Monetary Spillovers in Financial Markets:
Policymakers and Premia

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As global markets become more interlinked, there is growing concern about the Federal Reserve’s effects internationally. I examine the channels through which Fed monetary policy spills over into foreign developed financial markets. I establish a new fact on the asymmetric reactions of currencies and foreign bonds to Fed announcements, and show that it provides evidence against two leading channels of spillovers. Using high-frequency returns, I show that when the Fed tightens: (i) the dollar appreciates by more against high-rate currencies (e.g. the Australian dollar) than against low-rate currencies (e.g. the yen), and (ii) high-rate long-maturity bond yields rise more than low-rate long-maturity bond yields. The asymmetries across currency and bond markets provide evidence against theories in which foreign central banks react to the Fed, or in which foreign risk premia shift in a complete markets framework. Currency markets predict that when the Fed tightens, central banks tighten most or stochastic discount factors rise most in low-rate countries. By contrast, bond markets predict that central banks tighten most or stochastic discount factors rise most in high-rate countries. I further argue against these theories, by using evidence from a decomposition of foreign term structures and by showing unusual restrictions generated in models of complete markets. Only shifts in foreign risk premia under incomplete markets are consistent with these patterns, and I document the conditions such models must match. My results suggest that spillovers do not diminish the independence of central banks, but rather illustrate the importance of frictions in global markets.

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1 Introduction

Deepening international linkages between markets have intensified global spillovers of the Federal Reserve's monetary policy into foreign financial markets, as noted by Rey [2013]. Debates over spillovers between academics and between policymakers have intensified too. In the last few years, foreign central banks have concurrently blamed the Fed for chaos in local markets, decried their reduction in monetary independence, and called for a new regime of global monetary coordination. Various Fed governors, including Ben Bernanke, Stan Fischer, Jerome Powell, and Lael Brainard, have responded publicly, deflecting the blame and reaffirming the Fed’s commitment to domestic objectives. At the heart of the debate is an essential question for assessing welfare and designing policy: what are the channels of monetary spillovers by the Federal Reserve into foreign financial markets?

To answer this question, I establish a new fact on how currencies and bonds react asymmetrically to the Fed’s announcements and use it to test different channels. I divide explanations for monetary spillovers into three classes: ones in which foreign central banks react to the Fed, ones in which foreign risk premia (i.e. compensation for bearing risk) shift under complete markets, and ones in which foreign risk premia shift under incomplete markets. My fact provides evidence against the first two channels of spillovers, and is only consistent with the third channel.

I document asymmetric responses of foreign bonds and currencies to Fed announcements, using high-frequency data and methodologies robust to market noise. When the Fed tightens, the dollar appreciates more versus currencies of high-rate countries (e.g. Australia) and appreciates less versus currencies of low-rate countries (e.g. Japan). Moreover, when the Fed tightens, long-maturity bond yields of high-rate countries rise more than those of low-rate countries.

The observed patterns of asymmetries in this fact are inconsistent with theoretical patterns of asymmetries predicted by models in which central banks react to the Fed and by models in which risk premia shift under complete markets. For instance, the observed asymmetries in currency markets suggest that the central banks of low-rate countries tighten most when the Fed tightens, while the observed asymmetries in bond markets suggest that the central banks of high-rate countries tighten most. Moreover, the asymmetries in currency markets suggest that the stochastic discount factors of low-rate countries rise most when the Fed tightens, while the asymmetries in bond markets suggest that the stochastic discount factors of high-rate countries rise most. I offer further evidence from the term structures of foreign bonds to argue against theories in which central banks react to the Fed, and I discuss further how models with complete markets must incorporate complex and unusual assumptions to be consistent with this novel fact. The only theories that explain spillovers are ones in which risk premia shift under incomplete markets. Successful models in this class must match one additional condition: they must ensure that only a few pivotal central banks can generate spillovers, as I document that most central banks do not have effects on global markets.

To illustrate the fact on asymmetric spillovers, consider an example. At 12:30 PM on January 25, 2012, the Fed announced its intentions to keep interest rates low until 2014. The surprise monetary easing affected foreign assets in asymmetric ways, as sixty-minute windows around the
Yields on ten-year Australian bonds immediately fell whereas yields on ten-year Japanese bonds did not fall. Moreover, the dollar depreciated more against the Australian dollar than against the yen, or equivalently the Australian dollar appreciated against the yen. These asymmetric shifts in bond and currency markets must reflect one of three explanations: (i) changes in the expected paths of short rates by the Reserve Bank of Australia and Bank of Japan, (ii) changes in investors’ willingness to hold Australian and Japanese assets (i.e. shifts in risk premia under complete markets), or (iii) changes in investors’ abilities to hold Australian and Japanese assets (i.e. shifts in risk premia under incomplete markets). I stress that these hypotheses are exhaustive, and any other taxonomy of the channels of Fed spillovers into foreign markets can be mapped into these three. I use these asymmetric asset responses and other evidence to argue against the first two hypotheses, in favor of the third hypothesis.

Figure 1: Market Reactions to the Fed Easing on January 25, 2012

(a) Ten-Year Bonds

(b) Currencies

Notes: The figures depict reactions in foreign bond and currency markets in sixty-minute windows around the Fed’s surprise easing on January 25, 2012. In bond markets, Japanese yields do not move, while Australian yields fall by 30 basis points. In currency markets, the dollar depreciates by 100 basis points against the Australian dollar and by 60 basis points against the yen. Equivalent, the Australian dollar appreciates by 40 basis points against the yen.

To do so, I first establish the key empirical fact: across nine developed countries — Australia, Canada, the Eurozone, Japan, Norway, New Zealand, Sweden, Switzerland, and the United Kingdom — the dollar appreciates or depreciates most against currencies in countries with high rates historically (e.g. Australia) and least against currencies in countries with low rates historically (e.g. Japan) when the Fed tightens or eases respectively. Equivalently, following Fed tightening or easing, the Australian dollar depreciates or appreciates versus the yen respectively. Similarly, when the Fed tightens or eases, long-maturity bond yields from historically high-rate countries rise
or fall more than long-maturity bond yields from historically low-rate countries respectively.\(^1\) The example in Figure 1 illustrates these patterns, as the dollar depreciates more against the Australian dollar, and Australian yields fall more.

This fact owes its causal interpretation and its precision to four methodological components: high-frequency returns around Fed announcements, long-maturity instruments, non-announcement windows, and inferred monetary shocks. The methodology compares currency and long-maturity bond returns in sixty-minute and daily windows around Fed announcements to returns in windows outside Fed announcements, and estimates how assets react to latent monetary shocks from the differences. First, the combination of high frequency returns in announcement windows and returns in non-announcement windows allows me to isolate the effects of monetary policy. Market variation driven by monetary policy can be separated from market variation driven by other forms of news and from idiosyncratic market noise. Next, long-maturity bonds and currencies allows me to capture responses to Fed announcements at all horizons, as they reflect changes in the paths of short rates and risk premia over the bond’s horizon and the infinite horizon respectively. By contrast, approaches that link realized changes in short rates and risk premia to Fed announcements over long horizons through a vector autoregression framework suffer from power issues. Finally, the combination of high-frequency returns and inferred shocks allows me to identify precise asymmetries in asset reactions, by limiting the amount of idiosyncratic noise in the data and by capturing the entire paths of shocks. Differences in the movements of currencies and bonds can be subtle, and low-frequency data or noisier shocks will miss them altogether.

First, I use this fact on asymmetric reactions in currency and bond markets to argue against hypotheses in which central banks react to Fed announcements. These hypotheses are grounded both in theoretical work from the open-economy New Keynesian literature, such as Obstfeld and Rogoff [1996] and Corsetti and Pesenti [2001], and in empirical work on countries’ “fear of floating” freely by Calvo and Reinhart [2002]. To illustrate my argument against these hypotheses, consider the two possible scenarios in Figure 1. The first scenario is that the Reserve Bank of Australia is expected to ease more than the Bank of Japan. This explanation predicts that Australian yields should fall more than Japanese yields and that the Australian dollar should depreciate against the yen. While the observed responses of bond yields support this explanation, the observed responses of currencies do not. The second scenario is that the Reserve Bank of Australia is expected to ease less than the Bank of Japan. Now the observed responses of currencies support this explanation, but the observed responses of bond yields do not. Generalizing from my example, I show that no set of central bank reactions across the nine countries are consistent with asymmetries in both currency and bond markets concurrently.

I provide additional evidence against hypotheses in which central banks react to Fed announcements, using the term structures of foreign bonds. I decompose yields into the paths of short-term policy rates and term premia in each country, and show that term premia react systematically to Fed announcements in most countries whereas the paths of short-term policy rates do not. This

\(^1\)This is not driven by the zero lower bound for interest rates, as I discuss in Section 3.
decomposition is done in two ways. First, I use the edges of the yield curve (short maturity yields and distant forward yields) as proxies for the paths of rates and term premia. Second, I fit a Gaussian affine term structure model to explicitly decompose the yield curve into the paths of rates and term premia. The results from both methods further support the conclusion that central banks do not react to the Fed. Taken together, all of these findings suggest that central banks continue to exercise their independence.

Second, I use the original fact on asymmetric reactions to argue against hypotheses in which risk premia shift under complete markets following Fed announcements. These hypotheses have a long tradition in the international asset pricing literature, where the majority of theories use models of complete markets rather than models of incomplete markets, as surveyed by Engel [2014]. Again, I illustrate my argument against these hypotheses by considering the two possible scenarios. The first scenario is that the stochastic discount factor (i.e. marginal utility) of Japanese investors temporarily rose more than the stochastic discount factor of Australian investors. This explanation predicts that Japanese yields should rise more than Australian yields, as Japanese investors borrow more than Australian investors; and that the yen should appreciate against the Australian dollar, as exchange rates in complete markets reflect the ratio of stochastic discount factors. While the observed responses of bond yields support this explanation, the observed responses of currencies do not. The second scenario is that the stochastic discount factor of Japanese investors temporarily rose less than that of Australian investors. Now the observed responses of currencies support this explanation, but the observed responses of bond yields do not. (Explanations in which stochastic discount factors rise permanently generate no responses in bonds.) Across my sample, I show that basic shifts in risk premia under complete markets contradict the asymmetries in either currency or bond markets.

I argue further against hypotheses in which risk premia shift under complete markets by considering a more general framework with both temporary and permanent shocks to the stochastic discount factor as in Lustig et al. [2017], and by showing that such a framework generates implausible conditions. In models with complete markets, currencies respond to both shocks, while bonds only respond to temporary shocks. Introducing two sources of heterogeneity in the stochastic discount factor gives those models enough freedom to match my results, but in ways that are economically unusual. The underlying forms of heterogeneity must run in opposite directions — the permanent component of Japan’s stochastic discount factor must be more volatile than Australia’s, while the transitory component of Australia’s stochastic discount factor must be more volatile than Japan’s. I illustrate the implausibility of such a specification using a general model that nests several common models and incorporates rich cross-country heterogeneity. The findings suggest that models of complete markets are ill-suited to explain global financial markets around Fed announcements, despite their prevalence in explaining markets at other times.

Third, I argue the original fact on asymmetric reactions is only consistent with hypotheses in which risk premia shift under incomplete markets following Fed announcements. I do not elevate any one specific model, but I demonstrate using a model of segmented markets as in Gabaix and
Maggiori [2015]. In this setting, Fed announcements adjust constraints on leveraged intermediaries, causing them to resize trades that arbitrage rate differentials, and thus causing high-rate currencies and bonds to move together versus low-rate ones. Figure 1 can be interpreted as follows: following the Fed’s easing, intermediaries invest more in Australian assets than in Japanese assets, causing both the Australian dollar to appreciate versus the yen and Australian ten-year yields to fall relative to Japanese ten-year yields. There are other plausible models of frictions, and so instead I offer further restrictions for models in this class. Most importantly, any plausible model must give the Fed special sway over the frictions constraining investors. I examine the spillovers of other central banks, which is in itself novel to the literature. With the exception of the European Central Bank, which generates spillovers that especially affect non-Eurozone countries in continental Europe, no central bank affects foreign financial markets systematically. As such, successful models must go beyond incorporating asymmetries in the reaction of assets — they must incorporate asymmetries in central banks.

This paper provides evidence to debates on two central topics in international finance: the Mundell-Fleming trilemma and the carry trade. First, spillovers have revitalized the fear that central banks have limited independence, particularly if the real ramifications of spillovers are large enough to force the central banks of recipient countries into accommodating them. In particular, Rey [2013] argues that this threatens the trilemma of Mundell [1963] and Fleming [1962]. While the trilemma guarantees independent monetary policy if countries let their currencies float, sufficiently strong spillovers may constrain a central bank’s effective independence even with floating exchange rates.

Despite the importance of this topic to policymakers, answers have remained elusive. The leading approaches in the literature link the paths of short-term rates across countries in a vector autoregression framework, but they either raise identification concerns or suffer limited power. One approach, taken by Rey [2015], Caceres et al. [2016], Hofmann and Takats [2015], and Takats and Vela [2014], is to measure whether innovations to US interest rates predict future innovations to foreign interest rates. This raises the concern, articulated by Bernanke [2017] among others, that omitted factors such as global or regional growth shocks drive all innovations. These papers either impose timing restrictions or augment the specification with more controls, but concerns remain. A second approach, taken by Miranda-Agrippino and Rey [2015] and Rogers et al. [2016], is to identify only from innovations to US interest rates on Fed announcement days. This addresses the identification concerns but weakens power, and these papers find results only at the 68% confidence level. Linking future innovations in foreign rates to current ones in US rates is already challenging given market noise, and this is exacerbated when limited the sample to the Fed’s eight meetings per year.² My paper uses an alternate approach, which both retains power while addressing the identification concerns. As a result, I can show with confidence that foreign central banks retain and exercise their independence in the presence of large spillovers. The trilemma remains a valid

²Ilzetzki et al. [2017] take a third approach, in which they use overall exchange rate volatility and macroeconomic co-movement to assess which countries peg to the dollar. This approach focuses primarily on emerging markets.
framework for the international monetary system.

Second, my paper offers lessons for the literature explaining the carry trade, in which investors earn consistent returns for holding high-rate currencies over low-rate currencies. The international asset pricing literature typically explains its profits as compensation for bearing aggregate risk under complete markets. Within this framework, the explanations range widely: Verdelhan [2010] uses a model of habit, Farhi and Gabaix [2016] focus on rare disasters, Colacito and Croce [2011] present a framework of long-run risk, Hassan [2013] explains through variation in country size, and Ready et al. [2017] propose a setting with global differences in productivity. By contrast, the literature explaining carry trade returns through models of incomplete markets is newer and smaller. The main explanations are ones involving borrowing constraints as in Bruno and Shin [2017] or segmented markets as in Gabaix and Maggiori [2015]. My paper offers evidence from a well-identified setting that financial frictions explain asset returns better than models of risk-sharing. This informs the debate between complete and incomplete markets in explaining the overall carry trade.

The paper proceeds as follows. Section 2 reviews the literature on empirical spillover patterns. Section 3 discusses the empirical framework and the data. Section 4 introduces the main fact on asymmetries in currency and bond markets, and uses these asymmetries jointly to argue that neither central banks nor risk premia in complete markets react to the Fed. Section 5 shows further that central banks do not react to the Fed using term structures of foreign bond yields. Section 6 shows further that models of complete markets do not explain spillovers by showing the general modeling tensions. Section 7 discusses models with incomplete markets, and shows the divergence between spillovers emanating from the Fed and from other central banks. Section 8 concludes.

2 Literature Review on Spillovers

Beyond contributing to the two more theoretical debates identified in Section 1, I contribute to the much broader body of literature on documenting empirical patterns in spillovers. Papers here have found Fed spillovers in every conceivable asset. The most cited papers include Rey [2015] and Miranda-Agrippino and Rey [2015], which look at the Fed’s effects on a wide range of markets. However, there are many papers that look at more specific markets. To review a handful: Brusa et al. [2017] study equity markets, Fratzscher et al. [2017] study capital flows, Cetorelli and Goldberg [2012] and Morais et al. [2015] study bank liquidity and lending, and Gilchrist et al. [2016] study bond markets. All find strong spillovers emanating from the Fed.

However, my empirical findings are new to this literature in two regards. First, I characterize spillovers by their heterogeneous effects, and do so for each individual country. That allows me to link asymmetries in spillovers across currency and bond markets. By contrast, papers that do study heterogeneity in spillovers relate it to a single macroeconomic variable, which precludes linking asymmetries. (Moreover, most of these papers focus on emerging markets.) The two most consistent variables are proxies for a country’s fundamentals and measures of financial integration.
Georgiadis [2016], Chen and Chen [2012], Bowman et al. [2015], Mishra et al. [2014], Ahmed et al. [2015], and Aizenman, Binici and Hutchison [2016] for instance find that spillovers are muted when the recipient country has strong fundamentals. Hausman and Wongswan [2011], Miyajima et al. [2014], Eichengreen and Gupta [2015], and Aizenman, Chinn and Ito [2016] find that spillovers are stronger when recipient countries are more financially integrated with the US. Finally, Dedola et al. [2017] find no consistent macroeconomic indicators that explain heterogeneity.

Second, I study spillovers emanating from other central banks, in which there is far less work. The one exception is the ECB’s spillovers into European countries both inside and outside the Eurozone, which has been studied by Jardet and Monks [2014], Kucharcukova et al. [2016], Horvath and Voslarova [2017], McQuade et al. [2015], Ciarlone and Colabella [2016], and Bluwstein and Canova [2016]. However, only Fratzscher et al. [2016] and Kim and Nguyen [2009] study the ECB’s effects on non-European countries too, as I do. Beyond the ECB, coverage drops quickly. Craine and Martin [2008] study the effects of the Reserve Bank of Australia on American equities, and Gerko and Rey [2017] look at the effects of Bank of England spillovers on the US. Finally, Rogers et al. [2016] and Aizenman, Chinn and Ito [2016] briefly study spillovers from the Eurozone, Japan, and the UK.

3 Empirical Framework and Data

This section introduces the empirical framework and the data used to identify and characterize monetary spillovers. The empirical framework answers whether currencies and bonds react to Fed announcements in asymmetric ways. In this section, I outline the main equation and its components, explain the methodology used to identify the equation, and discuss how the data on Fed announcements and asset returns are collected.

The core components of the empirical framework are high-frequency windows, long-maturity assets, non-announcement windows, and inferred (i.e. latent) monetary shocks. These four components improve on the existing approaches in the literature in establishing how causal, comprehensive, and precise estimates. Existing approaches often measure asset returns over low-frequency windows, use short-maturity assets, fail to utilize non-announcement windows to correct for background noise, or use explicit but flawed measures of shocks.

High-frequency windows have two uses: they allow for causal interpretations, and they provide power. My windows are predominantly sixty-minute windows, although I use daily windows to replace any intraday windows with poor liquidity. These are improvements on methods that use exclusively daily windows or windows of lower frequencies. First, low-frequency windows run the risk that non-monetary news come out during the window, and so asset returns could reflect extraneous information. High-frequency windows ensure that assets are reacting to Fed announcements and not other global shocks, although I also explicitly check for other known shocks (including announcements from other central banks, and labor market and inflation releases).³ Second, low-

³Details of overlapping announcements and releases are provided in appendix A.
frequency windows are dominated by excessive idiosyncratic fluctuations, and this makes it hard to distinguish the effects of Fed announcements from noise. High-frequency windows mitigate this concern.

Long-maturity assets, namely countries’ ten-year sovereign bonds and currencies, are essential for capturing the entirety of the Fed’s effects in global markets. Particularly in the last decade, Fed announcements explicitly provide guidance over the moderate horizons, and it is possible that foreign central banks or investors respond in turn at even longer horizons. Measuring reactions to Fed announcements in short-maturity assets would not capture all changes in the paths of rates or in risk premia. However, ten-year bonds and currencies capture changes in these over a ten-year horizon and an infinite horizon, respectively.

Non-announcement windows (also known as non-event windows), which serve as reference points for announcement windows, keep my estimates conservative. Asset returns respond continuously to small idiosyncratic market shocks (e.g. market flows), and methods that do not use non-announcement windows would incorrectly ascribe these routine fluctuations to Fed statements, as event studies do. I observe asset returns through non-announcement windows, i.e. windows of equal duration on other days, to identify the ordinary variance and covariance of assets. As a result, only asset movements during announcement windows that exceed the ordinary patterns in non-announcement windows are linked to monetary policy.

Inferred shocks (i.e. latent factors) are important for precise identification. My shocks are estimated through a filtration of the asset returns themselves, rather than measured from observed data (e.g. movements in the Fed Funds futures). Observed shocks are problematic because assets in different currencies, in different asset classes, and at different maturities respond differentially to announcements. Regressing a given response asset on a distantly-related explanatory one weakens precision and can lead to incorrect inferences. Regardless, my inferred shocks correlate well with many observed measures of shocks, and my results are more noisily estimated but qualitatively unchanged when using measured shocks.

3.1 Framework

Before introducing the method or data, I first introduce the core equation that identifies asset responses to monetary announcements. I then provide details on the four key components of my empirical strategy: high-frequency returns, long-maturity assets, non-announcement windows, and inferred shocks.

3.1.1 Equation

Equation (1) is the main equation that I estimate. In this equation, I decompose a univariate or multivariate vector of long-maturity asset returns at time \( t \) \( (r_t) \) into a univariate or multivariate vector of constants \( \alpha \), the product of a univariate or multivariate vector of coefficients \( \beta \) and a univariate monetary shock \( m_t \), and a univariate or multivariate error \( \epsilon_t \). I primarily test whether assets react differentially to monetary shocks, i.e. test for equality between different elements of
the multivariate vector $\beta$. In Section 5, I also test for whether assets react to monetary shocks in the first place ($\beta = 0$).

$$r_t = \alpha + \beta m_t + \epsilon_t$$ (1)

I set $m_t$ to be a single-dimensional latent shock. This choice is made for both technical and practical reasons. From a technical point of view, a single factor explains the vast majority of asset returns. From a practical point of view, this choice largely eliminates the need to find an optimal factor rotation. A standard limitation of factor analysis is that $\beta m_t$ in Equation (1) is identified only up to a rotation, but there is only a single rotation of $(-1)$ with a single-dimensional shock. I thus normalize my shock to be positive during a monetary tightening, as defined when the dollar appreciates in specifications with currencies and when Treasury yields rise in specifications with bonds.

In this equation, the only observed data are the high-frequency and long-maturity asset returns $r_t$ during announcement windows. However, I let the errors $\epsilon_t$ take on a distribution learned from non-announcement windows, rather than assuming them to be homoskedastic white noise. Finally, $m_t$ refers to the inferred monetary shock. I discuss each of these sequentially.

### 3.1.2 High-Frequency Returns

High-frequency returns $r_t$ around announcements are essential for two reasons: arguing causality and identifying with power. For most assets, I measure returns from the fifteen minutes before a monetary announcement to the forty-five minutes after. For assets with poor intraday liquidity, I use daily windows, in which I measure asset returns from market close to market close over the day of the announcement.

Measuring returns at high frequencies is important for a causal interpretation of my results. As Bernanke [2017] notes, a major concern in the monetary spillovers literature is that asset reactions over lower-frequency windows do not measure reactions to monetary shocks, but rather to common global shocks. This concern is mitigated by using sixty-minute and daily windows, in which other global shocks are unlikely to dominate monetary shocks. I further address this concern by using non-announcement windows, and I also check explicitly for overlapping inflation releases, labor market releases, and monetary announcements from other central banks.

Moreover, high-frequency returns are important for power. In Table 1, I test whether I can detect Fed announcements in currency returns over different windows (as currency markets are open around-the-clock). Windows up to twelve hours long show consistent evidence of monetary shocks, but power drops quickly beyond the twelve-hour threshold. This drop-off in power is less stark in two cases: bond returns and domestic assets. Bonds exhibit approximately half the volatility as currencies (as they are driven primarily by shocks in one and not two countries), and so

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4The leading eigenvectors of currency returns and bond returns in sixty-minute windows around Fed announcements explains 89% and 96% of the variation respectively (whereas the second eigenvectors explain less than 5% each). Moreover, a parallel analysis procedure formally selects one factor for each specification.
retain power at daily frequencies. Moreover, domestic assets (e.g. the dollar) respond very strongly to Fed announcements, and so would also retain power at lower frequencies – although this paper largely focuses on foreign and not domestic responses to monetary policy.

Table 1: Power across Window Lengths

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Notes: The table tests whether currency returns (measured against an equal-weighted basket) by the column country are more volatile in windows around Fed announcements than in other windows of equal length, where the length of windows varies by row. If returns are statistically more volatile at the 1% level, the table records a checkmark. Intraday windows are needed to detect the Fed’s effects on foreign currency markets.

3.1.3 Long-Maturity Instruments

Long-maturity instruments are important for capturing the full reactions of monetary policy and of risk premia to Fed announcements. There is abundant evidence that the Fed guides monetary policy at moderate horizons in the US, e.g. the example in Figure 1 in which the Fed committed in early 2012 to policy actions until 2014. Paths of policy rates or risk premia in other countries may similarly respond at moderate or even longer horizons. For instance, Miranda-Agrippino and Rey [2015] and Rogers et al. [2016] estimate the largest responses to the Fed by the ECB, Bank of Japan, and Bank of England appear at the one to three year horizon, although these findings are not statistically significant at the 95% confidence level. Thus, reactions in short-maturity instruments around Fed announcements would likely miss much of its effects and be unable to show the channels of Fed spillovers conclusively.

I use two assets for most of this paper: exchange rates and ten-year bonds. In Equation (2), I show the asset pricing equation for currencies, using the dollar versus currency $j$. Returns in a given currency reflect changes in the paths of interest rates in the US and in country $j$ over the infinite horizon, plus changes in currency premia over the infinite horizon (or investors’ relative willingness to hold one currency over the other). In addition, currencies are driven by the infinite-horizon exchange rate, although I argue both in Section 4.2 and in Appendix C that this does not react to Fed announcements.

\[
\Delta s_j^t/\$ = \Delta E_t \sum_{k=1}^{\infty} i_{t+k-1}^j - \Delta E_t \sum_{k=1}^{\infty} i_{t+k-1}^i - \Delta E_t p_{t,\infty}^j + \Delta E_t s_{t,\infty}^j
\]  (2)
In Equation (3), I show the asset pricing equation for bond yields. Changes in yields on ten-year bonds reflect changes in the paths of interest rates over those ten years, plus changes in term premia over ten years (or investors’ relative willingness to hold long-maturity versus short-maturity assets).\(^5\)

\[
\frac{\Delta y_t^j(t, t + 10)}{10Y \text{ Yield}} = \Delta \mathbb{E}_t \sum_{k=1}^{10} \frac{\Delta r_{t+k-1}^j}{10Y \text{ Path of Rates}} + \frac{\Delta \mathbb{E}_t \gamma_{t,t+10}^j}{10Y \text{ Term Premia}}
\]  

(3)

The alternative to using long-maturity instruments would be to identify the effects of Fed announcements on foreign short-term rates in a vector autoregression framework. This is indeed the approach taken by Miranda-Agrippino and Rey [2015] and Rogers et al. [2016], but it suffers from power limitations. Given the general noise in financial markets, ascribing movements in interest rates to announcements two years ago is difficult. Moreover, I estimate a vector autoregression model in Appendix B following Wright [2012]. This maps closely to my empirical framework on all other dimensions, as it incorporates non-announcement windows, high-frequency returns, and inferred shocks, but it utilizes short-maturity rather than long-maturity instruments. I similarly find no reactions at conventional significance levels, but I take this as evidence of low power rather than as evidence of the channels behind Fed spillovers.

### 3.1.4 Non-Announcement Windows

Non-announcement windows, which are used to parameterize error term \(\epsilon_t\) in Equation (1), serve as reference points for announcement windows and are essential for ensuring that my results do not ascribe background noise during announcement windows to the Fed’s effects. I identify non-announcement windows the same way that I identify my announcement windows, as a combination of sixty-minute and daily windows. For each announcement, I measure eight non-announcement windows: the subsequent and preceding four Mondays through Thursdays. Chordia et al. [2001] find that Fridays have lower liquidity than other weekdays, and Cieslak et al. [2016] find that extraneous monetary news (through speeches by central bank governors, for instance) drops in the week preceding and following a monetary announcement. Of course, non-announcement windows that overlap with major news shocks, such as announcements by other central banks, are excluded. As an illustration, I show the yen and the Australian dollar against the dollar over one such non-announcement window in Figure 2, exactly one week prior to the introduction’s example announcement.

This approach exemplifies the spirit of Rigobon [2003], Rigobon and Sack [2003], and Rigobon and Sack [2004], which seeks to measure not variation over announcement windows as event studies do, but \textit{excess} variation over announcement windows. As with high-frequency returns, using non-announcement returns helps me establish a causal relationship between monetary announcements

---

\(^{5}\)For some countries, coupon bonds are more liquid at intraday frequencies than zero-coupon bonds. As a result, the effective duration on those bonds may be slightly shorter than ten years. For instance, futures markets often trade ten-year bonds with a 6% coupon, paid semi-annually. The duration on those assets is approximately eight years, and thus still represents a valid long-maturity asset.
and assets. Even the narrowest high-frequency window around monetary announcements have to contend with some idiosyncratic market fluctuations. Rather than ascribing such noise to monetary policy, I use non-announcement windows to establish those fluctuations as a baseline and look only for unusual responses.

3.1.5 Inferred Shocks

Inferred shocks or latent shocks, which are represented by $m_t$ in Equation (1), are used in lieu of measured shocks throughout this paper, and they are important for precise identification. Conventional measures of monetary shocks either fail to capture the entire monetary surprise (as relevant to currencies and foreign bonds) or capture it noisily, whereas inferred shocks do not suffer from either limitation.

In the literature, the most popular type of measured shock uses data from the US rates market, either at the short end (e.g. the Fed Funds futures market) or at the medium end (e.g. the one-year or two-year Treasury yield). The former approach is clearly problematic, as most Fed announcements in the past decade primarily revealed information about the path of future rates than about the imminent target. The latter approach mitigates this concern by using longer-maturity assets, but even that approach can miss the Fed’s full effects on currency and foreign bond markets. First, Fed announcements may have effects at even longer horizons than two years. Second, international assets likely react to different components of Fed announcements compared
to domestic assets, and so regressing international assets on domestic ones could lead to imprecise estimates and incorrect inferences (e.g. concluding that assets react symmetrically when they in fact react asymmetrically). Inferred shocks avoid the potential maturity mismatch and the extra layer of noise, by filtering for shocks that match the maturity and character of the response variables. Indeed, I replicate my results using the two-year Treasury yield in lieu of inferred shocks, and I find that my results are qualitatively the same but more noisily estimated. Those results are reported in Appendix B.

Inferred shocks do not come without costs of their own. First, they do not disentangle the Fed’s direct and indirect effects: they capture foreign assets reacting both to the changes in the path of domestic rates and to the changes in domestic premia. Second, they do not have units, making the coefficients generated by regressing foreign asset returns on them similarly free of units. (I can only estimate the units of the product of shocks and coefficients, and cannot disentangle the two.) For papers concerning the operations of monetary policy, these are fatal shortcomings: such questions strive to estimate the exact pass-through of Fed direct policy on foreign yields. For my paper, which tests whether assets respond at all and whether assets respond symmetrically or asymmetrically to a given announcement (regardless of its underlying size or components), these costs are small.

Finally, one additional advantage of inferred shocks over measured shocks is that they can be constructed at an intraday frequency for any central bank, which I study in Section 7. This is not noteworthy for the Fed and for other large central banks, but it is useful when looking at the central banks of smaller countries (e.g. the Reserve Bank of New Zealand or Norges Bank), which do not have liquid equivalents for the Fed Funds futures market or medium-term sovereign bond futures markets.

3.2 Methodology

To identify asymmetries, or whether a given asset responds more or less than another asset to Fed announcements, I find the maximum likelihood estimates of both \((\alpha, \beta)\) and \(m_t\) in the multivariate version of Equation (1). This is akin to estimating a factor model, or to estimating a model with interactive fixed effects, in which shocks vary across time and loadings vary across countries. In this section, I explain the estimation procedure and the two ways of identifying asymmetries in \(\beta\).

3.2.1 Estimation Procedure

The method explains the excess variation in announcement windows over non-announcement windows as a combination of time-varying shocks and asset-varying coefficients. To illustrate the estimation procedure, suppose I want to test whether Fed shocks pass symmetrically into three currency pairs: the euro, the pound, and the yen, all measured against the dollar. I write Equation
(1) in its multivariate form:

\[
\begin{bmatrix}
\Delta s_t^{\text{e}/\text{$/t$}} \\
\Delta s_t^{\text{f}/\text{$/t$}} \\
\Delta s_t^{\text{r}/\text{$/t$}}
\end{bmatrix} =
\begin{bmatrix}
\alpha^{\text{e}/\text{}$/t$} \\
\alpha^{\text{f}/\text{}$/t$} \\
\alpha^{\text{r}/\text{}$/t$}
\end{bmatrix} +
\begin{bmatrix}
\beta^{\text{e}/\text{}$/t$} \\
\beta^{\text{f}/\text{}$/t$} \\
\beta^{\text{r}/\text{}$/t$}
\end{bmatrix} m_t^+$
\begin{bmatrix}
\epsilon_t^{\text{e}/\text{}$/t$} \\
\epsilon_t^{\text{f}/\text{}$/t$} \\
\epsilon_t^{\text{r}/\text{}$/t$}
\end{bmatrix}
\]

(4)

I write the likelihood function associated with Equation (4) next. As before, the errors $\epsilon$ are assumed to have some covariance matrix $\Sigma$ learned from non-announcement windows, rather than being homoskedastic white noise, and so this likelihood function resembles the one a generalized least squares methodology optimizes.\(^6\)

\[
\max_{\alpha,\beta,\{m_t\}} -\frac{1}{2T} \sum_{t=1}^{T} \left[ (\Delta s_t - \alpha - \beta m_t)^T \Sigma^{-1} (\Delta s_t - \alpha - \beta m_t) \right]
\]

Since this term involves the product of estimated quantities $m_t$ and $\beta$, I cannot analytically solve the system of interlocking first-order conditions.\(^7\) Instead, I use the Expectation-Maximization algorithm. The approach alternately takes the expectation of log-likelihood function with respect to the monetary shocks $m_t$ and then maximizes the expression with respect to the parameters. Convergence is guaranteed since the EM algorithm improves the likelihood function on every iteration. This yields estimates for parameters $(\alpha, \beta, m_t)$.\(^8\)

In Appendix B, I discuss a more general procedure that handles complications arising from partially-missing data. In my paper, partially-missing data are a major concern for bond markets, which (unlike currency markets) are not always liquid and are not open around-the-clock. Dropping partially-missing observations would cut my sample dramatically, as at least one or two markets are illiquid or closed during any given announcement. Ignoring high-frequency bond returns in favor of daily bond returns (which are almost never missing) would reduce my power substantially. Instead, I take two steps to maintain power. First, I construct each term in my log-likelihood function as a function only of the data available at that time. Second, I incorporate both sixty-minute and daily windows concurrently when a specification has particularly severe issues with missing data, although I restrict the coefficients for any given asset to be the same across windows of different lengths. This ensures that all non-missing data is utilized.

\(^6\)I assume the errors have zero mean in this specification, but I demean the data by the non-announcement means first in practice. These non-announcement means are extremely close to zero.

\(^7\)Moreover, I cannot solve the system iteratively since convergence is neither guaranteed in theory nor achieved in practice.

\(^8\)Rohde and Cappe [2011] among many others argue that computing the marginal likelihood of continuous latent factors $m_t$ is intractable, and instead recommend a modified expectation step that uses a variational posterior distribution for $m_t \sim N(\mu_t, V_t)$. Specifically, I take the expectation of the log-likelihood function with respect to $m_t(\mu_t, V_t)$ initially, and also augment the log-likelihood function with the Kullback-Leibler divergence between the variational posterior and the prior $N(0, 1)$. I then maximize the expression alternately with respect to parameters $(\mu_t, V_t)$ and $(\alpha, \beta)$, which maps to the original two steps. When conducting robustness checks with the shocks themselves, I approximate for the shocks $m_t$ using the means of their posterior distributions $\mu_t$. Further details can be found in Appendix B.
I compute standard errors for \((\alpha, \beta)\) by bootstrap, sampling the set of monetary announcements and their associated asset responses with replacement. There are no analytic solutions for standard errors, given the adjustments with missing data. In Appendix B, I discuss alternative approaches to estimating Equation (4) and their limitations, including Identification by Heteroskedasticity by Rigobon [2003].

I next turn to testing asymmetries in \(\beta\) in Equation (4).

### 3.2.2 Average Coefficient

A natural way to test asymmetries in \(\beta\) would be a series of pairwise tests. This is reasonable if \(\beta\) has two or three elements, but is incomprehensible in practice as \(\beta\) has nine elements (which involves thirty-six pairwise tests). Moreover, this may yield qualitatively inconsistent results. To illustrate, consider an example in which I find significant evidence that \(\beta^{E/\$} > \beta^{Y/\$}\), but I cannot reject \(\beta^{E/\$} \neq \beta^{L/\$}\) or \(\beta^{E/\$} \neq \beta^{Y/\$}\). At least one of these tests must be wrong.

Instead, I test for asymmetries in a closely related way by comparing each element of \(\beta\) to an average of the other \((n - 1)\) elements. In this example, the test looking for asymmetries in how the euro reacts to Fed announcements, relative to how the yen or pound reacts, becomes:

\[
H_0 : \beta^{E/\$} = \frac{1}{2} \left( \beta^{L/\$} + \beta^{Y/\$} \right)
\]

### 3.2.3 Lower-Dimensional Structure

An alternate way to test for asymmetries is to cast the elements of \(\beta\) (and \(\alpha\)) to a lower-dimensional structure, in which different assets are encouraged to share coefficients unless they respond too differentially from each other.

To illustrate, consider the opening example of the Fed announcement on January 25, 2012. In Figure 3, I add the New Zealand dollar to the original plot of currencies. Visually, the monetary shock passes into the Australian dollar and New Zealand dollars comparably, but differently into the yen.

I formally test this by estimating Equation (4) with a lower-dimensional structure, in which assets with similar responses share coefficients. To find this optimal structure, I compute the (extended) Bayesian Information Criterion, a model selection criterion, for each possible permutation of assets sharing coefficients, and take the structure that scores best. As an illustration, if the estimates for \(\beta^{L/\$}\) are much closer to \(\beta^{Y/\$}\) than to \(\beta^{E/\$}\), one possible structure that may emerge is

---

9 Appendix B is more exhaustive, but the main limitation of Identification by Heteroskedasticity is that its solution algorithm, GMM, is not guaranteed to converge in high-dimensional spaces whereas the EM algorithm is. I do successfully implement it when utilizing currency data in Equation (1) and find very similar results to those generated by the EM algorithm. I find convergence to be an issue when utilizing bonds data.

10 Regardless, I show the p-values for all thirty-six pairwise tests in Appendix F for my main results.

11 This problem is closely related to clique cover problems in graph theory. However, since the number of assets is small, I iterate through every permutation without needing approximate algorithms, such as LASSO.
Figure 3: Currency Reactions to the Fed Easing, January 25, 2012

Notes: The figure depicts the reactions of three currencies in sixty-minute windows around the Fed’s surprise easing of January 25, 2012: the yen, the Australian dollar, and the New Zealand dollar, all measured against the dollar. The Australian and New Zealand dollars appreciate similarly (approximately 1%), while the yen appreciates by substantially less (60 basis points). There is little difference between treating the Australian and New Zealand dollars as distinct currencies or collapsing them into one currency area, but there is a big difference for the yen.

as follows:

\[
\begin{bmatrix}
\Delta s_t^{\text{E}/S} \\
\Delta s_t^{\text{£}/S} \\
\Delta s_t^{\text{¥}/S}
\end{bmatrix}
= \begin{bmatrix}
1 & 0 & 1 \\
0 & 1 & 0 \\
0 & 1 & 0
\end{bmatrix}
\begin{bmatrix}
\alpha^{\text{E}/S} \\
\alpha^{\text{£}/(¥, £)/S} \\
\beta^{\text{E}/S}
\end{bmatrix}
+ \begin{bmatrix}
1 & 0 & 1 \\
0 & 1 & 0 \\
0 & 1 & 0
\end{bmatrix}
\begin{bmatrix}
\beta^{\text{£}/(¥, £)/S} \\
\beta^{\text{¥}/S}
\end{bmatrix}
+ \begin{bmatrix}
e_t^{\text{E}/S} \\
e_t^{\text{£}/S} \\
e_t^{\text{¥}/S}
\end{bmatrix}
\]

Like other model selection criteria, the extended Bayesian Information Criterion trades off the likelihood of a given structure against penalties for the structure’s dimensionality, i.e. the number of coefficients needed. The optimal specification thus forces assets which react to Fed shocks similarly to share coefficients, as the improvements in the likelihood function from asset-specific coefficients are dwarfed by the penalties imposed for the higher dimensionality. Similarly, the optimal specification allows assets which react to Fed shocks asymmetrically to have different coefficients, as the resulting losses in the likelihood function are much greater than the savings from lower penalties. The extended Bayesian Information Criterion is more conservative than the widely-used Akaike Information Criterion and regular Bayesian Information Criterion, as it penalizes dimensionality more severely. Chen and Chen [2012] and Foygel and Drton [2011] recommend using these more conservative approaches when the number of parameters in the model is high, given the elevated risk of overfitting. Their recommendations seem particularly appropriate for my specification, which estimates both the parameters of the shocks and coefficients.

Thus, the optimal structure breaks assets into groups, in which assets react to Fed announcements similarly to other members of their group, but dissimilarly to assets in other groups.
3.3 Data

The two core pieces of data for this paper are the exact timestamps of Fed announcements from 2001 - 2016, and high-frequency and daily currency and bond returns across ten countries. This section provides details on these pieces of data, and on additional pieces of data occasionally used in the paper.

3.3.1 Monetary Announcements

I gather the 128 scheduled monetary announcements following Fed Open Market Committee meetings made from 2001 - 2016, in which the Fed announces the upcoming Fed Funds target and guidance about future targets. I do not incorporate unscheduled announcements (e.g. following September 11) to avoid cases in which the Fed may be releasing news about fundamentals concurrently with monetary news. I also exclude announcements made during the depths of the financial crisis, from September 2008 until March 2009.

Nakamura and Steinsson [2017] note that monetary announcements may actually be informational announcements, releasing the Fed’s private information about fundamentals. I offer two defenses. First, this interpretation yields a simple prediction: the market should digest Fed announcements similarly to the fundamentals announcements, such as the Bureau of Labor Statistics’ unemployment reports. In the appendix, I document starkly different patterns of asset asymmetries between Fed and BLS announcements. For instance, the yen and Canadian dollar have similar reactions to each other following Fed announcements, and strongly different reactions from each other following BLS announcements. Second, these post-FOMC announcements still represent the cleanest possible sources of monetary news. Speeches by Fed governors or releases of FOMC minutes, while informative about monetary policy, run greater risks of releasing private information too. Statements following FOMC meetings are succinct and brief, and designed to give monetary guidance only.

In addition, I collect the regularly scheduled rate announcements by the nine central banks of the nine other countries in my sample: the Reserve Bank of Australia, the Bank of Canada, the Swiss National Bank, the European Central Bank, the Bank of England, the Bank of Japan, the Norges Bank, the Reserve Bank of New Zealand, and the Riksbank. I employ my methodology to check for monetary spillovers for these central banks, and report the results in Section 7. Details can be found in the appendix.

3.3.2 Asset Returns

For exchange rates, I collect minute-by-minute currency data from the foreign exchange brokers Forexite and Olsen Data. Missing data are largely not prevalent, as currency markets are open and liquid for these currencies 24 hours per day, five days per week. Per the 2016 BIS Triennial Survey, the currencies of the ten countries in my sample constitute ten of the eleven most liquid floating currencies (along with the Mexican peso).
For bonds, I collect two types of data. First, I collect high-frequency ten-year bond futures through Thomson Reuters, listed on the various futures exchanges around the globe.\textsuperscript{12} For countries without bonds on a liquid futures exchange, I use an intraday benchmark rate published by Thomson Reuters based on reported transactions. However, some of these returns are illiquid around Fed announcements.

Second, I collect zero-coupon bond yields from Datastream, compiled by the world’s largest brokerage firm ICAP. These are measured at the daily frequency, and cover a cross-section of ten annual maturities (one-year to ten-year) in addition to twenty-year and thirty-year maturities for Australia, Canada, Germany, Japan, Switzerland, the United Kingdom, and the United States.

In addition to pruning observations that risk overlapping with inflation, unemployment, or foreign monetary releases, I also prune extreme observations, defined as returns over announcement windows whose distance from mean announcement returns exceed the most extreme 1% threshold. I specifically measure Mahalanobis distance, which generalizes Euclidean distance to multivariate and correlated data.

4 Currency and Bond Asymmetries

I examine asymmetries in currency markets and in bond markets following Fed announcements to show that monetary spillovers from the Fed are neither consistent with shifts in the paths of central banks, nor consistent with basic shifts in risk premia under complete markets. First, I introduce a new fact that establishes how currencies and bonds respond asymmetrically to monetary shocks from the Fed. Second, I show that these two classes of explanations are inconsistent with the fact by combining the results from currencies and bonds.

This section has two major contributions. First, I identify asymmetries in the responses of currencies and ten-year sovereign bonds to Fed announcements, using high-frequency currency returns and a mixture of high-frequency and daily bond returns across my ten countries. This is a new fact to the literature in its own right, and shows that spillovers are heterogeneous across countries. Asymmetric responses across countries can be subtle, and common methods in the literature that utilize noisier shocks or lower-frequency returns will miss differences as discussed in Section 2.

The results are that when the Fed tightens, the dollar appreciates more against currencies of high-rate countries (e.g. the Australian dollar) than currencies of low-rate countries (e.g. the Japanese yen); and ten-year bond yields of high-rate countries rise more than bond yields of low-rate countries. This is unique among central banks. In Section 7, I apply this methodology to other central banks and find that most central banks, with the exception of the European Central Bank, do not generate asymmetric spillovers. In Appendix F, I show that this finding is robust to different time periods (pre-crisis and post-crisis) and different states (recessionary and expansionary).

\textsuperscript{12}Examples include the Sydney Futures Exchange for Australian data, the Chicago Mercantile Exchange for American data, the Eurex Exchange for Swiss and German data, the London International Financial Futures and Options Exchange for British data, the Osaka Securities Exchange for Japanese data, etc.
The second major contribution of this section is to show that this fact is inconsistent with two major classes of explanations: central banks reacting to the Fed, and risk premia shifting in a complete markets framework. For both of these explanations, results from currency and bond markets suggest opposing stories. If central banks follow the Fed, then currency markets predict that low-rate countries tighten with the Fed, while bond markets predict that high-rate countries tighten with the Fed. If risk premia shift in a complete markets framework, then currency predict suggest that the stochastic discount factors of low-rate countries rise when the Fed tightens, while bond markets predict that the stochastic discount factors of high-rate countries rise when the Fed tightens.\footnote{In this section, I only consider transitory movements in stochastic discount factors. If the stochastic discount had only permanent movements, bond yields would never move, and so that case can be ignored. In Section 6, I consider a more complex framework with both transitory and permanent fluctuations in the stochastic discount factor.}

I formalize this argument by constructing a portfolio in which a US-based investor shorts a foreign long-maturity bond. This portfolio, which combines currencies and bonds, has two properties. First, it has no exposure to foreign monetary policy. Second, it has no exposure to transitory shifts in foreign complete markets-based premia. Even if the Fed causes foreign central banks to adjust their paths of policy rates or if the Fed triggers shifts in foreign risk premia, these portfolios should be insensitive to such movements. In fact, I show that this portfolio is highly sensitive to Fed announcements, proving that neither explanation can explain my results.

4.1 Baseline Fact: Asymmetric Asset Responses

Before testing the classes of explanations for spillovers, I first present the baseline fact, in which Fed announcements affect currencies and bonds of different countries differentially. In currencies, when the Fed tightens (eases), the dollar appreciates (depreciates) by most against high-rate currencies like the Australian dollar and least against low-rate currencies like the Japanese yen. In bonds, when the Fed tightens (eases), ten-year bond yields of high-rate countries rise (fall) more than yields of low-rate countries.

I utilize the Expectation Maximization algorithm to fit the multivariate version of Equation (1). I replicate the example with three currencies here. To recap the methodology, I fit country-specific coefficients \((\alpha, \beta)\) and time-varying shocks \(m_t\), where \(\epsilon_t\) is assumed to take on the distribution of non-announcement windows rather than being homoskedastic white noise. To identify asymmetries, I compute standard errors with respect to the average of other coefficients, and I look for a lower-dimensional structure that groups similar coefficients together.

\[
\begin{bmatrix}
\Delta s_{t}^{E/S} \\
\Delta s_{t}^{L/S} \\
\Delta s_{t}^{Y/S}
\end{bmatrix}
= 
\begin{bmatrix}
\alpha^{E/S} \\
\alpha^{L/S} \\
\alpha^{Y/S}
\end{bmatrix} 
+ 
\begin{bmatrix}
\beta^{E/S} \\
\beta^{L/S} \\
\beta^{Y/S}
\end{bmatrix} m_{t}^{S} 
+ 
\begin{bmatrix}
\epsilon_{t}^{E/S} \\
\epsilon_{t}^{L/S} \\
\epsilon_{t}^{Y/S}
\end{bmatrix}
\]

The following two figures show the results. I estimate Equation (1) separately for currencies using sixty-minute returns, and for bonds using a mixture of sixty-minute and daily returns, although
I estimate them jointly in the next section.

Consider currencies first. I plot the coefficients $\beta$ in Figure 4. These coefficients refer to the appreciation or depreciation against various currencies when the dollar appreciates or depreciates by 1% on average. In this figure, the dollar appreciates most against the Australian and New Zealand dollars, the Norwegian krone, and the Swedish krona when the Fed tightens, and appreciates least against the Japanese yen, the Canadian dollar, and the British pound. Both the standard errors and the lower-dimensional structure support this conclusion.

Figure 4: Currency Responses to US Monetary Shocks

Notes: The figure depicts by how much the dollar appreciates against a given reference currency when it appreciates by 1% on average, following a Fed tightening. Standard error bars are computed versus the average appreciation of 1%; and the shading of the coefficient bars refers to the lower-dimensional structure, whereby the dollar appreciates similarly against currencies of the same color and dissimilarly against currencies of different colors. The dollar appreciates by little against the Japanese yen among other currencies, and by a lot against the Australian dollar among other currencies.

Next consider bonds, in which coefficients $\beta$ are plotted in Figure 5. These coefficients refer to the annualized ten-year sovereign bond yield returns by country when the US yield rises by 1%. When the Fed tightens and US yields rise by 1%, Swiss and Japanese yields respond little and rise by 0.1-0.3%, while Australian and New Zealand yields respond strongly and rise by over 1%. Again, both the standard errors and the lower-dimensional structure support this conclusion.

These asymmetric responses point to one of three classes of explanations: central banks react to the Fed asymmetrically, premia react to the Fed asymmetrically under assumptions of complete markets, or premia react to the Fed asymmetrically under assumptions of incomplete markets. I rewrite the asset pricing equations for currencies and ten-year bonds, Equations 2 and (3) to illustrate this. Since the US components are common across currencies, differential appreciation and depreciation of the dollar points to differential movements by foreign central banks or in
Figure 5: Bond Responses to US Monetary Shocks

Notes: The figure depicts by how much yields of a given country’s ten-year bonds rise when US ten-year yields rise by 1%, following a Fed tightening. Standard error bars are computed versus the average rise in foreign yields; and the shading of the coefficient bars refers to the lower-dimensional structure, whereby bonds of the same color rise similarly when US yields rise and bonds of different colors rise dissimilarly when US yields rise. Australian yields among other bonds rise a lot, and Japanese yields among other bonds rise little, when US yields rise.

Currency premia. Differential responses in yields similarly point to differential movements by foreign central banks or in term premia.

\[
\Delta s_t^{j/S} = \Delta E_t \sum_{k=1}^{\infty} i_t^{j+k-1} - \Delta E_t \sum_{k=1}^{\infty} i_t^{j+k-1} + \Delta E_t^j p_t^{j/S} + \Delta E_t^j s_t^{j/S}
\]

\[
\Delta y_t^{j/(t, t+10)} = \Delta E_t \sum_{k=1}^{10} i_t^{j+k-1} + \Delta E_t^j y_{t,t+10}^{j}
\]

The Fed’s asymmetric spillovers into currency and bond markets are special among central banks. In Section 7 and in Appendix 7, I discuss other central banks in more detail, but I consider asymmetric spillovers from the Reserve Bank of New Zealand (RBNZ) here. I show that the Reserve Bank of New Zealand has no asymmetric effects on either currencies or bonds in Figure 6. Spillovers are only slightly differential into Australia, as Figure 6 shows that Australian yields react slightly to RBNZ announcements, and the NZD appreciates or depreciates slightly less against the Australian dollar following RBNZ announcements.

\footnote{I address differential movement in the infinite-horizon exchange rate in Appendix C.}
Figure 6: Market Reactions to NZ Monetary Shocks

(a) Currency Responses

(b) Bond Responses

Notes: The figures depict the reactions of currency and bond markets to announcements by the Reserve Bank of New Zealand. The left figure shows by how much the NZD appreciates against a given reference currency when it appreciates by 1% on average; and the right figure shows by how much foreign ten-year yields rise when New Zealand ten-year yields rise by 1%. Standard error bars in both pictures are computed versus the average reaction across currencies or foreign bonds; and the shading of the coefficient bars refers to the lower-dimensional structure, whereby assets of the same color react similarly and assets of different colors react dissimilarly following RBNZ announcements. The NZD appreciates symmetrically against all currencies and foreign yields do not rise when the RBNZ tightens, with the limited exception of Australian assets.

Although the empirical specifications are computed for each country, I confirm that the cross-sectional sorting of asymmetries by countries maps closely to the level of interest rates. In Figures 4 and 5, the canonical low-rate countries of Japan and Switzerland are found together, while the canonical high-rate countries of Australia, New Zealand, and Norway are found together. I test this formally by interacting the specification in Equation 1 with each country’s interest rate spread relative to the US, and by removing currency-specific parameters. Equation 5 shows the revision. I test whether $\beta_1$ in this model is significant, i.e. whether interest rates help predict variation across countries and across time in currency and bond returns. I find that the spread in interest rates at all maturities – from one-month rates to ten-year rates – are significant at the 5% level in all specifications. In Appendix D, I show that this result is robust to dropping any country from the specification.

$$\begin{bmatrix}
\Delta s_t / \epsilon_t \\
\Delta \ell_t / \ell_t \\
\Delta s_t / \ell_t
\end{bmatrix} = 
\begin{bmatrix}
1 & \ell_t - i_t \\
1 & \ell_t - i_t \\
1 & \ell_t - i_t
\end{bmatrix}
\begin{bmatrix}
\alpha_0 \\
\alpha_1 \\
\beta_1
\end{bmatrix} + 
\begin{bmatrix}
\epsilon_t / \epsilon_t \\
i_t / \ell_t \\
i_t / \ell_t
\end{bmatrix}
\begin{bmatrix}
\beta_0 \\
\beta_1
\end{bmatrix}
+ 
\begin{bmatrix}
\epsilon_t / \epsilon_t \\
i_t / \ell_t \\
i_t / \ell_t
\end{bmatrix}
$$

Interest rates are one of several significant predictors, and I test several others in Appendix D: local measures of equity and currency volatility, the cross-currency basis, trade flows, shares of trade flows invoiced in dollars, cross-border bank positions, cross-border portfolio debt positions,
cross-border portfolio equity positions, and cross-country distances. Among these, four are also statistically significant in all specifications: measures of currency skew extracted from currency options, cross-border bank positions, cross-border equity portfolio positions, and trade flows.

However, I focus on interest rates because of their relevance to risk premia and because of their ubiquity in the literature. Since Hansen and Hodrick [1980] and Fama [1984] first documented the excess returns to the carry trade, in which investors borrow in low-rate currencies and lend in high-rate currencies, a large body of work has posited various explanations for risk factors that correlate with the level of interest rates. For instance, Lustig and Verdelhan [2007] argue that aggregate consumption risk correlate with rates, Lettau et al. [2014] suggest that state-dependent market exposures correlate with rates, Colacito et al. [2017] argue that exposure to global shocks correlate with rates. Every paper in this literature correlates their explanation with the level of rates, either by assumption or endogenously, and so I do too.

4.2 Test: Cross-Border Bond Portfolio

With the baseline fact on asymmetries established, I use this fact to show that Fed spillovers are driven neither by shifts in the reactions of central banks, nor by basic shifts in risk premia under complete markets. I do so by constructing portfolios that combine the currency and bond of a given country in a way that removes dependence altogether on its central bank’s path of short rates, and that offsets its country-specific currency and term premia under complete markets. Specifically, a US-based investor shorts a long-maturity foreign bond in these portfolios. Under the asset pricing equations for currencies and bonds, these portfolios should be symmetrically exposed to the Fed if the underlying source of the asymmetry was one of these two explanations. Empirically, the portfolios have very strongly asymmetric exposure to the Fed. This suggests that fluctuations in risk premia under incomplete markets explain spillovers better.

I construct portfolios in which a US-based investor shorts a long-maturity foreign bond and invests in the US riskfree rate, and thus bears foreign currency and foreign interest rate risk. I show these statements mathematically in the next paragraph, but I first present the intuition of why these portfolios are agnostic to both foreign central banks and foreign shifts in risk premia. Suppose the portfolio is with respect to Japan, such that the investor shorts a ten-year Japanese bond and invests at the US riskfree rate. If the Bank of Japan tightens, the yen appreciates versus the dollar, and Japanese bond prices fall. The investor thus loses money through the portfolio’s currency exposure but makes money through the portfolio’s interest rate exposure. If the Japanese stochastic discount factor has a transitory positive innovation, again the yen appreciates and Japanese ten-year bond prices fall. Again, the investor makes money on one component of the portfolio and loses money on the other component. The two components offset perfectly under some weak assumptions.

To construct this portfolio, I take the following two steps. First, I add Equation (2), the decomposed exchange rate equation, to Equation (3), the decomposed long-maturity bond yield equation. This creates a portfolio that is equivalent to shorting a foreign long-maturity bond and investing in the US riskfree rate. Second, I assume that both the infinite-horizon exchange rate and the
path of foreign interest rates after a ten-year horizon are constant through Fed announcements. In Appendix C, I do a variety of tests to verify this assumption. Under models of long-run monetary neutrality, this assumption simplifies to arguing that expectations of foreign inflation do not react to the Fed at long horizons.\footnote{While monetary neutrality is only asymptotic, Gopinath [2015] and Carvalho and Dam [2010] survey the international literature and find that the median price duration across countries is approximately one year, and that most firms adjust prices at least once every two years.}  I find no evidence from foreign inflation-linked bonds to suggest otherwise, and indeed surveyed expectations of long-run inflation are so persistent that they fluctuate as much in an entire year as my assets do in a given sixty-minute window. Moreover, I check that my results are robust to using thirty-year bonds and to examining ten-year bonds in the pre-crisis era, as these are tests in which there is no risk that changes in monetary policy exceed the maturity of my bond instrument.

\[
\Delta s^j_t/s^l_t (l) = \Delta s^j_t/s^l_t + \Delta y^j_t(t, t + 10) \\
\approx \Delta \mathbb{E}_t \sum_{k=1}^{\infty} \tau^j_{t+k-1} + \Delta \mathbb{E}_t \mathbf{P}_{t, \infty}^{j/\infty} + \Delta \mathbb{E}_t \mathbf{P}_{t, t+10}^{j/10} \\
\text{Portfolio Return} \quad \text{Currency Exposure} \quad \text{Foreign Bond Exposure} \quad \text{US Path of Rates} \quad \text{Currency Premia} \quad \text{Foreign Term Premia}
\]

(6)

First, it is immediately apparent from Equation (6) that this portfolio does not depend on foreign monetary policy. Asymmetries in the cross-section of how foreign central banks react to the Fed do not generate asymmetries in the cross-section of these portfolios.

Second, although it is not immediately apparent from Equation (6), shifts in a country’s term premia offset with shifts in its share of currency premia. To show this, I rewrite Equations (2) and (3), or the asset decomposition equations for currencies and bonds, in terms of stochastic discount factors. Under complete markets, the exchange rate return is exactly the difference of innovations to the log stochastic discount factors.

\[
\Delta s^j_t/s^l_t = \Delta m^j_t - \Delta m^l_t \\
\text{Exchange Rate} \quad \text{US SDF} \quad \text{Foreign SDF}
\]

(7)

Moreover, changes in long-maturity bond yields are equal to the difference of two components: changes in the contemporaneous log stochastic discount factor and changes in expectations of the log long-run pricing kernel, when bonds pay off.

\[
\Delta y^j_t(t, t + 10) = \Delta m^l_t - \Delta \mathbb{E}_t \lambda^j_{t+10} \\
\text{10Y Yield} \quad \text{SDF} \quad \text{10Y Pricing Kernel}
\]

(8)

As such, I take the same two steps to construct the portfolio in this revised framework. First, I again add Equation (7) to Equation (8). Second, I assume that expectations of the pricing kernel at long horizons do not react to the Fed. Lustig et al. [2017] argue that there is indeed no heterogeneity in this component across countries using ten-year bonds; and as before, my results are robust using
thirty-year bonds. In Section 6, I relax this assumption altogether.

\[
\begin{align*}
\text{Portfolio Return} & = \Delta s_{t}^{j/s} + \Delta y_{t}^{j}(t, t + 10) \\
& \approx \Delta m_{t}^{S} \text{US SDF}
\end{align*}
\]

(9)

I illustrate with the example of power utility to make Equations (7) and (8) more tangible. Suppose the representative investor in country \(i\) has power utility with a stationary autoregressive log consumption process, and receives a shock \(\epsilon_{t}\) from the Fed’s announcement. The exchange rate equation (7) can be rewritten with this specific stochastic discount factor. In this expression, positive innovations to country \(i\)’s stochastic discount factor, which are equivalent to negative innovations in its consumption, cause the dollar to depreciate versus its currency. Equivalently, as the foreign consumption basket becomes more dear, its relative price rises.

\[
\Delta s_{t}^{j/s} = \left( -\gamma^{S} \Delta c_{t}^{S} \right) - \left( -\gamma^{j} \Delta c_{t}^{j} \right)
\]

The long-maturity bond yield equation (8) can also be rewritten with this specific stochastic discount factor. In this expression, transitory positive innovations to country \(i\)’s stochastic discount factor, which are equivalent to negative innovations in its consumption, cause yields to rise. When the marginal utility of country \(i\) is temporarily elevated, the country smooths away that variation by borrowing from the future, causing yields at long maturities to rise. The one mitigating factor is whether the country still expects to have elevated marginal utility in ten years time, in this example. If so, investors prefer to borrow at longer maturities. However, this requires substantial persistence of innovations, and I again invoke the results from both Lustig et al. [2017] about ten-year bonds and robustness from my own results at thirty-year horizons.

\[
\Delta y_{t}^{j}(t, t + 10) = \left( -\gamma^{j} \Delta c_{t}^{j} \right) + \left( \rho \right)^{10} \sigma^{j} c_{t}^{j} \approx \left( -\gamma^{j} \Delta c_{t}^{j} \right)
\]

Put together, if the representative investor in a foreign country receives a negative consumption shock, the consumption basket of that country becomes more expensive and so its currency relatively appreciates. At the same time, the representative investor borrows from the future to offset the shock and so its ten-year bond yields rise. These effects mathematically offset when added together.\(^{16}\)

Thus, I have two equivalent representations of this portfolio. Equation (6) makes clear that these portfolios do not depend on foreign monetary policy. Equation (9) makes clear that these

\(^{16}\)In this exposition, I disregard predictable movements in currencies and bonds. In practice, these predictable movements are close to zero over high frequencies, and anyways I use returns measured over non-announcement windows to control for such predictable movements.
portfolios do not depend on foreign stochastic discount factors under a complete markets framework. In short, I should not expect to find any heterogeneity in these portfolios across countries under either explanation. These portfolios should react symmetrically to Fed announcements.

I test to see if portfolios respond unequally to the Fed. Equation (4) becomes in this context:

\[
\begin{bmatrix}
\Delta s_t^e/\$, + \Delta y_t^e(t, t+10) \\
\Delta s_t^£/\$, + \Delta y_t^£(t, t+10) \\
\Delta s_t^Y/\$, + \Delta y_t^Y(t, t+10)
\end{bmatrix}
= \begin{bmatrix}
\alpha^e/\$ \\
\alpha^£/\$ \\
\alpha^Y/\$
\end{bmatrix} + \begin{bmatrix}
\beta^e/\$ \\
\beta^£/\$ \\
\beta^Y/\$
\end{bmatrix} m_t^\$ + \begin{bmatrix}
\epsilon_t^e/\$ \\
\epsilon_t^£/\$ \\
\epsilon_t^Y/\$
\end{bmatrix}
\]

I plot the coefficients \( \beta \) in Figure 7, and find that the portfolios are unequally responsive to the Fed. The portfolio with Japan reacts differently than the portfolios with Australia and New Zealand, while the portfolios of the other six countries exhibit smaller but still substantial asymmetries among themselves. The reactions of central banks cannot explain asymmetries in assets following Fed spillovers. The reactions of stochastic discount factors in a complete markets framework cannot explain asymmetries in assets following Fed spillovers either. Those asymmetries must stem from adjustments in risk premia, under an incomplete markets framework.

**Figure 7: Cross-Border Bond Portfolio Responses to US Monetary Shocks**

Notes: The figure depicts by how much a portfolio that shorts a given country’s ten-year bond and lends at the US riskfree rate rises when the average portfolio rises by 1%, following a Fed tightening. Standard error bars are computed versus the average portfolio rise of 1%; and the shading of the coefficient bars refers to the lower-dimensional structure, whereby portfolios of the same color react similarly and portfolios of different colors react dissimilarly following Fed announcements. The portfolio of Australian assets rise a lot, while the portfolio of Japanese assets rise little, when the Fed tightens.

Finally, although this test is a standalone one, I compare the coefficients in this plot to the coefficients for the portfolio that shorts the foreign riskfree bond, rather than the foreign long-
maturity bond. This portfolio only has exchange rate risk. I compare coefficients in Figure 8. Interestingly, the point estimates for the long-maturity portfolio become more asymmetric compared to the short-maturity portfolio; although the standard error bars also widen, such that both tests yield approximately the same degree of asymmetry once adjusting for noise. This illustrates the point that asymmetries deepen when currencies and bonds are linked, rather than being offset.

Figure 8: Cross-Border Portfolio Responses to US Monetary Shocks

(a) Short-Maturity Bond Portfolio  
(b) Long-Maturity Bond Portfolio

Notes: The figures compare the reactions of two types of portfolios to Fed announcements. The left figure shows by how much a portfolio that shorts a given country’s riskfree bond and lends at the US riskfree rate, i.e. the short-maturity portfolio, rises when the average portfolio rises by 1%. The right figure shows by how much a portfolio that shorts a given country’s ten-year bond and lends at the US riskfree rate, i.e. the long-maturity portfolio, rises when the average portfolio rises by 1%. Standard error bars are computed versus the average portfolio rise of 1%; and the shading of the coefficient bars refers to the lower-dimensional structure, whereby portfolios of the same color react similarly and portfolios of different colors react dissimilarly following Fed announcements. The responses of long-maturity portfolios are comparably or more heterogeneous than the responses of short-maturity portfolios. This duplicates Figures 4 and 7, except with common axes for comparison purposes.

5 Responses in Bond Yield Curves

Section 4 shows that asset reactions in currency and bond markets are inconsistent with a general explanation of central banks reacting to the Fed. However, this does not preclude explanations in which a few central banks react to the Fed while most countries see shifts in risk premia, or explanations in which central banks partially react to the Fed. In this section, I offer evidence for each country on its own, to show that the country’s risk premia rather than its path of short rates reacts to Fed announcements. I use data from a rich term structure of bond yields for each country. This test cannot distinguish between shifts in risk premia under complete or incomplete markets, and so it is only used to argue against explanations in which central banks react.

---

17The number of statistically significant pairwise differences among the short-maturity portfolios is almost exactly equal to that number among the long-maturity portfolios.
I make the argument in two ways: through a simplified approach and through an affine term structure model. First, I show that the parts of a country’s yield curve most dominated by short rates do not react to Fed announcements; while the parts of a country’s yield curve most dominated by term premia do react to Fed announcements. This approach relies on the fact that short yields are largely driven the paths of short-term policy rates, while distant forward yields are largely driven by term premia. Second, I use a Gaussian affine term structure model to decompose a country’s yields explicitly into the path of short rates and term premia, and I show again that term premia do react to Fed announcements while the paths of rates do not.

I augment the analysis in one crucial way, to mollify concerns that this approach lacks power. I show that the paths of short rates in foreign countries respond to announcements from their own central banks. While I cannot entirely rule concerns that short rates actually do react to the Fed but are too noisy to detect, this augmentation shows that at least some expected movements in short rates can be detected. Moreover, in doing so, I document new facts about the divergence between central banks. The Fed has effects on term premia around the globe, whereas most central banks do not. I return to this point further in Section 7.

Since the questions in this section are about the existence of reactions (i.e. whether a given bond yield reacts to the Fed or not), I utilize a methodology called inference by heteroskedasticity. This maintains the same core empirical principles of high-frequency returns around Fed announcements, long-maturity instruments, non-announcement windows, and inferred shocks. However, it requires fewer assumptions than the previous section as I answer a simpler question: whether bond yields react or do not react, rather than whether they react symmetrically or asymmetrically. I introduce this methodology first.

5.1 Methodology: Inference by Heteroskedasticity

The method employed in this section, Inference by Heteroskedasticity, is a method that uses fewer assumptions than before to test the existence of reactions: whether a given bond responds at all to Fed monetary shocks. Essentially, the method deduces that a bond reacts to Fed announcements if the variance of bond returns during announcement windows exceeds the variance during non-announcement windows.

The methodology only relies on the assumption that, in the absence of Fed announcements, announcement windows would have the same distribution as non-announcement windows. Beyond that, it is robust to misspecification on the dimensionality of Fed shocks; and it is agnostic to whether shocks into a given bond yield are highly idiosyncratic to that yield or common across multiple yields. Finally, it is transparent, as the test gets its power from one moment alone.

The test compares the variance of changes for one yield during announcement and non-announcement windows, and tests whether the former is larger. To illustrate, suppose I want to test whether Australian bond yields react to Fed announcements. I write Equation (1) and its non-announcement
counterpart as follows, where all parameters are unidimensional for simplicity.

\[
\text{Announcement Windows: } \Delta y_t^{\text{AUD}} = \alpha + \beta m_t^S + \epsilon_t \\
\text{Non-announcement Windows: } \Delta \tilde{y}_t^{\text{AUD}} = \epsilon_t
\]

This methodology tests \( H_0 : \beta = 0 \), or whether Australian bond yields do or do not respond to Fed shocks. To answer this, I simply take the variance of both sides and link the two equations through the variance of the error.

\[
\text{Announcement Windows: } \mathbb{V}(\Delta y_t^{\text{AUD}}) = \beta^2 \mathbb{V}(m_t^S) + \mathbb{V}(\epsilon_t) \\
\text{Non-announcement Windows: } \mathbb{V}(\Delta \tilde{y}_t^{\text{AUD}}) = \mathbb{V}(\epsilon_t) \\
\mathbb{V}(\Delta y_t^{\text{AUD}}) > \mathbb{V}(\Delta \tilde{y}_t^{\text{AUD}}) \implies \beta \neq 0 \tag{10}
\]

If Australian bond yields react to Fed announcements (if \( \beta \neq 0 \)), those yields should be more volatile around Fed announcements than otherwise. Although the F-test is the most common test for equality of variances, I employ the Brown-Forsythe test to mitigate concerns about non-normal distributions. The Brown-Forsythe test looks at median absolute deviations, rather than mean squared deviations as the F-test does. In Appendix B, I show that this test strongly outperforms the F-test on simulated data with high kurtosis.

This section almost entirely utilizes daily data, as bond yields are not priced at sixty-minute intervals for many countries, with the exception of the benchmark ten-year yield.

5.2 Baseline Fact: Spillovers in Yields

Before decomposing the term structure of bond yields into the paths of rates and term premia, I first demonstrate that foreign ten-year bond yields react to the Fed and react to their own central banks. In addition, some foreign bond yields react to the ECB. This is the only table to use intraday data.

In Table 2, I test Equation (10) for each country’s ten-year bond returns over each central bank’s announcements. Rows refer to central banks, and columns to that country’s local ten-year bond. Consider the example in the previous section, of whether Australian bond yields react to Fed announcements. I report the result in the bottom-left corner of Table 2. I present the ratio of excess standard deviations when statistically significant at the 1% level, and leave the cell blank if not statistically significant. In this example, Australian yields are 207% more volatile around Fed announcements than they are at that time otherwise.

\[
\frac{\sigma(\Delta y_t^{\text{AUD}})}{\sigma(\Delta \tilde{y}_t^{\text{AUD}})} - 1
\]

The results indicate that the Fed has strong effects on most bond markets. The ECB also has strong effects, although the magnitudes are smaller, possibly because its announcements are less
Table 2: Excess Volatility in 10Y Bond Returns

<table>
<thead>
<tr>
<th></th>
<th>AUD</th>
<th>CAD</th>
<th>CHF</th>
<th>EUR</th>
<th>GBP</th>
<th>JPY</th>
<th>NOK</th>
<th>NZD</th>
<th>SEK</th>
<th>USD</th>
</tr>
</thead>
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<tr>
<td>Australia</td>
<td>144%</td>
<td></td>
<td></td>
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<td></td>
<td></td>
<td></td>
<td></td>
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<tr>
<td>Canada</td>
<td>30%</td>
<td>82%</td>
<td></td>
<td></td>
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<tr>
<td>Switzerland</td>
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<td></td>
<td></td>
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<tr>
<td>Euro</td>
<td>34%</td>
<td>39%</td>
<td>88%</td>
<td>37%</td>
<td>133%</td>
<td>90%</td>
<td>24%</td>
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<td></td>
<td></td>
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<tr>
<td>United Kingdom</td>
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<td></td>
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<td>Japan</td>
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<td></td>
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<tr>
<td>Norway</td>
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<td></td>
<td></td>
<td></td>
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<td>82%</td>
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<tr>
<td>United States</td>
<td>207%</td>
<td>144%</td>
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<td>216%</td>
<td>46%</td>
<td></td>
<td></td>
<td>54%</td>
<td></td>
<td>233%</td>
</tr>
</tbody>
</table>

Notes: The table tests whether the ten-year bond of the column country is more volatile in the sixty minutes around announcements by the row central bank than in other sixty-minute intervals. (Daily windows are used in light grey, if returns are illiquid at intraday frequencies.) If returns are more volatile at the 1% level using the Brown-Forsythe test, the cell records the excess ratio of standard deviations (announcement window to non-announcement window standard deviations, minus 100%). If returns are not statistically more volatile, the cell is left blank. The full table, including ratios significant at the 5% level and insignificant ratios, can be found in the Appendix as Table 12. The Fed and ECB have spillover effects, but most other central banks only affect their own ten-year bonds.

surprising. Few other central banks have systematic effects, which is a finding that is new in the monetary spillovers literature but consistent with results in other international finance settings.

As before, bond returns can be decomposed into two components: changes in the paths of short rates or changes in the term premia that investors demand for those bonds. I now turn to identifying which component in Equation (3) (rewritten here) is largely responsible.

\[
\Delta y_{jt}^{10Y}(t,t+10) = \Delta E_t^{10Y} \sum_{k=1}^{10} \hat{y}_{t+k-1}^j + \Delta E_t^{10Y} \gamma_{t+10}^j \\
\text{10Y Yield} \quad \text{10Y Path of Rates} \quad \text{10Y Term Premia}
\]

5.3 Test 1: Extremities of Yield Curve

With the baseline fact established in bond markets, I next utilize a transparent approach to argue that Fed spillovers are driven by shifts in term premia, rather than shifts in the reactions of central banks. I look at the extremities of the yield curves for my ten countries, and argue that the short end is driven by central bank policy while the long end is driven by term premia. Fed spillovers affect the long ends of yield curves across the globe, while the short ends of these curves respond to their own countries’ monetary policies.

Before showing the results, I first explain the logic. Consider a typical yield curve. Short yields (e.g. one-year) have little term premia. Several papers, such as Hamilton [2009] show this formally, by noting that compensation for maturity risk shrinks to zero as the maturity of the bond shrinks.
Thus, movements in short yields are largely driven by fluctuations in local monetary policy.

\[ \Delta y^j_{t}(t, t + T_0) = \left( \Delta \sum_{k=1}^{T_0} E_t i^j_{t+k-1} + \Delta E_t \gamma^j_{t,t+T_0} \right) \approx \Delta E_t \sum_{k=1}^{T_0} i^j_{t+k-1} \]

On the other end, movements in long-maturity forward yields (future rates that can be guaranteed today) are driven primarily by term premia, not by changes in the paths of short rates. Most New Keynesian models find long-run monetary neutrality with real rates, as nominal rigidities are reversed over time. In Section 4 and in Appendix C, I argue that long-run inflation forecasts are extremely stable over time and estimates of inflation from foreign inflation-linked bonds do not react to the Fed. As a result, I argue that long-run conditional monetary neutrality holds with nominal rates over Fed announcement windows. This assumption does not need to hold precisely, but only needs to hold approximately. Estimates by Adrian et al. [2013] suggest that over 80% of variation in US long forward yields on Fed announcement days are driven by adjustments in term premia.

\[ \Delta y^j_{t}(T_1, t + 10) = \left( \Delta E_t \sum_{k=1}^{10-T_1} i^j_{T_1+k-1} + \Delta E_t \gamma^j_{T_1,t+10} \right) \approx \Delta E_t \gamma^j_{T_1,t+10} \]

There is a tradeoff in power and in contamination in selecting short and long cutoffs \((T_0, T_1)\). For short yields, changes in the paths of rates beyond \(T_0\) are not captured; but setting \(T_0\) to be too distant means that these yields include term premia. For long forward yields, changes in term premia before \(T_1\) are not captured; but setting \(T_1\) too close means that those yields include changes in the paths of rates. Adrian et al. [2013] note that 83% of the variation in the US one-year yield on Fed announcement days is driven by the path of rates, and so I set \(T_0\) to be 1 year. For forward yields, the authors estimate that a forward rate that starts between five and six years from now has approximately 83% of its variation driven by term premia on Fed announcement days. I thus set \(T_1\) to 6 years, and focus on the six-year ahead, four-year yield (i.e. a yield that can be locked into in 2017 to borrow and lend between 2023 to 2027).

I utilize the Inference by Heteroskedasticity method to measure how the measures of foreign yields react to announcements by the Fed (and by other central banks). The results are presented in Tables 3 and 4.

The findings are stark. Table 3 shows that the Fed does not affect other countries’ paths of short-term policy rates, while the central banks of those countries do. Table 4 shows that the Fed has strong effects on other countries’ term premia, while the central banks of other countries do not. This points to a story in which term premia, rather than the policies of central banks, adjust to announcements by the Fed for each individual country.

The results are robust to different cutoffs. I illustrate with one extreme example: ten-year forward twenty-year rates, or rates that can be locked into in 2017 for borrowing and lending between 2027 and 2047. It is implausible that central banks regularly release guidance at such horizons. Table 5 shows the results for countries that issue thirty-year bonds.
### Table 3: Excess Volatility in Daily 1Y Bond Returns

<table>
<thead>
<tr>
<th>Country</th>
<th>AUD</th>
<th>CAD</th>
<th>CHF</th>
<th>EUR</th>
<th>GBP</th>
<th>JPY</th>
<th>NOK</th>
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<tr>
<td>Switzerland</td>
<td>108%</td>
<td>36%</td>
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<tr>
<td>Euro</td>
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<tr>
<td>New Zealand</td>
<td>53%</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td>133%</td>
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<tr>
<td>Sweden</td>
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<td></td>
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<td></td>
<td></td>
<td>79%</td>
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<tr>
<td>United States</td>
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</tbody>
</table>

Notes: The table tests whether the one-year bond of the column country is more volatile around announcements by the row central bank than at other times, using daily returns. If returns are more volatile at the 1% level using the Brown-Forsythe test, the cell records the excess ratio of standard deviations (announcement window to non-announcement window standard deviations, minus 100%). If returns are not statistically more volatile, the cell is left blank. The full table, including ratios significant at the 5% level and insignificant ratios, can be found in the Appendix as Table 13. Central banks affect their own one-year bonds, but the Fed does not affect other countries’ one-year bonds.

### Table 4: Excess Volatility in Daily 6F4Y Bond Returns

<table>
<thead>
<tr>
<th>Country</th>
<th>AUD</th>
<th>CAD</th>
<th>CHF</th>
<th>EUR</th>
<th>GBP</th>
<th>JPY</th>
<th>NOK</th>
<th>NZD</th>
<th>SEK</th>
<th>USD</th>
</tr>
</thead>
<tbody>
<tr>
<td>Australia</td>
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<tr>
<td>Canada</td>
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<tr>
<td>Switzerland</td>
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<tr>
<td>Euro</td>
<td>19%</td>
<td>32%</td>
<td></td>
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<td></td>
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<tr>
<td>United Kingdom</td>
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<tr>
<td>Japan</td>
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<td></td>
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<td></td>
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<tr>
<td>Norway</td>
<td>30%</td>
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<tr>
<td>New Zealand</td>
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<tr>
<td>Sweden</td>
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<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>United States</td>
<td>38%</td>
<td>36%</td>
<td>30%</td>
<td>50%</td>
<td>29%</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Notes: The table tests whether the six-year forward four-year bond of the column country is more volatile around announcements by the row central bank than at other times, using daily returns. If returns are more volatile at the 1% level using the Brown-Forsythe test, the cell records the excess ratio of standard deviations (announcement window to non-announcement window standard deviations, minus 100%). If returns are not statistically more volatile, the cell is left blank. The full table, including ratios significant at the 5% level and insignificant ratios, can be found in the Appendix as Table 14. The Fed affects most other countries’ six-year forward four-year bonds.
The Fed once again has strong effects, affecting three other yields at a 1% level, and Australian and British yields at a 5% level as documented in Appendix F. This confirms that Fed announcements induce strong shifts in term premia for each country.

Table 5: Excess Volatility in Daily 10F20Y Bond Returns

<table>
<thead>
<tr>
<th>Country</th>
<th>AUD</th>
<th>CAD</th>
<th>CHF</th>
<th>EUR</th>
<th>GBP</th>
<th>JPY</th>
<th>USD</th>
</tr>
</thead>
<tbody>
<tr>
<td>Australia</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td>54%</td>
</tr>
<tr>
<td>Canada</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Switzerland</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Euro</td>
<td>32%</td>
<td>36%</td>
<td>32%</td>
<td></td>
<td></td>
<td></td>
<td>51%</td>
</tr>
<tr>
<td>United Kingdom</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Japan</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>United States</td>
<td>32%</td>
<td>36%</td>
<td>32%</td>
<td></td>
<td></td>
<td></td>
<td>51%</td>
</tr>
</tbody>
</table>

Notes: The table tests whether the ten-year forward twenty-year bond of the column country is more volatile around announcements by the row central bank than at other times, using daily returns. If returns are more volatile at the 1% level using the Brown-Forsythe test, the cell records the excess ratio of standard deviations (announcement window to non-announcement window standard deviations, minus 100%). If returns are not statistically more volatile, the cell is left blank. The full table, including ratios significant at the 5% level and insignificant ratios, can be found in the Appendix as Table 15. This rate cannot be constructed for New Zealand, Norway, and Sweden as they do not issue thirty-year bonds, and so they are omitted. The Fed affects many other countries’ ten-year forward twenty-year bonds.

5.4 Test 2: Affine Term Structure Model

I next use an affine term structure to decompose yield curves explicitly across different countries into their two components: paths of rates and term premia. I again show that the foreign premia react to the Fed, while the paths of rates react only to their own central banks.

This approach complements the previous approach for two reasons. First, it captures the entire paths of rates and the entire term premia, whereas the previous approach could only examine fractions of those quantities. Second, because the approach estimates the price of risk directly from the curvature and co-movement in nominal yields, it does not require long-run monetary neutrality or constant inflation targets. Of course, it imposes alternate assumptions through its model structure.

The model I utilize is the five-factor Gaussian affine term structure model of Adrian et al. [2013]. This belongs to the class of models in which yields are affine in state variables, through setting the pricing kernel to be exponentially affine in shocks, setting prices of risk to be affine in state variables, and setting innovations to be Gaussian. In the original paper, this specific model fits the US yield curve only, but I apply it on my international yield curves.

I choose this model over other choices for two reasons. First, this model can conduct the decomposition at a daily frequency, in contrast to other models applied to international yield curves, such as in Wright [2011], that operate at monthly or quarterly frequencies. In this model, state variables are principal components of the yield curve itself, measured at daily frequencies;
whereas most other models use macro state variables, such as inflation or GDP, that are measured at lower frequencies. Second, this model incorporates five factors, which the authors argue offers substantial improvements over models with fewer factors. I apply the methodology almost exactly as described by the original paper, with only one small modification to estimate eigenvectors more robustly given data limitations for some countries (e.g., New Zealand). Details of the procedure are offered in Appendix E.

I again utilize the Inference by Heteroskedasticity method to measure how these foreign primitives react to announcements by the Fed (and by other central banks) at a daily frequency. The results are presented in Tables 6 and 7.

Table 6: Excess Volatility in Daily 10Y Rate Returns

<table>
<thead>
<tr>
<th>Country</th>
<th>AUD</th>
<th>CAD</th>
<th>CHF</th>
<th>EUR</th>
<th>GBP</th>
<th>JPY</th>
<th>NOK</th>
<th>NZD</th>
<th>SEK</th>
<th>USD</th>
</tr>
</thead>
<tbody>
<tr>
<td>Australia</td>
<td>97%</td>
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<td></td>
<td></td>
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<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Canada</td>
<td></td>
<td>82%</td>
<td></td>
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<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Switzerland</td>
<td>28%</td>
<td>91%</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Euro</td>
<td></td>
<td>38%</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>United Kingdom</td>
<td>44%</td>
<td></td>
<td></td>
<td></td>
<td>92%</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td>30%</td>
</tr>
<tr>
<td>Japan</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td>92%</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td>21%</td>
</tr>
<tr>
<td>Norway</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td>132%</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>New Zealand</td>
<td>35%</td>
<td></td>
<td></td>
<td></td>
<td>102%</td>
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<tr>
<td>Sweden</td>
<td></td>
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<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td>127%</td>
</tr>
<tr>
<td>United States</td>
<td></td>
<td>40%</td>
<td></td>
<td></td>
<td>30%</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td>97%</td>
</tr>
</tbody>
</table>

Notes: The table tests whether the model-estimated ten-year path of rates of the column country is more volatile around announcements by the row central bank than at other times, using daily returns. If returns are more volatile at the 1% level using the Brown-Forsythe test, the cell records the excess ratio of standard deviations (announcement window to non-announcement window standard deviations, minus 100%). If returns are not statistically more volatile, the cell is left blank. The full table, including ratios significant at the 5% level and insignificant ratios, can be found in the Appendix as Table 16.

The Fed affects few other countries’ estimated paths of rates.

While the results are noisier, they show the same basic trends: term premia across the globe react to the Fed, while countries’ paths of rates react primarily to their own central banks. This confirms that term premia drive monetary spillovers emanating from the Fed for all countries in my sample, and that the class of explanations around central banks reacting should be discarded.

6 Models of Complete Markets

Section 4 shows that asset reactions in currency and bond markets are also inconsistent with simple stories in which stochastic discount factors react to the Fed under complete markets. However, this does not preclude more complex stories in which the stochastic discount factor has multiple forms of heterogeneity. In this section, I address that by proving the restrictions that more complex stochastic discount factors must obey to match my results, both in two commonly used models and in a general and preference-free framework. These restrictions are unusual if not impossible to
Table 7: Excess Volatility in Daily 10Y Term Returns

<table>
<thead>
<tr>
<th>Country</th>
<th>AUD</th>
<th>CAD</th>
<th>CHF</th>
<th>EUR</th>
<th>GBP</th>
<th>JPY</th>
<th>NOK</th>
<th>NZD</th>
<th>SEK</th>
<th>USD</th>
</tr>
</thead>
<tbody>
<tr>
<td>Australia</td>
<td>29%</td>
<td></td>
<td></td>
<td></td>
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<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Canada</td>
<td></td>
<td>38%</td>
<td>58%</td>
<td></td>
<td></td>
<td></td>
<td>61%</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Switzerland</td>
<td></td>
<td></td>
<td>61%</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Euro</td>
<td></td>
<td>28%</td>
<td></td>
<td>21%</td>
<td></td>
<td></td>
<td></td>
<td></td>
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<td></td>
</tr>
<tr>
<td>United Kingdom</td>
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<td></td>
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<tr>
<td>Japan</td>
<td></td>
<td></td>
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<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td>22%</td>
<td></td>
</tr>
<tr>
<td>Norway</td>
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<td></td>
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<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td>46%</td>
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<tr>
<td>New Zealand</td>
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<td></td>
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<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td>20%</td>
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<tr>
<td>Sweden</td>
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<td></td>
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<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td>46%</td>
</tr>
<tr>
<td>United States</td>
<td>27%</td>
<td>46%</td>
<td>25%</td>
<td>35%</td>
<td>32%</td>
<td>30%</td>
<td>17%</td>
<td>68%</td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Notes: The table tests whether the model-estimated ten-year path of term premia of the column country is more volatile around announcements by the row central bank than at other times, using daily returns. If returns are more volatile at the 1% level using the Brown-Forsythe test, the cell records the excess ratio of standard deviations (announcement window to non-announcement window standard deviations, minus 100%). If returns are not statistically more volatile, the cell is left blank. The full table, including ratios significant at the 5% level and insignificant ratios, can be found in the Appendix as Table 17. The Fed affects many other countries’ estimated paths of term premia.

satisfy in many standard models of complete markets.

The general tension from Section 4 remains valid. Asymmetries in the exchange rate market and bond rate market imply opposite conclusions about the stochastic discount factor. In this section, I use the results of Figures 4 and 5 (the asymmetries in currencies and in bonds) separately, rather than linking them together as in Figure 7 (the cross-border bond portfolio). Using currencies and bonds separately gives me two conditions, whereas using the combined portfolios gives me only one.

I show that asymmetries in the exchange rate market imply that the low-rate stochastic discount factor is more volatile than the high-rate one following Fed announcements; and asymmetries in the bond market imply the opposite. These tensions can be remedied only by making stochastic discount factors heterogeneous in complex and unusual ways, as a simple source of heterogeneity is insufficient. A useful organizing framework from Lustig et al. [2017] is to decompose the stochastic discount factor into two components: a permanent component and a transitory component, on which exchange rates and bonds load differentially. Under this decomposition, the low-rate permanent component must be made more volatile and the low-rate transitory component made less volatile than their high-rate counterparts. This gives the stochastic discount factors enough mathematical freedom to match my results, but economically these two restrictions are highly unusual and difficult to disentangle economically.

This result may shed light on broader debates in global markets. Most models explain international asset returns with models of complete markets that have one source of heterogeneity. For instance, Hassan [2013] argues that country size explains variation, and Japan has a more volatile stochastic discount factor than Australia because shocks to its consumption are harder to offset. Colacito et al. [2017] present an international model of long-run risk and suggest that Australia
loads less on global shocks than Japan. Farhi and Gabaix [2016] present an international model with rare disasters, and propose that Japan has stronger resilience during rare disasters than Australia. Finally, Ready et al. [2017] argue that commodity producers like Australia are less exposed to global shocks than producers of finished goods like Japan. Every one of these models predicts that Japanese stochastic discount factor is more volatile, matching results from the currency market but failing to match results from bond markets.

These models aim at matching results on the unconditional carry trade. However, I show that either more complex models of complete markets or models of incomplete markets are important for matching currency premia earned around Fed announcements. For instance, Lustig et al. [2017] incorporate multiple sources of heterogeneity in several classic models of complete markets. This may be more promising, although justifying those multiple sources ex ante are necessary too.

In this section, I first illustrate the tension between how currencies and bonds react to Fed announcements using a simple example with power utility. Second, I show the divergence in a fairly general and richly heterogeneous model with Epstein-Zin utility. Finally, I provide an organizing conceptual framework for these restrictions by decomposing the stochastic discount factor in a preference-free framework into its transitory and permanent components.

6.1 Example 1: Power Utility and Simple Dynamics

I illustrate the tension between currency movements and bond movements using power utility and a simple process for consumption. This follows closely the example in Section 4.2, although I generate conditions for currency and bond markets separately rather than combining them. In this framework, currency markets predict that the Japanese stochastic discount factor is more volatile, while bond markets predict that the Australian stochastic discount factor is more volatile. Although this section showcases a simple and real model, I show in subsequent sections that this tension remains with complex and nominal models.

In this framework, log consumption follows an AR(1) process. The shocks hitting Australia and Japan have different volatilities, but I restrict all other parameters \((\rho, \beta, \gamma)\) to be common. I relax this assumption later.

\[
\log C_{t+1}^i = \rho \log C_t^i + \sigma^i \epsilon_{t+1}^i \quad \text{where} \quad \rho \in [0, 1] \quad \text{and} \quad \epsilon_{t+1}^i \sim N(0, 1)
\]

I first consider currency markets, and define the exchange rate \(S\) to be yen per Australian dollars. Under complete markets and in the absence of arbitrage opportunities, there is a unique stochastic discount factor that prices any asset; and so the Australian stochastic discount factor and the Japanese stochastic discount factor equal one another having adjusted for exchange rates.

\[
\beta \left( \frac{C_{t+1}^A}{C_t^A} \right)^{-\gamma} = \beta \left( \frac{C_{t+1}^J}{C_t^J} \right)^{-\gamma} \frac{S_{t+1}}{S_t}
\]

As a result, exchange rate returns reflect the relative innovations to the stochastic discount
factor. Since the stochastic discount factors are conditionally lognormal, Backus et al. [2001] show that the excess currency return equals half the difference in variance between the log stochastic discount factors. Since Australia and Japan represent high-rate and low-rate countries, this excess return corresponds to the returns of the carry trade.

The carry trade earns positive returns, as has been established by a large body of literature starting with Hansen and Hodrick [1980] and Fama [1984]. Moreover, not only does the carry trade earn excess returns unconditionally, but Mueller et al. [2017] show that it earns excess returns through Fed announcements specifically.\footnote{This pattern holds up in my sample with varying significance, but I have a shorter sample and thus less power than Mueller et al. [2017].} This generates my first condition.

\[ \mathbb{E}_t \Delta s_{t+1} + r^A_{f,t} - r^J_{f,t} = \frac{1}{2} \gamma^2 ((\sigma^J)^2 - (\sigma^A)^2) > 0 \] (11)

I second consider long-maturity bond markets. The current value of a zero-coupon bond paying off at \( t + k \) is simply the expectation of the stochastic discount factor linking today with that future time period. I can again simplify this since the stochastic discount factor is conditionally lognormal.

\[ \mathbb{E}_t \left( \beta^k \left( \frac{C_{t+k}^i}{C^i_t} \right)^{-\gamma} \frac{1}{P^i_t(t+k)} \right) = 1 \implies \log P^i_t(t+k) = k \log \beta - \gamma (\mathbb{E}_t c_{t+k} - c_t) + \gamma^2 \nu c_{t+k} \]

Moreover, the change in yields in long-maturity zero-coupon bonds are simply the (inverse) difference in log prices.

\[ \Delta y^j_{t+1}(t+k) = \log P^i_t(t+k) - \log P^i_{t+1}(t+k) \]

Finally, I draw directly on evidence from Figure 5, in which the bond yields of high-rate countries are more volatile than the bond yields of low-rate countries around Fed announcements. This generates my second condition.

\[ \nu_t (\Delta y^j_{t+1}(t+k)) - \nu_t (\Delta y^A_{t+1}(t+k)) = \gamma^2 \left( 1 - \rho^{k-1} \right)^2 ((\sigma^J)^2 - (\sigma^A)^2) < 0 \] (12)

Equations (11) and (12) are exactly contradictory. The former condition requires Japanese shocks to be larger than Australian shocks, to match the excess returns investors earn for holding Australian assets over Japanese assets. The latter condition requires Australian shocks to be larger than Japanese shocks, to match the volatility in Australian yields that stem from Australian investors constantly readjusting their portfolios. These conditions continue to contradict each other even if parameters \((\beta, \gamma)\) are heterogeneous across countries.

The only possible resolution is for \( \rho \) to vary by countries. However, I argue that ten-year bonds approximate infinite-maturity bonds in my sample, such that \( \rho^k \approx 0 \). I rely upon two pieces of
First, Lustig et al. [2017] use the term structure model of Joslin et al. [2011] to argue that the approximation of ten-year to infinite-maturity bonds for the same sample of ten countries is reasonable. Second, in Appendix H, I find that the variance of changes in yields for thirty-year bonds is higher in Australia and other high-rate countries than in Japan and other low-rate countries.

6.2 Example 2: Epstein-Zin Utility and Complex Dynamics

I show that the tension between currency movements and bond movements persists, using a much richer model with Epstein-Zin utility and with a dynamic consumption process. In this model, I show that a single source of heterogeneity is insufficient. Two sources of heterogeneity are sufficient, but they are jointly implausible and require countries to have two countervailing forms of heterogeneity. For instance, to match the empirical results, Japan must be strongly exposed to idiosyncratic consumption shocks and but weakly exposed to trend consumption shocks from Fed announcements, while Australia must be weakly exposed to idiosyncratic shocks and strongly exposed to trend shocks.

In this model, consumption growth has both an idiosyncratic component and a persistent component; and the volatility of the shocks to these two components itself is stochastic. This nests many common models. Under some calibrations (e.g. $\phi = 0$ and $\sigma_w = 0$), this is the base case of Epstein-Zin utility; under others (e.g. $\phi = 0$), this is the model of stochastic volatility; and under yet others (e.g. high $\phi$ and $\rho$), this is the model of long-run risk by Bansal and Yaron [2004]. I present the utility functions and the consumption dynamics for country $i$, although I modify the consumption dynamics to incorporate heterogeneity next.

$$U_t(i) = \left((1 - \delta) C_t(i)^{1-1/\psi} + \delta E_t \left(U_{t+1}(i)^{(1-1/\psi)}((1-1/\psi)/(1-\gamma))\right)\right)^{1/(1-\gamma)}$$

$$\Delta c_{t+1} = \mu + \phi x_{t+1} + \sigma_t \eta_{t+1}$$

$$x_{t+1} = \rho x_t + \varphi e_{t+1}$$

$$(\sigma_{t+1}^2) = \sigma^2 + v \left( (\sigma_t^2 - \sigma^2) \right) + \sigma_w w_{t+1}$$

To incorporate heterogeneity across countries, I use a modeling innovation developed by Colacito et al. [2017] in the long-run risk literature. They decompose the shock $e_{t+1}$ into two independent components: a global component $e_{t+1}^z$ and an idiosyncratic component $e_{t+1}^i$. Different countries $i$ have differential loadings $1 + \beta_i$ on the global components of shocks.\(^{19}\) I utilize that modeling innovation, and in fact decompose all shocks ($\eta, e, w$) into both global and idiosyncratic components. Global components of ($\eta, e$) have constant global volatility, while idiosyncratic components continue to have idiosyncratic stochastic volatility. Finally, shocks decompose into these two components.

\(^{19}\)Without loss of generality, I restrict $1 + \beta_i \geq 0$. 

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with weighting \((\alpha_\eta, \alpha_e, \alpha_w)\). I present the updated dynamics.

\[
\Delta c_{t+1}^i = \mu + \phi x_t^i + (\sqrt{\alpha_\eta \sigma (1 + \beta_\eta^i)} \eta_{t+1}^i + \sqrt{1 - \alpha_\eta} \sigma_i^i \eta_{t+1}^i) \\
\Delta x_{t+1}^i = \rho x_t^i + \varphi_e (\sqrt{\alpha_e \sigma (1 + \beta_e^i)} \epsilon_{t+1}^i + \sqrt{1 - \alpha_e} \sigma_i^i \epsilon_{t+1}^i) \\
(\sigma_{t+1}^i)^2 = \sigma^2 + \nu \left( (\gamma^i_t)^2 - \sigma^2 \right) + \sigma_w (\sqrt{\alpha_w (1 + \beta_w^i)} w_{t+1}^i + \sqrt{1 - \alpha_w} w_{t+1}^i)
\]

Finally, as before, the empirical results in currencies require that returns in the Japanese stochastic discount factor \(m_t^J\) be more volatile than returns in the Australian one \(m_t^A\); and as before, the empirical results in bonds require changes in Australian yields to be more volatile than changes in Japanese yields.

\[
\mathbb{V}_t (\Delta m_{t+1}^J) > \mathbb{V}_t (\Delta m_{t+1}^A) \quad \text{and} \quad \mathbb{V}_t (\Delta y_{t+1}^A) > \mathbb{V}_t (\Delta y_{t+1}^J) \tag{13}
\]

In Appendix H, I derive the expressions for the variance of returns in stochastic discount factors \(m_t^J\) and of yields for long-maturity (specifically, infinite-maturity) bonds under this model. Shocks are assumed to be conditionally lognormal. Terms that average out across countries over time, namely the idiosyncratic components of shocks, are only described in the appendix and suppressed here in the ellipses.

\[
\mathbb{V}_t (\Delta m_{t+1}^J) = \alpha_\eta (\gamma \sigma)^2 (1 + \beta_\eta^i)^2 + \alpha_e (1 - \rho)^{-1} (\gamma - 1/\psi) \phi \varphi \sigma^2 (1 + \beta_e^i)^2 \\
+ \alpha_w (1 - v)^{-1} (\gamma - 1/\psi)(1 - \gamma) K_0 \sigma_w (1 + \beta_w^i)^2 + \ldots
\]

\[
\mathbb{V}_t (\Delta y_{t+1}^i) = \alpha_e (1 - \rho)^{-1} (1/\psi) \phi \varphi \sigma^2 (1 + \beta_e^i)^2 \\
+ \alpha_w (1 - v)^{-1} (1/\psi - \gamma - \gamma/\psi) K_0 \sigma_w (1 + \beta_w^i)^2 + \ldots
\]

where \(K_0 = \frac{1}{2} \left( (1 - \alpha_\eta) + (1 - \alpha_e) \phi^2 \left( \frac{\varphi \phi}{1 - \rho} \right)^2 \right)\)

These expressions make stark the difficulties this model faces in matching my empirical findings. First, they show that a single source of heterogeneity is insufficient. For instance, suppose only the global loading \(1 + \beta_\eta^i\) varies across countries, while the other global loadings are equal across countries. In this case, Japan’s loading must dominate Australia’s loading in order to match the currency restrictions in Equation (13). However, in this setup Japan and Australia have the same variance in bond yields, violating the bond restrictions in Equation (13). A related argument applies to \(1 + \beta_e^i\) or \(1 + \beta_w^i\): Japan needs to have the higher loading to match the currency restrictions, while Australia needs to have the higher loading to match the bond restrictions.

Two sources of heterogeneity are sufficient, but they are unusual. Broadly, Japan dominates Australia in whichever loading has a larger relative coefficient in the variance of stochastic discount
factors; and Australia dominates Japan in whichever loading has a larger relative coefficient in the variance of bond yields. To make this concrete, suppose I allow $1 + \beta_i^\eta$ and $1 + \beta_i^e$ to vary across countries. Since idiosyncratic consumption shocks do not affect bond yields, this forces $1 + \beta_J^\eta > 1 + \beta_A^\eta$ to match the currency restriction in Equation (13). In turn, this requires $1 + \beta_A^e > 1 + \beta_J^e$ to match the bond restriction in Equation (13).

This is qualitatively unusual, as it implies that Japanese idiosyncratic consumption growth is more sensitive to the Fed than Australian idiosyncratic consumption growth; but Australian trend consumption growth is more sensitive to the Fed than Japanese trend consumption growth. Few models would generate this prediction easily. Countries that are exposed to the Fed through trade flows, bank linkages, and other real channels would likely be exposed in all dimensions of consumption too. Permutations involving heterogeneity in the volatility loading $1 + \beta_i^w$ fare no better: one country is more exposed to the Fed in consumption growth, while the other is more exposed to the Fed in consumption volatility. For instance, I calibrate the model per Bansal et al. [2012], and find that Japan is more exposed in consumption growth while Australia is more exposed in consumption volatility, as both shocks $\eta$ and $e$ load relatively more on the stochastic discount factors more while shocks $w$ load relatively more on bonds. This illustrates the general tension that models of complete markets must confront when matching the asymmetries I document, even when rich heterogeneity is incorporated.

6.3 General Restrictions on Stochastic Discount Factors

In this section, I characterize the tension between currency and bond movements in a general preference-free framework with complete markets, allowing for higher order moments and deriving results applicable to both nominal or real stochastic discount factors. To show the tension, I decompose a general stochastic discount factor into transitory and permanent components. The results from currencies require the permanent component of stochastic discount factors to be more volatile through Fed announcements in Japan than in Australia. By contrast, the results from bonds require the transitory components of stochastic discount factors to be more volatile through Fed announcements in Australian than in Japan. Although the permanent and transitory components are mathematically different objects, they are economically highly related, and so these two restrictions seem unusual. This section follows the approach taken by Lustig et al. [2017] closely.

I make three adjustments to the prior approaches. First, I derive my conditions using entropy rather than variance, denoted by operator $L_t$.\(^{21}\) Entropy captures higher-order moments, although there remains an open debate over the importance of higher-order moments to currency risk, with estimates ranging from under 20% by Jurek and Xu [2014] to as high as 40% by Chernov et al. [2014]. Second, in keeping with the notation of Alvarez and Jermann [2005], I distinguish between the pricing kernel $\Lambda_t$ and the stochastic discount factor $M_t$, where $M_{t+k}$ is the ratio of pricing kernels.

\(^{20}\)These conditions are necessary but not sufficient; the actual minimum gap between the two loadings depends on the exact parameterization of the model.

\(^{21}\)This is equivalent to half the conditional variance of the log of a random variable when working with lognormal random variables.
\( \Lambda_{t+k} \) and \( \Lambda_t \), i.e. the growth rate of pricing kernels between \( t + k \) and \( t \). Third, I assume that each pricing kernel is the product of two components: a martingale permanent component, and a residual transitory component, following the decomposition of Alvarez and Jermann [2005]. Alvarez and Jermann [2005] discuss the regularity conditions behind this decomposition, but broadly the conditions correspond to pricing kernels that neither explode nor collapse in the infinite-horizon limit.

\[
\frac{M_{t+1}}{\Lambda_t} = \frac{\Lambda_{t+1}^{i,P} \Lambda_{t+1}^{i,T}}{\Lambda_t^{i,P} \Lambda_t^{i,T}}
\]

I first turn to my results from currencies. Backus et al. [2001] generalizes the expression in Equation (11) from variance to entropy. Specifically, the excess currency return between two countries is equal to the differences in entropy in each country’s stochastic discount factor under complete markets. In addition to high-rate currencies earning excess returns over low-rate currencies unconditionally, I again use the evidence from Mueller et al. [2017] to argue that they earn excess returns through Fed announcements specifically. This yields my first restriction.

\[
\begin{align*}
\mathbb{E}_t \Delta s_{t+1} + r_{A,t} - r_{J,t} &= L_t \left( \frac{\Lambda_{t+1}^J}{\Lambda_t^J} \right) - L_t \left( \frac{\Lambda_{t+1}^A}{\Lambda_t^A} \right) > 0 \\
L_t \left( \frac{\Lambda_{t+1}^{J,P} \Lambda_{t+1}^{J,T}}{\Lambda_t^{J,P} \Lambda_t^{J,T}} \right) - L_t \left( \frac{\Lambda_{t+1}^{A,P} \Lambda_{t+1}^{A,T}}{\Lambda_t^{A,P} \Lambda_t^{A,T}} \right) &> 0
\end{align*}
\]

I next turn to bond markets, and continue to assume that my results for ten-year and thirty-year bonds would extend to infinite-maturity bonds. The value of an infinite-horizon zero-coupon bond in country \( i \) is the expectation of the stochastic discount factor spanning those two periods.

\[
P_t^i(\infty) = E_t \left( \frac{\Lambda_{\infty}}{\Lambda_t^i} \right) = E_t \left( \frac{\Lambda_{\infty}^{i,P} \Lambda_{\infty}^{i,T}}{\Lambda_t^{i,P} \Lambda_t^{i,T}} \right)
\]

Alvarez and Jermann [2005] argue that at infinite maturities, there is no transitory component of the pricing kernel \( \Lambda_{\infty}^{i,T} \); there is only the permanent component. Moreover, since the permanent component is a martingale, today’s expectations of the infinite-horizon permanent component equal today’s permanent component. This yields a simplification in the expression for prices. In turn, I use this to compute gross yields as a function of innovations in the transitory component.

\[
P_t^i(\infty) = \frac{1}{\Lambda_t^{i,P}} \Rightarrow 1 + \Delta y_{t+1}^i(\infty) = \frac{P_t^i(\infty)}{P_{t+1}^i(\infty)} = \frac{\Lambda_{t+1}^{i,T}}{\Lambda_t^{i,T}}
\]

I next compute the empirical entropy of bond returns following Fed announcements in my sample, and confirm in Appendix H that it is statistically higher in high-rate countries.\(^{22}\) This

\(^{22}\)A long literature discusses the existence of “peso problems” in currency markets, i.e. the possibility that limited historical samples do not include observations of extreme events and thus bias calculations of risk. Burnside et al. [2011] estimate returns from currency portfolios that hedge such extreme events through options, and they find
yields the second restriction.

\[ L_t \left( \Delta y_{t+1}(\infty) \right) = L_t \left( \frac{\Lambda_{t+1}^{A,\tau}}{\Lambda_t^{A,\tau}} \right) > L_t \left( \Delta y_{t+1}(\infty) \right) = L_t \left( \frac{\Lambda_{t+1}^{J,\tau}}{\Lambda_t^{J,\tau}} \right) \]  

Equations (14) and (15) illustrate this tension again. The transitory component of the stochastic discount factor today must be more volatile in Australia than in Japan, but the overall stochastic discount factor today must be more volatile in Japan than in Australia. As a result, either the permanent component must be more volatile in Japan than in Australia, or the correlation between the Japanese components must be higher than in Australia. These results are preference-free and encompass higher moments. Models with a single source of heterogeneity cannot match these equations simultaneously.

To make this discussion less abstract, I decompose the stochastic discount factors in the power utility and Epstein-Zin examples into permanent and transitory components. First consider the power utility example. The permanence of a consumption shock is driven by \( \rho \). If \( \rho = 1 \), the shock is permanent. Agents cannot smooth away a permanent shock, and so both bond yields and the transitory component of the stochastic discount factor remain constant. If \( \rho < 1 \), the shock is transitory. Agents want to borrow or lend against the future to smooth their transient fluctuations in marginal utility, causing both bond yields and the transitory component of the stochastic discount factor to move. There is no parameterization whatsoever that allows power utility to match my results.

Next consider the example with Epstein-Zin utility. I derive these expressions in Appendix H, using the approach of Hansen and Scheinkman [2009]. I again abstract from the components of entropy that are common across all countries. Each shock has a different loading. Level shocks to consumption \( \eta \) are permanent, and so they only appear when computing the permanent components of entropy. By contrast, shocks to the trend and volatility of consumption growth have both permanent and transitory elements. Consumption itself follows a random walk, but the trend component of consumption growth is a stationary autoregressive process. As a result, shocks that affect that process both dissipate over time but affect consumption permanently in the meanwhile.

While this is possible mathematically, it is highly unusual economically. As before, these restrictions require Japan and Australia to be exposed to different parts of consumption in strongly asymmetric and countervailing ways. Models in which one country is more integrated with the US
would not typically distinguish between these two components.

\[
\begin{bmatrix}
L_t \left( \Lambda_{t+1} / \Lambda_t \right) \\
L_t \left( \Lambda_{t+T}^{iT} / \Lambda_t^{iT} \right) \\
L_t \left( \Lambda_{t+1}^{iP} / \Lambda_t^{iP} \right)
\end{bmatrix}
= A
\begin{bmatrix}
(1 + \beta_i^L)^2 \\
(1 + \beta_i^e)^2 \\
(1 + \beta_i^w)^2
\end{bmatrix}
\]

where

\[
A = \begin{bmatrix}
\alpha_\eta (\gamma \sigma)^2 & \alpha_e \left( (1 - \rho)^{-1} (\gamma - 1/\psi) \phi \varphi \sigma \right)^2 & \alpha_w \left( (1 - v)^{-1} (\gamma - 1/\psi)(1 - \gamma) K_0 \sigma_w \right)^2 \\
0 & \alpha_e \left( (1 - \rho)^{-1} (1/\psi) \phi \varphi \sigma \right)^2 & \alpha_w \left( (1 - v)^{-1} (1/\psi - \gamma / \psi) K_0 \sigma_w \right)^2 \\
\alpha_\eta (\gamma \sigma)^2 & \alpha_e \left( (1 - \rho)^{-1} \gamma \phi \varphi \sigma \right)^2 & \alpha_w \left( (1 - v)^{-1} \gamma^2 K_0 \sigma_w \right)^2
\end{bmatrix}
\]

7 Models of Incomplete Markets

Classes of explanations around central banks reacting or risk premia shifting under complete markets cannot easily match my empirical findings. This leaves models of incomplete markets. Whereas my empirical findings are inconsistent or inconvenient for the first categories of explanations, they fit many models of incomplete markets naturally. In this section, I do two things. I first illustrate with a simple model of segmented markets in the style of Gabaix and Maggiori [2015]. Second, I document one final restriction that these models must obey: they must be asymmetric in exposure to central banks. Specifically, I show that the Fed and ECB generate spillovers in foreign financial markets, but few other central banks do, as already seen briefly in Table 2.

Where models of complete markets need multiple forms of countervailing heterogeneity, models of incomplete markets can use forms of complementary heterogeneity. For instance, in models with segmented markets such as Gabaix and Maggiori [2015] or Alvarez et al. [2009], bonds and exchange rates are only linked through their positions in the portfolios of intermediaries. As such, high-rate currencies such as the Australian dollar depreciate against low-rate currencies such as the yen precisely when high-rate bond prices fall most, consistent with my empirical observations. Alternatively, in models with leverage constraints such as Maggiori [2017] or Bruno and Shin [2015], bonds and exchange rates are exposed to the same forms of heterogeneity. In Maggiori [2017], for instance, cyclical borrowing constraints generate volatility in the bond market while cyclical trade financing costs generate volatility in exchange rates, and the model is consistent with my empirical findings if Australia is less financially developed than Japan. Sandulescu et al. [2017] offers a more general take on restrictions that stochastic discount factors must face in an incomplete markets setting.

However, such models must obey one additional restriction. Frictions must be asymmetrically sensitive to the central banks of the Fed and perhaps the ECB, and not sensitive to the central banks of other countries. In this section, I extend my methodology with inferred shocks to the announcements by the other nine central banks in my sample, to test whether assets react to those central banks, and to test whether those assets react asymmetrically. I find that foreign assets react to the ECB, with continental European assets reacting differentially than non-European assets.
Foreign assets do not broadly react to other central banks. These are noteworthy findings in their own right, since very few papers examine spillovers from other central banks, possibly because of prior limitations in finding measured shocks.

7.1 Example Model of Segmented Markets

I illustrate that even the simplest models of incomplete markets can match my empirical findings, by using the model of Gabaix and Maggiori [2015]. In this setting, currencies and bonds are only linked through the positions of intermediaries. Currencies and bonds that intermediaries have large positive positions in (e.g. those in high-rate countries) depreciate together when constraints on intermediaries rise, and currencies and bonds that intermediaries have small or negative positions in (e.g. those in low-rate countries) appreciate together. This exactly matches my empirical findings.

This example follows from the multi-asset generalization of Gabaix and Maggiori [2015]. In this story, intermediaries’ positions $\theta$ are defined by three terms: their (scalar) risk aversion $\Gamma$, the variance-covariance matrix $V$, and their expected profits $\pi$.

$$\theta = \Gamma^{-1}V^{-1}\pi$$

An interpretation of the Fed’s tightening and easing is that constraints on intermediaries are respectively tightened or eased, as in Bruno and Shin [2015]. Drechsler et al. [2017] provide one possible channel, arguing that the cost of leverage is directly tied to the Fed’s nominal rate. Another channel relies on the Fed’s abilities to induce panic in market participants. Suppose the Fed tightens and raises $\Gamma$ in this story. I make the simplifying assumptions that external demand for assets and the variance of assets remains the same through Fed announcements, yielding:

$$\frac{\partial \pi}{\partial \Gamma} = \Gamma^{-1}\pi$$

Thus, the fluctuations in prices are exactly proportional to expected profits, through the intermediaries’ positions. When constraints tighten, intermediaries with large existing positions in a given asset are induced to continue holding that asset only if returns rise substantially (i.e. prices fall more). In practice, bonds with high rates, such as Australian bonds, are likely more lucrative than bonds with low rates, such as Japanese bonds. When the Fed tightens, yields on Australian bonds rise more than yields on Japanese bonds; and the Australian dollar depreciates versus the yen.

This stylized story immediately matches the core findings in this paper, and gives new meaning to the high-rate and low-rate dichotomy. In Appendix D, I further look at direct measures of intermediary positions: cross-border bank positions from the Bank of International Settlements, and cross-border portfolio positions from the IMF. These also strongly correlate with the asymmetries.
7.2 Spillovers by Other Central Banks

Models of incomplete markets must obey one additional restriction: asymmetry in the link between financial frictions and central banks. I study all ten central banks in my sample, and find that only the Fed and ECB can generate international spillovers. Specifically, European assets react differently than non-European assets to the ECB, with the euro appreciating less against European currencies and European yields rising more when the ECB tightens. Currencies and bonds do not broadly react to other central banks, although occasionally a country’s assets react to a neighboring central bank (e.g. New Zealand assets respond to the Reserve Bank of Australia).

While I present my results in the context of guiding successful models, these are noteworthy results alone. As discussed further in Section 2, there is a paucity of literature studying spillovers emanating from other central banks (except for the ECB’s effects on European assets). This is likely due to limitations in measuring monetary shocks. However, my method sidesteps that issue by estimating inferred shocks and can be applied to any central bank.

The results of Fed and ECB dominance echo work in other domains, where the dollar and euro are similarly dominant. For instance, Gopinath [2015] finds that trade is invoiced overwhelmingly in dollars and euros, Bruno and Shin [2015] find cross-border bank loans are most commonly in dollars and euros, and Maggiori et al. [2017] find that the foreign currency exposures of international portfolio positions are biased toward dollars and euros.

I first show results from currency markets in Table 8, to complement results from bond markets in Table 2. The previous pattern holds up: the Fed continues to have strong effects on most currency markets. For instance, the Australian dollar (measured against an equal-weighted basket of currencies, except for the US dollar) is 73% more volatile through Fed announcement windows than non-announcement windows, and this is statistically significant at the 1% level. Moreover, central banks have the power to affect their own currencies.

Among other central banks, the ECB stands out: it also has strong effects on most other currency markets. Few central banks otherwise have effects, although there are occasional results involving neighboring countries. The Reserve Bank of Australia and the Reserve Bank of New Zealand respectively affect each others’ currencies, and the Swiss National Bank affects the Swedish krona.

I next employ the factor model to each central bank, to identify asymmetries in how currencies and bonds react to that central bank’s monetary shocks. The most stark results come from the ECB, plotted in Figure 9. When the ECB tightens, the euro appreciates less against other continental European currencies (e.g. the Norwegian krone) than against non-European currencies, and European bond yields rise more than non-European bond yields. This is noteworthy, although the results from currency and bond markets do not rule out central banks reacting to the ECB; nor do they rule out shifts in risk premia under complete markets. If the Norges Bank follows the ECB, both Norwegian bond yields would rise and the euro would appreciate less against the Norwegian

23These empirical results are robust to excluding the Swiss franc and Swiss bond yields during the three years during which the Swiss National Bank formally capped the franc’s exchange rate against the euro.
## Table 8: Excess Volatility in 60M Currency Returns

<table>
<thead>
<tr>
<th></th>
<th>AUD</th>
<th>CAD</th>
<th>CHF</th>
<th>EUR</th>
<th>GBP</th>
<th>JPY</th>
<th>NOK</th>
<th>NZD</th>
<th>SEK</th>
<th>USD</th>
</tr>
</thead>
<tbody>
<tr>
<td>Australia</td>
<td>293%</td>
<td>41%</td>
<td>16%</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
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<tr>
<td>Canada</td>
<td>185%</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Switzerland</td>
<td>152%</td>
<td>37%</td>
<td>37%</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Euro</td>
<td>29%</td>
<td>67%</td>
<td>125%</td>
<td>37%</td>
<td>30%</td>
<td>17%</td>
<td>35%</td>
<td>50%</td>
<td></td>
<td></td>
</tr>
<tr>
<td>United Kingdom</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td>137%</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Japan</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td>94%</td>
<td>18%</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Norway</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td>227%</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>New Zealand</td>
<td>48%</td>
<td></td>
<td>531%</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
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</tr>
<tr>
<td>Sweden</td>
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<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td>140%</td>
</tr>
<tr>
<td>United States</td>
<td>73%</td>
<td>75%</td>
<td>48%</td>
<td>82%</td>
<td>102%</td>
<td>120%</td>
<td>23%</td>
<td>60%</td>
<td>70%</td>
<td>317%</td>
</tr>
</tbody>
</table>

Notes: The table tests whether the currency of the column country is more volatile in the sixty minutes around announcements by the row central bank than in other sixty-minute intervals. (Currencies are measured against an equal-weighted basket of all other currencies except the one whose central bank is making announcements.) If returns are more volatile at the 1% level using the Brown-Forsythe test, the cell records the excess ratio of standard deviations (announcement window to non-announcement window standard deviations, minus 100%). If returns are not statistically more volatile, the cell is left blank. The full table, including ratios significant at the 5% level and insignificant ratios, can be found in the Appendix as Table 18. The Fed and ECB have spillover effects, but most other central banks only affect their own currencies.

I also test whether assets react asymmetrically to other central banks. Unsurprisingly, I find few strong asymmetries: if assets aren’t reacting to other central banks as in Tables 2 and 8, they cannot react asymmetrically. For instance, consider the Bank of Japan in Figure 10, whose standard error bars and lower-dimensional structure show no asymmetry. When the yen appreciates, it does so symmetrically against all currencies; and when Japanese bond yields move, other bond yields do not move asymmetrically (or at all).

The exceptions are again with neighboring central banks. In Section 4, I illustrated that Australian assets have some asymmetric exposure to the Reserve Bank of New Zealand. I similarly show that New Zealand assets have some asymmetric exposure to the Reserve Bank of Australia in Figure 11. When either reserve bank tightens, its currency depreciates by slightly less against its counterpart than other currencies. In this example, the Australian dollar appreciates or depreciates less against the New Zealand dollar (and Canadian dollar), and New Zealand bond yields rise and fall with Australian yields. I show the full results for all ten central banks in Appendix G.

These results contrast with the results of Lustig and Richmond [2017]. Their paper finds asymmetries throughout currency markets (driven by distance), whereas I largely find symmetries, with only slight distance-driven asymmetries outside the ECB. The key difference is the shocks: I use...
Figure 9: Market Reactions to EU Monetary Shocks

(a) Currency Responses

(b) Bond Responses

Notes: The figures depict the reactions of currency and bond markets to announcements by the European Central Bank. The left figure shows by how much the EUR appreciates against a given reference currency when it appreciates by 1% on average; and the right figure shows by how much foreign ten-year yields rise when German ten-year yields rise by 1%. Standard error bars in both pictures are computed versus the average reaction across currencies or foreign bonds; and the shading of the coefficient bars refers to the lower-dimensional structure, whereby assets of the same color react similarly and assets of different colors react dissimilarly following ECB announcements. The EUR appreciates by less against continental European currencies, and by more against all other currencies when the ECB tightens. Moreover, European yields rise more than non-European yields when the ECB tightens.

Figure 10: Market Reactions to JP Monetary Shocks

(a) Currency Responses

(b) Bond Responses

Notes: The figures depict the reactions of currency and bond markets to announcements by the Bank of Japan. The left figure shows by how much the JPY appreciates against a given reference currency when it appreciates by 1% on average; and the right figure shows by how much foreign ten-year yields rise when Japanese ten-year yields rise by 1%. Standard error bars in both pictures are computed versus the average reaction across currencies or foreign bonds; and the shading of the coefficient bars refers to the lower-dimensional structure, whereby assets of the same color react similarly and assets of different colors react dissimilarly following BoJ announcements. The JPY appreciates symmetrically against all currencies and foreign yields do not rise asymmetrically when the BoJ tightens.
currency returns that are only exposed to monetary shocks, whereas their currency returns are measured at monthly frequencies and so are exposed to all global shocks, monetary and fundamental. Given our divergent findings, shocks to fundamentals seem to affect currencies differently than shocks to monetary policy.

8 Conclusion

The asymmetries in currency and bond markets around the globe following Fed announcements are both puzzling and illuminating. They suggest that the central banks of developed markets do not adjust their monetary policies to follow the Fed, and they further suggest that models of complete markets are unlikely to explain shifts in risk premia. In short, they provide unequivocal negative answers to entire classes of explanations. The only explanation for which my findings fit naturally are models with incomplete markets.

This offers an important lesson for policymakers. Despite concerns of the trilemma being replaced by a “dilemma,” monetary independence in developed markets survives. Central banks do not yet feel the need to offset the Fed’s actions on their conventional macroeconomic indicators. However, foreign monetary policy has changed in one key respect: its abilities to influence its own financial sector is dwarfed by the Fed’s abilities to do so, likely stemming from frictions in markets. As financial markets grow in size and complexity, and as they become more intertwined globally, policymakers may want to consider employing macroprudential tools in order to wrest influence
over local investors back from the Fed. Otherwise, spillovers may spark the financial and real crises of tomorrow.
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Appendix

A. Central Bank Announcements: I document announcements for the ten central banks in my paper, by describing the schedules and nuances of each central bank. I also explain why Fed announcements represent monetary news more than news about fundamentals.

B. Empirical Framework: I document the fully generalized versions of my main methodologies (the factor model and inference by heteroskedasticity), and I discuss the limitations of other methodologies, including identification by heteroskedasticity, regression with measured shocks, etc.

C. Long-Run Inflation: I document that long-run inflation forecasts at distant horizons do not change through Fed announcements, which confirms the paper’s interpretations of the empirical findings.

D. Characterizing Asymmetric Responses: I document the country-varying and time-varying indicators which best explain asymmetric reactions to the Fed in currency and bond markets, as suggestive evidence of the underlying mechanisms.

E. Gaussian Affine Term Structure Model: I document the structural model employed to decompose yield curves in each of the ten countries into the paths of rates and the paths of term premia.

F. Robustness Checks: I document the full results for the tests in bond and currency markets using inference by heteroskedasticity, and I document that the asymmetric patterns in currency and bond markets following Fed announcements hold up across different time periods and different economic conditions.

G. Other Central Banks: I document how currency and bond markets react to announcements by all ten central banks.

H. Models of Complete Markets: I document both empirical results on bond entropy and theoretical proofs behind the model of complete markets.
A Central Bank Announcements

For each of the ten central banks, I collect the date and exact time of rate announcements that follow regularly scheduled meetings of the monetary policy committee, from 2001-2016. Unscheduled meetings and post-announcement press conferences (with one limited exception) are omitted, as they may impart news about fundamentals. I remove announcements that coincide with inflation or unemployment releases by that country’s statistical agencies. The data are sourced from Bloomberg and from the websites of each central bank, and also from Ranaldo and Rossi [2016] for Switzerland and Lucca and Moench [2015] for the United States. Details for each central bank are provided below.

A.1 Australia

Australian rate announcements target the Official Cash Rate and are made by the Reserve Bank of Australia. Until December 2007, announcements were made at 9:30 AM AEST the day following a meeting; and starting in January 2008, announcements are made at 2:30 PM AEST the day of a meeting. Announcements were made on Wednesday mornings until December 2007 and are made on Tuesday afternoons starting in January 2008. There are eleven regularly scheduled announcements per year for a total of 176 regularly scheduled announcements. There are no unscheduled announcements.

Until December 2007, rate announcements were only made following a change in the rate. This is problematic if, before 9:30 AM, the bank reveals to the market that no announcement will be made that day. However, the Australian dollar (measured against an equal-weighted basket of the dollar, euro, yen, and pound) is 50% more volatile in the 60 minutes around 9:30 AM AEST on days in which no announcement is made versus other days, and the finding is statistically significant via the Brown-Forsythe test. Thus the surprise seems to be digested at 9:30 AM until December 2007, and so the entire sample is retained.

A.2 Canada

Canadian rate announcements target the Key Interest Rate and are made by the Bank of Canada. Until December 2012, announcements were made at 9:00 AM EST; and starting in January 2013, announcements are made at 10:00 AM EST. Until December 2012, announcements were generally made on Tuesday mornings and are always made on Wednesday mornings starting in January 2013. There are eight regularly scheduled announcements per year for a total of 128 regularly scheduled announcements. The Bank of Canada has made several unscheduled rate announcements (e.g. following September 11, 2001 or during the financial crisis), which are excluded from the sample.

A.3 Eurozone

European rate announcements primarily target the Main Refinancing Rate (although they concurrently target other rates too) and are made by the European Central Bank. Announcements are made at 1:45 PM CET. Announcements are generally made on Thursday afternoons and occasionally on Wednesday afternoons. Until November 2001, regularly scheduled announcements were made twice a month; from November 2001 - December 2014, regularly scheduled announcements were made monthly; and starting in January 2015, there are eight regularly scheduled announcements per year. This leads to a total of 193 regularly scheduled announcements (21 in 2001, 12 in 2002 - 2014, and 8 in 2015 - 2016). The ECB has made two unscheduled rate announcements (following September 11, 2001, and during the financial crisis), which are excluded from the sample.
European rate announcements discuss only the current rate and do not discuss the future path of monetary policy.\textsuperscript{24} Particularly in the zero-rate era, the future path and not the current rate deliver most of the surprise. For instance, consider the January 2016 announcement, presented in its entirety:

At today’s meeting the Governing Council of the ECB decided that the interest rate on the main refinancing operations and the interest rates on the marginal lending facility and the deposit facility will remain unchanged at 0.05%, 0.30% and -0.30% respectively.

The President of the ECB will comment on the considerations underlying these decisions at a press conference starting at 14:30 CET today.

By contrast, the opening statement of the associated press conference provides guidance on the future path of monetary policy and thus resembles the rate announcements of other countries more closely. Consider the opening paragraph\textsuperscript{25} of the January 2016 press conference:

Based on our regular economic and monetary analyses, and after the recalibration of our monetary policy measures last month, we decided to keep the key ECB interest rates unchanged and we expect them to remain at present or lower levels for an extended period of time. Regarding our non-standard monetary policy measures, the asset purchases are proceeding smoothly and continue to have a favourable impact on the cost and availability of credit for firms and households.

For the ECB, it seems prudent to include the opening statement of the press conference to ensure surprises are captured and to make shocks comparable across countries. Following the rate announcement at 1:45 PM CET, press conferences occur at 2:30 PM CET. The opening statement takes around ten minutes to read; and so I measure a 60-minute window from 1:40 PM - 2:40 PM. Ten of the 2001 announcements and one of the 2002 announcements were not followed by conferences, but the methodology is not altered to be consistent.

\subsection{A.4 Japan}

Japanese rate announcements target the Overnight Call Rate and are made by the Bank of Japan. Announcements immediately follow the conclusion of meetings of the Monetary Policy Committee, and typically occur between 11:00 AM and 2:00 PM JST. Starting in January 2006, both meeting conclusion times and announcement times are recorded; but until December 2005, only meeting conclusion times are recorded. The post-2006 records show that 98\% of announcements occur within ten minutes of a meeting’s conclusion, and all announcements occur within 20 minutes. Thus, pre-2006 announcements are assumed to occur five minutes (the modal delay) after a meeting conclusion. Announcements occur on any day of the week, although infrequently on Mondays. From 2001 - 2005, regularly scheduled announcements were made fifteen or sixteen times annually; from 2006 - 2015, regularly scheduled announcements were made fourteen times annually; and starting in January 2016, there are eight regularly scheduled announcements per year. One regularly scheduled announcement that immediately follows the 2011 Tohoku Earthquake by coincidence is omitted, leaving 225 announcements. The Bank of Japan has made several unscheduled rate announcements (e.g. following September 11, 2001, at the start of the Iraq War, or during the financial crisis), which are excluded from the sample.

\textsuperscript{24}The ECB has changed this as of the middle of 2016, but this only affects a few announcements.

\textsuperscript{25}Technically, the opening paragraph bids Happy New Year; this is the first paragraph with content.
Since the Bank of Japan does not release announcements at a preset time, market participants might anticipate unusually late announcements as being unusually important. As such, I check that the core results are robust to excluding the fourteen announcements that take place after 2:00 PM JST.

### A.5 New Zealand

New Zealand rate announcements target the Official Cash Rate and are made by the Reserve Bank of New Zealand. Announcements are made at 9:00 AM NZDT, with the exception of a few meetings in 2001 that were made at either 8:00 AM or 10:00 AM NZDT. Announcements are made predominantly on Thursdays and otherwise on Wednesdays. From 2001 - 2015, there were eight regularly scheduled meetings per year; and starting in 2016, there are seven regularly scheduled meetings per year, for a total of 127 regularly scheduled meetings. The Reserve Bank of New Zealand made one unscheduled announcement following September 11, 2001, which is excluded from the sample.

Since 2009, the Reserve Bank of New Zealand holds press conferences to discuss its rate announcement four times annually (but only twice in 2009). From early 2009 until mid-2016, those conferences started at 9:00 AM NZDT too. In mid-2016, the conferences were moved to later in the day; but as a result, 27 announcement windows from 2009 - 2016 include both the effects of rate announcements and the associated press conferences. Since excluding them outright would reduce the sample by 21%, I instead check that my core results are robust to their exclusion.

### A.6 Norway

Norwegian rate announcements target the Key Policy Rate and are made by the Norges Bank. Until December 2012 and including one announcement in May 2013, announcements were made at 2:00 PM CET; and starting January 2013, announcements are made at 10:00 AM CET. Until December 2012, announcements were made on Wednesdays or Thursdays and are always made on Thursdays starting in January 2013. Until December 2008, regularly scheduled meetings were held approximately every six weeks (for eight or nine meetings annually); from 2009 - 2011, there were eight regularly scheduled meetings per year; and starting in 2012, there are six regularly scheduled meetings per year. One regularly scheduled announcement in October 2016 that coincides with Statistics Norway’s unemployment release is omitted, leaving a total of 124 announcements. One unscheduled rate announcement during the financial crisis is excluded from the sample.

The Norges Bank holds press conferences alongside the rate announcements. These press conferences are brief (sometimes just ten minutes), but some conferences are held concurrently with rate announcements. Comprehensive records are unavailable, but archived pages of Norway’s calendar, Bloomberg’s calendar, and Bloomberg’s Nordic Report give some indications. From 2004 - 2009, conferences were held 45 minutes later than the announcement and from 2015 - 2016, conferences were held 30 minutes later than the announcement. From 2011 - 2014, conferences were held concurrently with the rate announcements. To be conservative, I assume conferences in 2010 were held concurrently with rate announcements too. As a result, 34 announcement windows from 2010 - 2014 may include both the effects of rate announcements and the associated press conferences. Since excluding them outright would reduce the sample by 27%, I instead check that my core results are robust to their exclusion.
A.7 Sweden

Swedish rate announcements target the Repo Rate and made by the Sveriges Riksbank. Around 60% of the announcements from 2001 - 2006 and all announcements starting in January 2007 are made at 9:30 AM CET, although 40% of the announcements from 2001 - 2006 were made at 8:00 AM, 9:00 AM, or 11:00 CET. Announcements are made on any day of the week except Mondays. From 2001 - 2004, regularly scheduled announcements were made eight times annually; from 2005 - 2007, regularly scheduled announcements were made seven times annually; and starting in January 2008, there are six regularly scheduled announcements per year. Seven regularly scheduled announcements that coincide with Statistics Sweden’s inflation releases are omitted, leaving a total of 100 announcements. The Riksbank has made four unscheduled rate announcements (e.g. following September 11, 2001 or during the financial crisis), which are excluded from the sample.

A.8 Switzerland

Swiss rate announcements target the 3-Month Libor Target Rate and are made by the Swiss National Bank. Until December 2010, announcements were made at 9:30 AM or 2:00 PM CET; and starting in January 2011, all announcements are made at 9:30 AM CET. Announcements are almost always made on Thursdays, except for five announcements from 2001 - 2003 made on Fridays. There are four regularly scheduled announcements per year. One regularly scheduled announcement that was moved in response to the events of September 11, 2001 is omitted, leaving a total of 63 regularly scheduled announcements. The SNB has made several unscheduled rate announcements (e.g. to counteract the financial crisis, to implement a cap in 2011, and to remove that cap in 2015), which are excluded from the sample. Three regularly scheduled announcements coincide with releases on import and producer prices by the Swiss Federal Statistics Office. Since these are not the benchmark inflation reports, I do not exclude the announcements outright but check that the core results are robust to their exclusions.

Importantly, from September 2011 until January 2015, the Swiss franc was capped versus the Euro. In response to appreciation pressures, the Swiss National Bank stated on September 6, 2011:

> With immediate effect, [the SNB] will no longer tolerate a EUR/CHF exchange rate below the minimum rate of CHF 1.20. The SNB will enforce this minimum rate with the utmost determination and is prepared to buy foreign currency in unlimited quantities.

Equally unexpectedly, the SNB abandoned the peg on January 15, 2015. While the franc was not officially pegged to the Euro over the intervening forty months, strong appreciation pressures combined with the cap led to far lower volatility. The time series of the EUR/CHF exchange rate depicts this.

While these forty months witnessed lower volatility, rate announcements still delivered shocks. The Swiss franc (measured against the Euro) is over twice as volatile in the 60 minutes around SNB announcements versus other periods, and this finding is statistically significant via the Brown-Forsythe test. I therefore include rate announcements over the capped era. However, since it is possible that the SNB’s objectives diverged from its traditional objectives over this period, I check that the core results are robust to this period’s exclusion.

A.9 United Kingdom

British rate announcements target the Official Bank Rate and are made by the Bank of England. Announcements are always made at 12:00 PM BST. Announcements are almost always made on
Notes: The figure depicts the reactions of three currencies in sixty-minute windows around the Fed’s surprise easing of January 25, 2012: the yen, the Australian dollar, and the New Zealand dollar, all measured against the dollar. The Australian and New Zealand dollars appreciate similarly (approximately 1%), while the yen appreciates by substantially less (60 basis points). There is little difference between treating the Australian and New Zealand dollars as distinct currencies or collapsing them into one currency area, but there is a big difference for the yen.

Thursdays, with only 2% of announcements being made on Mondays or Wednesdays. Until late 2016, regularly scheduled announcements were made twelve times annually, but have now switched to eight regularly scheduled announcements annually. One regularly scheduled announcement that was moved in response to the financial crisis is omitted, leaving a total of 190 regularly scheduled announcements. The Bank of England made one unscheduled announcement following September 11, 2001, which is excluded from the sample.

Until recently, British rate announcements discussed only the current rate and did not discuss the future path of monetary policy. For instance, consider the January 2015 announcement, presented in its entirety:

The Bank of England’s Monetary Policy Committee at its meeting today voted to maintain Bank Rate at 0.5%. The Committee also voted to maintain the stock of purchased assets financed by the issuance of central bank reserves at £375 billion, and so to reinvest the £4.35 billion of cash flows associated with the redemption of the January 2015 gilt held in the Asset Purchase Facility.

The minutes of the meeting will be published at 9.30 a.m. on Wednesday 21 January.

Unlike the European Central Bank, the Bank of England’s brevity is not easily remedied. Press conferences happen on different days as the scheduled rate announcement, following the release of the Inflation Report. Minutes are also released on different days. Including other days is feasible but dangerous: it represents a fundamental change in my methodology by including events that are distinctly different than scheduled rate announcements.

\footnote{The Bank of England has changed this as of August 2015, but this only affects a few announcements.}
Thus, I choose to be conservative and only record scheduled rate announcements. Fortunately, this is not problematic for the paper. The primary concern is that a conservative approach would limit surprises and thus limit power. In fact, the pound (measured against an equal-weighted basket of the dollar, euro, and yen) is over twice as volatile in the 60 minutes around Bank of England announcements versus other periods, and this finding is statistically significant via the Brown-Forsythe test.

A.10 United States

American rate announcements (known as FOMC announcements) target the Federal Funds Rate and are made by the Federal Reserve. Until early 2013, announcements were made at 2:15 PM EST; and starting in January 2013, announcements are made at 2:00 PM EST. Some announcements from 2011 - 2012 are made at 12:30 PM EST. The Federal Reserve prior to August 2006 does not record the time of its announcements, but Lucca and Moench [2015] estimate the times by looking at time-stamped newswire releases, and find all announcements made within a few minutes of 2:15 PM. Announcements are made on Tuesdays and Wednesdays, and occasionally on Thursdays. There are eight regularly scheduled announcements per year, for a total of 128 regularly scheduled announcements. The Federal Reserve has made several unscheduled rate announcements (e.g. following September 11, 2001 or during the financial crisis), which are excluded from the sample.

One regularly scheduled announcement on January 31, 2006 closely overlaps with a Senate confirmation vote for Chairman Bernanke. Since the vote was largely anticipated, the announcement is not excluded outright but I check that the core results are robust to its exclusion.

Nakamura and Steinsson [2017] have raised concerns that FOMC announcements signal the Fed’s private information about fundamentals, even though announcements do not officially release new information about fundamentals. If true, monetary announcements should look qualitatively like official announcements about fundamentals: both types of announcements would concurrently release information about the state of the economy and the Fed’s reaction to it. As such, I collect unemployment releases by the Bureau of Labor Statistics, and look at the reaction of currencies and bonds globally to that. I compare such reactions to the reactions following Fed announcements in Figure 13. There is a sharp difference: the patterns that emerge following monetary announcements look little like the patterns that emerge following labor announcements. For instance, whereas the Canadian dollar and yen react similarly to each other following Fed announcements, they react very differently following BLS announcements; and whereas Australian and New Zealand bonds react similarly to each other following Fed announcements, they react very differently following BLS announcements. It seems safe to assume that FOMC announcements are not just announcements about fundamentals.
Notes: The figures depict the reactions of currency and bond markets to announcements in the US. The two top figures are reactions to Fed announcements (which are presented in the main paper also), and the bottom two are reactions to releases of the Bureau of Labor Statistics’s monthly Employment Report. The left figures show reactions in currency markets: specifically, they show by how much the dollar appreciates a given reference currency when it appreciates by 1% on average; and the right figures show reactions in ten-year bond markets: specifically, they show how much the foreign yields of other countries rise when US yields rise by 1%. Standard error bars in both pictures are computed versus the mean reaction across all foreign currencies or foreign bonds; and the shading of the coefficient bars refers to the lower-dimensional structure, whereby currencies or bonds of the same color react similarly and currencies or bonds of different colors react differently.
B Empirical Framework

This section discusses the empirical framework behind the paper’s core equation. First, it introduces the notation for the fully general equation. Second, it discusses the two leading methodologies used to identify its parameters: a simplified variance test and a maximum likelihood-based approach (i.e. the Expectation Maximization algorithm). The section then provides details of the approach used to find the model’s lower-dimensional mapping. Finally, the section discusses the implementations and highlights the shortcomings of alternate methodologies: a generalized method of moments-based approach (i.e. Identification by Heteroskedasticity), a fixed effects approach, and approaches that rely on measured shocks.

B.1 Model

Given panel data of asset returns $r$ around announcements from a given central bank at times $t = 1, ..., T$, I identify parameters and shocks $(\alpha, \beta, m)$ in the following specification:

$$r_t(i_t) = X^\alpha_t(i_t)\alpha + X^\beta_t(i_t)\beta m_t + \epsilon_t(i_t) \quad \forall t \tag{16}$$

Each component of Equation (16) is defined as follows:

- $r_t$ is a $C_r \times 1$ vector of asset returns at time $t$.
- $i_t$ is a $C_r \times 1$ vector of indicators as to whether the underlying returns are present or missing at time $t$. As such, $r_t(i_t)$ refers to the subset of asset returns that are non-missing at time $t$; and $X^\alpha_t(i_t)$, $X^\beta_t(i_t)$, and $\epsilon_t(i_t)$ are defined analogously, in which the rows of these matrices are shrunken to the indicators marked present, but columns are maintained.
- $\alpha$ is a $C_\alpha \times 1$ vector of constants.
- $\beta$ is a $C_\beta \times C_m$ matrix of coefficients (i.e. factor loadings).
- $X^\alpha_t$ and $X^\beta_t$ are covariate matrices at time $t$ with dimensions $C_r \times C_\alpha$ and $C_r \times C_\beta$ respectively.
- $m_t$ is a $C_m \times 1$ vector of factors at time $t$ with underlying distribution $m_t \sim N(\bar{m}, \Omega)$.
- $\epsilon_t$ is a $C_r \times 1$ vector of residuals at time $t$ with underlying distribution $\epsilon_t \sim N(\bar{0}, \Sigma)$. $\Sigma$ is learned from asset returns through non-event windows (denoted as $\tilde{r}_t$). To ensure that the mean of the residuals is zero in this estimation, I subtract the mean asset return through non-event windows from asset returns through event windows ($r_t$) before fitting this model to those returns.

To give a concrete illustration, consider a specification that regresses movements in the dollar against three currencies pairs (euro, yen, and pound) on both currency-specific coefficients and on a common coefficient on the local bond yield $y$ in a single-factor model with UK data missing at time $t$. In this specification, $X^\alpha_t = X^\beta_t$.

$$\begin{bmatrix} r^{e/\$} \\ r^{\$/\$} \\ r^{L/\$} \\ r^{\$/t} \\ r^{L/\$} \\ r^{t/\$} \end{bmatrix} \begin{bmatrix} 1 \\ 0 \\ 0 \\ 1 \\ 0 \\ 0 \\ 0 \\ 0 \\ 1 \\ \alpha^e \\ \alpha^y \end{bmatrix} \begin{bmatrix} 1 \\ 0 \\ 0 \\ \gamma^e \\ \gamma^y \end{bmatrix} + \begin{bmatrix} 1 \\ 0 \\ 0 \\ \gamma^e \\ \gamma^y \end{bmatrix} \begin{matrix} \alpha^e \\ \alpha^y \end{matrix} + \begin{bmatrix} 1 \\ 0 \\ 0 \\ \gamma^e \\ \gamma^y \end{bmatrix} \begin{matrix} \beta^e \\ \beta^y \end{matrix} \times m_t + \begin{bmatrix} \epsilon^{e/\$} \\ \epsilon^{\$/\$} \\ \epsilon^{L/\$} \\ \epsilon^{t/\$} \end{bmatrix} \begin{bmatrix} 1 \\ 0 \\ 0 \\ 0 \end{bmatrix}$$

66
In addition, it is worth noting that the different asset returns of $r_t$ need not be returns of different assets; in many cases, they are returns of the same assets measured over different time intervals. This is because some markets can be illiquid over short windows, and so the model makes use of returns measured over both intraday and daily frequencies (whereby the former returns provide more power when present, whereas the latter returns are more likely to be available). However, both sets of returns will embed the same underlying shocks and the same underlying coefficients. As such, it is important to parameterize the $X_\alpha$ and $X_\beta$ matrix to enforce this. For instance, consider a variant of the example above in which the model has two types of dollar-yen series, but only wants to fit one set of yen parameters.

$$\begin{bmatrix} r_{t,1}^e \r_{t,2}^e \end{bmatrix} = \begin{bmatrix} 1 & 0 & 0 \ 0 & 1 & 0 \end{bmatrix} \times \begin{bmatrix} \alpha_1^e \\ \alpha_2^e \end{bmatrix} + \begin{bmatrix} 1 & 0 & 0 \ 0 & 1 & 0 \end{bmatrix} \times \begin{bmatrix} \beta_1^e \\ \beta_2^e \end{bmatrix} \times m_t + \begin{bmatrix} \epsilon_{t,1}^e \\ \epsilon_{t,2}^e \end{bmatrix}$$

Since this methodology can handle multiple representations of the same asset and missing elements of observations, it is highly robust to missing data. It utilizes every available piece of data without exception, and it measures how assets react to every Fed announcement (switching between intraday returns when possible and daily returns when not).

### B.2 Variance Tests

Consider a single-asset variant of Equation (16) without time-varying covariates (i.e. $C_r = 1$ and $X_\beta = X_\beta^0$) in which I want to test $H_0 : \beta = \vec{0}$. This corresponds to testing whether that asset has any exposure to monetary shocks $m_t$ emanating from a given central bank. In this simplified setting, it is overkill to make various structural assumptions, to employ sophisticated algorithms to fit the model, or even to derive point estimates for $\beta$. Instead, this hypothesis can be tested by a simple variance test. Taking the variance of Equation (16) yields the following, where missing returns are simply dropped:

$$\mathbb{V}(r_t) = X_\beta \Omega \beta^T X_\beta^T + \mathbb{V}(\epsilon_t)$$

Notice that unless $\beta = \vec{0}$, monetary shocks will increase the variance of asset returns around announcements relative to the variance of residuals (ignoring trivial cases, e.g. $X_\beta = 0$). As such, the test simplifies to:

$$\mathbb{V}(r_t) > \mathbb{V}(\epsilon_t) \implies \beta \neq \vec{0}$$

In the paper, the variance of residuals itself is estimated from non-event windows ($\hat{r}_t$), which are windows without monetary announcements. Thus, the actual test is:

$$\mathbb{V}(r_t) > \mathbb{V}(\hat{r}_t) \implies \beta \neq \vec{0}$$
Although the F-test for equality of variances is the best known variance test, I employ the Brown-Forsythe test for equality of variances instead. This choice is discussed further below.

This methodology is advantageous for two reasons. First, it is transparent: identification comes from a single moment alone. Second, it is assumption-lite. The main assumption is that non-event windows \( \tilde{r}_t \) and event windows \( r_t \) are identical apart from the event itself. The paper discusses the ways in which non-event windows are chosen to mirror the liquidity of event windows, to ensure this assumption holds.

By contrast, there are several assumptions that this methodology does not need to make. First, shocks need not be common across assets (e.g. shocks to the euro market need not look like shocks to the yen market), as this test is conducted for each asset in isolation. Second, the dimensionality of \( m_t \) need not be specified. While the paper often discusses \( m_t \) as though it is univariate, in fact the test is robust to a multivariate \( m_t \). For instance, consider a simple case where \( m_t \) has two imperfectly correlated components and \( X^\beta = I_2 \) for simplicity.

\[
\mathbb{V}(r_t) - \mathbb{V}(\tilde{r}_t) = \beta_1 \sigma_1^2 + \beta_2 \sigma_2^2 + 2\beta_1 \beta_2 \rho \sigma_1 \sigma_2 \begin{cases} = 0 & \text{if } \beta_1 \text{ and } \beta_2 = 0 \\ > (\beta_1 \sigma_1 - \beta_2 \sigma_2)^2 & \geq 0 \text{ if } \beta_1 \text{ or } \beta_2 \neq 0 \end{cases}
\]

Third, the choice of the Brown-Forsythe test over the F-test means that asset returns need not be normal. Broadly, the Brown-Forsythe test computes its test statistic via absolute deviations from the median, rather than squared deviations from the mean as the F-test does, and both of these adjustments ensure the test remains robust to fat-tailed data. To further illustrate its relative advantages, I simulate repeated samples of fat-tailed data with unitary variance and varying excess kurtosis, and test a random subset of each sample against another random subset under both the Brown-Forsythe and the F-test. Table 9 shows the percentage of time that these tests reject at the 5% level. When the data has no excess kurtosis (i.e. a standard normal), both tests correctly reject 5% of the time. However, for moderate or high values of excess kurtosis, the F-test rejects far too frequently; whereas the Brown-Forsythe test continues to reject 5% of the time correctly.

Table 9: Test Rejection Percentages on Simulated Data

<table>
<thead>
<tr>
<th>Kurtosis</th>
<th>BF Test</th>
<th>F-Test</th>
</tr>
</thead>
<tbody>
<tr>
<td>0</td>
<td>0.05</td>
<td>0.05</td>
</tr>
<tr>
<td>4</td>
<td>0.04</td>
<td>0.23</td>
</tr>
<tr>
<td>8</td>
<td>0.05</td>
<td>0.39</td>
</tr>
</tbody>
</table>

Notes: The table tests whether returns of the column country’s ten-year sovereign bonds are more volatile around announcements by the row central bank than at other times. Dark grey refers to returns in sixty-minute windows; and light grey refers to returns in daily windows when bond markets of that country are too illiquid at that time to accurately compute returns over sixty-minute windows. The cell shows the excess ratio of standard deviations for that asset (announcement window standard deviation over non-announcement window standard deviation, minus 100%). Significance is assessed at the 1% (**) and 5% (*) level by the Brown-Forsythe test.

### B.3 Expectation Maximization Algorithm

To identify the parameters \((\alpha, \beta, m)\) in the general setting of Equation (16), I use the Expectation Maximization algorithm. This approach finds the maximum likelihood estimates of the parameters. The likelihood function is written below, but the maximum likelihood parameter estimates cannot
be solved directly, due to the product of $\beta$ and $m$. The Expectation Maximization algorithm, by contrast, updates each of $(\alpha, \beta, m)$ sequentially until the algorithm converges on the optimal estimates.

$$
\max_{\alpha, \beta, (m_t)} \sum_{t=1}^{T} \left[ \left( r_t(i_t) - X_t^\alpha(i_t)\alpha - X_t^\beta(i_t)\beta m_t \right)^T \Sigma(i_t, i_t)^{-1} \left( r_t(i_t) - X_t^\alpha(i_t)\alpha - X_t^\beta(i_t)\beta m_t \right) \right]
$$

This specification is a latent Gaussian model (a type of hidden Markov model), and is common in the signal processing literature. Such models are frequently solved in that literature through the Expectation Maximization algorithm.

To solve this model, I augment the log-likelihood expression with a variational posterior distribution for $m_t \sim N(\mu_t, V_t)$. This allows us to take the expectation of the log-likelihood with respect to $m_t$. These two steps are below, and I can subsequently iterate over the new parameter space $(\alpha, \beta, \mu_t, V_t)$. The main assumption imposed here is that the ex ante variance of monetary shocks is set at $I$. Without this assumption, $\beta m_t$ will be indeterminate; and it is a standard assumption in factor models. Note that the additional terms in the log-likelihood are the negative of the Kullback-Leibler divergence between the variational Gaussian posterior and the Gaussian prior for $m_t$.

$$
\max_{\alpha, \beta, (\mu_t, V_t)} \sum_{t=1}^{T} \left( -\frac{1}{2} (r_t(i_t) - X_t^\alpha(i_t)\alpha - X_t^\beta(i_t)\beta m_t)^T \Sigma(i_t, i_t)^{-1} (r_t(i_t) - X_t^\alpha(i_t)\alpha - X_t^\beta(i_t)\beta m_t) + \frac{1}{2} \log |V_t| - \frac{1}{2} \text{Tr}(V_t) - \frac{1}{2} \mu_t^T \mu_t \right)
$$

I now take first order conditions with respect to the underlying parameters, and update them iteratively until the algorithm converges. The rearranged conditions for $(V_t, \mu_t, \alpha)$ are presented first.

$$
V_t = \left( \beta^T X_t^\beta(i_t)\Sigma(i_t, i_t)^{-1} X_t^\beta(i_t) + I_{C_m} \right)^{-1} \quad \forall t
$$

$$
\mu_t = V_t \left( \beta^T X_t^\beta(i_t)\Sigma(i_t, i_t)^{-1} (r_t(i_t) - X_t^\alpha(i_t)\alpha) \right) \quad \forall t
$$

$$
\alpha = \left( \frac{1}{T} \sum_{t=1}^{T} X_t^\alpha(i_t)\Sigma(i_t, i_t)^{-1} X_t^\alpha(i_t) \right)^{-1} \left( \frac{1}{T} \sum_{t=1}^{T} X_t^\alpha(i_t)\Sigma(i_t, i_t)^{-1} \left( r_t(i_t) - X_t^\beta(i_t)\beta m_t \right) \right)
$$

Rearranging the first order condition for $\beta$ presents some complications, as $\beta$ is not easily isolated:
\[ \frac{1}{T} \sum_{t=1}^{T} X_t^\beta (i_t)\Sigma(i_t, i_t)^{-1} X_t^\beta (i_t) \beta (\mu_t \mu_t^T + V_t) = \frac{1}{T} \sum_{t=1}^{T} X_t^\beta (i_t)\Sigma(i_t, i_t)^{-1} (r_t(i_t) - X_t^\alpha \alpha) \mu_t^T \]

As such, there are three approaches. The first and main approach, which is utilized throughout the paper, is to assume the factor is univariate. If so, I can rearrange the expression more easily, since \((\mu_t \mu_t^T + V_t)\) is a scalar:

\[ \beta = \left( \frac{1}{T} \sum_{t=1}^{T} (\mu_t^2 + V_t) X_t^\beta (i_t)\Sigma(i_t, i_t)^{-1} X_t^\beta (i_t) \right)^{-1} \left( \frac{1}{T} \sum_{t=1}^{T} \mu_t X_t^\beta (i_t)\Sigma(i_t, i_t)^{-1} (r_t(i_t) - X_t^\alpha (i_t) \alpha) \right) \]

There are two other approaches, which I do not implement in this paper but describe for completeness. The second approach is to attempt to solve \(\beta\) implicitly from this expression. The third expression is to remove the variation in \(X_t^\beta (i_t)\) and \(\Sigma(i_t, i_t)\) through time. For instance, this requires that the model have no missing data, so that \(i_t\) can be removed. Under these assumptions, factors can remain multivariate; and \(\beta\) can be isolated:

\[ \beta = \left( X_t^{\beta T} \Sigma^{-1} X_t^\beta \right)^{-1} \left( X_t^{\beta T} \Sigma^{-1} \right) \left( \frac{1}{T} \sum_{t=1}^{T} (r_t - X_t^\alpha \alpha) \mu_t^T \right) \left( \frac{1}{T} \sum_{t=1}^{T} (\mu_t \mu_t^T + V_t) \right)^{-1} \]

Standard errors for \(\alpha\) and \(\beta\) are computed by bootstrap, sampling vectors of asset returns at time \(t = 1, ..., T\) with replacement. This is safer than analytic expressions for standard errors, as those formulas typically do not address missing data.

Compared to the previous approach, this one imposes more assumptions and structure. It posits that the factor dimensionality is known, and that shocks are common across assets. Moreover, identification comes from both the diagonal and the off-diagonal elements of the variance matrix; whereas the previous approach only uses the diagonal elements for identification. Finally, this approach is more sensitive to deviations from normality, although I prune outliers to ensure normality holds approximately.

### B.3.1 Markov Chain Monte Carlo

This section will be written soon. Please check soon, and ensure you are reading the most recent version of the paper.

### B.4 Lower-Dimensional Mapping

This paper routinely maps the factor model to the lowest-dimensional structure possible, in which the assets of partner countries have similar reactions to monetary shocks, and thus share coefficients (e.g. Norwegian and Swedish assets often react similarly). An example lower-dimensional structure is below, in which the yen and pound share a coefficient.

\[
\begin{bmatrix}
\epsilon_t^E/S \\
\epsilon_t^X/S \\
\epsilon_t^E/S \\
\epsilon_t^X/S \\
\end{bmatrix} = \begin{bmatrix} 1 & 0 \\ 0 & 1 \\ 0 & 1 \\ 0 & 1 \\ \end{bmatrix} \times \begin{bmatrix} \alpha^E \\ \alpha^X \\ \end{bmatrix} \times \begin{bmatrix} \beta^E \\ \beta^X \\ \end{bmatrix} \times m_t + \begin{bmatrix} \epsilon_t^E/S \\ \epsilon_t^X/S \\ \epsilon_t^E/S \\ \epsilon_t^X/S \\ \end{bmatrix}
\]

To find the optimal lower-dimensional structure, I proceed in three steps. First, I fit the parameters of each lower-dimensional structure through the Expectation Maximization algorithm.
Second, I evaluate the model’s likelihood at these parameters, using Equation (17). Third, I find the best model fit using the (extended) Bayesian Information Criterion, which trades off the model log likelihood $L(\theta)$ against the model’s dimensionality $|\theta|$. The expression, written below, differs from the (vanilla) Bayesian Information Criterion in that it is more conservative and penalizes parameters more severely than usual. Chen and Chen [2012] and Foygel and Drton [2011] argue in favor of this more conservative model criteria when the number of parameters is high, to combat the heightened risk of overfitting. My model, which fits not only $\alpha$ and $\beta$ but also the variational posterior parameters for each monetary shock, is a prime candidate for such an approach.

$$EBIC(\theta) = 2L(\theta) - \log(n)|\theta| - 2\log(|\theta|)|\theta|$$

Since I have nine counterpart countries (excluding the country from which the monetary shocks emanate), there are 21,147 permutations of a lower-dimensional structure. This number is computed as the solution to the counting problem of the number of ways to place nine distinguishable countries in up to nine indistinguishable groups. The solution to this problem involves Stirling numbers of the second kind, which count the number of ways to place $n$ distinguishable objects in $k$ indistinguishable boxes, with no empty boxes. I thus sum up the Stirling numbers for $k = 1, \ldots, 9$. The formula for Stirling numbers of the second kind is:

$$S(n, k) = \frac{1}{k!} \sum_{i=0}^{k-1} (-1)^i \binom{k}{i} (k - i)^n$$

While this problem is computationally intensive, it is not outlandishly so. (By contrast, if I were placing nine distinguishable countries in up to nine distinguishable groups, I would have $9^9 \approx 400$ million permutations over which to iterate.) As such, I do not need to find the optimal structure through heuristics or approximations; but can actually compute each permutation. I use Harvard’s Odyssey computing cluster for this task.

Finally, it is important to stress that assets in different asset classes never share coefficients; only assets in the same asset class would. As an example, consider a model that fits both currency and bond returns from the Eurozone, Japan, and the UK; and as before, Japanese and British assets share coefficients. The lower-dimensional representation of this system would be written as follows:

$$\begin{bmatrix}
\epsilon_e,$$
$\epsilon_c,$$
$\epsilon_b,$$
$\epsilon_t$
\end{bmatrix} = 
\begin{bmatrix}
1 & 0 & 0 & 0 \\
0 & 1 & 0 & 0 \\
0 & 1 & 0 & 0 \\
0 & 0 & 1 & 0 \\
0 & 0 & 0 & 1 \\
0 & 0 & 0 & 1 \\
\end{bmatrix} \times 
\begin{bmatrix}
\alpha_e \\
\alpha_c \\
\alpha_b \\
\alpha_t \\
\beta_e \\
\beta_c \\
\beta_b \\
\beta_t \\
\end{bmatrix} + 
\begin{bmatrix}
1 & 0 & 0 & 0 \\
0 & 1 & 0 & 0 \\
0 & 1 & 0 & 0 \\
0 & 0 & 1 & 0 \\
0 & 0 & 0 & 1 \\
0 & 0 & 0 & 1 \\
\end{bmatrix} \times m_t + 
\begin{bmatrix}
\epsilon_e,$$
$\epsilon_c,$$
$\epsilon_b,$$
$\epsilon_t$
\end{bmatrix}$$

### B.5 Identification by Heteroskedasticity

An alternate approach to identify the parameter $\beta$ for Equation (16) is to use Identification by Heteroskedasticity. This approach is a generalized methods of moments (GMM) estimator. While this approach gives equivalent results to the Expectation Maximization algorithm when it converges, it has poor convergence properties in the first place due to the severe non-linearities and high dimensionality of the estimator. As such, it is not utilized except as a robustness check. This section discusses both the implementation and the issues.
To understand the principle behind Identification by Heteroskedasticity in this context, I first present an extremely simplified version of Equation (16) to build intuition. In this example, asset returns are demeaned, monetary shocks $m_t$ are univariate, there is no missing data, there is one coefficient per asset ($C_r = C_\beta$), and $X_t^\beta = I_{C_r}$. In this simplified framework, asset returns are measured around the set of event windows $E$ and the set of non-event windows $N$:

$$r_t = \beta m_t + \epsilon_t \quad \forall \, t \in E$$
$$\tilde{r}_t = \epsilon_t \quad \forall \, t \in N$$

Taking the second moments of these systems yields the following:

$$\mathbb{E}r_tr_t^T = \beta\beta^T\mathbb{E}m_t^2 + \Sigma$$
$$\mathbb{E}\tilde{r}_t\tilde{r}_t^T = \Sigma$$

I make the same normalization as in the Expectation Maximization approach: that the variance of monetary shocks is set at one. As a result, $\beta$ can be estimated from the difference in the implied variance-covariance matrices of asset returns through event windows and through non-event windows.

$$\mathbb{E}r_tr_t^T - \mathbb{E}\tilde{r}_t\tilde{r}_t^T = \beta\beta^T$$

I now present the more general framework and the exact operational steps to fit $\beta$ via GMM. This framework allows for missing data, shared coefficients between series, and non-trivial covariate matrices $X^\beta$; but it continues to insist that asset returns are demeaned, that covariate matrices are constant across time, and that monetary shocks are univariate.

$$r_t(i_t) = X^\beta(i_t)\beta m_t + \epsilon_t(i_t) \quad \forall \, t \in E$$
$$\tilde{r}_t(i_t) = \epsilon_t(i_t) \quad \forall \, t \in N$$

Since there may be missing data, I do not present the expression for a general variance-covariance matrix but instead focus on element $(j,k)$ of this matrix. As before, I compute the second moment of both event and non-event windows, and take the difference.

$$\sum_{t \in E, (j,k) \in i_t} \frac{1}{t \in E, (j,k) \in i_t} \sum_{t \in E, (j,k) \in i_t} r_t(j)r_t(k) = \sum_{t \in E, (j,k) \in i_t} \frac{1}{t \in E, (j,k) \in i_t} \sum_{t \in E, (j,k) \in i_t} X^\beta(j)\beta^T X^\beta(k)\beta^T m_t^2 + \sigma(j,k)$$

$$\sum_{t \in N, (j,k) \in i_t} \frac{1}{t \in N, (j,k) \in i_t} \sum_{t \in N, (j,k) \in i_t} \tilde{r}_t(j)\tilde{r}_t(k) = \sigma(j,k)$$

$$\sum_{t \in E, (j,k) \in i_t} \frac{1}{t \in E, (j,k) \in i_t} \sum_{t \in E, (j,k) \in i_t} r_t(j)\tilde{r}_t(k) - \sum_{t \in N, (j,k) \in i_t} \frac{1}{t \in N, (j,k) \in i_t} \sum_{t \in N, (j,k) \in i_t} \tilde{r}_t(j)\tilde{r}_t(k) = X^\beta(j)\beta^T X^\beta(k)^T$$

There are $\frac{1}{2} \times C_r \times (C_r + 1)$ unique equations to estimate $C_\beta$ parameters. This is an over-determined system, and so the Generalized Method of Moments is utilized to find the best fit. The sample moment condition for asset returns at time $t$ and variance-covariance entry $(j,k)$ is defined as follows:
\( g_{t,(j,k)}(\beta) = \begin{cases} 
\left( \sum_{t \in E,(j,k) \in i_t} \right)^{-1} r_t(j) r_t(k) - X^\beta(j) \beta^T X^\beta(k)^T & \text{if } t \in E,(j,k) \in i_t \\
\left( \sum_{t \in N,(j,k) \in i_t} \right)^{-1} \tilde{r}_t(j) \tilde{r}_t(k) - X^\beta(j) \beta^T X^\beta(k)^T & \text{if } t \in N,(j,k) \in i_t \\
0 & \text{otherwise}
\end{cases} \)

The sample moment conditions are aggregated across time and stacked via the usual definitions; and parameter estimates are solved through the standard framework, which estimate deviations from zero under some weighting matrix \( W \):

\[
g(\beta) = \left[ \sum_t g_{t,(j=1,k=1)}(\beta) \cdots \sum_t g_{t,(j=1,k=C_r)}(\beta) \cdots \sum_t g_{t,(j=C_r,k=C_r)}(\beta) \right]^T = \arg \min_{\beta} g(\beta)^T W g(\beta)
\]

There are large computational issues that make this methodology largely inappropriate for the paper. In general, this is a very high-dimensional and a very non-linear problem. With respect to the former concern of dimensionality: for my benchmark specification with nine currencies, I have 45 moments; and for my bonds specification (in which I combine both less liquid intraday returns and more liquid daily returns), I have between 100-200 moments. By contrast, many papers in the literature that use Identification by Heteroskedasticity (e.g. Rigobon [2003], Rigobon and Sack [2003], Rigobon and Sack [2004], Craine and Martin [2008], Nakamura and Steinsson [2017], Hebert and Schreger [2017]) have between two and six moments. With respect to the latter concern of non-linearity: because I minimize quadratic deviations in moments and because those are in turn quadratic functions of parameters, my parameters are raised to the fourth power. The combination of these two means that convergence can no longer be taken for granted. In my specification with nine currencies (45 moments) and with the identity weighting matrix, my estimator converges – with results virtually identical to that of the Expectation Maximization algorithm. In my specifications with bonds (100+ moments) or with the optimal weighting matrix, it fails to converge.

There are some smaller reasons to prefer the Expectation Maximization approach to this approach. One is that convergence tends to be quicker, which can be useful when computing the many thousands of high-dimensional mapping permutations. Another is that time-varying covariate matrices are no longer problematic. A third is that the Expectation Maximization algorithm estimates the shocks alongside the coefficients, which can be useful for validating the approach.

### B.6 Fixed Effects

An alternate approach to identify the parameter \( \beta \) for Equation (16) would be to use fixed effects for coefficients, and ignore variation in monetary shocks over time. Again, consider a simplified version of Equation (16) in which asset returns are demeaned, monetary shocks \( m_t \) are univariate, there is no missing data, there is one coefficient per asset \( (C_r = C_\beta) \), and \( X^\beta_t = I_{C_r} \):

\[
\begin{bmatrix}
\epsilon_t / $ \\
r_t / $ \\
\tilde{r}_t / $
\end{bmatrix}
= \begin{bmatrix}
\beta \epsilon_t / $ \\
\beta r_t / $ \\
\beta \tilde{r}_t / $
\end{bmatrix}
\begin{bmatrix}
m_s \\
\epsilon_t / $ \\
\epsilon_t / $
\end{bmatrix}
\]

Since monetary shocks have zero mean, this approach regresses the absolute value of asset returns on a set of currency fixed effects (while continuing to parameterize the variance-covariance matrix of the errors, as in generalized least squares). Assets that systematically respond more to
monetary shocks will have larger absolute movements, and therefore will have larger coefficients.

\[
\begin{bmatrix}
\frac{\epsilon_t}{S} \\
\frac{\epsilon_t}{L} \\
\frac{\epsilon_t}{Y}
\end{bmatrix}
= \begin{bmatrix}
\beta \frac{\epsilon_t}{S} \\
\beta \frac{\epsilon_t}{L} \\
\beta \frac{\epsilon_t}{Y}
\end{bmatrix} + \begin{bmatrix}
\epsilon_t/S \\
\epsilon_t/L \\
\epsilon_t/Y
\end{bmatrix}
\]

This approach is similar in spirit to identification by heteroskedasticity, in that it identifies coefficients from the absolute asset return (instead of the squared asset return). However, it suffers from weak power. This is easy to show by simulation. I simulate 500 samples in which asset returns are driven by monetary shocks and by other noise. I compare the mean squared error of the coefficients estimated under the Expectation Maximization algorithm to those estimated under the Fixed Effects estimator. For “large” shocks (where the shocks are four times as large as the background noise), the two approaches are comparable, although the Expectation Maximization algorithm performs better. For “small” shocks (where shocks are twice as large as the background noise), the Expectation Maximization algorithm performs vastly better.

B.7 Vector Autoregression

This section will be written soon. Please check soon, and ensure you are reading the most recent version of the paper.

B.8 Measured Shocks

B.8.1 Traditional Measures

An alternate approach is to use traditional measures of monetary shocks, rather than estimating or inferring them. There are five popular ones: movements in short-term rates (e.g. Fed Funds futures) around announcements, movements in medium-term rates (e.g. the two-year Treasury) around announcements, differences between actual policy and surveyed expectations, shocks constructed through the narrative method (i.e. Romer and Romer shocks), and policy deviations from the Taylor Rule.
Four of these are immediately problematic; the exception is medium-term rates. Short-term measures and surveys focus only on the surprises delivered to the short end of the yield curve; but especially in the past decade, monetary surprises are most commonly delivered to the medium end of the yield curve. For instance, I graph how often surveys corrected anticipated the monetary announcement over my sample below. They were accurate 80+% of the time, and so a measure of surprises constructed from survey data would involve retaining 20% of the data at most.

Figure 14: Surveys and Monetary Shocks

Furthermore, the Romer & Romer narrative approaches are strongly sensitive to the methodology used; and most central banks have deviated greatly from the Taylor Rule in the past decade, making an approach that relies on it highly suspect. As such, I avoid these four traditional measures of monetary shocks. Regardless, they are still reasonably correlated with monetary shocks inferred from the Expectation Maximization algorithm. Below, I present the correlations between my inferred shocks and surveys, short rates, and (in the case of the United States), shocks from the Fed Funds futures, Nakamura and Steinsson [2017] shocks, and Romer & Romer shocks (updated by Coibion et al. [2017]).

The one exception are medium-term rates (e.g. two-year rates), which may solve the maturity and methodology concerns. There are still reasons to prefer inferred shocks to these measured shocks. A simple argument is that medium-term rates may still be insufficient to capture the entire path of shocks. A subtle but more important argument is that medium-term rates capture one component of Fed shocks, as applied to domestic medium-term assets. Currencies and foreign bonds capture other components of shocks, as applied to foreign and long-term assets. Correlating these two types of shocks results in weakened power; there are components of Fed announcements that are strongly relevant to foreign currencies but irrelevant to medium-term domestic bonds. I am not interested in the magnitude of $\beta$; I am only interested in its comparative properties, and so inferred shocks suits this study better. Regardless, I still utilize two-year Treasury yield shocks.
Table 10: Comparing Monetary Shocks

<table>
<thead>
<tr>
<th>Country</th>
<th>Survey</th>
<th>01M</th>
<th>01Y</th>
<th>FFR</th>
<th>N-S</th>
<th>R-R</th>
</tr>
</thead>
<tbody>
<tr>
<td>Australia</td>
<td>0.56**</td>
<td>0.55**</td>
<td>0.66**</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Canada</td>
<td>0.41**</td>
<td>0.33**</td>
<td>0.56**</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Switzerland</td>
<td></td>
<td>0.21</td>
<td>0.34**</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Euro</td>
<td>0.30**</td>
<td>-0.01</td>
<td>0.33**</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>United Kingdom</td>
<td>0.33**</td>
<td>0.00</td>
<td>0.28**</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Japan</td>
<td>-0.09</td>
<td>-0.02</td>
<td>0.26**</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Norway</td>
<td>0.46**</td>
<td>0.23*</td>
<td>0.51**</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>New Zealand</td>
<td>0.23*</td>
<td>0.37**</td>
<td>0.70**</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Sweden</td>
<td>0.45**</td>
<td>0.05</td>
<td>0.25*</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>United States</td>
<td>0.17</td>
<td>0.27**</td>
<td>0.49**</td>
<td>0.28**</td>
<td>0.50**</td>
<td>0.11</td>
</tr>
</tbody>
</table>

as a robustness check.

B.8.2 Average Return

An alternate approach to identify the parameter $\beta$ for Equation (16) would be to use take the average of $r_t$ as a measure of the monetary shock at time $t$. This is equivalent to regressing equity returns on the market return (i.e. their average return), as is common in the finance literature.

There are two concerns here: bias in coefficients and bias in standard errors. Consider a trivial example with a single asset return that is subsequently regressed on itself (i.e. the mean of itself). That regression would have a coefficient of one and standard errors of zero. As the number of assets increases, these two problems – coefficients biased to one and standard errors biased down – dissipate but do not vanish. Many of my specifications have nine assets, and there is a concern that nine is indeed insufficient. Moreover, since the errors of those assets are correlated, nine assets may be overstating the effective independence.

To show each of these biases in more depth, I simulate 500 samples and nine assets as before, and estimate coefficients using an oracle estimator (that has access to the true underlying shocks), the Expectation Maximization algorithm, and the Average Return method. In each sample, coefficients take a randomly chosen value $\in \{0.75, 1.00, 1.25\}$. I first explore the coefficient bias by restricting to the estimates where the underlying true coefficient is 0.75 or 1.25. I plot the density of coefficients obtained by the oracle estimator against the Expectation Maximization, and against the Average Return. There is a clear bias towards one in the case of the latter.
Second, I focus on the coefficients that are actually one, and plot the density of t-statistics for each method’s estimates. If standard errors are computed correctly, these should follow a standard normal distribution (e.g. 5% of the t-statistics should make a Type I error). For the Expectation Maximization algorithm, that happens. However, for the Average Return method, that does not happen – the standard errors are too small, and so the Type I errors made are far too numerous.
Figure 15: Market Reactions to US Monetary Shocks (2Y Treasury)

(a) Currencies
(b) Bonds
(c) Cross-Border Bond Portfolios

Notes: The figures depict the reactions of currencies, bonds, and cross-border bond portfolios to monetary announcements in the US, where monetary shocks are high-frequency movements in the two-year Treasury rather than the inferred shocks used throughout the paper. The currency figure shows by how much the dollar appreciates against a given reference currency when it appreciates by 1% on average. The bond figure shows by how much the foreign yields of other countries rise when US yields rise by 1%. Finally, the cross-border bond portfolio figure shows by how much a portfolio that shorts a given country’s ten-year bond and lends at the US riskfree rate appreciates when the average portfolio appreciates by 1%. Standard error bars in all pictures are computed versus the mean reaction across all currencies, bonds, or portfolios; and the shading of the coefficient bars refers to the lower-dimensional structure, whereby currencies, bonds, or portfolios of the same color (different colors) react similarly (dissimilarly).
Figure 16: Market Reactions to Monetary Shocks (EM vs GMM)

(a) Currency Responses to US Shocks (EM)
(b) Currency Responses to US Shocks (GMM)
(c) Currency Responses to EU Shocks (EM)
(d) Currency Responses to EU Shocks (GMM)

Notes: The figures depict the reactions of currencies to monetary announcements in the US and in the Eurozone, solved through either the EM algorithm (the paper’s preferred methodology) or through GMM per an Identification by Heteroskedasticity setup. The currency figures show by how much the dollar (euro) appreciates against a given reference currency when it appreciates by 1% on average following announcements by the Fed (ECB). Standard error bars in all pictures are computed versus the mean reaction across all currencies.
C Long-Run Inflation

In this section, I test the validity of the approximations in Equation (6). The exact equation for the cross-border bond portfolio is below, and it includes three additional terms: nominal interest rates in the US and in the foreign country at horizons longer than ten years, and the infinite-horizon nominal exchange rate. Potentially, asymmetries in the cross-border bond portfolio could be driven by these additional terms, within the framework of central banks reacting or premia shifting under complete markets. I argue in this section that these additional terms are unlikely to react at all to the Fed.

\[
\Delta E_t \sum_{k=1}^{10} \gamma^{i/S}_{t+k-1} - \Delta E_t \sum_{k=1}^{10} \gamma^\$_{t+k-1} + \Delta E_t \sum_{k=1}^\infty \pi^{i/S}_{t+k-1} + \Delta E_t \sum_{t=1}^\infty \rho^{i/S}_{t+k-1} + \Delta E_t \sum_{k=1}^\infty \rho^\$_{t+k-1}
\]

Cross-Border Bond Portfolio

While the assumption I rely on is that foreign components of these three terms do not move following Fed announcements, I will argue more generally that the path of interest rates beyond ten years or infinite-horizon exchange rates are unchanged regardless. First, there is no evidence to suggest central banks set policy ten or more years in advance. There are three known statements during the forward guidance period in which the Fed explicitly provided calendar-based guidance over its future actions. In 2011, the Fed promised low rates until 2013; in 2012, the Fed promised low rates until 2014; and later in 2012, the Fed promised low rates until 2015. None of these statements come close to a ten-year horizon. Other central banks are more coy, and do not cite specific dates. Nevertheless, I find a few vague but key phrases: “extended period of time” for the ECB, “over the next few years” for the Bank of England, and “for the time being” for the Bank of Japan. None of these phrases suggests ten-year horizons either. Finally, in the appendix, I show that my empirical results are robust when limiting to the pre-crisis era, during which central banks definitely did not provide guidance at long horizons.

To test this assumption further, I break the three additional terms into two sets of components: real factors and price factors.

\[
\Delta E_t \sum_{k=1}^\infty \pi^\$_{t+k-1} - \Delta E_t \sum_{k=1}^\infty \pi^{i/S}_{t+k-1} + \Delta E_t \sum_{t=1}^\infty \rho^{i/S}_{t+k-1} + \Delta E_t \sum_{k=1}^\infty \rho^\$_{t+k-1}
\]

US Path of Rates Foreign Path of Rates

First, consider the real factors: real interest rates ten years from now, and real exchange rates. Although real interest rates may undergo secular shifts and although purchasing power parity
might fail in the long run, these properties are driven by fundamentals (e.g. demographic shifts, technological improvements, etc) and not by monetary news. Models commonly show monetary neutrality to real variables in the long run, once prices adjust. Even Nakamura and Steinsson [2017], which argue for monetary non-neutrality at distant horizons, still note that neutrality is restored at the ten-year horizon.

Price factors could be more troublesome in principle, as these fall directly under the remits of central banks. I will argue that empirically, long-run inflation forecasts vary too little to explain my effects. Furthermore, I introduce high-frequency evidence from the TIPS (Treasury inflation-protected securities) bond market on Fed announcement days in the next section, to show that long-run inflation expectations do not react to Fed announcements.

There are few distant inflation forecasts, but the existing ones do not vary enough to explain my results. These forecasts change a few basis points per year, whereas my results find that bond yields move a few basis points per announcement. The Fed makes eight announcements per year and releases the majority of its monetary news outside of announcement windows (e.g. through speeches and meeting minutes, and through anticipatory forecasts following inflation and unemployment releases).

First, I consider the IMF’s World Economic Outlook, which has been making five-year inflation forecasts for my ten countries for the previous decade. I look at the median absolute revision in the forecast as the five-year ahead forecast this year becomes the four-year ahead forecast next year. This revision is approximately three basis points per year. Second, I look at the Fed’s Survey of Professional Forecasters and find that their median absolute revision for the five-year ahead forecast is six basis points per year. Finally, I consider the European Central Bank’s forecast, and their median absolute revision for the five-year ahead forecast is zero basis points.

These results are consistent with the literature. Nakamura and Steinsson [2017] and Hanson and Stein [2015] both find quantitatively small expected inflation responses on Fed announcement days in the US data. My paper relies on a weaker condition: quantitatively small expected inflation responses on Fed announcement days ten years away in foreign markets.

I collect returns on TIPS (Treasury inflation-protected securities) from the Federal Reserve Economic Data database.

In the previous section, I argue that inflation expectations at long horizons do not adjust routinely, and so in this context, it is unlikely that movements in forward yields reflect adjustments in expected inflation. I further argue this using returns in the Treasury inflation-protected securities (TIPS) market. I show that Fed announcements drive nominal US yields and real US yields, but fail to drive the difference, which reflects expectations of inflation (and inflation premia). Ideally, I would test this using foreign real yields rather than US real yields; but those markets are less liquid. Since US monetary policy should have the strongest effects on domestic assets, I consider my findings conservative. Table 11 shows that the Fed affects nominal and real yields for seven-year forward three-year and ten-year forward twenty-year maturities, but fails to affect the difference.27

---

27The seven-year TIPS yield is a benchmark rate that is updated daily, so I utilize that instead of the less liquid six-year TIPS yield.
Table 11: Excess Volatility in Nominal and Real Returns

<table>
<thead>
<tr>
<th></th>
<th>Nominal</th>
<th>Real (TIPS)</th>
<th>Inflation</th>
</tr>
</thead>
<tbody>
<tr>
<td>7F3Y</td>
<td>42%</td>
<td>24%</td>
<td></td>
</tr>
<tr>
<td>10F20Y</td>
<td>51%</td>
<td>33%</td>
<td></td>
</tr>
</tbody>
</table>

Notes: The table tests whether six types of bonds are more volatile around announcements by the Fed than at other times, using daily returns. The six bonds are the nominal yield, the TIPS yield, and the difference between the two, for both the seven-year forward three-year bond and the ten-year forward twenty-year bond. If returns are more volatile at the 1% level using the Brown-Forsythe test, the cell records the excess ratio of standard deviations (announcement window to non-announcement window standard deviations, minus 100%). If returns are not statistically more volatile, the cell is left blank. The Fed affects both nominal and TIPS (i.e. real) yields, but does not affect the differences (proxies for expected inflation), at long forward maturities.

D Characterizing Asymmetric Responses

This section discusses different indicators that successfully the asymmetric responses in currency markets, bond markets, and long-maturity carry portfolios. The paper occasionally characterizes the cross-section of responses through the heuristic of high-rate and low-rate countries, but this section tests more rigorously that heuristic and other possible explanations. The results here are suggestive of the actual mechanism behind asymmetric responses to Fed announcements, and should guide future work in identifying the exact channels. This section first introduces the methodology used, and then discusses the various indicators employed.

D.1 Methodology

The methodology takes the Factor Model discussed in Section 3, and replaces currency-specific coefficients with coefficients interacted with predictor variables. In the three-currency example used throughout the paper, I estimate $\beta_1$ from the following equation for a given measured indicator $X^i_t$:

$$
\begin{bmatrix}
\Delta s^\varepsilon_t/\$ \\
\Delta s^\varepsilon_t/L \\
\Delta s^L_t/\$ \\
\end{bmatrix}
= 
\begin{bmatrix}
1 & X^\varepsilon_t/\$ \\
1 & X^{L,\$} \\
1 & X^{L,\$} \\
\end{bmatrix}
\begin{bmatrix}
\alpha_0 \\
\alpha_1 \\
\end{bmatrix}
+ 
\begin{bmatrix}
1 & X^\varepsilon_t/\$ \\
1 & X^{L,\$} \\
1 & X^{L,\$} \\
\end{bmatrix}
\begin{bmatrix}
\beta_0 \\
\beta_1 \\
\end{bmatrix}
m^L_t + 
\begin{bmatrix}
\varepsilon_t/\$ \\
\varepsilon_t/L \\
\varepsilon_t/\$ \\
\end{bmatrix}
$$

A statistically significant $\beta_1$ means that the indicator $X^i_t$ is useful primarily for predicting the cross-sectional variation in responses to Fed monetary announcements, as well as the (much smaller) time variation in responses. I estimate Equation (18) for the three groups of assets discussed in the paper: currencies, foreign bonds, and long-maturity portfolios (that combine both a short position in a foreign currency with a short position in the foreign bonds). I consider an indicator variable to be predictive if it is statistically significant at the 5% level in each of these three specifications.

D.2 Results

I conduct the analysis on nine classes of predictors that vary across countries: the level of interest rates, measures of local volatility, deviations in CIP arbitrage, trade flows versus the US, dollar invoicing of trade, bank positions versus the US, portfolio debt positions versus the US, portfolio equity positions versus the US, and distance to the US. Of these, interest rates and some financial quantities are the most promising.
D.2.1 Interest Rates

I let $X_i^t$ measure an interest rate differential against the US for country $i$ at time $t$, and I do so using different maturities: one-month rates, one-year rates, five-year rates, and ten-year rates. All of these measures are significant in all specifications, and so this nomenclature is used throughout the paper. When the Fed makes an announcement, currencies of high-rate countries move against currencies of low-rate countries, yields of high-rate countries move more than yields of low-rate countries, and the returns on long-maturity portfolios of high-rate countries move more than the returns on long-maturity portfolios of low-rate countries.

Because of the central importance of this indicator, I check robustness by dropping each individual currency sequentially and recomputing the estimate. Across all maturities and specifications with one exception, the coefficient remains highly significant. For the specification with currencies and for very short maturities (e.g. one-month), significance can be only found at the 10% level on occasion.

D.2.2 Volatility

I let $X_i^t$ measure some metric of volatility compared to the US for country $i$ at time $t$, and I do so using four different metrics. The first metric is historical 30-day local equity market volatility minus historical 30-day US equity market volatility, and it is the historical analog to the VIX. (This data is sourced from Datastream.) The second, third, and fourth metrics use currency options data, sourced from Bloomberg. I extract implied volatility from the 25-delta call, the 50-delta call, and the 75-delta call (which roughly correspond to calls that have a 25%, 50%, and 75% chance of expiring in the money). The second metric is the implied volatility from the 50-delta call, which is an even-handed measure of expected currency volatility. The third metric is the difference in implied volatility between the 25-delta and 75-delta call, which is the volatility skew (or difference in volatility between options that appreciate and depreciate in value). The final metric is the difference in volatility between the average of the 25-delta and 75-delta call and the 50-delta call, which is the volatility kurtosis (or difference in volatility on the tails and volatility in the center of the distribution).

Of these four measures, the volatility skew is significant in all specifications. When the Fed makes an announcement, currencies with high skew move against currencies with low skew, yields of high-skew currencies move more than yields of low-skew currencies, and the returns on long-maturity portfolios of high-skew currencies move more than the returns of long-maturity portfolios of low-skew currencies. In particular, Australia, New Zealand, and Norway have high skew while Japan and Switzerland have low skew, as the former are seen as risky currencies with strong downside volatility (relative to upside), while the latter are seen as safe-haven currencies with limited downside volatility relative to upside.

D.2.3 Limits to Arbitrage

I let $X_i^t$ measure the cross-currency basis for that currency versus the dollar over ten years, sourced from Bloomberg. This measures the deviation from covered interest parity, and so should proxy for limited arbitrage in those currency markets. However, this measure is insignificant with respect to currencies and so not pursued further. Moreover, my empirical findings can be seen in both the pre-crisis and post-crisis eras, whereas the cross-currency basis was virtually zero prior to the financial crisis.
D.2.4 Trade Flows

I let $X^i_t$ measure trade flows against the US, sourced from the IMF’s Direction of Trade Statistics (DOTS) database. I construct two measures: one is a gross measure of how important trade with the US is for a given country, and a second is a net measure of trade against the US, comparing exports to imports. The gross measure is constructed as follows for a given country $i$:

$$X^i_t = \frac{\text{Exports}^i_{\to \text{US}} + \text{Imports}^i_{\to \text{US}}}{\sum_c (\text{Exports}^i_{\to c} + \text{Imports}^i_{\to c})}$$

The net measure is constructed as follows for a given country $i$:

$$X^i_t = \frac{\text{Exports}^i_{\to \text{US}} - \text{Imports}^i_{\to \text{US}}}{\text{Exports}^i_{\to \text{US}} + \text{Imports}^i_{\to \text{US}}}$$

The gross measure is statistically insignificant, but the net measure is significant. When the Fed makes an announcement, currencies whose countries have high bilateral imports versus the US move against currencies whose countries have high bilateral exports, yields of high-import countries move more than yields of high-export countries, and the returns on long-maturity portfolios of high-import countries move more than the returns of long-maturity portfolios of high-export countries. In particular, Australia and New Zealand import heavily from the US, while most other countries primarily export to the US.

D.2.5 Dollar Invoicing

I let $X^i_t$ measure the fraction of a given country $i$’s trade invoiced in dollars, based on work by Gopinath [2015]. Time variation in these series are both limited and poorly estimated, and so I only look at the cross-sectional variation here. However, this is insignificant for at least one of the three assets, whether I look at dollar invoicing in exports, dollar invoicing in imports, or total dollar invoicing. In general, the Pacific countries of Australia, Canada, and Japan, as well as Norway (which trades substantially in oil) have relatively high dollar shares, while the European countries have relatively low dollar shares. In my results, Australian assets react more strongly than European assets, which in turn react more strongly than Japanese assets. This creates a non-monotonic pattern and thus dollar invoicing is not a statistically significant predictor.

D.2.6 Bank Positions

I let $X^i_t$ measure bank positions against the US, sourced from the BIS’s Locational Banking Statistics (LBS) database. I construct two measures that parallel the measures in trade: a gross measure and a net measure. The gross measure is constructed to be the share of a given country’s assets and liabilities held against the US over total assets and liabilities:

$$X^i_t = \frac{\text{Assets}^i_{\to \text{US}} + \text{Liabilities}^i_{\to \text{US}}}{\sum_c (\text{Assets}^i_{\to c} + \text{Liabilities}^i_{\to c})}$$

The net measure is constructed to be the difference between assets and liabilities against the US, over the total position against the US:

$$X^i_t = \frac{\text{Assets}^i_{\to \text{US}} - \text{Liabilities}^i_{\to \text{US}}}{\text{Assets}^i_{\to \text{US}} + \text{Liabilities}^i_{\to \text{US}}}$$
The gross measure is statistically insignificant, but the net measure is significant. When the Fed makes an announcement, currencies whose countries have high liability positions versus the US move against currencies whose countries have high asset positions, yields of high-liability countries move more than yields of high-asset countries, and the returns on long-maturity portfolios of high-liability countries move more than the returns of long-maturity portfolios of high-asset countries. Consistent with what we might expect from an interest rate spread, Australia has high liability positions and Japan has high asset positions.

D.2.7 Portfolio Debt Positions

I let $X^d_i$ measure portfolio debt positions against the US, sourced from the IMF’s Coordinated Portfolio Investment Survey (CPIS) database. I construct two measures that parallel the measures in trade and in bank positions: a gross measure and a net measure. As before, the gross measure is the share of a given country’s assets and liabilities held against the US over total assets and liabilities, while the net measure is the difference between assets and liabilities against the US over the total position against the US.

Both the gross and net measures are statistically insignificant for at least one of the specifications. For example, Australia and Canada alike have strongly negative net positions in portfolio debt versus the US, despite having strongly different reactions in currency markets.

D.2.8 Portfolio Equity Positions

I let $X^e_i$ measure portfolio equity positions against the US, sourced from the IMF’s Coordinated Portfolio Investment Survey (CPIS) database. Once more, I construct two measures that parallel the measures in trade, in bank positions, and in portfolio debt positions: a gross measure and a net measure. As before, the gross measure is the share of a given country’s assets and liabilities held against the US over total assets and liabilities, while the net measure is the difference between assets and liabilities against the US over the total position against the US.

The gross measure is statistically insignificant, but the net measure is significant. When the Fed makes an announcement, currencies whose countries have high asset positions versus the US move against currencies whose countries have high liabilities positions, yields of high-asset countries move more than yields of high-liability countries, and the returns on long-maturity portfolios of high-asset countries move more than the returns of long-maturity portfolios of high-liability countries. Surprisingly, many of the investment currencies in the carry trade (e.g. Norway and New Zealand) actually have strong asset positions in US equities, while Switzerland and Japan have large liability positions. This may be because the countries with higher interest rates are smaller, and have underdeveloped local equity markets.

D.2.9 Distance

Finally, I let $X^d_i$ measure the distance between the US and a given country, based on work by Lustig and Richmond [2017]. There is no time variation in this measure. This is statistically significant: the closest countries are Canada and the UK, and the furthest Australia and New Zealand. (Japan is an outlier in this regard.) Regardless, when the Fed makes an announcement, currencies of distant countries move against currencies of close countries, yields of distant countries move more than yields of close countries, and the returns on long-maturity portfolios of distant countries move more than the returns of long-maturity portfolios of close countries.

However, Lustig and Richmond [2017] find distance to be a significant predictor of asymmetric responses in currencies for all countries. As such, I extend this methodology to all other central
banks, for versions of Equation (18) involving currencies only. The results are decidedly more mixed. In Australia, the Eurozone, and New Zealand, distance is a significant predictor. Australia and New Zealand have small effects on each other’s currencies, and the ECB has strong effects on other continental European currencies. But for the UK and Canada, the results go the opposite way: both these countries have small effects on the Australian dollar and New Zealand dollar, despite being very distant. Finally, the effects are insignificant for the central banks of the other four countries: Japan, Norway, Sweden, and Switzerland. More generally, currency asymmetries beyond the Fed and ECB are extremely limited, and so the effects that Lustig and Richmond [2017] identify may be more related to other global shocks than to tightly identified monetary shocks.
E Gaussian Affine Term Structure Model

This section documents the steps used to decompose the bond yield curve into the path of short rates and the path of term premia, following the structural model of Adrian et al. [2013], hereafter ACM. ACM use a Gaussian affine term structure model for this decomposition. This article references their paper’s equations as a guide for the reader.

1. Assume we have risk neutral bond pricing parameters $A^{RF}_n$ and $B^{RF}_n$. In the final paragraph of Section 2.4, ACM note that the time average of future short rates over the next $n$ months is $-\frac{1}{n-1}(A^{RF}_n + B^{RF}_n X_t)$. Similarly, the difference between yields and these rates is the average of term premia over the next $n$ periods.

2. In the final paragraph of Section 2.4, ACM notes that setting the risk parameters $\lambda_0$ and $\lambda_1$ in Equations 25 and 26 (written below) will generate the risk-neutral bond pricing parameters $A^{RF}_n$ and $B^{RF}_n$.

$$A_n = A_{n-1} + B'_{n-1} (\mu - \lambda_0) + \frac{1}{2} (B'_{n-1} \Sigma B_{n-1} + \sigma^2) - \delta_0, n = 1, ..., 120$$

$$B'_n = B'_ {n-1} (\Phi - \lambda_1) - \delta'_1, n = 1, ..., 120$$

As such, I need to estimate $(\mu, \Sigma, \sigma^2, \delta_0, \delta_1, \Phi)$ to build this relation recursively.

3. Start with $\delta_0$ and $\delta_1$. In the final paragraph of Section 3.1, ACM suggest regressing the one-month Treasury bill on pricing factors $X_t$, where $\delta_0$ is the constant and $\delta_1$ the coefficients. Implicitly, they assume that there is no risk premia at the shortest end of the maturity curve, and so a regression of observed yields directly on the state variable (without needing to adjust for risk) will uncover $A_1$ and $B_1$:

$$r_{t+1} = -A_1 - B_1 X_t$$

Moreover, ACM note that $A_1 = -\delta_0$ and $B_1 = -\delta_1$ when delivering the exposition for Equations 25 and 26.

4. Now consider $\mu$ and $\Phi$. These are estimated from Equation 1, as the parameters in a VAR regression of state variables $X_t$ (a $K \times 1$ dimensional vector at time $t$):

$$X_{t+1} = \mu + \Phi X_t + v_{t+1}$$

I discuss the method of gathering the state variables shortly.

5. $\Sigma$ is defined as the variance-covariance matrix from which $v_{t+1}$ (errors in the previous estimation) are drawn, such that $\Sigma = \frac{1}{T} \sum_{t=1}^{T} v_{t+1}v_{t+1}'$.

6. $\sigma^2$ is defined through a more complex process, described in the second step of the three-step procedure described in Section 2.2. After conducting the following regression, $\hat{\sigma}^2 = \text{Tr}(\hat{E}\hat{E}')/(NT)$; that is, $\sigma^2$ is measured as the average volatility in the errors.

$$rx = a^{\prime}_T + \beta' \hat{V} + cX_- + E$$

In turn, each of these components is defined as follows:

- $i_T$ is a vector of ones.
• $\hat{V}$ is the stacked errors $v_{t+1}$ over time (explained in the first step of Section 2.2’s three-step procedure).

• $X_-$ is the lagged state variable.

• $a, \beta, c$ are the parameters in this regression, and are estimated.

• $rx$ refers to the excess returns to holding a longer maturity bond over the one-month bond, over a one month holding pattern. In Section 3.1, ACM calculate excess returns from holding the $n = 6, 12, 18, 24, ..., 60, 84, 120$ month zero-coupon bonds for one month (at which point they are $n = 5, 11, ..., 119$ month bonds), versus holding the $n = 1$ month zero-coupon bond.

It is important to note that this regression is done at a monthly level, and so only the end-of-the-month values for state variables $X_-$ and errors $\hat{V}$ are used to predict end-of-month excess returns. It is also worthwhile noting that this parameter seems to have little impact on the final results, despite the relative complexity in generating it.

7. Finally, I define the state variables matrix $X_t$. ACM report in Section 3.1 that, in their baseline specification, they extract principal components from the cross-section of yields every three months; and repeatedly, they stress the importance of five factors. Furthermore, in Section 4.4, they discuss using measuring this cross-section of yields at the monthly frequency to get the principal components, and then applying these weights to the cross-section of yields measured at a daily frequency to get daily factors. I replicate their methodology almost exactly, computing the five leading eigenvectors of the cross-section of the yield curve measured at a lower frequency, and projecting the cross-section measured at a daily frequency across those eigenvectors. However, I make one small difference. Because I have less data than ACM (they have twice as many years as me), I use yields measured at the weekly (rather than monthly) frequency to extract the eigenvectors. Since I am estimating a $40 \times 40$ covariance matrix, it is safer to use 800 rather than 200 cross-sectional observations.

Below, I present the results for the decomposition of international yield curves for my ten countries into the average rate and average term premia over a ten-year horizon, annualized. They all show similar trends – a sharp drop in the path of rates during the financial crisis that largely persists until today, and increases in term premia at that time that was slowly undone over the subsequent decade.
F Robustness Checks

Table 12: Excess Volatility in 10Y Bond Returns

<table>
<thead>
<tr>
<th>Country</th>
<th>AUD</th>
<th>CAD</th>
<th>CHF</th>
<th>EUR</th>
<th>GBP</th>
<th>JPY</th>
<th>NOK</th>
<th>NZD</th>
<th>SEK</th>
<th>USD</th>
</tr>
</thead>
<tbody>
<tr>
<td>Australia</td>
<td>144**</td>
<td>2</td>
<td>1</td>
<td>6</td>
<td>10*</td>
<td>41*</td>
<td>12</td>
<td>19*</td>
<td>6</td>
<td>2</td>
</tr>
<tr>
<td>Canada</td>
<td>30**</td>
<td>82**</td>
<td>22</td>
<td>24*</td>
<td>4</td>
<td>0</td>
<td>−15</td>
<td>3</td>
<td>21*</td>
<td>3</td>
</tr>
<tr>
<td>Switzerland</td>
<td>−1</td>
<td>−2</td>
<td>124**</td>
<td>−5</td>
<td>−20</td>
<td>85*</td>
<td>27*</td>
<td>96**</td>
<td>14</td>
<td>3</td>
</tr>
<tr>
<td>Euro</td>
<td>34**</td>
<td>24*</td>
<td>39**</td>
<td>88**</td>
<td>37**</td>
<td>8</td>
<td>133**</td>
<td>7</td>
<td>90**</td>
<td>24**</td>
</tr>
<tr>
<td>United Kingdom</td>
<td>2</td>
<td>−12</td>
<td>15*</td>
<td>10</td>
<td>79**</td>
<td>−11</td>
<td>−3</td>
<td>3</td>
<td>26*</td>
<td>−11</td>
</tr>
<tr>
<td>Japan</td>
<td>15</td>
<td>−1</td>
<td>−8</td>
<td>7</td>
<td>−1</td>
<td>34**</td>
<td>4</td>
<td>−8</td>
<td>−2</td>
<td>27**</td>
</tr>
<tr>
<td>Norway</td>
<td>14</td>
<td>17</td>
<td>8</td>
<td>8</td>
<td>2</td>
<td>7</td>
<td>26**</td>
<td>34**</td>
<td>22</td>
<td>21*</td>
</tr>
<tr>
<td>New Zealand</td>
<td>27</td>
<td>1</td>
<td>9</td>
<td>29</td>
<td>9</td>
<td>27*</td>
<td>−5</td>
<td>26**</td>
<td>−11</td>
<td>−1</td>
</tr>
<tr>
<td>Sweden</td>
<td>4</td>
<td>19</td>
<td>12</td>
<td>39**</td>
<td>7</td>
<td>−1</td>
<td>25</td>
<td>−1</td>
<td>82**</td>
<td>3</td>
</tr>
<tr>
<td>United States</td>
<td>207**</td>
<td>144**</td>
<td>40**</td>
<td>216**</td>
<td>46**</td>
<td>2</td>
<td>21*</td>
<td>54**</td>
<td>25**</td>
<td>233**</td>
</tr>
</tbody>
</table>

Notes: The table tests whether returns of the column country’s ten-year sovereign bonds are more volatile around announcements by the row central bank than at other times. Dark grey refers to returns in sixty-minute windows; and light grey refers to returns in daily windows when bond markets of that country are too illiquid at that time to accurately compute returns over sixty-minute windows. The cell shows the excess ratio of standard deviations for that asset (announcement window standard deviation over non-announcement window standard deviation, minus 100%). Significance is assessed at the 1% (**) and 5% (*) level by the Brown-Forsythe test.

Table 13: Excess Volatility in Daily 1Y Bond Returns

<table>
<thead>
<tr>
<th>Country</th>
<th>AUD</th>
<th>CAD</th>
<th>CHF</th>
<th>EUR</th>
<th>GBP</th>
<th>JPY</th>
<th>NOK</th>
<th>NZD</th>
<th>SEK</th>
<th>USD</th>
</tr>
</thead>
<tbody>
<tr>
<td>Australia</td>
<td>86**</td>
<td>−20</td>
<td>−2</td>
<td>−16</td>
<td>−10</td>
<td>1</td>
<td>−10</td>
<td>18</td>
<td>−18</td>
<td>−13</td>
</tr>
<tr>
<td>Canada</td>
<td>−12</td>
<td>67**</td>
<td>5</td>
<td>−1</td>
<td>−6</td>
<td>18</td>
<td>11</td>
<td>13</td>
<td>−11</td>
<td>−5</td>
</tr>
<tr>
<td>Switzerland</td>
<td>14</td>
<td>−17</td>
<td>108**</td>
<td>−13</td>
<td>13</td>
<td>11</td>
<td>67</td>
<td>36**</td>
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<td>12</td>
</tr>
<tr>
<td>Euro</td>
<td>4</td>
<td>−17</td>
<td>45**</td>
<td>53**</td>
<td>−4</td>
<td>−27</td>
<td>12*</td>
<td>18</td>
<td>22*</td>
<td>12</td>
</tr>
<tr>
<td>United Kingdom</td>
<td>28**</td>
<td>−23</td>
<td>29**</td>
<td>−16</td>
<td>22</td>
<td>−11</td>
<td>21*</td>
<td>9</td>
<td>13</td>
<td>2</td>
</tr>
<tr>
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<td>−3</td>
<td>−8</td>
<td>5</td>
<td>−2</td>
<td>104**</td>
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<td>1</td>
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<td>2</td>
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<tr>
<td>Norway</td>
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<td>−16</td>
<td>12</td>
<td>1</td>
<td>−3</td>
<td>9</td>
<td>131**</td>
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<td>−18</td>
<td>50*</td>
</tr>
<tr>
<td>New Zealand</td>
<td>53**</td>
<td>−3</td>
<td>−5</td>
<td>4</td>
<td>9</td>
<td>4</td>
<td>−33</td>
<td>92**</td>
<td>−25</td>
<td>−4</td>
</tr>
<tr>
<td>Sweden</td>
<td>14</td>
<td>6</td>
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<td>−1</td>
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<td>−1</td>
<td>12</td>
<td>−20</td>
<td>10</td>
<td>79**</td>
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</tbody>
</table>

Notes: The table tests whether daily returns of the column country’s one-year sovereign bonds are more volatile around announcements by the row central bank than at other times. The cell shows the excess ratio of standard deviations for that asset (announcement window standard deviation over non-announcement window standard deviation, minus 100%). Significance is assessed at the 1% (**) and 5% (*) level by the Brown-Forsythe test.
### Table 14: Excess Volatility in Daily 6F4Y Bond Returns

<table>
<thead>
<tr>
<th>Country</th>
<th>AUD</th>
<th>CAD</th>
<th>CHF</th>
<th>EUR</th>
<th>GBP</th>
<th>JPY</th>
<th>NOK</th>
<th>NZD</th>
<th>SEK</th>
<th>USD</th>
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</tr>
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<td>−18</td>
<td>14</td>
<td>−19</td>
<td>−8</td>
</tr>
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<td>30*</td>
<td>18</td>
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<td>17</td>
<td>8</td>
<td>15</td>
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<td>28**</td>
<td>41*</td>
</tr>
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<td>3</td>
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<td>32**</td>
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<td>0</td>
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<td>−4</td>
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</tr>
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<td>−17</td>
<td>−12</td>
<td>13</td>
<td>−12</td>
<td>−6</td>
</tr>
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<td>3</td>
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<td>−6</td>
<td>1</td>
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</tr>
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<td>−5</td>
<td>15*</td>
<td>38**</td>
<td>6</td>
<td>−6</td>
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<td>10</td>
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<td>0</td>
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<td>−2</td>
<td>15*</td>
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</tr>
<tr>
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<td>36**</td>
<td>30**</td>
<td>50**</td>
<td>29**</td>
<td>7</td>
<td>18</td>
<td>57**</td>
<td>29**</td>
<td>51**</td>
</tr>
</tbody>
</table>

Notes: The table tests whether daily returns of the column country’s six-year forward four-year sovereign bonds (e.g., the rate one can guarantee from 2023 to 2027, in 2017) are more volatile around announcements by the row central bank than at other times. The cell shows the excess ratio of standard deviations for that asset (announcement window standard deviation over non-announcement window standard deviation, minus 100%). Significance is assessed at the 1% (***) and 5% (*) level by the Brown-Forsythe test.

### Table 15: Excess Volatility in Daily 10F20Y Bond Returns

<table>
<thead>
<tr>
<th>Country</th>
<th>AUD</th>
<th>CAD</th>
<th>CHF</th>
<th>EUR</th>
<th>GBP</th>
<th>JPY</th>
<th>SEK</th>
<th>USD</th>
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<td>7</td>
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</tr>
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<td>Canada</td>
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<td>−14</td>
<td>−1</td>
<td>−14</td>
<td>−24</td>
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<td>−7</td>
<td></td>
</tr>
<tr>
<td>Switzerland</td>
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<td>6</td>
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<td>−12</td>
<td>18</td>
<td>−1</td>
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<tr>
<td>Euro</td>
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<td>8</td>
<td>11*</td>
<td>21*</td>
<td>6</td>
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<td></td>
</tr>
<tr>
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<td>−1</td>
<td>−6</td>
<td>−10</td>
<td>2</td>
<td>−7</td>
<td>−4</td>
<td></td>
</tr>
<tr>
<td>Japan</td>
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<td>3</td>
<td>5</td>
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<td>−7</td>
<td>−6</td>
<td>9*</td>
<td></td>
</tr>
<tr>
<td>United States</td>
<td>18*</td>
<td>32**</td>
<td>36**</td>
<td>32**</td>
<td>25*</td>
<td>−9</td>
<td>51**</td>
<td></td>
</tr>
</tbody>
</table>

Notes: The table tests whether daily returns of the column country’s ten-year forward twenty-year sovereign bonds are more volatile around announcements by the row central bank than at other times. The cell shows the excess ratio of standard deviations for that asset (announcement window standard deviation over non-announcement window standard deviation, minus 100%). Significance is assessed at the 1% (***) and 5% (*) level by the Brown-Forsythe test. Norway, Sweden, and New Zealand do not issue thirty-year bonds and are omitted.
Table 16: Excess Volatility in Daily 10Y Rate Returns

<table>
<thead>
<tr>
<th>Country</th>
<th>AUD</th>
<th>CAD</th>
<th>CHF</th>
<th>EUR</th>
<th>GBP</th>
<th>JPY</th>
<th>NOK</th>
<th>NZD</th>
<th>SEK</th>
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<td>7</td>
<td>−15</td>
<td>−5</td>
</tr>
<tr>
<td>Canada</td>
<td>0</td>
<td>82**</td>
<td>−18</td>
<td>0</td>
<td>−7</td>
<td>−2</td>
<td>3</td>
<td>19</td>
<td>3</td>
<td>−12</td>
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<tr>
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<td>91**</td>
<td>25</td>
<td>22*</td>
<td>40*</td>
<td>92**</td>
<td>19</td>
<td>24*</td>
<td>30</td>
</tr>
<tr>
<td>Euro</td>
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<td>−1</td>
<td>17*</td>
<td>38**</td>
<td>−1</td>
<td>−5</td>
<td>11*</td>
<td>−7</td>
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<td>13</td>
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<td>92**</td>
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<td>−1</td>
<td>21**</td>
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<td>20</td>
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<td>−13</td>
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<td>9</td>
</tr>
<tr>
<td>New Zealand</td>
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<td>−9</td>
<td>8</td>
<td>−17</td>
<td>−6</td>
<td>28*</td>
<td>−20</td>
<td>102**</td>
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<td>3</td>
<td>14*</td>
<td>97**</td>
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</table>

Notes: The table tests whether daily returns of the column country’s model-estimated ten-year path of rates are more volatile around announcements by the row central bank than at other times. The cell shows the excess ratio of standard deviations for that asset (announcement window standard deviation over non-announcement window standard deviation, minus 100%). Significance is assessed at the 1% (**) and 5% (*) level by the Brown-Forsythe test.

Table 17: Excess Volatility in Daily 10Y Term Returns

<table>
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<tr>
<th>Country</th>
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<th>CHF</th>
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<th>GBP</th>
<th>JPY</th>
<th>NOK</th>
<th>NZD</th>
<th>SEK</th>
<th>USD</th>
</tr>
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<td>0</td>
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<td>18</td>
<td>−2</td>
<td>−4</td>
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<td>−18</td>
<td>−14</td>
<td>11</td>
<td>−27</td>
<td>21</td>
<td>−18</td>
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<td>−6</td>
<td>0</td>
<td>−1</td>
<td>5</td>
<td>21**</td>
</tr>
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<td>7</td>
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<td>−6</td>
<td>7</td>
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<td>−1</td>
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<td>7</td>
<td>21*</td>
<td>15</td>
<td>−6</td>
<td>46**</td>
<td>20**</td>
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<td>−27</td>
</tr>
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<td>19</td>
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<td>−27</td>
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<td>14</td>
<td>30**</td>
<td>17**</td>
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</table>

Notes: The table tests whether daily returns of the column country’s model-estimated ten-year path of term premia are more volatile around announcements by the row central bank than at other times. The cell shows the excess ratio of standard deviations for that asset (announcement window standard deviation over non-announcement window standard deviation, minus 100%). Significance is assessed at the 1% (**) and 5% (*) level by the Brown-Forsythe test.
Table 18: **Excess Volatility in 60M Currency Returns**

<table>
<thead>
<tr>
<th>Country</th>
<th>AUD</th>
<th>CAD</th>
<th>CHF</th>
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<th>GBP</th>
<th>JPY</th>
<th>NOK</th>
<th>NZD</th>
<th>SEK</th>
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<td>41**</td>
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<td>15**</td>
</tr>
<tr>
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<td>185**</td>
<td>5</td>
<td>0</td>
<td>6</td>
<td>0</td>
<td>11</td>
<td>12</td>
<td>9</td>
<td>5</td>
</tr>
<tr>
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<td>−12</td>
<td>152**</td>
<td>−16</td>
<td>−12</td>
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<td>30**</td>
<td>17**</td>
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</tr>
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<td>−1</td>
<td>−14</td>
<td>−8</td>
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<td>5</td>
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<td>23**</td>
<td>60**</td>
<td>70**</td>
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</tbody>
</table>

Notes: The table tests whether returns of the column country’s currency are more volatile around announcements by the row central bank than at other times. The cell shows the excess ratio of standard deviations for that asset (announcement window standard deviation over non-announcement window standard deviation, minus 100%). Significance is assessed at the 1% (**) and 5% (*) level by the Brown-Forsythe test. Currencies are measured against an equal-weighted basket of all other currencies, but the one whose central bank is making announcements.

Table 19: **Excess Volatility in 60M Currency Returns (Kolmogorov-Smirnov test)**

<table>
<thead>
<tr>
<th>Country</th>
<th>AUD</th>
<th>CAD</th>
<th>CHF</th>
<th>EUR</th>
<th>GBP</th>
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<th>NOK</th>
<th>NZD</th>
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<td>23**</td>
<td>0</td>
<td>41**</td>
<td>4</td>
<td>15</td>
</tr>
<tr>
<td>Canada</td>
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<td>−12</td>
<td>152**</td>
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<td>−12</td>
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<td>37</td>
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<td>125**</td>
<td>13</td>
<td>37**</td>
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<td>17**</td>
<td>35</td>
<td>50</td>
</tr>
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<td>−1</td>
<td>9</td>
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<td>94**</td>
<td>6</td>
<td>18*</td>
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<td>9</td>
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</tr>
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<td>82**</td>
<td>102**</td>
<td>120**</td>
<td>23**</td>
<td>60**</td>
<td>70</td>
<td>317**</td>
</tr>
</tbody>
</table>

Notes: The table tests whether returns of the column country’s currency are statistically different around announcements by the row central bank than at other times. The cell shows the excess ratio of standard deviations for that asset (announcement window standard deviation over non-announcement window standard deviation, minus 100%). Significance is assessed at the 1% (**) and 5% (*) level by the Kolmogorov-Smirnov test. Currencies are measured against an equal-weighted basket of all other currencies, but the one whose central bank is making announcements.
Figure 17: Market Reactions to US Monetary Shocks, across Time

Notes: The figures depict the reactions of currencies, bonds, and cross-border bond portfolios to monetary announcements in the US in two different periods: the pre-crisis sample (2001 until mid-2008) on the left, and the post-crisis sample (mid-2009 until 2016) on the right. The currency figures show by how much the dollar appreciates against a given reference currency when it appreciates by 1% on average. The bond figures show by how much the foreign yields of other countries rise when US yields rise by 1%. Finally, the cross-border bond portfolio figures show by how much a portfolio that shorts a given country’s ten-year bond and lends at the US riskfree rate appreciates when the average portfolio appreciates by 1%. Standard error bars in all pictures are computed versus the mean reaction across all currencies, bonds, or portfolios; and the shading of the coefficient bars refers to the lower-dimensional structure, whereby currencies, bonds, or portfolios of the same color (different colors) react similarly (dissimilarly).
Figure 18: Market Reactions to US Monetary Shocks, across States

(a) Currencies, Expansions

(b) Currencies, Recessions

(c) Bonds, Expansions

(d) Bonds, Recessions

(e) Cross-Border Bond Portfolios, Expansions

(f) Cross-Border Bond Portfolios, Recessions

Notes: The figures depict the reactions of currencies, bonds, and cross-border bond portfolios to monetary announcements in the US in two different states: expansionary states (with above-average GDP growth) on the left, and recessionary states (below-average GDP growth) on the right. The currency figures show by how much the dollar appreciates against a given reference currency when it appreciates by 1% on average. The bond figures show by how much the foreign yields of other countries rise when US yields rise by 1%. Finally, the cross-border bond portfolio figures show by how much a portfolio that shorts a given country’s ten-year bond and lends at the US risk-free rate appreciates when the average portfolio appreciates by 1%. Standard error bars in all pictures are computed versus the mean reaction across all currencies, bonds, or portfolios; and the shading of the coefficient bars refers to the lower-dimensional structure, whereby currencies, bonds, or portfolios of the same color (different colors) react similarly (dissimilarly).
Figure 19: Cross-Border 30Y Bond Portfolio Reactions to US Monetary Shocks

(a) Standard Methodology

(b) Modified Methodology

Notes: The figures depict the reactions cross-border bond portfolios involving thirty-year bonds (rather than ten-year bonds) to monetary announcements in the US under two methodologies: the standard one in this paper on the left, and a modified one on the right. Both figures show by how much a portfolio that shorts a given country’s thirty-year bond and lends at the US riskfree rate appreciates when the average portfolio appreciates by 1%. Standard error bars in all pictures are computed versus the mean reaction across all currencies, bonds, or portfolios; and the shading of the coefficient bars refers to the lower-dimensional structure, whereby currencies, bonds, or portfolios of the same color (different colors) react similarly (dissimilarly). The left figure uses the methodology throughout the paper, in which foreign portfolios are regressed on monetary shocks and in which a more conservative version of the Bayesian information criterion is used to identify the lower-dimensional structure. However, since high-frequency data for thirty-year bonds is virtually nonexistent outside the US, the right figure modifies the methodology to offset the loss of power. Specifically, it both uses US thirty-year bonds as an additional variable in the regression to offer some high-frequency identification, and it uses the regular version of the Bayesian information criterion to identify the lower-dimensional structure.

Table 20: Pairwise Comparisons on Currency Responses to US Monetary Shocks

<table>
<thead>
<tr>
<th></th>
<th>AUD</th>
<th>CAD</th>
<th>CHF</th>
<th>EUR</th>
<th>GBP</th>
<th>JPY</th>
<th>NOK</th>
<th>NZD</th>
<th>SEK</th>
</tr>
</thead>
<tbody>
<tr>
<td>AUD</td>
<td>0.000</td>
<td>0.047</td>
<td>0.132</td>
<td>0.000</td>
<td>0.003</td>
<td>0.736</td>
<td>0.329</td>
<td>0.680</td>
<td></td>
</tr>
<tr>
<td>CAD</td>
<td>0.000</td>
<td>0.013</td>
<td>0.000</td>
<td>0.812</td>
<td>0.664</td>
<td>0.000</td>
<td>0.000</td>
<td>0.000</td>
<td></td>
</tr>
<tr>
<td>CHF</td>
<td>0.047</td>
<td>0.013</td>
<td>0.132</td>
<td>0.007</td>
<td>0.004</td>
<td>0.006</td>
<td>0.021</td>
<td>0.010</td>
<td></td>
</tr>
<tr>
<td>EUR</td>
<td>0.132</td>
<td>0.000</td>
<td>0.132</td>
<td>0.000</td>
<td>0.003</td>
<td>0.004</td>
<td>0.065</td>
<td>0.001</td>
<td></td>
</tr>
<tr>
<td>GBP</td>
<td>0.000</td>
<td>0.812</td>
<td>0.007</td>
<td>0.000</td>
<td>0.498</td>
<td>0.000</td>
<td>0.000</td>
<td>0.000</td>
<td></td>
</tr>
<tr>
<td>JPY</td>
<td>0.003</td>
<td>0.664</td>
<td>0.004</td>
<td>0.003</td>
<td>0.498</td>
<td>0.000</td>
<td>0.001</td>
<td>0.001</td>
<td></td>
</tr>
<tr>
<td>NOK</td>
<td>0.736</td>
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<td>0.006</td>
<td>0.004</td>
<td>0.000</td>
<td>0.000</td>
<td>0.405</td>
<td>0.885</td>
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</tr>
<tr>
<td>NZD</td>
<td>0.329</td>
<td>0.000</td>
<td>0.021</td>
<td>0.065</td>
<td>0.000</td>
<td>0.001</td>
<td>0.405</td>
<td>0.389</td>
<td></td>
</tr>
<tr>
<td>SEK</td>
<td>0.680</td>
<td>0.000</td>
<td>0.010</td>
<td>0.001</td>
<td>0.000</td>
<td>0.001</td>
<td>0.885</td>
<td>0.389</td>
<td></td>
</tr>
</tbody>
</table>

The table supports Figure 4 by implementing pairwise comparisons among the coefficients associated with each currency. Figure 4 depicts by how much the dollar appreciates against a given reference currency when it appreciates by 1% on average, following a Fed tightening. This table shows the p-values that emerge from a two-sided two-sample t-test between the relative appreciation for two reference currencies.
Table 21: **Pairwise Comparisons on Bond Responses to US Monetary Shocks**

<table>
<thead>
<tr>
<th></th>
<th>AUD</th>
<th>CAD</th>
<th>CHF</th>
<th>EUR</th>
<th>GBP</th>
<th>JPY</th>
<th>NOK</th>
<th>NZD</th>
<th>SEK</th>
</tr>
</thead>
<tbody>
<tr>
<td>AUD</td>
<td>0.000</td>
<td>0.000</td>
<td>0.000</td>
<td>0.008</td>
<td>0.000</td>
<td>0.000</td>
<td>0.356</td>
<td>0.000</td>
<td>0.000</td>
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<tr>
<td>CAD</td>
<td>0.000</td>
<td>0.048</td>
<td>0.032</td>
<td>0.887</td>
<td>0.000</td>
<td>0.144</td>
<td>0.005</td>
<td>0.194</td>
<td></td>
</tr>
<tr>
<td>CHF</td>
<td>0.000</td>
<td>0.048</td>
<td>0.076</td>
<td>0.247</td>
<td>0.000</td>
<td>0.439</td>
<td>0.000</td>
<td>0.721</td>
<td></td>
</tr>
<tr>
<td>EUR</td>
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<td>0.032</td>
<td>0.076</td>
<td>0.059</td>
<td>0.066</td>
<td>0.430</td>
<td>0.000</td>
<td>0.175</td>
<td></td>
</tr>
<tr>
<td>GBP</td>
<td>0.008</td>
<td>0.887</td>
<td>0.247</td>
<td>0.059</td>
<td>0.000</td>
<td>0.181</td>
<td>0.012</td>
<td>0.229</td>
<td></td>
</tr>
<tr>
<td>JPY</td>
<td>0.000</td>
<td>0.000</td>
<td>0.000</td>
<td>0.066</td>
<td>0.000</td>
<td>0.028</td>
<td>0.000</td>
<td>0.001</td>
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</tr>
<tr>
<td>NOK</td>
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<td>0.144</td>
<td>0.439</td>
<td>0.430</td>
<td>0.181</td>
<td>0.028</td>
<td>0.000</td>
<td>0.653</td>
<td></td>
</tr>
<tr>
<td>NZD</td>
<td>0.356</td>
<td>0.005</td>
<td>0.000</td>
<td>0.000</td>
<td>0.012</td>
<td>0.000</td>
<td>0.000</td>
<td>0.001</td>
<td></td>
</tr>
<tr>
<td>SEK</td>
<td>0.000</td>
<td>0.194</td>
<td>0.721</td>
<td>0.175</td>
<td>0.229</td>
<td>0.001</td>
<td>0.653</td>
<td>0.001</td>
<td></td>
</tr>
</tbody>
</table>

The table supports Figure 5 by implementing pairwise comparisons among the coefficients associated with each country’s bond yield. Figure 5 depicts by how much yields of a given country’s ten-year bonds rise when US ten-year yields rise by 1%, following a Fed tightening. This table shows the p-values that emerge from a two-sided two-sample t-test between the relative rise for the bonds of two countries.

---

Table 22: **Pairwise Comparisons on Portfolio Responses to US Monetary Shocks**

<table>
<thead>
<tr>
<th></th>
<th>AUD</th>
<th>CAD</th>
<th>CHF</th>
<th>EUR</th>
<th>GBP</th>
<th>JPY</th>
<th>NOK</th>
<th>NZD</th>
<th>SEK</th>
</tr>
</thead>
<tbody>
<tr>
<td>AUD</td>
<td>0.000</td>
<td>0.000</td>
<td>0.000</td>
<td>0.001</td>
<td>0.000</td>
<td>0.011</td>
<td>0.555</td>
<td>0.028</td>
<td></td>
</tr>
<tr>
<td>CAD</td>
<td>0.000</td>
<td>0.022</td>
<td>0.883</td>
<td>0.203</td>
<td>0.002</td>
<td>0.157</td>
<td>0.000</td>
<td>0.011</td>
<td></td>
</tr>
<tr>
<td>CHF</td>
<td>0.000</td>
<td>0.022</td>
<td>0.090</td>
<td>0.587</td>
<td>0.000</td>
<td>0.968</td>
<td>0.010</td>
<td>0.148</td>
<td></td>
</tr>
<tr>
<td>EUR</td>
<td>0.000</td>
<td>0.883</td>
<td>0.090</td>
<td>0.486</td>
<td>0.016</td>
<td>0.178</td>
<td>0.002</td>
<td>0.012</td>
<td></td>
</tr>
<tr>
<td>GBP</td>
<td>0.001</td>
<td>0.203</td>
<td>0.587</td>
<td>0.486</td>
<td>0.000</td>
<td>0.681</td>
<td>0.004</td>
<td>0.186</td>
<td></td>
</tr>
<tr>
<td>JPY</td>
<td>0.000</td>
<td>0.002</td>
<td>0.000</td>
<td>0.016</td>
<td>0.000</td>
<td>0.001</td>
<td>0.000</td>
<td>0.000</td>
<td></td>
</tr>
<tr>
<td>NOK</td>
<td>0.011</td>
<td>0.157</td>
<td>0.968</td>
<td>0.178</td>
<td>0.681</td>
<td>0.001</td>
<td>0.019</td>
<td>0.303</td>
<td></td>
</tr>
<tr>
<td>NZD</td>
<td>0.555</td>
<td>0.000</td>
<td>0.010</td>
<td>0.001</td>
<td>0.004</td>
<td>0.000</td>
<td>0.019</td>
<td>0.079</td>
<td></td>
</tr>
<tr>
<td>SEK</td>
<td>0.028</td>
<td>0.011</td>
<td>0.148</td>
<td>0.012</td>
<td>0.186</td>
<td>0.000</td>
<td>0.303</td>
<td>0.079</td>
<td></td>
</tr>
</tbody>
</table>

The table supports Figure 7 by implementing pairwise comparisons among the coefficients associated with portfolios for each country. Figure 5 depicts by how much a portfolio that shorts a given country’s ten-year bond and lends at the US riskfree rate rises when the average portfolio rises by 1%, following a Fed tightening. This table shows the p-values that emerge from a two-sided two-sample t-test between the relative rises for the portfolios of two countries.
Figure 20: Market Reactions to Australian Monetary Shocks

(a) Currencies

(b) Bonds

Notes: The figures depict the reactions of currency and bond markets to announcements by the Reserve Bank of Australia. The left figure shows by how much the AUD appreciates against a given reference currency when it appreciates by 1% on average; and the right figure shows by how much foreign ten-year yields rise when Australian ten-year yields rise by 1%. Standard error bars in both pictures are computed versus the average reaction across currencies or foreign bonds; and the shading of the coefficient bars refers to the lower-dimensional structure, whereby assets of the same color react similarly and assets of different colors react dissimilarly following RBA announcements. The AUD appreciates symmetrically against all currencies and foreign yields do not rise asymmetrically when the RBA tightens, with the primary exception of New Zealand assets. This is a duplicate of Figure 11.
Figure 21: Market Reactions to Canadian Monetary Shocks

(a) Currencies

(b) Bonds

Notes: The figures depict the reactions of currency and bond markets to announcements by the Bank of Canada. The left figure shows by how much the CAD appreciates against a given reference currency when it appreciates by 1% on average; and the right figure shows by how much foreign ten-year yields rise when Canadian ten-year yields rise by 1%. Standard error bars in both pictures are computed versus the average reaction across currencies or foreign bonds; and the shading of the coefficient bars refers to the lower-dimensional structure, whereby assets of the same color react similarly and assets of different colors react dissimilarly following BoC announcements. The CAD appreciates symmetrically against all currencies and foreign yields do not rise asymmetrically when the BoC tightens, with the possible exception of Australian and New Zealand assets.

Figure 22: Market Reactions to Swiss Monetary Shocks

(a) Currencies

(b) Bonds

Notes: The figures depict the reactions of currency and bond markets to announcements by the Swiss National Bank. The left figure shows by how much the CHF appreciates against a given reference currency when it appreciates by 1% on average; and the right figure shows by how much foreign ten-year yields rise when Swiss ten-year yields rise by 1%. Standard error bars in both pictures are computed versus the average reaction across currencies or foreign bonds; and the shading of the coefficient bars refers to the lower-dimensional structure, whereby assets of the same color react similarly and assets of different colors react dissimilarly following SNB announcements. The CHF appreciates symmetrically against all currencies and foreign yields do not rise when the SNB tightens.
Figure 23: Market Reactions to European Monetary Shocks

(a) Currencies

Notes: The figures depict the reactions of currency and bond markets to announcements by the European Central Bank. The left figure shows by how much the EUR appreciates against a given reference currency when it appreciates by 1% on average; and the right figure shows by how much foreign ten-year yields rise when German ten-year yields rise by 1%. Standard error bars in both pictures are computed versus the average reaction across currencies or foreign bonds; and the shading of the coefficient bars refers to the lower-dimensional structure, whereby assets of the same color react similarly and assets of different colors react dissimilarly following ECB announcements. The EUR appreciates by less against continental European currencies, and by more against all other currencies when the ECB tightens. Moreover, European yields rise more than non-European yields when the ECB tightens. This is a duplicate of Figure 9.

Figure 24: Market Reactions to British Monetary Shocks

(a) Currencies

Notes: The figures depict the reactions of currency and bond markets to announcements by the Bank of England. The left figure shows by how much the GBP appreciates against a given reference currency when it appreciates by 1% on average; and the right figure shows by how much foreign ten-year yields rise when British ten-year yields rise by 1%. Standard error bars in both pictures are computed versus the average reaction across currencies or foreign bonds; and the shading of the coefficient bars refers to the lower-dimensional structure, whereby assets of the same color react similarly and assets of different colors react dissimilarly following BoE announcements. The GBP appreciates symmetrically against all currencies and foreign yields do not rise asymmetrically when the BoE tightens.
Figure 25: Market Reactions to Japanese Monetary Shocks

(a) Currencies

(b) Bonds

Notes: The figures depict the reactions of currency and bond markets to announcements by the Bank of Japan. The left figure shows by how much the JPY appreciates against a given reference currency when it appreciates by 1% on average; and the right figure shows by how much foreign ten-year yields rise when Japanese ten-year yields rise by 1%. Standard error bars in both pictures are computed versus the average reaction across currencies or foreign bonds; and the shading of the coefficient bars refers to the lower-dimensional structure, whereby assets of the same color react similarly and assets of different colors react dissimilarly following BoJ announcements. The JPY appreciates symmetrically against all currencies and foreign yields do not rise asymmetrically when the BoJ tightens. This is a duplicate of Figure 10.

Figure 26: Market Reactions to Norwegian Monetary Shocks

(a) Currencies

(b) Bonds

Notes: The figures depict the reactions of currency and bond markets to announcements by the Norges Bank. The left figure shows by how much the NOK appreciates against a given reference currency when it appreciates by 1% on average; and the right figure shows by how much foreign ten-year yields rise when Norwegian ten-year yields rise by 1%. Standard error bars in both pictures are computed versus the average reaction across currencies or foreign bonds; and the shading of the coefficient bars refers to the lower-dimensional structure, whereby assets of the same color react similarly and assets of different colors react dissimilarly following Norges Bank announcements. The NOK appreciates symmetrically against all currencies and foreign yields do not rise asymmetrically when the Norges Bank tightens, with the possible exception of Swedish assets.
Figure 27: Market Reactions to New Zealand Monetary Shocks

(a) Currencies

(b) Bonds

Notes: The figures depict the reactions of currency and bond markets to announcements by the Reserve Bank of New Zealand. The left figure shows by how much the NZD appreciates against a given reference currency when it appreciates by 1% on average; and the right figure shows by how much foreign ten-year yields rise when New Zealand ten-year yields rise by 1%. Standard error bars in both pictures are computed versus the average reaction across currencies or foreign bonds; and the shading of the coefficient bars refers to the lower-dimensional structure, whereby assets of the same color react similarly and assets of different colors react dissimilarly following RBNZ announcements. The NZD appreciates symmetrically against all currencies and foreign yields do not rise when the RBNZ tightens, with the limited exception of Australian assets. This is a duplicate of Figure 6.

Figure 28: Market Reactions to Swedish Monetary Shocks

(a) Currencies

(b) Bonds

Notes: The figures depict the reactions of currency and bond markets to announcements by the Swedish Riksbank. The left figure shows by how much the SEK appreciates against a given reference currency when it appreciates by 1% on average; and the right figure shows by how much foreign ten-year yields rise when Swedish ten-year yields rise by 1%. Standard error bars in both pictures are computed versus the average reaction across currencies or foreign bonds; and the shading of the coefficient bars refers to the lower-dimensional structure, whereby assets of the same color react similarly and assets of different colors react dissimilarly following Riksbank announcements. The SEK appreciates symmetrically against all currencies and foreign yields do not rise when the Riksbank tightens, with the possible exception of Norwegian assets.
Figure 29: Market Reactions to American Monetary Shocks

(a) Currencies

(b) Bonds

Notes: The figures depict the reactions of currency and bond markets to announcements by the Federal Reserve. The left figure shows by how much the USD appreciates against a given reference currency when it appreciates by 1% on average; and the right figure shows by how much foreign ten-year yields rise when American ten-year yields rise by 1%. Standard error bars in both pictures are computed versus the average reaction across currencies or foreign bonds; and the shading of the coefficient bars refers to the lower-dimensional structure, whereby assets of the same color react similarly and assets of different colors react dissimilarly following Fed announcements. The USD appreciates by less against low-rate currencies, and by more against high-rate currencies when the Fed tightens. Moreover, yields of high-rate countries rise more than yields of low-rate countries when the Fed tightens. This is a duplicate of Figures 4 and 5.
H  Models of Complete Markets

This appendix supports Section 6, on models with complete markets, in two ways. First, it confirms that high-rate (Australian) bonds have higher entropy than low-rate (Japanese) bonds, at different maturities and using both returns and inverse returns. Second, it provides derivations for the full model of complete markets, relating returns in bonds and in the stochastic discount factors to the underlying shocks in the model.

H.1 Bond Entropy

To argue that bonds of high-rate countries have higher entropy following Fed announcements than bonds of low-rate countries, I first estimate the entropy induced by the Fed; and second, I correlate that estimated quantity with the level of interest rates. I find that the correlation is high and statistically significant, across both ten-year and thirty-year bonds; and using both bond yields and bond returns.

I first write bond returns for a given country $i$ as a function of a Fed-driven component and an idiosyncratic component.

$$ r_i = m_i + \epsilon_i $$

I apply the entropy operator to both sides of the expression for bond returns from a given country. Since the Fed-driven component and the idiosyncratic component are independent, they can be decoupled; and the Fed-driven entropy can thus be isolated.

$$ L(m_i) = L(r_i) - L(\epsilon_i) $$

As before, I use bond returns from announcement windows to estimate $L(r_i)$ and bond returns from non-announcement windows to estimate $L(\epsilon_i)$, and thus estimate $L(m_i)$. Once I generate the Fed-driven entropy for a given country’s bond returns, I then compute a cross-sectional correlation of this term and the average level of interest rates, across countries. I bootstrap across time intervals to generate standard errors for the correlation.

I conduct the analysis across four specifications: ten-year bond yields, ten-year bond returns, thirty-year bond yields, and thirty-year bond returns. (Six foreign countries issue thirty-year bonds: Australia, Canada, Germany, Japan, Switzerland, and the United Kingdom). The point estimates for the correlation range between 0.7 to 0.8; and all ten-year estimates are significantly different from zero at the 1% level and thirty-year estimates are significantly different from zero at the 5% level.

H.2 Solving the General Model of Complete Markets

This section documents the steps needed to derive expressions for currency and bond returns in a general model of complete markets. I first set up the model and Euler equation. I next derive the expression for returns of the (unobserved) consumption asset, which is important for deriving returns in other terms. I finally derive the expression for both returns in the stochastic discount factors and in bonds, as functions of underlying shocks.
H.2.1 Model Setup

In the baseline model, the representative consumer starts with Epstein-Zin utility:

$$ U_t = \left( (1 - \delta)C_t^{1-1/\psi} + \delta \mathbb{E}_t \left( U_{t+1}^{1-\gamma} \right)^{(1-1/\psi)/(1-\gamma)} \right)^{1/(1-1/\psi)} $$

Consumption follows the following process, with both a trend component and an idiosyncratic component:

$$ \Delta c_{t+1} = \mu + \phi x_t + \sigma_t \eta_{t+1} $$

In turn, the trend consumption follows a persistent process; and all errors themselves have stochastic volatility:

$$ x_{t+1} = \rho x_t + \varphi e_{t+1} $$

$$ (\sigma_{t+1})^2 = \sigma^2 + v \left( (\sigma_t)^2 - \sigma^2 \right) + \sigma_w w_{t+1} $$

All errors in these log processes are normal, making the underlying variables lognormal.

To extend this model to a multi-country setting and to incorporate heterogeneity, I look at the long-run risk literature, where Colacito and Croce [2011] and Colacito et al. [2017] transform the long-run risk model of Bansal and Yaron [2004] similarly. First, the papers make each process specific to country $i$. Second, the papers decompose the shock $e_{t+1}$ into two components: a global component $e_{t+1}^g$ and an idiosyncratic component $e_{t+1}^i$. To incorporate structured heterogeneity, different countries $i$ have differential loadings $1 + \beta_i^g$ on the global components of shocks. For simplicity, $1 + \beta_i^g \geq 0$ but this is not actually necessary for any results.

I utilize these innovations, and I decompose all shocks into global and idiosyncratic components. The global components have constant global volatility, while idiosyncratic components have idiosyncratic stochastic volatility. I weight the global and idiosyncratic components by parameters $\alpha$, yielding the updated utility function and dynamics:

$$ U_t(i) = \left( (1 - \delta)C_t(i)^{1-1/\psi} + \delta \mathbb{E}_t \left( U_{t+1}(i)^{1-\gamma} \right)^{(1-1/\psi)/(1-\gamma)} \right)^{1/(1-1/\psi)} $$

$$ \Delta c_{t+1}^i = \mu + \phi x_t^i + \left( \sqrt{\alpha_y} \sigma (1 + \beta_y^i) \eta_{t+1}^i + \sqrt{1 - \alpha_y} \sigma^i e_{t+1}^i \right) $$

$$ x_{t+1}^i = \rho x_t^i + \varphi e \left( \sqrt{\alpha_e} \sigma (1 + \beta_e^i) e_{t+1}^i + \sqrt{1 - \alpha_e} \sigma^i e_{t+1}^i \right) $$

$$ (\sigma_{t+1}^i)^2 = \sigma^2 + v \left( (\sigma_t^i)^2 - \sigma^2 \right) + \sigma_w (1 + \beta_w^i) w_{t+1}^i + \sqrt{1 - \alpha_w} w_{t+1}^i $$

To solve for the entropy of stochastic discount factors and entropy of long-maturity bond returns, I use the approximation tools of Campbell and Shiller [1988]. One-period ahead returns have the following process:

$$ r_{t+1} \approx \kappa_0 + \chi(z) z_{t+1} - z_t + g_{t+1} $$

where $z_t = p_t - d_t$, i.e. the log price-to-dividend ratio, and where $g_{t+1}$ is the log growth rate in dividends. It is worth noting that the coefficient on $z_{t+1}$, a function of the long-term stationary price-to-dividend ratio $z$, is effectively one in my setting. Campbell and Shiller [1988] note that $\chi(z) = (1 + e^{-z})^{-1}$, and find $z = 2.68$ in annual data and thus $\chi(z) = 0.936$ in annual data. In shorter windows, the log price-to-dividend ratio escalates rapidly, as prices stay the same over any unit of time but dividends fall. An annual ratio of 2.68 becomes a daily ratio of 8.20 (excluding weekends), and so $\chi(z) = 0.9997$. In the derivations below, I generate expressions that include the coefficient $\chi(z)$, but then approximate it to one in the final simplification.
As shown in Epstein and Zin [1989], this utility function yields the following Euler equation for any asset \( j \):

\[
E_t \left( \delta^\theta \left( \frac{C_{t+1}^i}{C_t^i} \right)^{-\theta/\psi} R_{a,t+1}(i)^-(1-\theta) R_{j,t+1} \right) = 1
\]

where \( \theta = \frac{1-\gamma}{1-1/\psi} \) and where \( R_a \) is the (unobservable) gross return on an asset that pays out consumption in country \( i \). Since asset returns and the SDF are assumed to be jointly lognormal, I use the following Euler equation:

\[
E_t \exp \left( \theta \log \delta - \frac{\theta}{\psi} \Delta c_{t+1}^i + (\theta - 1)r_{a,t+1}^i + r_{j,t+1} \right) = 1
\]  

which makes the log SDF:

\[
m_{t+1}^i = \theta \log \delta - \frac{\theta}{\psi} \Delta c_{t+1}^i + (\theta - 1)r_{a,t+1}^i
\]  

**H.2.2 The Consumption Asset**

First, I price the (unobserved) asset that pays off aggregate consumption. This is needed to price the stochastic discount factor and in turn bond returns. To price the consumption asset, I start with a Campbell-Shiller approximation:

\[
r_{a,t+1}^i \approx \kappa_{a,0} + \chi(z_a) z_{a,t+1}^i - z_{a,t}^i + \Delta c_{t+1}^i
\]

Second, I conjecture that the price-dividend ratio \( z_{a,t} \) is a linear function of the state variables \( x_t^i \) and \( (\sigma_t^i)^2 \), as in Bansal and Yaron [2004].

\[
z_{a,t}^i = A_0 + A_1 x_t^i + A_2 (\sigma_t^i)^2
\]

To solve the coefficients, I use Equation (19), which is the log Euler equation, and price the consumption asset itself \( j = a \):

\[
\theta \log \delta - \frac{\theta}{\psi} E \Delta c_{t+1}^i + \theta E r_{a,t+1} + \frac{1}{2} \nu \left( -\frac{\theta}{\psi} \Delta c_{t+1}^i + \theta r_{a,t+1} \right) = 0
\]  

where:

\[
r_{a,t+1} \approx \kappa_{a,0} + \chi(z_a) \left( A_0 + A_1 (\rho x_t + \varphi_e (\sqrt{\alpha_e} \sigma (1 + \beta_e^i) e_{t+1}^i + \sqrt{1 - \alpha_e \sigma_t^i e_{t+1}^i}) + A_2 \left( \sigma^2 + v \left( (\sigma_t^i)^2 - \sigma^2 \right) + \sigma_w (\sqrt{\alpha_w} (1 + \beta^i_w) w_{t+1}^i + \sqrt{1 - \alpha_w w_{t+1}^i}) \right) - \left( A_0 + A_1 x_t^i + A_2 (\sigma_t^i)^2 \right) + (\mu + \phi x_t^i + (\sqrt{\alpha_w} \sigma (1 + \beta^i_w) x_{t+1}^i + \sqrt{1 - \alpha_w x_{t+1}^i}) \right) + (1 - \gamma) \left( A_0 + A_1 x_t^i + A_2 (\sigma_t^i)^2 \right) - \left( A_0 + A_1 x_t^i + A_2 (\sigma_t^i)^2 \right) \right) + \frac{1}{2} \theta^2 \chi(z_a) A_2^2 \sigma_w^2 \left( \alpha_w (1 + \beta^i_w)^2 + (1 - \alpha_w) \right) + \frac{1}{2} (1 - \gamma)^2 \left( \alpha_\eta \sigma^2 (1 + \beta^i_\eta)^2 + (1 - \alpha_\eta) (\sigma_t^i)^2 \right) = 0
\]
This expression must hold for any arbitrary value of the state variables \( x^i_t \) and \( (\sigma^i_t)^2 \), and so this will pin down \( A_1 \) and \( A_2 \). \((A_0 \) is a constant and so there is no need to identify it.) As such, I group all terms involving each state variable, and impose the restriction that their coefficients must equal zero:

\[
x^i_t : \quad \phi (1 - \gamma) + \theta \chi(z_a) A_1 \rho - \theta A_1 = 0
\]

\[
(\sigma^i_t)^2 : \quad \theta \chi(z_a) A_2 v - \theta A_2 + \frac{1}{2} \theta^2 \left(1 - \frac{1}{\psi}\right)^2 (1 - \alpha_v) + \frac{1}{2} \theta^2 \chi(z_a)^2 A_1^2 \psi^2 (1 - \alpha_e) = 0
\]

This yields the following solutions for the coefficients, using the approximation that \( \chi(z_a) = 1 \) over my short windows:

\[
A_1 = \phi (1 - \rho)^{-1} \left(1 - \frac{1}{\psi}\right)
\]

\[
A_2 = \frac{1}{2} (1 - \gamma) \left(1 - \frac{1}{\psi}\right) \frac{(1 - \alpha_v) + (1 - \alpha_e) \phi^2 \left(\frac{\varphi}{1 - \rho}\right)^2}{1 - v}
\]

With the coefficients, I can now return to Equation (22) to simplify it. I group together all constants as \( K_a \).

\[
\begin{align*}
\frac{\Delta r^i_{a,t+1}}{\psi} &= K_a + \frac{1}{\psi} x^i_t - (\sigma^i_t)^2 A_2 (1 - v) + A_1 \varphi_e \left(\sqrt{\alpha_v \sigma^2 (1 + \beta^i_e) e^i_{t+1}} + \sqrt{1 - \alpha_e \sigma^i_t e^i_{t+1}}\right) \\
&\quad + A_2 \varphi_w \left(\sqrt{\alpha_w (1 + \beta^i_w) w^i_{t+1}} + \sqrt{1 - \alpha_w w^i_{t+1}}\right) + \sqrt{\alpha_v \sigma^2 (1 + \beta^i_v) \eta^i_{t+1}} + \sqrt{1 - \alpha_v \sigma^i_t \eta^i_{t+1}}
\end{align*}
\]

H.2.3 Stochastic Discount Factor

With the expression for the consumption asset, I return to Equation (20) to generate an expression for the stochastic discount factor. That expression is written below:

\[
m^i_{t+1} = \theta \log \delta - \frac{\theta}{\psi} \Delta c^i_{t+1} + (\theta - 1) r^i_{a,t+1}
\]

I plug Equation (23) (along with the dynamics for consumption) to get an expression relating the stochastic discount factor to underlying shocks:

\[
m^i_{t+1} = \theta \log \delta - \frac{\theta}{\psi} \left(\mu + \phi x^i_t + \left(\sqrt{\alpha_v \sigma^2 (1 + \beta^i_v) \eta^i_{t+1}} + \sqrt{1 - \alpha_v \sigma^i_t \eta^i_{t+1}}\right)\right)
\]

\[
+ (\theta - 1) \left(\frac{1}{\psi} x^i_t - (\sigma^i_t)^2 A_2 (1 - v) + A_1 \varphi_e \left(\sqrt{\alpha_v \sigma^2 (1 + \beta^i_e) e^i_{t+1}} + \sqrt{1 - \alpha_e \sigma^i_t e^i_{t+1}}\right)\right)
\]

\[
+ A_2 \varphi_w \left(\sqrt{\alpha_w (1 + \beta^i_w) w^i_{t+1}} + \sqrt{1 - \alpha_w w^i_{t+1}}\right) + \sqrt{\alpha_v \sigma^2 (1 + \beta^i_v) \eta^i_{t+1}} + \sqrt{1 - \alpha_v \sigma^i_t \eta^i_{t+1}}
\]

This expression can be simplified, as follows.

\[
m^i_{t+1} = K_m - \frac{1}{\psi} x^i_t + (\gamma - 1/\psi)(1 - \gamma) K_0 (\sigma^i_t)^2
\]

\[
- (1 - \rho)^{-1} (\gamma - 1/\psi) \phi \varphi_e \left(\sqrt{\alpha_v \sigma^2 (1 + \beta^i_e) e^i_{t+1}} + \sqrt{1 - \alpha_e \sigma^i_t e^i_{t+1}}\right)
\]

\[
- (1 - v)^{-1} (\gamma - 1/\psi) (1 - \gamma) K_0 \sigma_w \left(\sqrt{\alpha_w (1 + \beta^i_w) w^i_{t+1}} + \sqrt{1 - \alpha_w w^i_{t+1}}\right)
\]

\[
- \gamma \left(\sqrt{\alpha_v \sigma^2 (1 + \beta^i_v) \eta^i_{t+1}} + \sqrt{1 - \alpha_v \sigma^i_t \eta^i_{t+1}}\right)
\]

In this expression, I define two additional constants \( K_0 \) and \( K_m \); the exact specification of \( K_m \)
is unimportant, but I represent the specification of \( K_0 \).

\[
K_0 = \frac{1}{2} \left( (1 - \alpha_\eta) + (1 - \alpha_e) \phi^2 \left( \frac{\varphi_e}{1 - \rho} \right)^2 \right)
\]

H.2.4 Long-Maturity Bonds

Long-maturity bonds can be priced similarly to the consumption asset. The major difference is that the dividend process is not a function of shocks; in this case, I set it to be a constant \( \mu^i_b \). As before, I begin with the Campbell-Shiller approximation for returns, where the log bond price-bond dividend ratio is a linear function of state variables \( x^i_t \) and \( (\sigma_t^i)^2 \):

\[
r^i_{b,t+1} \approx \kappa_{b,0} + \chi(z_b) z^i_{b,t+1} - z^i_{b,t} + \mu^i_b
\]

\[
z^i_{b,t} = B_0 + B_1 x^i_t + B_2 (\sigma^i_t)^2
\]

I combine these expressions with laws of motion for the state variables to get the full expression for the long-maturity bond return:

\[
r^i_{b,t+1} \approx \kappa_{b,0} + \chi(z_b) \left( B_0 + B_1 (\rho x^i_t + \varphi_e (\sqrt{\alpha_e} \sigma (1 + \beta_e^i) e^i_{t+1} + \sqrt{1 - \alpha_e} \sigma_i e^i_{t+1})) + B_2 \left( \sigma^2 + v \left( (\sigma^i_t)^2 - \sigma^2 \right) + \sigma_w (\sqrt{\alpha_w} (1 + \beta_w^i) w^i_{t+1} + \sqrt{1 - \alpha_w} w^i_{t+1}) \right) \right) - \left( B_0 + B_1 x^i_t + B_2 (\sigma^i_t)^2 \right) + \mu^i_b
\]

As before, I use the Euler equation, Equation (19), to identify the coefficients \( B_1 \) and \( B_2 \):

\[
\mathbb{E} m^i_{t+1} + \mathbb{E} r^i_{b,t+1} + \frac{1}{2} \psi (m^i_{t+1} + r^i_{b,t+1}) = 0
\]

which expands to the following:

\[
\kappa_{b,0} + \chi(z_b) \left( B_0 + B_1 (\rho x^i_t + B_2 \left( \sigma^2 + v \left( (\sigma^i_t)^2 - \sigma^2 \right) \right) \right) - \left( B_0 + B_1 x^i_t + B_2 (\sigma^i_t)^2 \right) + \mu^i_b + K_m
\]

\[
\phi \psi x_t + (\gamma - 1/\psi) (1 - \gamma) K_0 (\sigma^i_t)^2 + \frac{1}{2} \left( \alpha_e (\chi(z_b) B_1 - (1 - \rho)^{-1}(\gamma - 1/\psi) \phi) \varphi_e^2 \sigma^2 (1 + \beta_e^i)^2 \right.
\]

\[
+(1 - \alpha_e) (\chi(z_b) B_1 - (1 - \rho)^{-1}(\gamma - 1/\psi) \phi) \varphi_e^2 (\sigma^i_t)^2 + \alpha_{\eta} \gamma^2 \sigma^2 (1 + \beta_{\eta}^i)^2 + (1 - \alpha_{\eta}) \gamma^2 (\sigma^i_t)^2
\]

\[
+\alpha_w (\chi(z_b) B_2 - (1 - v)^{-1}(\gamma - 1/\psi) (1 - \gamma) K_0^2) \sigma^2 (1 + \beta_{w}^i)^2
\]

\[
+(1 - \alpha_w) (\chi(z_b) B_2 - (1 - v)^{-1}(\gamma - 1/\psi) (1 - \gamma) K_0^2) \sigma^2 (1 + \beta_{w}^i)^2 \right)
\]

As before, since this must hold regardless of \( x^i_t \) and \( (\sigma^i_t)^2 \), I group the respective coefficients and equate them to zero. As before, I also impose \( \chi(z_b) = 1 \):

\[
x^i_t: \quad \chi(z_b) B_1 \rho - B_1 - \phi \frac{1}{\psi} = 0 \quad \Rightarrow \quad B_1 = -\phi (1 - \rho)^{-1} \frac{1}{\psi}
\]

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\[
\begin{aligned}
\sigma_t^2 : \quad & \chi(z_b)B_2v - B_2 + (\gamma - 1/\psi)(1 - \gamma)K_0 \\
& + \frac{1}{2}(1 - \alpha_e)(\chi(z_b)B_1 - (1 - \rho)^{-1}(\gamma - 1/\psi)\phi)^2 \varphi_e^2 + \frac{1}{2}(1 - \alpha_e)\gamma^2 = 0
\end{aligned}
\]

\[
B_2 = \frac{\left(\gamma - \frac{1}{\psi} + \frac{\gamma}{\psi}\right)K_0}{1 - v}
\]

Thus, I can represent bond returns as a function of the underlying shocks. This expression again utilizes \(K_b\) (which does not need to be defined) and \(K_0\) (defined previously).

\[
\begin{aligned}
r_{b,t+1}^i = & K_b + \frac{\phi}{\psi}x_t^i - (\gamma - 1/\psi + \gamma/\psi)K_0 \sigma_t^2 \\
& - (1 - \rho)^{-1}(1/\psi)\phi\varphi_e \left(\sqrt{\alpha_e\sigma(1 + \beta_e^i)e_{t+1}^z} + \sqrt{1 - \alpha_e\sigma_t^i e_{t+1}^z}\right) \\
& - (1 - v)^{-1}(1/\psi - \gamma - \gamma/\psi)K_0\sigma_w \left(\sqrt{\alpha_w(1 + \beta_w^i)w_{t+1}^z} + \sqrt{1 - \alpha_w w_{t+1}^z}\right)
\end{aligned}
\]

**H.2.5 Permanent and Transitory Components**

This section will be written soon. Please check soon, and ensure you are reading the most recent version of the paper.