 2017 for all economies with the exception of Hong Kong and Russia which are from 7 Nov 2016 – 30 Dec 2016

Source: Bloomberg.

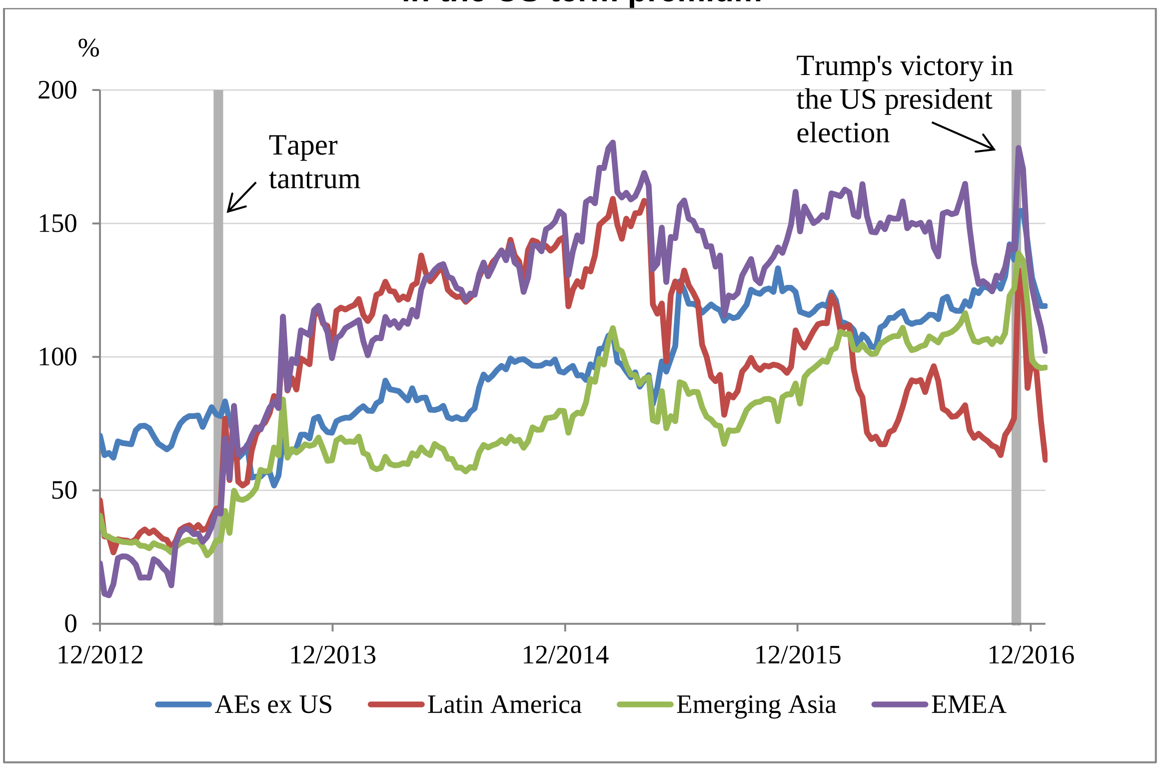
Figure 3: The 10-week cumulative responses to a 200-basis-point increase in the US term premium



Group	Pre-	Post-
EMEA	70	146
Lat.Am.	79	123
Em.Asia	53	96
AE	81	117
All econ.	70	117

Note: pre-tapering period denotes January 2011 to May 2013. Post-tapering period denotes June 2013 to December 2016.
Source: Staff estimates.

Figure 4: The time-varying responses to a 200-basis-point increase in the US term premium



Note: Each line represents the regional average of interest rate responses from respective economies. Results based on a 100-week rolling window estimation of Eq. (1).
Source: Staff estimates.