Essays in International Finance

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Abstract

This dissertation discusses central banking as world markets become more interlinked, as the Federal Reserve generates international shocks, and as the dollar becomes more pervasive globally.

In the first chapter, I establish a new fact on how the Fed's monetary policy spills over into foreign currency and bond markets asymmetrically. Using high-frequency data, I show that when the Fed tightens: (i) the dollar appreciates 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 yields. Apart from the European Central Bank, other countries do not generate spillovers of their own. My results illustrate the unique potency of the Fed and the heterogeneity in global markets.

In the second chapter, I examine the channels through which the Fed's monetary policy spills over into foreign financial markets. Specifically, I show that the fact identified in the first chapter provides evidence against two leading channels of spillovers. The asymmetries across currency and bond markets reject theories in which foreign central banks react to the Fed, and reject models with full risk-sharing in which foreign risk premia shift, as currency and bond markets contradict each other under these two channels. Shifts in risk premia under models with incomplete markets are most consistent with these patterns. My results suggest that the Fed's spillovers do not diminish the independence of central banks, but rather illustrate the importance of frictions.

In the third chapter, I document and explain the accumulation of large dollar portfolios by foreign central banks, by arguing that these reserve portfolios hedge liquidity shocks to dollarized financial systems. First, I extract currency shares for the foreign reserves of seventy-seven countries. The dollar shares are large and well-explained by the dollar shares of their financial systems' liabilities, particularly for countries that cannot borrow from the Fed directly. Second, I generate a

model in which central banks use dollar reserves to mitigate liquidity shocks to their dollarized financial systems, particularly when foreign exchange transaction costs are high.

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To my grandmothers

Introduction

My dissertation examines a broad and important topic at the intersection of financial economics and international macroeconomics: how hegemonic institutions, like the Federal Reserve and the US dollar, shape the international financial system, and how foreign monetary authorities react and adapt. These American pillars have remained enduringly influential in global finance despite the shrinking American presence in global trade, and this tension poses new challenges for policymakers. In addition, my dissertation is methodologically noteworthy for its use of high-frequency data and statistical models to extract agents' expectations and actions from financial markets, instead of relying on noisily-reported surveys and measures of quantities as most other papers do.

The first chapter of my dissertation, "Documenting Monetary Spillovers in Financial Markets," documents a new fact on how the Federal Reserve's monetary policy affects foreign currency and bond markets in asymmetric ways. Specifically, it shows disparities between the way in which the Federal Reserve affects the currencies and long-maturity bonds of countries with high short-term interest rates, like Australia and New Zealand, versus those with low short-term interest rates, like Japan. This paper uses latent factor models and and high-frequency data to identify subtle asymmetries that simpler models and low-frequency data, as used in the literature, miss. In addition, the paper establishes the relative importance of American monetary policy in global markets. The existing literature has not found evidence of other central banks generating spillovers, but it is again unclear whether their results are driven by low statistical power. I find that most central banks have only small regional effects, and that the European Central Bank has large regional and small global effects. These results show that American institutions are qualitatively different from those of other countries in the international financial system.

The second chapter of my dissertation, "Understanding the Channels of Monetary Spillovers in Financial Markets," examines the channels by which the Federal Reserve's monetary policy affects foreign financial markets. I use the fact identified in the first chapter to argue against the two leading channels of spillovers: ones in which foreign central banks react to the Federal Reserve, and ones in which risk premia shift under models of full-risksharing. Currency markets predict that when the Federal Reserve tightens, central banks tighten most or stochastic discount factors rise most in Japan and other low-rate countries. By contrast, bond markets predict that central banks tighten most or stochastic discount factors rise most in Australia and other high-rate countries. This contradiction means that neither channel convincingly explains the Federal Reserve's spillovers. Instead, I argue that the Federal Reserve alters the underlying risk-bearing abilities of global investors. This suggests that the Federal Reserve's global reach operates by constraining financial intermediaries, rather than by altering foreign fundamentals. Methodological differences explain the divergence between my results and those in the literature. The monetary spillovers literature commonly uses sensitive vector autoregression methods to link interest rates across long horizons and concludes (at low significance levels) that central banks react to US monetary policy. By contrast, I extract long-horizon expectations from asset returns around Federal Reserve announcements and thus benefit from greater statistical power. The international asset pricing literature explains key puzzles in currency markets (e.g. the carry trade) using models of complete markets, but I synthesize results from bond markets in addition to currency markets to show inconsistencies in this explanation for the Federal Reserve's spillovers.

The third chapter of my dissertation, "Foreign Dollar Reserves and Financial Stability," explains the large foreign reserve holdings of central banks as hedges to liquidity shocks in dollarized financial systems. Central banks conventionally only release information about the sizes of their foreign reserves, which limits the literature's ability to adjudicate between different motivations. Officially, central banks keep the *compositions* of these portfolios strictly confidential. I circumvent this limitation and extract hidden information about portfolios' compositions by projecting fluctuations in the portfolio sizes of seventy-seven developed and emerging countries onto the returns of commonly traded currencies, through a Bayesian dynamic linear model. This new information best supports an explanation in which concerns over financial stability drives large foreign reserve holdings, as central banks' portfolio dollar shares track the dollar shares of their banking systems' liabilities — and only for the set of countries that cannot borrow from the Federal Reserve directly, through swap lines. I formally model this story, and show that high foreign exchange transaction costs during crises are important to justify large foreign reserve portfolios. This paper establishes a causal link between two well-known trends: the dollarization of financial systems and the dominance of the dollar in governments' reserve portfolios.

This dissertation studies topics that are important for informing the design of monetary policy and financial regulation, both in the US and abroad. As the disparity between the United States' presence in financial and trade markets grows, and as the consequences of the American-led financial regime intensify, policymakers everywhere will have to reckon with these new dynamics.

Chapter 1

Documenting Monetary Spillovers in Financial Markets

1.1 Introduction

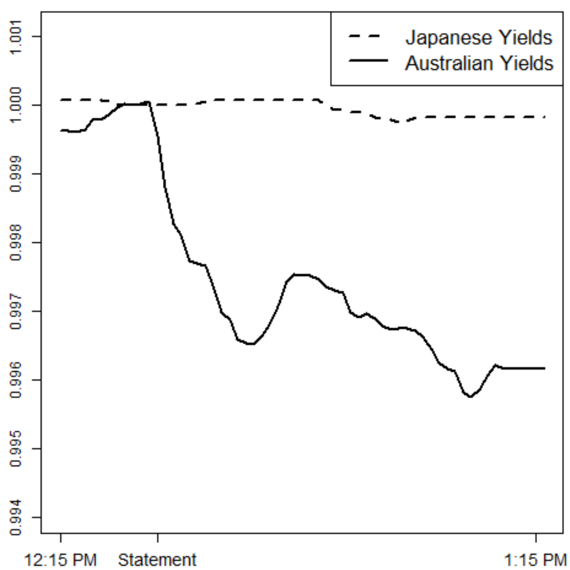
Textbooks have long taught that a currency is determined by monetary policy in the two countries that it spans, or that a bond is driven by domestic events. But episodes in which currencies react to monetary policy in third countries, deemed *external monetary policy*, have become more ubiquitous recently. One widely-acknowledged example is the Taper Tantrum of June 2013, in which currency pairs not including the dollar and foreign bonds reacted to the Fed's change in policy. Such events are being documented more frequently, which raises broader questions on how foreign assets respond to external monetary policies.

In this paper, I establish a novel fact on how currencies and bonds react *asymmetrically* to the Fed's announcements. This fact is identified using high-frequency data and methodologies robust to market noise, and so it has the statistical power to overcome limitations that have hampered much of the literature. I find that when the Fed tightens, the dollar appreciates more against currencies of high-interest rate countries (e.g. Australia) than against currencies of low-interest 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. This is unique: I document that the monetary policies of most other countries do *not* spill over into foreign markets.

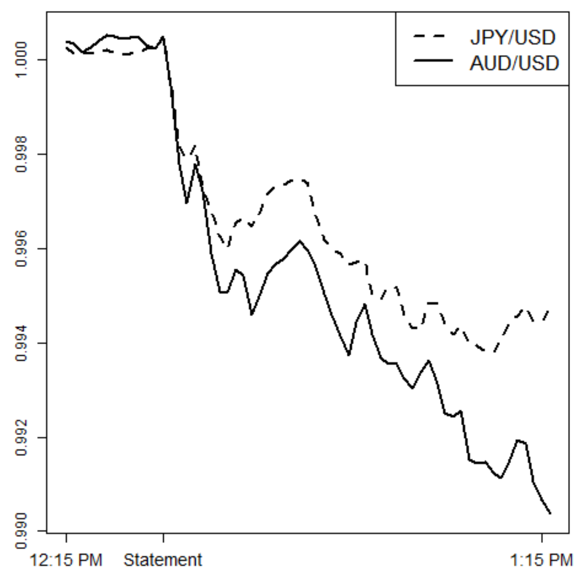
To illustrate the paper's finding, consider an example. At 12:30 PM on January 25, 2012, the Fed announced its intentions to keep interest rates low until 2014. The surprise monetary easing affected foreign assets in asymmetric ways, as sixty-minute windows around the announcement show in Figure 1.1. Yields on ten-year Australian bonds immediately fell whereas yields on ten-year Japanese bonds did not. Moreover, the dollar depreciated more against the Australian dollar than against the yen, or equivalently the Australian dollar appreciated against the yen.

Figure 1.1: *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.

This example is representative of a more general pattern on the asymmetric responses of currencies and bonds to Fed announcements. 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 interest rates historically (e.g. Australia) and least against currencies in countries with low interest rates historically (e.g. Japan) when the Fed tightens or eases respectively. At the same time, 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 fact applies to Fed announcements from 2001 - 2016, and describes markets both before and after the financial crisis. The example in Figure 1.1 illustrates these patterns, as the dollar depreciates more against the Australian dollar, and Australian yields fall more.

Such findings are special to the Fed. I document that the central banks of the other nine countries cannot generate spillovers in global financial markets — with the exception of the European Central Bank, which has particularly strong effects on non-Eurozone countries in continental Europe. This finding is in itself novel to the literature both in scope and in statistical power, and it underscores the importance of designing models that show heterogeneity among central banks.

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 reactions to Fed announcements at all horizons, as they reflect forecasted 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 weak statistical power. 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

[1]This is not driven by the zero lower bound for interest rates, as I discuss in Section 1.4.

miss them altogether.[2] I embed these components within a latent factor model.

The paper proceeds as follows. Section 1.2 reviews the literature on empirical spillover patterns. Section 1.3 discusses the empirical framework and the data. Section 1.4 introduces the main fact on asymmetries in currency and bond markets. Section 1.5 shows the lack of asymmetries emanating from other central banks. Section 1.6 concludes.

1.2 Review of Literature on Spillovers

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 comprehensive papers include Rey (2015) and Miranda-Agrippino and Rey (2015), which look at the Fed's effects on a wide range of markets.[3] However, there are many papers that examine more specific markets. To review a handful: Brusa *et al.* (2017) study equity markets; Fratzscher *et al.* (2017), Burger *et al.* (2017), and Chari *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, although some new literature such as Cerutti *et al.* (2017) challenges that claim.

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 country. By contrast, papers that study heterogeneity in spillovers relate it to a set of macroeconomic variables, which precludes contrasting asymmetries in such ways. (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 *et al.* (2016a) 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 *et al.*

[2] As one prominent example, Rey (2015) and Cerutti *et al.* (2017) reach opposing conclusions on the foundational question of whether the Fed's monetary policy affects foreign capital flows, as capital flows data are noisy and measured at low frequencies.

[3] The one exception that Rey (2015) and Miranda-Agrippino and Rey (2015) find is foreign direct investment.

(2016b) find that spillovers are stronger when recipient countries are more financially integrated with the US. Finally, Zhang (2017) argues that the share of a country's trade invoiced in dollars drives cross-country heterogeneity in spillovers.[4]

Second, I study spillovers emanating from other central banks, a topic on 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. 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 *et al.* (2016b) briefly study spillovers from the Eurozone, Japan, and the UK.

1.3 Empirical Framework and Data

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

The core components of the empirical framework are high-frequency windows around Fed announcements, 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 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 flawed observed measures of shocks.

High-frequency windows around Fed announcements have two uses: they allow for causal interpretations, and they provide power. My windows are predominantly sixty-minute windows,

[4] However, Dedola *et al.* (2017) find no consistent macroeconomic indicators that explain heterogeneity.

although I use daily windows to replace any intraday windows with poor liquidity. Building off work by Gurkaynak *et al.* (2005) and Gertler and Karadi (2015) in domestic markets, high-frequency windows in international markets improve on low-frequency windows in two ways. First, low-frequency windows run the risk that non-monetary news comes out during the window, and so asset returns could reflect extraneous information. Second, low-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 both concerns.

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 moderate horizons, and foreign central banks or investors may respond at long horizons too. Measuring reactions to Fed announcements in short-maturity assets would not capture all changes in the paths of rates or in risk premia. In contrast, 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, serve as reference points for announcement windows, and they keep my estimates conservative. I observe asset returns through windows of equal duration on other days, to identify the ordinary variance and covariance of assets. Thus, only asset movements during announcement windows that exceed the ordinary patterns in non-announcement windows are linked to monetary policy. 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 in event studies.

Inferred shocks (i.e. latent factors) are important for precise identification. I estimate my shocks as the common factor in asset returns, rather than measuring them from observed data such as movements in the Fed Funds futures. This methodology ensures that the entire common monetary surprise is retained, by not narrowly restricting monetary surprises to those affecting shorter-maturity US yields. In turn, this methodology ensures that my estimates remain capable of detecting subtle asymmetries. Indeed, my results are more noisily estimated, although qualitatively unchanged, when using measured shocks.

1.3.1 Framework

Before describing the method or data, I first introduce the core equation that identifies asset responses to Fed 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.

Equation

Equation (1.1) is the main equation that I estimate. 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 (α), the product of a univariate or multivariate vector of coefficients (β) and a univariate monetary shock (m_t), and a univariate or multivariate error ϵ_t. I primarily test whether assets react differentially to monetary shocks, i.e. test for equality between different elements of the multivariate vector β.

$$r_t = \alpha + \beta m_t + \epsilon_t \tag{1.1}$$

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.[5] 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 βm_t in Equation (1.1) 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 by the dollar appreciating and/or Treasury yields rising.

In Equation (1.1), the only observed data are the high-frequency and long-maturity asset returns r_t during announcement windows. However, I let the errors ϵ_t take on a distribution parameterized from asset returns over non-announcement windows, rather than assuming them to be homoskedastic white noise. As discussed, m_t refers to the inferred monetary shock. I discuss each of these sequentially in more depth.

[5]The 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.

High-Frequency Returns

High-frequency returns r_t around Fed announcements are essential for two reasons: arguing causality and identifying with power. In most cases, I measure returns from the fifteen minutes before the Fed's announcement to the forty-five minutes after. For returns with poor intraday liquidity, I use daily windows, in which I measure returns over the day of the announcement.[6]

First, 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 Fed 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.

Second, high-frequency returns are important for statistical power. Idiosyncratic noise in currency markets overwhelms monetary shocks in windows beyond sixteen hours. Bond markets are approximately half as noisy as currency markets, and so daily windows work when necessary.

Assets do not revert after announcements, and so my results do not reflect transient phenomena. Appendix A.4 shows there is no correlation between returns during and after sixty-minute windows.

Long-Maturity Instruments

Long-maturity instruments are important for capturing the full reactions of foreign monetary policy and of risk premia to Fed announcements. There is abundant evidence that the Fed releases guidance on its own policy over moderate horizons, e.g. the example in Figure 1.1 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 to long horizons. For instance, Miranda-Agrippino and Rey (2015) and Rogers *et al.* (2016) estimate that the largest responses by foreign central banks to the Fed occur at the one- to three-year horizon, although at low confidence levels. Thus,

[6]Currency markets are liquid and traded around-the-clock, but bond markets in smaller countries often have poor liquidity or limited hours.

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 sovereign ten-year bonds.[7] In Equation (1.2), I present the standard currency equation, which defines currency premia as the residuals from a repeated cross-border carry trade. Changes 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 (i.e. investors' relative willingness to hold one currency over the other).[8] In addition, currencies are driven by the infinite-horizon exchange rate.

$$\underbrace{\Delta s_t^{j/\$}}_{\text{Exchange Rate}} = \underbrace{\sum_{k=1}^{\infty} \Delta i_{t+k-1}^{\$}}_{\text{US Path of Rates}} - \underbrace{\sum_{k=1}^{\infty} \Delta i_{t+k-1}^{j}}_{\text{Foreign Path of Rates}} + \underbrace{\sum_{k=1}^{\infty} \Delta p_{t+k-1}^{j/\$}}_{\text{Currency Premia}} + \underbrace{\Delta s_{\infty}^{j/\$}}_{\text{Long-Run Exchange Rate}} \tag{1.2}$$

In Equation (1.3), I present the standard equation for bond yields, which defines term premia as the residuals from a cross-maturity trade. 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 (i.e. investors' relative willingness to hold long-maturity versus short-maturity assets).[9]

$$10\,\underbrace{\Delta y_t^j(t, t+10)}_{\text{10Y Yield}} = \underbrace{\sum_{k=1}^{10} \Delta i_{t+k-1}^{j}}_{\text{10Y Path of Rates}} + \underbrace{\sum_{k=1}^{10} \Delta \gamma_{t+k-1}^{j}}_{\text{10Y Term Premia}} \tag{1.3}$$

Note that throughout this dissertation, the operator Δ is defined as the change in expectations (i.e. the innovation) to a random variable, *not* as the first difference in that variable. This does not make any empirical or qualitative difference, as over high-frequency windows, currencies and

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[7]Exchange rates and bonds are preferred to other liquid instruments, such as equities or corporate bonds, as these have uncertain payouts owing to variation in dividends or corporate default risk respectively. Sovereign default risks for the ten developed countries in my sample are largely negligible ($< 1\%$), as Damodaran (2017) shows.

[8]Anderson *et al.* (2003) argue that returns in the exchange rate market match with rational expectations theory, as currencies react to a wide range of unanticipated announcements but not anticipated ones.

[9]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 they still represent valid long-maturity assets.

bonds are effectively martingales.[10] In this paper, shocks are realized at time t.[11]

$$\Delta x_t \equiv \mathbb{E}_t x_t - \mathbb{E}_{t-1} x_t \quad (\text{not } x_t - x_{t-1}) \tag{1.4}$$

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. The confidence intervals are wide, and so it is difficult to either take this as evidence for or against any given hypothesis. Given the general noise in financial markets, ascribing movements in interest rates to announcements several years ago is difficult.

Non-Announcement Windows

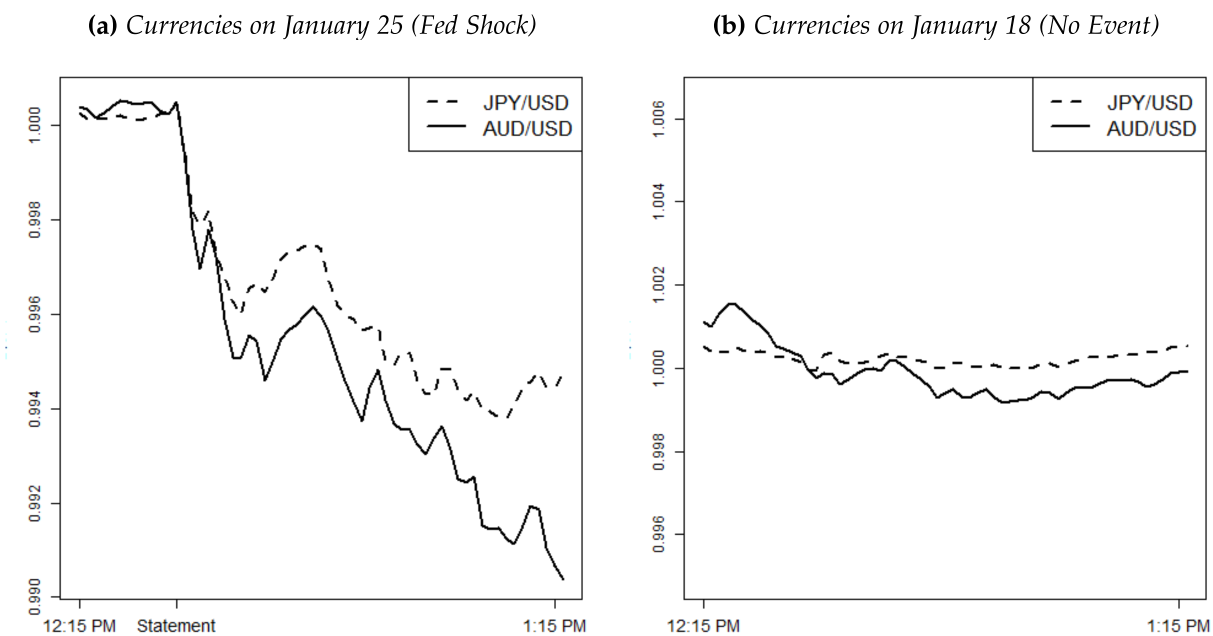
Non-announcement windows, which are used to parameterize error term ϵ_t in Equation (1.1), serve as reference points for announcement windows and are essential for ensuring that my results do not ascribe background noise during announcement windows to the Fed's effects. I identify non-announcement windows in the same way that I identify my announcement windows, as a combination of sixty-minute and daily windows. The sample consists of windows that fall within one week before and after Fed announcements, measured at the same time of day. Cieslak *et al.* (2016) find that extraneous monetary shocks, e.g. speeches by Fed governors, drop in the week preceding and week following an announcement. I remove non-announcement windows that overlap with other news, such as announcements by other central banks; and I remove windows on Fridays, as Chordia *et al.* (2001) find lower market liquidity then. As an illustration, I show the yen and the Australian dollar against the dollar over one such non-announcement window in Figure 1.2, 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). Rather than identifying from variation over announcement windows as event

[10] However, my empirical specifications are exact rather than approximate. For instance, although I observe $\mathbb{E}_t x_t - \mathbb{E}_{t-1} x_{t-1}$ for empirical specifications, the constant in the specifications formally controls for the (typically negligible) expected change in the response variable $\mathbb{E}_{t-1} x_t - \mathbb{E}_{t-1} x_{t-1}$, leaving Δx_t to be explained by Fed shocks.

[11] Formally, time periods are annual and prior expectations of shocks, i.e. $\mathbb{E}_{t-1} x_t$, are assumed to settle just prior to the Fed's announcement.

Figure 1.2: *Currency Movements over Announcement and Non-Announcement Windows*



Notes: The figures contrast reactions in currency markets in two sixty-minute windows. The left figure depicts currency returns around the Fed's surprise easing of January 25, 2012; and the right figure depicts currency returns exactly one week prior on January 18, 2012, when no news was released. While currencies react more strongly to the Fed, they also fluctuate and co-move by a few basis points against the dollar during the non-announcement window.

studies do, it identifies from *excess* variation over announcement windows.[12] As with high-frequency returns, using non-announcement returns helps me establish a causal relationship between Fed announcements and assets. Even narrow 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 a counterfactual benchmark.

Inferred Shocks

Inferred shocks or latent shocks, which are represented by m_t in Equation (1.1), are important for precise and comprehensive identification. Since m_t is estimated, this turns Equation (1.1) into a factor model. Inferred shocks have two advantages over measured shocks. First, inferred shocks capture the entirety of monetary surprises relevant in my specifications. Second, I do not impose additional data burdens on the estimation, and so I can extend my analysis to spillovers

[12]These seminal papers work with observed shocks, rather than latent shocks. My approach follows even more closely to the spirit of Craine and Martin (2008), which utilize such methods with latent shocks.

emanating from smaller central banks. Conceptually, my specification with inferred shocks closely resembles one with interactive fixed effects.[13]

First, an inferred shock captures the entirety of the Fed's announcement as relevant to my specification. Fed announcements have different effects across different maturities, asset classes, and countries. The inferred shock captures the entire common surprise across assets r_t in Equation (1.1). This methodology contrasts with popular approaches in the literature, in which monetary shocks are measured as movements in US rates — either at short maturities (e.g. the Fed Funds futures as in Kuttner (2001) and a large subsequent literature) or at medium maturities (e.g. the one-year Eurodollar curve or two-year Treasury, as in Nakamura and Steinsson (2017) and Hanson and Stein (2015)).[14] Such approaches lose some of the common factor affecting response variables. This concern is most salient for short-maturity rate shocks, as Fed announcements often reveal information about future rates alongside the imminent target, noted by Gurkaynak *et al.* (2005). This concern is mitigated but still present for medium-maturity rate shocks for two reasons. First, Fed announcements may either intentionally reveal information about future rates at horizons longer than two years (especially in the last decade), or otherwise affect long-horizon assets differently than medium-horizon assets, as Boyarchenko *et al.* (2017) find. Second, currencies and foreign bonds may systematically react to Fed announcements in different ways than domestic bonds. For both reasons, I would lose some of the common surprise by projecting assets onto measured shocks, and this informational loss can lead to imprecise estimates and inferences.Indeed, I replicate my results using the two-year Treasury yield instead of inferred shocks, and I find that my results are qualitatively the same but more noisily estimated.[15]

Second, inferred shocks do not require additional data beyond the response variables r_t in Equation (1.1), and so they can be constructed at an intraday frequency for all central banks, which I do in Section 1.5. This feature is especially useful when looking at the central banks of smaller countries, such as the Reserve Bank of New Zealand or the Norges Bank, which do not

[13]Indeed, Bai (2009) notes that algorithms for estimating latent factor models can be used to estimate models with interactive fixed effects, depending on the specification's dimensionality.

[14]However, my inferred shocks correlate reasonably with existing measures of monetary shocks including these and differences from surveyed expectations, as documented in Appendix A.2.

[15]Those results are reported in Figure A.7 in Appendix A.2.

have liquid equivalents for the Fed Funds futures market or other rate futures markets.

The primary limitation of inferred shocks is that they do not have an operational interpretation, for two reasons. First, my methodology cannot disentangle how assets respond to the Fed's direct effects on the path of US rates from how assets respond to the Fed's indirect effects on US risk premia. Second, my methodology cannot estimate meaningful units for coefficients, as I only estimate the units of the product of shocks and coefficients. This limitation, applicable to all factor models, is highly problematic for some questions but not problematic for my question. Specifically, this limitation is fatal for papers that quantify the pass-through of Fed policy on assets. This limitation is minor for my paper, which tests whether assets respond at all and whether they respond symmetrically or asymmetrically to an underlying monetary surprise, regardless of its size or its components.

1.3.2 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 (α, β) and m_t in the multivariate version of Equation (1.1). 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 β.

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 with a simple example, 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 (1.1) in its multivariate form:

$$\begin{bmatrix} \Delta s_t^{€/\$} \\ \Delta s_t^{£/\$} \\ \Delta s_t^{¥/\$} \end{bmatrix} = \begin{bmatrix} \alpha^{€/\$} \\ \alpha^{£/\$} \\ \alpha^{¥/\$} \end{bmatrix} + \begin{bmatrix} \beta^{€/\$} \\ \beta^{£/\$} \\ \beta^{¥/\$} \end{bmatrix} m_t^{\$} + \begin{bmatrix} \epsilon_t^{€/\$} \\ \epsilon_t^{£/\$} \\ \epsilon_t^{¥/\$} \end{bmatrix} \tag{1.5}$$

I present the Gaussian likelihood function associated with Equation (1.5) next. Errors ϵ are assumed to have some covariance matrix Σ, where Σ is learned from non-announcement windows rather than being homoskedastic white noise. As a result, this likelihood function resembles the one a generalized least squares approach optimizes.[16]

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

Since this term involves the product of estimated quantities m_t and β, I cannot analytically solve the system of interlocking first-order conditions.[17] 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 (α, β, m_t).[18] I make one scaling assumption, $\mathbb{V}m_t = 1$, since I cannot identify the magnitudes of m_t and β separately.

In Appendix A.2, I discuss a more general procedure that handles complications arising from partially-missing data. 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 the statistical power of my methodology substantially. Instead, I take two steps to maintain power. First, I reformulate

[16]In this exposition I implicitly assume the errors to have zero mean, but I demean the data by the non-announcement means first in practice. These non-announcement means are extremely close to zero. In addition, I treat Σ in this specification as fixed because of the overwhelming amount of non-announcement data available to estimate it. In Appendix A.2, I treat Σ as estimated when using the Identification by Heteroskedasticity methodology instead, and find the same results.

[17]Moreover, I cannot solve the system iteratively since convergence is neither guaranteed in theory nor achieved in practice.

[18]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 (μ_t, V_t) and (α, β), which maps to the original two steps. When conducting robustness checks with the shocks themselves, I set the shocks m_t to be their MAP estimates μ_t, i.e. the means of their posterior distributions. Further details can be found in Appendix A.2. Moreover, I check that the results are not sensitive to utilizing variational methods to solve for parameters. In Appendix A.2, I use MCMC methods instead to solve for parameters in the main specifications. The results, in Figure A.3, indicate that both approaches yield the same results.

each observation 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 all windows. This ensures that *all* non-missing data are utilized.

I compute standard errors for (α, β) by bootstrap, sampling the set of Fed announcements and their associated asset responses with replacement. There are no analytic solutions for standard errors, given the missing data adjustments. In Appendix A.2, I discuss alternative approaches to estimating Equation (1.5), including Identification by Heteroskedasticity by Rigobon (2003).[19]

I next turn to the two ways I concurrently test asymmetries in β in Equation (1.1).

Average Coefficient

A natural way to test asymmetries in β would be a series of pairwise tests. This is reasonable if β has two or three elements, but it is incomprehensible in practice as β has nine elements (which requires thirty-six pairwise tests). Moreover, this may yield qualitatively inconsistent results. To illustrate, consider an example in which I find significant evidence that $\beta^{€/\$} > \beta^{¥/\$}$, but I cannot reject $\beta^{€/\$} \neq \beta^{£/\$}$ or $\beta^{¥/\$} \neq \beta^{£/\$}$. At least one of these tests must be wrong.[20]

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

$$H_0 : \beta^{€/\$} = \frac{1}{2}\left(\beta^{£/\$} + \beta^{¥/\$}\right)$$

Lower-Dimensional Structure

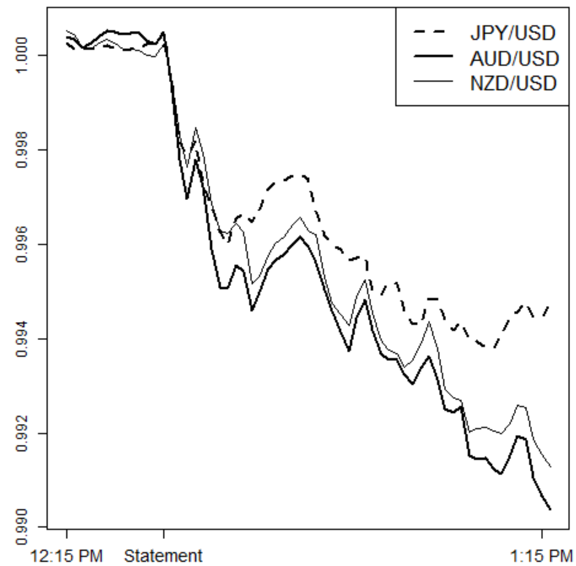
An alternate way to test for asymmetries is to cast the elements of β (and α) to a lower-dimensional structure, in which different assets are encouraged to share coefficients unless they respond too differently from each other to Fed shocks.

[19] Appendix A.2 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 (1.1) and find very similar results to those generated by the EM algorithm. I find convergence to be an issue when utilizing bonds data.

[20] Regardless, I show the p-values for all thirty-six pairwise tests in Appendix A.4 for my main results.

To illustrate, consider the opening example of the Fed announcement on January 25, 2012. In Figure 1.3, I add the New Zealand dollar to the original plot of currencies. Visually, the shock passes into the Australian and New Zealand dollars comparably, but differently into the yen.

Figure 1.3: *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 large difference for the yen.

I formally test this by estimating Equation (1.1) 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 shared coefficients, and take the structure that scores highest.[21] As an illustration, if the estimates for $\beta^{£/\$}$ are much closer to $\beta^{¥/\$}$ than to $\beta^{€/\$}$, one possible structure that may emerge in Equation (1.5) is as follows:

$$\begin{bmatrix} \Delta s_t^{€/\$} \\ \Delta s_t^{£/\$} \\ \Delta s_t^{¥/\$} \end{bmatrix} = \begin{bmatrix} 1 & 0 \\ 0 & 1 \\ 0 & 1 \end{bmatrix} \begin{bmatrix} \alpha^{€/\$} \\ \alpha^{(£,\,¥)/\$} \end{bmatrix} + \underbrace{\begin{bmatrix} 1 & 0 \\ 0 & 1 \\ 0 & 1 \end{bmatrix}}_{\text{Proposed Lower-Dimensional Structure}} \begin{bmatrix} \beta^{€/\$} \\ \beta^{(£,\,¥)/\$} \end{bmatrix} m_t^{\$} + \begin{bmatrix} \epsilon_t^{€/\$} \\ \epsilon_t^{£/\$} \\ \epsilon_t^{¥/\$} \end{bmatrix}$$

[21]This problem is closely related to clique cover problems in graph theory. Since the number of assets is small, I iterate through every permutation without needing approximate algorithms, such as LASSO.

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.[22]

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. This approach complements the test against average coefficients, and both are used concurrently to establish asymmetries.

1.3.3 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 data used in the paper.

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 continue

[22]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 — as in my specification — given the elevated risk of overfitting.

with my approach for two reasons. First, this interpretation yields a simple prediction: the market should digest Fed announcements similarly to 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. 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 guidance only on what the committee plans to implement.

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 1.5. Collection details can be found in the appendix.

Asset Returns

I collect exchange rate and bond returns for ten countries: Australia, Canada, the Eurozone (represented by Germany in bond markets), Japan, Norway, New Zealand, Sweden, Switzerland, the United Kingdom, and the United States. These ten developed markets have the most liquid assets, compared to smaller developed markets or emerging markets.[23]

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.[24] For

[23]In addition, these countries have negligible sovereign default risks, keeping the interpretation of bond returns straightforward. Emerging markets have much higher default risks, as Damodaran (2017) shows.

[24]Examples include the Sydney Futures Exchange for Australian data, the Chicago Mercantile Exchange for American

countries without bonds on a liquid futures exchange, I use an intraday benchmark rate published by Thomson Reuters based on reported transactions. In general, liquidity on established futures exchanges remains high through Fed announcements, as shown in Appendix A.4 — although liquidity for Thomson Reuters' benchmark rates can be problematic in some cases, necessitating the modifications for partially-missing data discussed earlier.

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 to ten-year) for all countries, in addition to twenty-year and thirty-year maturities for all countries except New Zealand, Norway, and Sweden.

In addition to removing 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.

1.4 Currency and Bond Asymmetries

I examine asymmetries in currency markets and in bond markets following Fed announcements, and establish a new fact. This is novel 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 miss these differences.

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. These effects are unique among central banks. In Section 1.5, 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 A.4, I show that this

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.

finding is robust to different time periods (pre-crisis and post-crisis), different states (recessionary and expansionary), and different shocks (tightening and easing).

I utilize the Expectation Maximization algorithm to fit the multivariate version of Equation (1.1), with country-specific coefficients (α, β) and time-varying shocks m_t. I replicate the example with three currencies here. To identify asymmetries, I compute standard errors with respect to the average of other coefficients, and I also look for a lower-dimensional structure that groups similar coefficients together.

$$\begin{bmatrix} \Delta s_t^{€/\$} \\ \Delta s_t^{£/\$} \\ \Delta s_t^{¥/\$} \end{bmatrix} = \begin{bmatrix} \alpha^{€/\$} \\ \alpha^{£/\$} \\ \alpha^{¥/\$} \end{bmatrix} + \begin{bmatrix} \beta^{€/\$} \\ \beta^{£/\$} \\ \beta^{¥/\$} \end{bmatrix} m_t^{\$} + \begin{bmatrix} \epsilon_t^{€/\$} \\ \epsilon_t^{£/\$} \\ \epsilon_t^{¥/\$} \end{bmatrix}$$

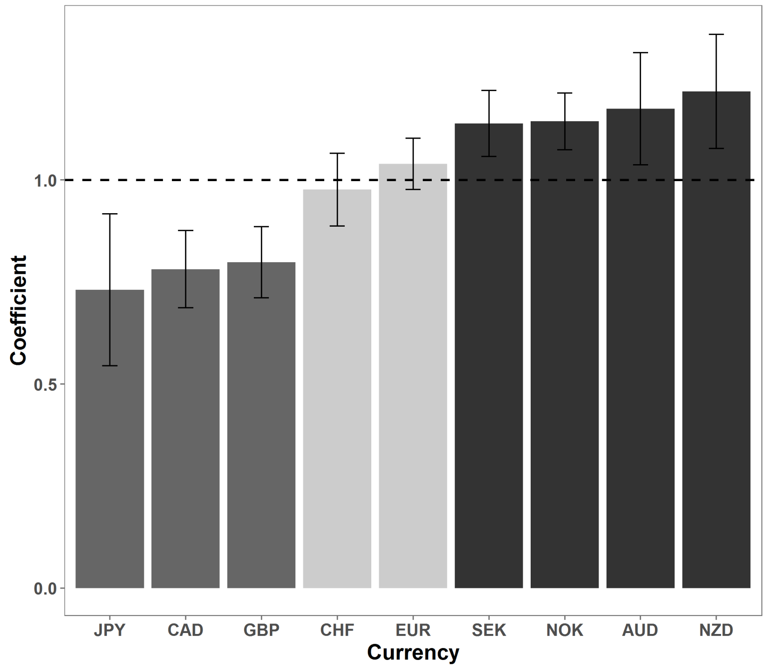
I estimate Equation (1.1) separately for currencies and for bonds in this section.

Consider currencies first. I plot the coefficients β in Figure 1.4. These coefficients refer to the dollar's appreciation (depreciation) against various currencies when it appreciates (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 finding of asymmetry.

Consider bonds next. I plot the coefficients β in Figure 1.5. These coefficients refer to the annualized rises in ten-year sovereign bond yields by country when US yields rise 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 finding of asymmetry.

On their own, these asymmetric responses show that any successful explanation of monetary spillovers must incorporate heterogeneity. To demonstrate, I rewrite the definitions for currencies and ten-year bonds, Equations (1.2) and (1.3). Since the US components are common across currencies, differential appreciation and depreciation of the dollar against different currencies points to differential movements by foreign central banks or in currency premia. Differential responses in yields similarly point to differential movements by foreign central banks or in term

Figure 1.4: *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 against 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.

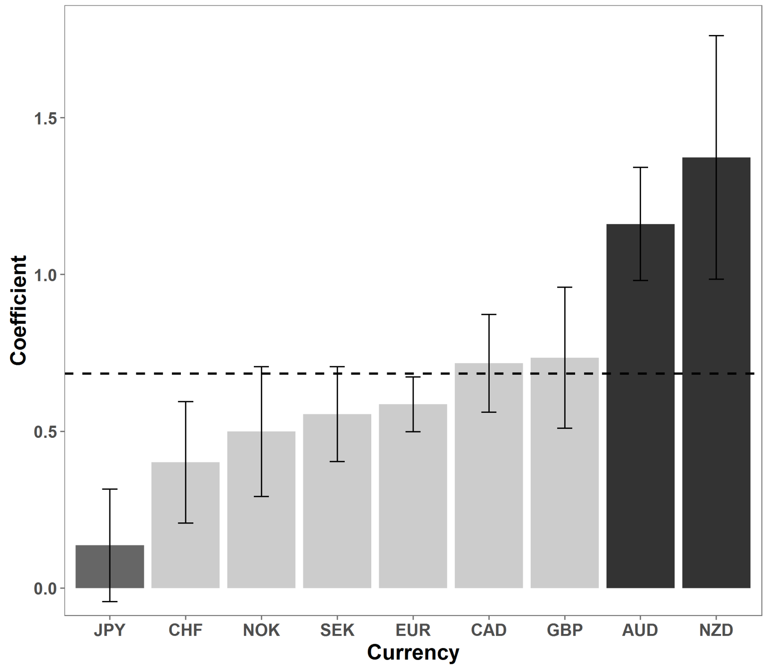
premia.

$$\underbrace{\Delta s_t^{j/\$}}_{\text{Exchange Rate}} = \underbrace{\sum_{k=1}^{\infty} \Delta i_{t+k-1}^{\$}}_{\text{US Path of Rates}} - \underbrace{\sum_{k=1}^{\infty} \Delta i_{t+k-1}^{j}}_{\text{Foreign Path of Rates}} + \underbrace{\sum_{k=1}^{\infty} \Delta p_{t+k-1}^{j/\$}}_{\text{Currency Premia}} + \underbrace{\Delta s_{\infty}^{j/\$}}_{\text{Long-Run Exchange Rate}}$$

$$10\,\underbrace{\Delta y_t^j(t, t+10)}_{\text{10Y Yield}} = \underbrace{\sum_{k=1}^{10} \Delta i_{t+k-1}^{j}}_{\text{10Y Path of Rates}} + \underbrace{\sum_{k=1}^{10} \Delta \gamma_{t+k-1}^{j}}_{\text{10Y Term Premia}}$$

Although the empirical specifications are computed on a country-by-country basis, I confirm that the cross-sectional sorting of asymmetries by countries maps closely to the level of interest rates in those countries. In Figures 1.4 and 1.5, 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 (1.1) with each country's pre-shock interest rate spread relative to the US, and by removing country-specific parameters. Equation (1.6) shows an illustration of the revised equation. I test

Figure 1.5: *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 against 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.

whether β_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 spreads 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 A.3, I show that this result is robust to dropping any country from the specification.

$$\begin{bmatrix} \Delta s_t^{€/\$} \\ \Delta s_t^{£/\$} \\ \Delta s_t^{¥/\$} \end{bmatrix} = \begin{bmatrix} 1 & i_{t-1}^{€} - i_{t-1}^{\$} \\ 1 & i_{t-1}^{£} - i_{t-1}^{\$} \\ 1 & i_{t-1}^{¥} - i_{t-1}^{\$} \end{bmatrix} \begin{bmatrix} \alpha_0 \\ \alpha_1 \end{bmatrix} + \underbrace{\begin{bmatrix} 1 & i_{t-1}^{€} - i_{t-1}^{\$} \\ 1 & i_{t-1}^{£} - i_{t-1}^{\$} \\ 1 & i_{t-1}^{¥} - i_{t-1}^{\$} \end{bmatrix}}_{\text{Interest Rate Spread}} \begin{bmatrix} \beta_0 \\ \beta_1 \end{bmatrix} m_t^{\$} + \begin{bmatrix} \epsilon_t^{€/\$} \\ \epsilon_t^{£/\$} \\ \epsilon_t^{¥/\$} \end{bmatrix} \tag{1.6}$$

Interest rates are one of several significant predictors, and I test others in Appendix A.3: 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, I find four that 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 they are meaningful in the context of risk premia in Section 1.5, and because they are similarly ubiquitous 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 global risk factors that align 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, and Colacito *et al.* (2017) show that exposure to global shocks correlate with rates. In other words, every paper in this literature relates its chosen explanation to the level of interest rates, either by assumption or endogenously, and I do too.

1.5 Spillovers by Other Central Banks

The asymmetric spillovers documented in Section 1.4 illustrate the uniqueness of the Fed. 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. Currencies and bonds do not react to other central banks, with only economically small effects from a central bank to its neighbors' assets.

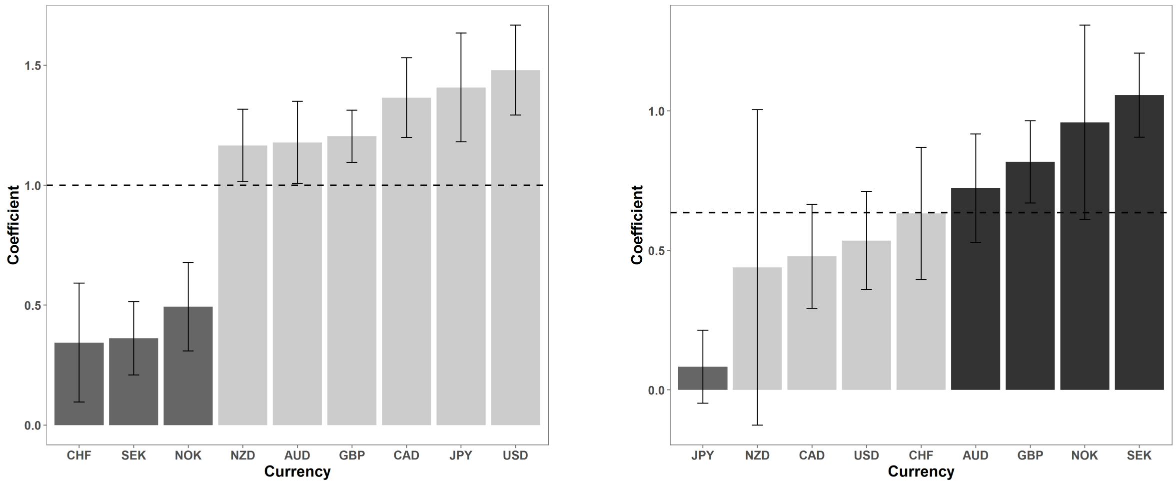
As before, I employ the latent factor model to each central bank, to identify asymmetries in how currencies and bonds react to that central bank's monetary shocks. The most stark results come from the ECB, plotted in Figure 1.6. When the ECB tightens, the euro appreciates less against other continental European currencies (e.g. the Norwegian krone) than against non-European currencies, and European bond yields rise more than non-European bond yields.

I also test whether assets react asymmetrically to other central banks. Surprisingly, I find very few asymmetries. For instance, consider the Bank of Japan in Figure 1.7. 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).

There are some statistical results regarding neighboring central banks and assets, but these are

Figure 1.6: *Market Reactions to EU Monetary Shocks*

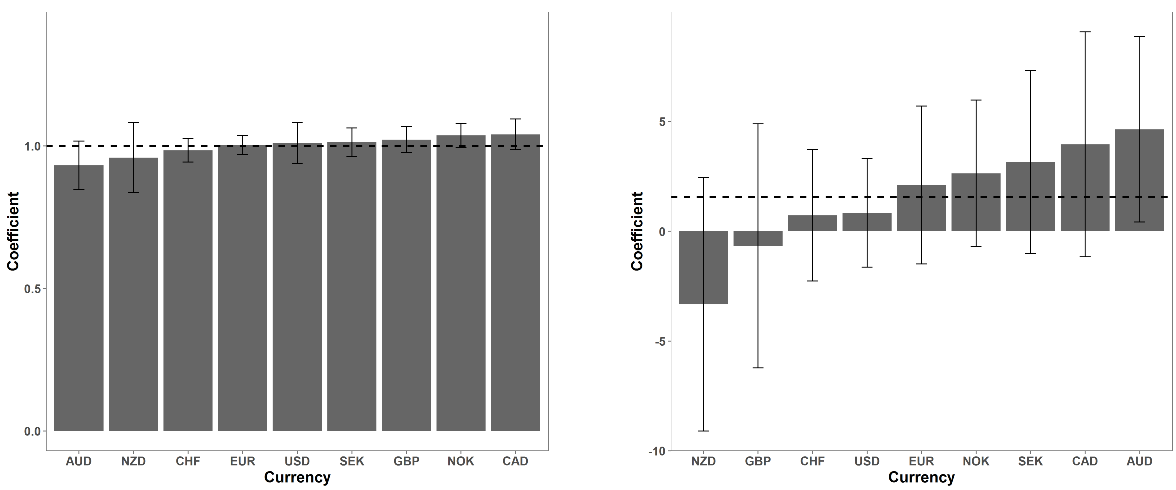
(a) *Currency Responses* **(b)** *Bond Responses*



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 1.7: *Market Reactions to JP Monetary Shocks*

(a) *Currency Responses* **(b)** *Bond Responses*

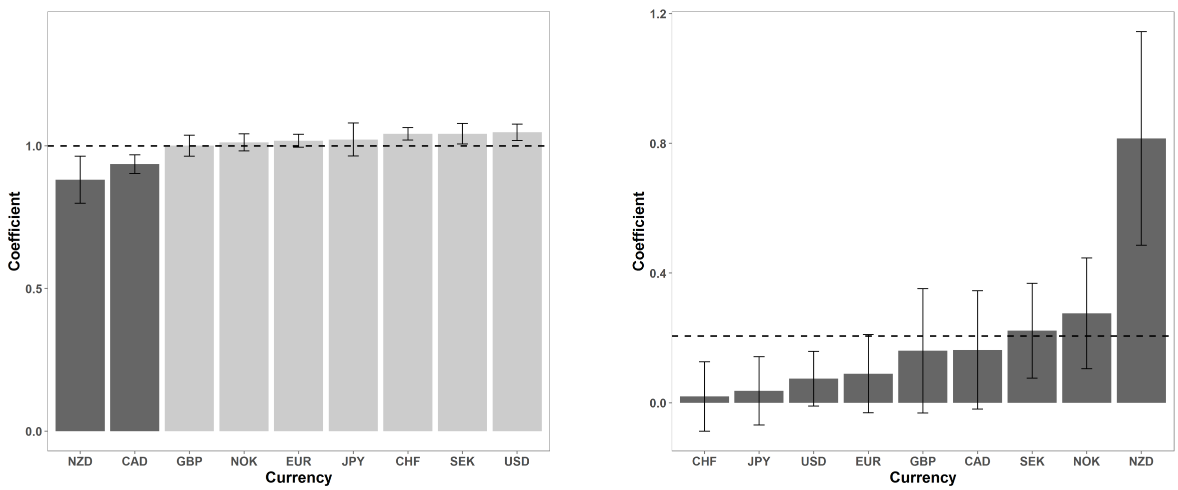


Notes: The figures depict the reactions of currency and bond markets to announcements by the Bank of Japan. The left figure shows by how much the JPY appreciates against a given reference currency when it appreciates by 1% on average; and the right figure shows by how much foreign ten-year yields rise 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.

economically small. For instance, I show that New Zealand assets have some small exposure to the Reserve Bank of Australia in Figure 1.8. When the RBA changes policy, 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 A.5. Particularly with currencies, the confidence intervals indicate statistically non-zero but small spillovers.

Figure 1.8: *Market Reactions to AU 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 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.

These results contrast with the results of Lustig and Richmond (2017), with the exception of the ECB. Their paper finds asymmetries throughout currency markets (driven by distance), whereas I largely find symmetries, with only slight distance-driven asymmetries. Once again, the key difference is the shocks — I use currency returns that are exposed to monetary shocks, whereas their currency returns are exposed to all global shocks.[25] As such, the divergent findings could once again be explained by the hypothesis that shocks to fundamentals affect markets differently than shocks to monetary policy.

[25] In addition, Lustig and Richmond (2017) have a wider sample, that includes many emerging markets.

1.6 Conclusion

The asymmetries in currency and bond markets around the globe following Fed announcements are unique. They show the potency of the Fed, in a way that few other central banks can mirror. High-rate currencies and bonds diverge from their low-rate counterparts following Fed announcements.

The next challenge involves interpreting the underlying dynamics behind such spillovers. But even without a full interpretation, the policy implications of such spillovers are important. Large central banks may wish to account for foreign effects when setting policy: a policy tightening that simultaneously tightens credit conditions in foreign markets may well amplify monetary effects. Central banks that are the recipients of spillovers (such as the Bank of Canada or Swiss National Bank) may similarly wish to forecast and account for foreign central bank policy actions. In a world with increasingly interconnected financial markets, the old norms of monetary policy are fast crumbling.

Chapter 2

Understanding the Channels of Monetary Spillovers in Financial Markets

2.1 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 among academics and policymakers have escalated 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. 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 designing policy: what are the channels of monetary spillovers by the Federal Reserve into foreign financial markets?

To answer this question, I interpret the fact established in Chapter 1 on how currencies and bonds react *asymmetrically* to the Fed's announcements, and use it to test different channels of spillovers. This fact is that when the Fed tightens, the dollar appreciates more against currencies of high-interest rate countries (e.g. Australia) than against currencies of low-interest 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. These two forms of heterogeneity in how countries receive the Fed's spillovers, while suggestive when each is studied in isolation as in Chapter 1, are potent when studied together.

I divide explanations for monetary spillovers into three broad classes of explanations, and show that my fact provides evidence against two of the channels. The first class of explanations covers ones in which spillovers operate through foreign central banks reacting to the Fed, and it is the channel most discussed by the monetary spillovers literature. However, 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. The second class of explanations covers ones in which foreign risk premia (i.e. compensation for bearing risk) react to the Fed per models with full risk-sharing (i.e. complete markets), which describe the majority of international finance models. However, the observed asymmetries in currency markets suggest that the stochastic discount factors of low-rate countries rise most when the Fed tightens, while the observed asymmetries in bond markets suggest that the stochastic discount factors of high-rate countries rise most. In short, the cross-sectional sorting of foreign countries' currencies contradicts the sorting of bonds under either class of explanations, and so these two channels do not explain monetary spillovers. I offer further evidence from the term structures of foreign bonds to argue against theories in which central banks react, and I discuss further how models with complete markets must place complex and economically implausible restrictions on stochastic discount factors to match this fact. Instead, my fact is most consistent with the third class of explanations for spillovers, in which risk premia react to the Fed under market incompleteness.

To illustrate the paper's argument, consider the example given in Figure 1.1 in Chapter 1. Upon the Fed's announcement of easing on January 25, 2012, yields on ten-year Australian bonds immediately fell whereas yields on ten-year Japanese bonds did not. Moreover, the dollar depreciated more against the Australian dollar than against the yen, or equivalently the Australian dollar appreciated against the yen. Such patterns are shown more broadly, with the dollar depreciating (appreciating) most against countries with high interest rates and long-maturity bond yields from high-rate countries falling most (least) when the Fed eases (tightens). Moreover, these asymmetric shifts in bond and currency markets must reflect one of three explanations: (i) changes in the expected paths of policy 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). These hypotheses are exhaustive for bond markets, and virtually any other taxonomy of the channels of Fed spillovers into financial 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.

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, as these are the leading explanations for monetary spillovers. Such hypotheses are grounded both in theoretical work from the open-economy macroeconomics 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 again these hypotheses, consider the two possible scenarios in this specific example. 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. This now violates the observed responses of bond yields. More generally, I show that no hypothesized set of central bank reactions across the nine countries can be consistent with asymmetries in both currency and bond markets concurrently, as these markets sort the reactions of foreign central banks in opposing ways.

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 whereas the paths of short-term policy rates do not. This 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. This analysis studies each recipient country individually, and so complements the previous analysis which studies recipient countries together. Taken together, all of my findings suggest that central banks do not react to the Fed, and 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 most models embed full risk-sharing, as surveyed by Engel (2014). Again, I illustrate my argument against these hypotheses by considering the two possible scenarios in this specific example. The first scenario is that the stochastic discount factor (e.g. 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 to smooth out intertemporal utility fluctuations. Moreover, this explanation predicts the yen should appreciate against the Australian dollar, as exchange rates under market completeness 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, and this now violates the observed responses of bond yields. More generally, I show that no hypothesized set of shifts in risk premia across the nine countries can be consistent with asymmetries in both currency and bond markets concurrently when using models with full risk-sharing, as these markets sort the reactions of foreign risk premia in opposing ways.

I argue further against hypotheses in which risk premia shift under complete markets by considering a preference-free and distribution-free framework, and showing that such a framework generates economically implausible constraints on stochastic discount factors. Following Lustig *et al.* (2017), I allow unrestricted international stochastic discount factors to receive both temporary and permanent shocks. Under market completeness, currencies respond to both shocks, while long-maturity bonds only respond to temporary shocks. Introducing two types of heterogeneity in the stochastic discount factor gives such models enough mathematical 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 qualitative implausibility (and occasionally, impossibility) of such a specification using some commonly used models. The findings suggest that models with full risk-sharing are ill-suited to explain global financial markets around Fed announcements, despite their prevalence in explaining international markets unconditionally.

Third, I argue the original fact on asymmetric reactions is most plausibly 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 this channel's potential explanatory power 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 interest rate differentials, and thus causing high-rate currencies and bonds to move together versus low-rate ones. Such an explanation rationalizes the sorting of countries by the level of interest rates. This specific example can be interpreted as follows: following the Fed's easing, less-constrained 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. This is one of many plausible models that can explain spillovers through either market frictions or incomplete spanning, and future research is needed to evaluate them.

The conclusions from this paper are relevant for 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 in the literature due to methodological limitations. 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 from limited statistical power. One approach, taken by Rey (2015), Caceres *et al.* (2016), Hofmann and Takats (2015), and Takats and Vela (2014), measures whether innovations to US rates predict current or future innovations to foreign rates. This raises the concern of omitted factors such as global or regional growth shocks, articulated by Bernanke (2017) among others. A second approach, taken by Miranda-Agrippino and Rey (2015) and Rogers *et al.* (2016), identifies only from innovations to US rates on Fed announcement days. This addresses the identification concerns but weakens statistical power, due to the combination of overwhelming market noise at long horizons and a small sample (the Fed has eight annual meetings), and so the results come with confidence levels well below 95%.[1] 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 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 involve borrowing constraints and segmentation as in Bruno and Shin (2017) or Gabaix and Maggiori (2015). My paper offers evidence from a well-identified setting that financial frictions can explain asset returns better than models of frictionless risk-sharing, and it thematically resembles the Backus-Smith puzzle or the low exchange rate volatility puzzle in showing an empirical contradiction under frictionless models. This informs the debate between complete and incomplete markets frameworks in explaining the overall carry trade and other international finance phenomena.

The paper proceeds as follows. Section 2.2 introduces the main fact on asymmetries in currency

[1] Ilzetzki *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 and to other anchor currencies. This approach focuses primarily on emerging markets.

and bond markets, and uses these asymmetries jointly to argue against two main channels of spillovers: central banks reacting to the Fed and risk premia shifting under complete markets. Section 2.3 uses the term structures of foreign bond yields to provide further evidence that central banks do not react to the Fed. Section 2.4 argues further that models of complete markets do not explain spillovers by showing the formal modeling tensions. Section 2.5 discusses models with incomplete markets, focusing on an intermediary-based model. Section 2.6 concludes.

2.2 Currency and Bond Asymmetries

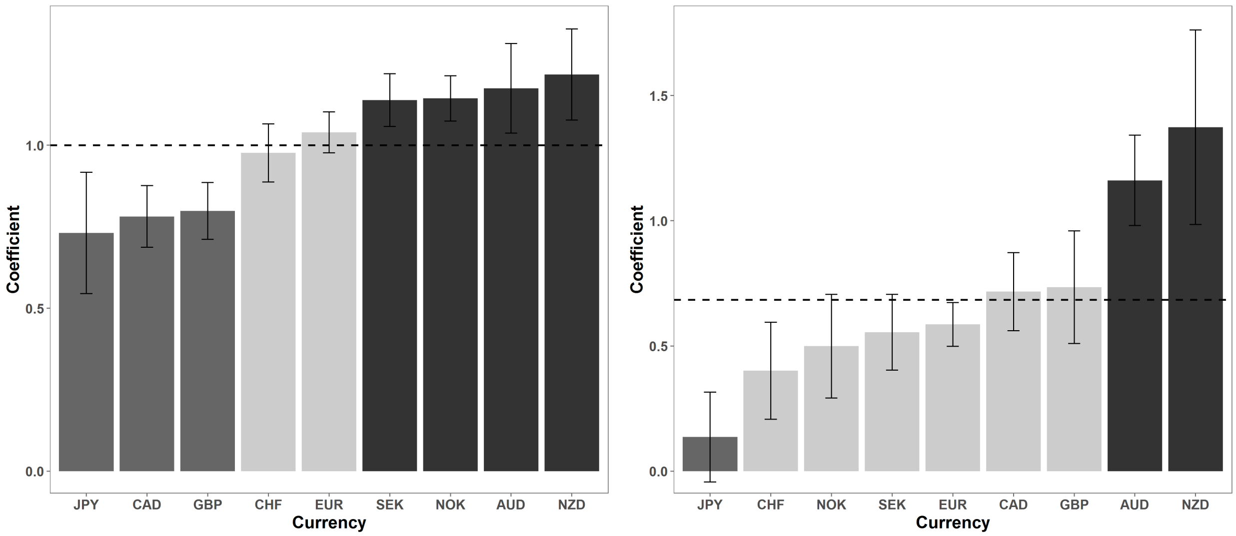
I examine asymmetries in currency markets and in bond markets following Fed announcements, documented in Chapter 1, to show that monetary spillovers from the Fed are neither consistent with shifts in the paths of foreign central banks, nor consistent with transitory shifts in foreign risk premia under complete markets. First, I review the fact in Chapter 1, that establishes how currencies and bonds from different countries respond asymmetrically to monetary shocks from the Fed. Second, I show that these two classes of explanations are inconsistent with the fact, by exploiting inconsistencies in the cross-sectional sorting of countries between currency markets and bond markets.

First, consider the fact. 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. I depict these results, discussed in depth in Chapter 1, in Figure 2.1, which respectively show currencies and bonds react to an average 1% appreciation in the US dollar and a 1% increase in the US ten-year yield.

Now, consider interpreting this figure under the two major classes of explanations: central banks reacting to the Fed, and risk premia shifting under market completeness. For both of these explanations, results from currency and bond markets suggest opposing stories in the cross-section of foreign countries. 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, then currency markets suggest that the stochastic discount factors of low-rate countries rise when the Fed tightens, while bond markets suggest that the stochastic

Figure 2.1: *Market Reactions to US Monetary Shocks*

(a) *Currency Responses* **(b)** *Bond Responses*



Notes: The figures depict the reactions of currency and bond markets to announcements by the Fed. 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 US ten-year yields rise by 1%. The dollar appreciates by little against the Japanese yen among other currencies, and by a lot against the Australian dollar among other currencies. Australian yields among other bonds rise a lot, and Japanese yields among other bonds rise little, when US yields rise.

discount factors of high-rate countries rise. This may already be suggestively apparent when looking at the equations for currencies, Equation (2.1), and bonds, Equation (2.2).

$$\underbrace{\Delta s_t^{j/\$}}_{\text{Exchange Rate}} = \underbrace{\sum_{k=1}^{\infty} \Delta i_{t+k-1}^{\$}}_{\text{US Path of Rates}} - \underbrace{\sum_{k=1}^{\infty} \Delta i_{t+k-1}^{j}}_{\text{Foreign Path of Rates}} + \underbrace{\sum_{k=1}^{\infty} \Delta p_{t+k-1}^{j/\$}}_{\text{Currency Premia}} + \underbrace{\Delta s_{\infty}^{j/\$}}_{\text{Long-Run Exchange Rate}} \tag{2.1}$$

$$10\,\underbrace{\Delta y_t^j(t,t+10)}_{\text{10Y Yield}} = \underbrace{\sum_{k=1}^{10} \Delta i_{t+k-1}^{j}}_{\text{10Y Path of Rates}} + \underbrace{\sum_{k=1}^{10} \Delta \gamma_{t+k-1}^{j}}_{\text{10Y Term Premia}} \tag{2.2}$$

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 premia under models with full risk-sharing. (I focus on transient shifts in risk premia here, and relax the assumption in Section 2.4.[2]) Therefore, if the Fed causes foreign

[2]The assumption on transitory shocks is viewed differently across fields. In much of the macroeconomics literature,

central banks to adjust their paths of policy rates or if the Fed triggers shifts in foreign risk premia, these portfolios should be equally insensitive to Fed announcements in the cross-section, i.e. portfolios for high-rate and low-rate countries should react indistinguishably to the Fed. In fact, I show that these portfolios are differentially sensitive to Fed announcements in the cross-section, proving that neither explanation can fully explain its spillovers.

Specifically, 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 subsequent paragraphs, 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. In both cases, the two components offset perfectly under some weak assumptions.

To construct this portfolio, I take the following two steps. First, I add the equation for exchange rates to the equation for bond yields. This creates a portfolio that is equivalent to shorting a foreign long-maturity bond and investing in the US riskfree rate. Second, I restrict movements in the portfolio's long-horizon terms around Fed announcements.

First, consider explanations in which central banks react to the Fed. I add Equation (2.1) to Equation (2.2), such that portfolio returns are expressed in Equation (2.3). Moreover, 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. Under models of long-run monetary neutrality, this assumption simplifies to one in which expectations of foreign inflation do not react to the

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monetary policy is neutral with respect to real variables at sufficiently long horizons, across a range of methodologies such as Uhlig (2005) and Nakamura and Steinsson (2017), and so this assumption is reasonable as long as I check long-run inflation dynamics. In the finance literature, Alvarez and Jermann (2005) find that permanent shocks are quantitatively necessary to match unconditional returns earned in financial markets under models of complete markets, and so this assumption is not reasonable. As such, I both take this assumption seriously in this section and relax it later.

Fed at long horizons.[3] I find no evidence from either domestic or foreign inflation-linked bonds that contradicts this assumption, and indeed surveyed expectations of long-run inflation are so persistent that they fluctuate approximately as much in an entire year as foreign assets do in a given sixty-minute window around announcements.[4]

$$\begin{aligned}\underbrace{\Delta r_t^{j/\$}}_{\text{Portfolio Return}} &= \underbrace{\Delta s_t^{j/\$}}_{\text{Currency Exposure}} + \underbrace{10\Delta y_t^j(t, t+10)}_{\text{Foreign Bond Exposure}} \\ &\approx \underbrace{\sum_{k=1}^{\infty} \Delta i_{t+k-1}^{\$}}_{\text{US Path of Rates}} + \underbrace{\sum_{k=1}^{\infty} \Delta p_{t+k-1}^{j/\$}}_{\text{Currency Premia}} + \underbrace{\sum_{k=1}^{10} \Delta \gamma_{t+k-1}^{j}}_{\text{10Y Term Premia}}\end{aligned} \tag{2.3}$$

It is immediately apparent that portfolio returns in Equation (2.3) do not depend directly on foreign monetary policy. Without other channels of spillovers at work concurrently, 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, consider explanations in which risk premia react to the Fed per models with full risk-sharing. Although it is not immediately apparent in Equation (2.3), shifts in a country's term premia offset with shifts in its share of currency premia. To make this point more clearly, I rewrite the definitions for currencies and bonds in terms of stochastic discount factors in Equations (2.4) and (2.5) respectively. Under the two assumptions of complete markets (complete spanning and no frictions), the exchange rate return is exactly the difference of innovations to the log stochastic discount factors. Moreover, changes in long-maturity bond yields are equal to the difference of two components: changes in the contemporaneous log stochastic discount factor and both expected and entropic changes in the long-run pricing kernel, or the stochastic discount factor used to price bond payoffs in the future. These expressions are standard, but I discuss their

[3] 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.

[4] I perform other tests, such as checking that my results are robust to thirty-year bonds and examining ten-year bonds in the pre-crisis era, as these are setting in which it is very unlikely that changes in monetary policy exceed the maturity of my bond instrument. I discuss these and other tests in Appendix B.2.

derivation in Section 2.4.[5]

$$\underbrace{\Delta s_t^{j/\$}}_{\text{Exchange Rate}} = \underbrace{\Delta m_t^{\$}}_{\text{US SDF}} - \underbrace{\Delta m_t^{j}}_{\text{Foreign SDF}} \tag{2.4}$$

$$10\,\underbrace{\Delta y_t^j(t,t+10)}_{\text{10Y Yield}} = \underbrace{\Delta m_t^j}_{\text{SDF}} - \underbrace{\left(\Delta \log \Lambda_{t+10}^j + \Delta L_t\left(\Lambda_{t+10}^j\right)\right)}_{\text{10Y Pricing Kernel}} \tag{2.5}$$

As such, I take the same two steps to construct the portfolio in this revised framework. First, I again add Equation (2.4) to Equation (2.5). Second, I assume that properties of the long-horizon pricing kernel do not react differentially to the Fed across countries, which is equivalent to assuming that variation in shocks across countries are transitory.[6]

$$\begin{aligned} \underbrace{\Delta r_t^{j/\$}}_{\text{Portfolio Return}} &= \underbrace{\Delta s_t^{j/\$}}_{\text{Currency Exposure}} + \underbrace{10\Delta y_t^j(t,t+10)}_{\text{Foreign Bond Exposure}} \\ &\approx \underbrace{\Delta m_t^{\$}}_{\text{US SDF}} \end{aligned} \tag{2.6}$$

I illustrate with the example of power utility and a stationary autoregressive log consumption process (with correlation ρ) to make Equations (2.4) and (2.5) more tangible. Negative innovations to country i's consumption basket, i.e. positive innovations to its stochastic discount factor, cause the dollar to depreciate versus its currency. Intuitively, countries with depreciated currencies should receive transfers to exploit cheap consumption, linking relatively high marginal utility and an appreciated exchange rate. At the same time, yields in that country rise, as the country smooths away temporarily elevated marginal utility by borrowing. These two effects offset when put together.

$$\underbrace{\Delta s_t^{j/\$}}_{\text{Exchange Rate}} = \underbrace{\left(-\gamma^{\$}\Delta c_t^{\$}\right)}_{\text{US SDF}} - \underbrace{\left(-\gamma^{j}\Delta c_t^{j}\right)}_{\text{Foreign SDF}}$$

$$10\,\underbrace{\Delta y_t^j(t,t+10)}_{\text{10Y Yield}} = \underbrace{\left(-\gamma^{j}\Delta c_t^{j}\right)}_{\text{SDF}} + \underbrace{\gamma\left(\rho^j\right)^{10}\sigma^j\epsilon_t^j}_{\text{10Y Pricing Kernel}} \approx \left(-\gamma^{j}\Delta c_t^{j}\right)$$

Thus, I have two equivalent representations of this portfolio, and they show the portfolio is not

[5]Equations (2.1) and (2.4) are equivalent representations of exchange rates, and Equations (2.2) and (2.5) are equivalent representations of bonds. For instance, in Equation (2.4), the actual realizations of the stochastic discount factors embed all future changes in expectations of rates and risk premia in Equation (2.1).

[6]An alternate concern is that shocks are transitory but have not decayed away in ten years. I show in Appendix B.4 that my results are robust when using thirty-year bonds.

directly exposed to two channels of spillovers. Equation (2.3) makes clear that these portfolios do not depend directly on foreign monetary policy. Equation (2.6) makes clear that these portfolios do not depend on foreign stochastic discount factors under market completeness. In short, I should not expect to find any heterogeneity in these portfolios across countries under either explanation alone — these portfolios should react symmetrically to Fed announcements.

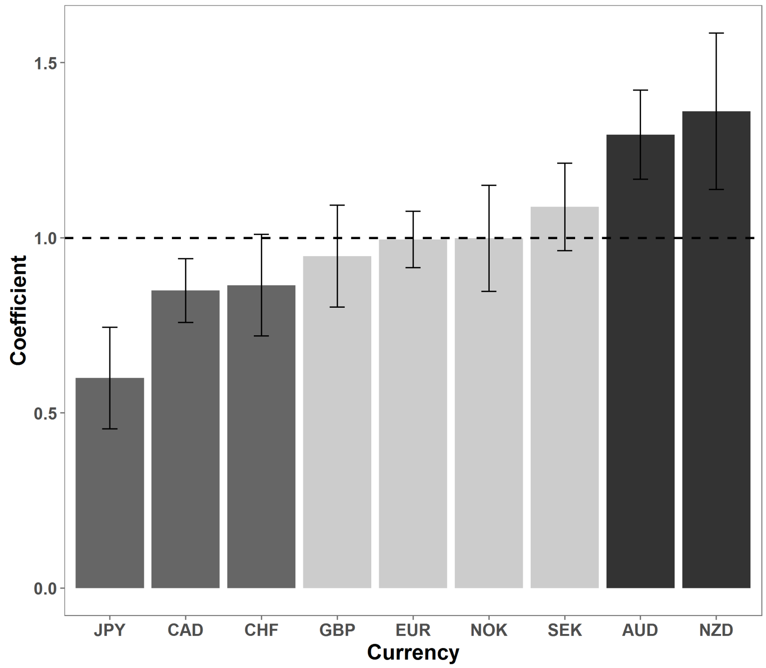
I test to see whether portfolios respond unequally to the Fed in the cross-section of countries. Following the methodology in Chapter 1, I regress these portfolios on monetary shocks, and plot the coefficients from that estimation in Figure 2.2. Indeed, I find that the portfolios are unequally responsive to the Fed across countries, rejecting two key channels of spillovers. The portfolio with Japan is less exposed to the Fed 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 Fed spillovers, as Figure 2.2 would have been symmetric across countries. For the same reasons, the reactions of stochastic discount factors in a complete markets framework cannot explain Fed spillovers either. Those asymmetries must stem from adjustments in risk premia, under market incompleteness.

This methodology is robust to concerns of countries at the zero-lower bound, for both methodological and empirical reasons. Methodologically, I compare *relative* movements in countries. Australian assets still react inconsistently with Japanese assets even if Japan's monetary policy is "stuck," and so my methods rule out the possibility that the Reserve Bank of Australia reacts to the Fed. Empirically, few countries are likely at the true zero-lower bound at all horizons of monetary policy; even Japanese long-maturity yields are weakly positive. Moreover, my results are robust both to excluding Japan altogether and to using data from the pre-crisis era only.

Finally, I also compare the coefficients in Figure 2.2 to the coefficients generated from the portfolio that shorts the foreign riskfree bond, rather than the foreign long-maturity bond, to illustrate the tension between currency and bond markets further. This alternate portfolio only has exchange rate risk, and does not have interest rate risk. I compare coefficients between the long-maturity and short-maturity portfolios in Figure 2.3. Interestingly, the point estimates for the long-maturity portfolio become more asymmetric compared to the short-maturity portfolio.[7]

[7]However, the standard error bars also widen, such that both figures have almost exactly the same number of statistically significant pairwise differences among the portfolios.

Figure 2.2: *Cross-Border Bond Portfolio Responses to US Monetary Shocks*

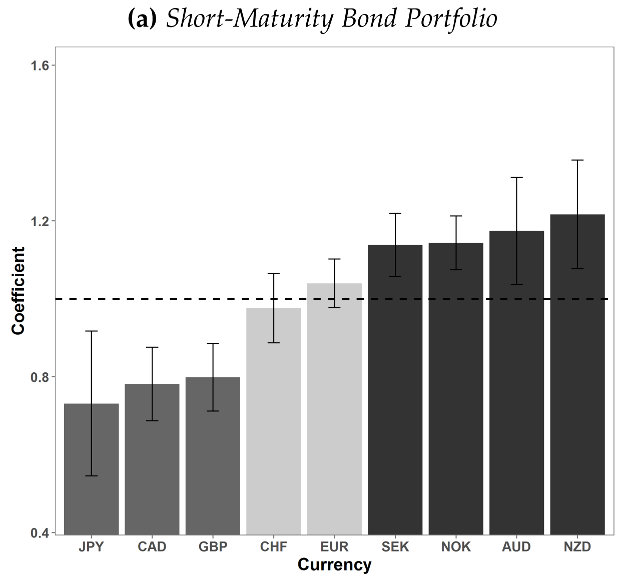


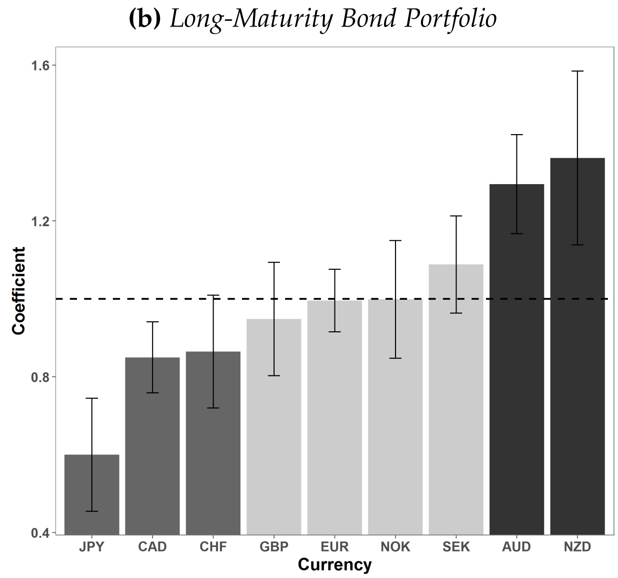
Notes: The figure depicts by how much a portfolio that shorts a given country's ten-year bond and lends at the US riskfree rate rises when the average portfolio rises by 1%, following a Fed tightening. Standard error bars are computed against the average portfolio rise of 1%; and the shading of the coefficient 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.

Figure 2.3 illustrates the argument that asymmetries deepen when currencies and bonds are linked, rather than being offset. In other words, it underscores that currency and bond markets are qualitatively rather than just quantitatively inconsistent under the leading explanations. These results contrast with the results of Lustig *et al.* (2017), who find that fluctuations in currencies and long-maturity bonds offset when linked in unconditional data. The key difference is the shocks: I use asset returns that are only exposed to monetary shocks, whereas their returns are measured at monthly frequencies and so reflect monetary and fundamental shocks alike.[8] Given the divergent findings, shocks to fundamentals may affect global markets differently than shocks to monetary policy. Unlike other shocks to international markets, Fed shocks are best explained by models of incomplete markets.

[8]In addition, Lustig *et al.* (2017) use data over sixty years, and currencies and long-maturity bonds offset better in the first few decades, than in the last few decades (which overlaps with my sample).

Figure 2.3: *Cross-Border Portfolio Responses to US Monetary Shocks*





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 against 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 2.1(a) and 2.2, except with common axes for comparison purposes.

2.3 Evidence from Bond Term Structures

I show further that central banks do not react to the Fed, using evidence from each country's bond term structure individually. Section 2.2 shows that asset reactions in currency and bond markets are inconsistent with a general explanation of central banks reacting to the Fed. This does not preclude explanations in which a few central banks react to the Fed while most countries see shifts in risk premia. This section addresses those concerns by offering evidence for each recipient country on its own, to show that the country's risk premia rather than its short rates react to Fed announcements, using its term structure of bond yields. Since these tests cannot distinguish between complete markets or incomplete markets, I use them only to argue against explanations in which central banks react.

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 each country's yield curve most exposed to its short rates, i.e. short-maturity yields, do not respond to Fed announcements. Second, I use a

Gaussian affine term structure model to decompose each country's yields explicitly into its path of short rates and term premia, and show again that the paths of rates globally do not react to the Fed. I contrast these results with two other sets of results. First, term premia — computed both through the simplified approach and the term structure model — do react to the Fed worldwide. Second, I apply the same methodology to announcements from all other central banks in my sample. I show that other central banks only affect the paths of their own short rates, and do not affect term premia or other countries' short rates.

These additional dimensions mollify concerns that my approach lacks statistical power, by showing that the methodology both detects the Fed affecting foreign yield curves and detects foreign central banks affecting their own yield curves in expected ways. Thus, the methodology could plausibly detect the Fed affecting foreign paths of rates, and its failure to do so is evidence against this class of explanations. Moreover, these additional dimensions illustrate a divergence in how central banks generate spillovers, as the Fed can shift term premia globally whereas other central banks cannot. This point is discussed further in Chapter 1.

2.3.1 Methodology: Inference by Heteroskedasticity

To conduct the analysis, I use an empirical methodology called inference by heteroskedasticity. Inference by heteroskedasticity is a method that tests the existence of reactions, i.e. whether a given bond responds at all to Fed announcements. The method infers that a bond does react to the Fed if the variance of its returns during announcement windows exceeds the variance during non-announcement windows. I first describe the method, and then I illustrate it by examining spillovers by all ten central banks on all ten-year yields.

Consider Equation (2.7), which describes a series of univariate asset returns r_t. In this case, I test whether $\beta = 0$, or whether an asset reacts to Fed announcements. The only assumption it requires is that returns during announcement windows would have the same distribution as returns during non-announcement windows, in the absence of Fed announcements. However, it is robust to misspecification on the dimensionality of Fed shocks, as covered further in Appendix B.1. The methodology is also transparent, as it gets all its statistical power from a single moment.

$$r_t = \alpha + \beta m_t + \epsilon_t \tag{2.7}$$

As an example, suppose I want to test whether Australian bond yields react to Fed announcements. I write Equation (2.7) and its non-announcement counterpart as follows.

$$\begin{aligned} \text{Announcement Windows: } \Delta y_t^{\text{AUD}} &= \alpha + \beta m_t^{\$} + \epsilon_t \\ \text{Non-announcement Windows: } \Delta \tilde{y}_t^{\text{AUD}} &= \epsilon_t \end{aligned}$$

To test $H_0 : \beta = 0$, i.e. whether Australian bond yields respond to Fed announcements, I take the variance of both sides and link the two equations through the variance of the error.

$$\begin{aligned} \text{Announcement Windows: } \mathbb{V}_{t-1}\left(\Delta y_t^{\text{AUD}}\right) &= \beta^2 \mathbb{V}_{t-1}\left(m_t^{\$}\right) + \mathbb{V}_{t-1}\left(\epsilon_t\right) \\ \text{Non-announcement Windows: } \mathbb{V}_{t-1}\left(\Delta \tilde{y}_t^{\text{AUD}}\right) &= \mathbb{V}_{t-1}\left(\epsilon_t\right) \end{aligned}$$

$$\mathbb{V}_{t-1}\left(\Delta y_t^{\text{AUD}}\right) > \mathbb{V}_{t-1}\left(\Delta \tilde{y}_t^{\text{AUD}}\right) \quad \Longrightarrow \beta \neq 0 \tag{2.8}$$

If Australian bond yields react to Fed announcements (if $\beta \neq 0$), those yields should be more volatile around Fed announcements than otherwise. I employ the Brown-Forsythe test to test for equality of variances. This test looks at median absolute deviations, rather than mean squared deviations as done by the F-test, another common test; and so it is robust to non-normal data.[9]

I illustrate this method by testing how all ten central banks affect the benchmark ten-year bond yields of all ten countries. The results, in Table 2.1, show that foreign ten-year bond yields globally react to the Fed and to their own central banks, while some foreign ten-year yields react to the ECB. (Rows refer to central banks and columns to that country's local ten-year bond, and I report the ratio of excess standard deviations when the test is significant at the 1% level and leave it blank otherwise.) The divergence between central banks is new in the monetary spillovers literature, and it is addressed further in Chapter 1. Continuing with the example, Australian yields are 207% more volatile around Fed announcements than they are otherwise.

As before, changes in ten-year bond yields can be decomposed into two components: changes in the paths of short rates or changes in the term premia over ten years. I next turn to examining

[9]For instance, in Appendix B.1, I show that the Brown-Forsythe test strongly outperforms the F-test on simulated data with high kurtosis. However, in Appendix B.4, I still show that my results are qualitatively unchanged when using both the F-test and the Kolmogorov-Smirnov test, which checks for equality of distributions between announcement returns and non-announcement returns.

Table 2.1: *Excess Volatility in 10Y Bond Returns*

	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK	USD
Australia	144%									
Canada	30%	82%								
Switzerland			124%					96%		
Euro	34%		39%	88%	37%		133%		90%	24%
United Kingdom					79%					
Japan						34%				27%
Norway							26%	34%		
New Zealand								26%		
Sweden				39%					82%	
United States	207%	144%	40%	216%	46%			54%	25%	233%

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 B.5. The Fed and ECB have spillover effects, but most other central banks only affect their own ten-year bonds.

which component reacts to Fed announcements.

$$10\,\underbrace{\Delta y_t^j(t,t+10)}_{\text{10Y Yield}} = \underbrace{\sum_{k=1}^{10} \Delta i_{t+k-1}^j}_{\text{10Y Path of Rates}} + \underbrace{\sum_{k=1}^{10} \Delta \gamma_{t+k-1}^j}_{\text{10Y Term Premia}}$$

2.3.2 Test 1: Extremities of Yield Curve

I first argue that Fed spillovers are driven by shifts in foreign term premia, rather than by the reactions of foreign central banks, by examining the ends of foreign yield curves. Movements at the short end of a yield curve are driven by the policies of its central bank, while movements at the long end of a yield curve are driven by shifts in term premia. Empirically, Fed announcements affect the long ends of yield curves across the globe, but do not affect the short ends of those curves, suggesting that spillovers work through term premia. The short ends of foreign yield curves respond only to their own countries' monetary announcements.

The logic behind this test is as follows. First, short yields (e.g. one-year yields) embed small term premia. Hamilton (2009) shows 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.

$$T_0 \underbrace{\Delta y_t^j(t, t+T_0)}_{T_0 \text{ Yield}} = \left(\sum_{k=1}^{T_0} \Delta i_{t+k-1}^j + \sum_{k=1}^{T_0} \Delta \gamma_{t+k-1}^j \right) \approx \underbrace{\sum_{k=1}^{T_0} \Delta i_{t+k-1}^j}_{T_0 \text{ Path of Rates}}$$

Second, 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 Appendix B.2, 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 too. This assumption seems to hold well in domestic data, as Adrian *et al.* (2013) find that over 80% of variation in US long forward yields on Fed announcement days are driven by term premia shifts.

$$(10 - T_1) \underbrace{\Delta y_t^j(T_1, t+10)}_{\text{Forward Yield over } (T_1, 10\text{Y})} = \left(\sum_{k=1}^{10-T_1} \Delta i_{T_1+k-1}^j + \sum_{k=1}^{10-T_1} \Delta \gamma_{T_1+k-1}^j \right) \approx \underbrace{\sum_{k=1}^{10-T_1} \Delta \gamma_{T_1+k-1}^j}_{\text{Term Premia over } (T_1, 10\text{Y})}$$

There is a tradeoff in power and in contamination in selecting 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. I set $T_0 = 1$ year, as that represents the shortest maturity in my dataset. 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. They similarly 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 $T_1 = 6$ years, and focus on the six-year ahead, four-year yield (i.e. a yield that can be locked into in 2018 to borrow and lend between 2024 to 2028).

I utilize the Inference by Heteroskedasticity method to measure how these measures of foreign yields react to announcements by the Fed (and by other central banks). The results are presented in Tables 2.2 and 2.3.

The findings are stark. Table 2.2 shows that the Fed does not affect other countries' paths of short-term policy rates. However, the central banks of those countries do affect their own paths of policy rates. Table 2.3 shows that the Fed has strong effects on other countries' term premia, and interestingly the central banks of other countries do not. This points to an explanation in which term premia in *each* individual country, rather than its central bank's plans, adjust around Fed announcements.

Table 2.2: *Excess Volatility in Daily 1Y Bond Returns*

	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK	USD
Australia	86%									
Canada		67%								
Switzerland			108%					36%		
Euro			45%	53%						
United Kingdom	28%		29%							
Japan						105%				
Norway							131%			
New Zealand	53%							92%		
Sweden									133%	
United States										79%

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 B.6. Central banks affect their own one-year bonds, but the Fed does not affect other countries' one-year bonds.

The results are robust to different cutoffs, as shown in Appendix B.4. Consider one extreme example: ten-year forward twenty-year rates, or rates that can be locked into in 2018 for borrowing and lending between 2028 and 2048. It is implausible that central banks regularly release guidance at such horizons.

Table 2.4 shows the results, for the sample of countries that issue bonds with thirty-year maturities. 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 B.4. This confirms that Fed announcements induce strong shifts in term premia for each country.

Table 2.3: *Excess Volatility in Daily 6F4Y Bond Returns*

	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK	USD
Australia										
Canada										
Switzerland								87%	28%	
Euro			19%	32%						
United Kingdom										
Japan										
Norway		30%						38%		
New Zealand										
Sweden										
United States	38%	36%	30%	50%	29%			57%	29%	51%

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 B.7. The Fed affects most other countries' six-year forward four-year bonds.

Table 2.4: *Excess Volatility in Daily 10F20Y Bond Returns*

	AUD	CAD	CHF	EUR	GBP	JPY	USD
Australia							
Canada							
Switzerland	54%						
Euro							
United Kingdom							
Japan							
United States		32%	36%	32%			51%

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 B.8. This rate cannot be constructed for New Zealand, Norway, and Sweden as they do not issue liquid thirty-year bonds, and so they are omitted. The Fed affects many other countries' ten-year forward twenty-year bonds.

2.3.3 Test 2: Affine Term Structure Model

I next argue that Fed spillovers are driven by shifts in foreign term premia, rather than by the reactions of foreign central banks, by using an affine term structure model to decompose yield curves explicitly into the paths of rates and term premia. As before, the results again show that foreign term premia respond to the Fed, while the paths of rates do not respond; the paths of rates only respond 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 to international yield curves from the ten countries in my sample.

I choose this model over other choices for two reasons. First, this model can decompose yields at a daily frequency, in contrast to international models that operate at monthly or quarterly frequencies, such as Wright (2011). In this model, state variables are principal components of the yield curve itself, measured at daily frequencies; whereas most other models use macroeconomic 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 B.3.

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 2.5 and 2.6.

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 — although it cannot distinguish between models of complete or incomplete markets — and that explanations around central banks do not seem plausible. The results from this method align with those by Bauer and Neely (2014), who use dynamic term structure models to show that quantitative easing by the Fed affected term premia in four other countries.

Table 2.5: *Excess Volatility in Daily 10Y Rate Returns*

	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK	USD
Australia	97%									
Canada		82%								
Switzerland	28%		91%				92%			
Euro				38%						
United Kingdom	44%						30%			
Japan						92%				21%
Norway							132%			
New Zealand	35%							102%		
Sweden									127%	
United States		40%		30%						97%

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 B.10. The Fed affects few other countries' estimated paths of rates.

2.4 Models of Complete Markets

I allow for more complex stochastic discount factors under market completeness, and show that the restrictions that my results require are either impossible or economically implausible for such models to match. Section 2.2 shows that asset reactions in currency and bond markets are inconsistent with transitory responses of stochastic discount factors to the Fed. This does not preclude more complex models in which stochastic discount factors have multiple forms of heterogeneity. In this section, I derive the restrictions that more complex stochastic discount factors

Table 2.6: *Excess Volatility in Daily 10Y Term Returns*

	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK	USD
Australia	29%									
Canada										
Switzerland		38%	58%							61%
Euro				28%						21%
United Kingdom										
Japan										22%
Norway							46%	20%		
New Zealand										
Sweden									59%	
United States	27%	46%	25%	35%	32%			30%	17%	68%

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 B.11. The Fed affects many other countries' estimated paths of term premia.

must obey to match my results, both in two commonly-used models and in a preference-free and distribution-free framework. Even in the preference-free framework, the restrictions are jointly difficult to match.

I show this using the tension in the cross-sectional sorting of countries between currency markets and bond markets, as shown in Section 2.2. As before, asymmetries in the currency market imply that stochastic discount factors in low-rate countries (e.g. Japan) are more volatile than ones in high-rate countries (e.g. Australia) following Fed announcements, and asymmetries in the bond market imply the opposite. This tension is resolved only by making stochastic discount factors heterogeneous in multiple ways, as a single form of heterogeneity is insufficient. Following the framework of Alvarez and Jermann (2005) and Lustig *et al.* (2017), I decompose the stochastic discount factor into two components: a permanent component and a transitory component, on which exchange rates and bonds load differentially. To align with currency and bond markets concurrently, the low-rate permanent component must be made more volatile and the low-rate transitory component made less volatile than their high-rate counterparts.[10] This

[10] While the correlation between currencies and bonds could add a third restriction, I find that its insight largely duplicates the insights from currencies and bonds separately. Specifically, this third potential restriction is that $\rho(\Delta m_t^J - \Delta m_t^A, \Delta y_t^J - \Delta y_t^A) < 0$. This is a useful restriction on its own, as shown in Section 2.2, but it does not

gives the stochastic discount factors enough mathematical freedom to match my results, but economically these two restrictions are highly unusual. I show using two common models that it is difficult if not impossible to disentangle transitory and permanent components in meaningful economic terms.

This result sheds light on models of global markets, by offering results that integrate two asset markets and stem from a well-identified shock. Most models assume market completeness and use one form of heterogeneity to explain currency markets, and so altogether miss the second form of heterogeneity that bond markets require. Lustig *et al.* (2017) use models of complete markets to study both currency and bond markets together, but similarly find one type of heterogeneity to be sufficient using low-frequency data. However, returns at low frequencies reflect many different shocks. By isolating one specific shock in two markets concurrently, I find that two types of heterogeneity are necessary. This imposes burdens on models with full risk-sharing that may be too great.[11] Thematically, this relates to Backus and Smith (1993) or Brandt *et al.* (2006), who respectively show that models with complete markets generate cross-sectional predictions for consumption and exchange rates that are inconsistent with the data, or generate implied correlations of stochastic discount factors that are unrealistic. But whereas these puzzles document tension using low-frequency and unconditional shocks, I document tension using high-frequency and identified shocks.

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 and distribution-free framework into its transitory and permanent components. For narrative purposes, I continue to focus on Japan and Australia as my representative examples of low-rate and high-rate countries.

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meaningfully strengthen the distinct results on currencies and bonds unless strong modeling assumptions are made. I thus do not consider it in this section.

[11] Zhang (2017) uses a model with complete markets to explain monetary spillovers in currency and bond markets, although the paper focuses on short-maturity rather than long-maturity bonds; and so the stresses placed on such models are not apparent.

2.4.1 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 2.2, although I generate restrictions imposed by currency and bond markets separately. 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, and it is *impossible* to resolve this tension as the model does not permit multiple forms of heterogeneity. 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, where shocks are realized at t. The shocks hitting Australia and Japan have different volatilities, but I restrict all other parameters (ρ, β, γ) to be common. I relax this assumption later.

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

I first consider currency markets, and define the exchange rate S to be yen per Australian dollars. Under the two assumptions behind complete markets — full spanning and no frictions — 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 state-by-state having adjusted for exchange rates.

$$\underbrace{\beta \left(\frac{C_t^A}{C_{t-1}^A} \right)^{-\gamma}}_{\text{Australian SDF}} = \underbrace{\beta \left(\frac{C_t^J}{C_{t-1}^J} \right)^{-\gamma}}_{\text{Japanese SDF}} \frac{S_t}{S_{t-1}}$$

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 more generally, 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.[12] I use this finding to generate my first restriction.

$$\underbrace{\mathbb{E}_{t-1}\left(s_t - s_{t-1}\right) + r^A_{f,t-1} - r^J_{f,t-1}}_{\text{Excess Currency Return}} = \frac{1}{2}\gamma^2\left((\sigma^J)^2 - (\sigma^A)^2\right) > 0 \tag{2.9}$$

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 expression since the stochastic discount factor is conditionally lognormal.

$$\mathbb{E}_t\left(\beta^k\left(\frac{C^i_{t+k}}{C^i_t}\right)^{-\gamma}\frac{1}{P^i_t(t,t+k)}\right) = 1 \quad \Rightarrow \quad \log P^i_t(t,t+k) = k\log\beta - \gamma\left(\mathbb{E}_t c_{t+k} - c_t\right) + \gamma^2\mathbb{V}_t c_{t+k}$$

Moreover, the innovation to yields in zero-coupon bonds is the (negative) innovation to log prices.

$$k\Delta y^i_t(t,t+k) = -\Delta\log P^i_t(t,t+k) = \gamma\left(\Delta c_{t+k} - \Delta c_t\right)$$

Finally, I draw directly on evidence from Figure 2.1(b), 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.

$$\underbrace{\mathbb{V}_{t-1}\left(\Delta y^J_t(t,t+k)\right) - \mathbb{V}_{t-1}\left(\Delta y^A_t(t,t+k)\right)}_{\text{Difference in Bond Yield Volatility}} = \frac{\gamma^2}{k^2}\left(1-\rho^k\right)^2\left((\sigma^J)^2 - (\sigma^A)^2\right) < 0 \tag{2.10}$$

Equations (2.9) and (2.10) 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 readjusting their portfolios. These conditions continue to contradict each other even if parameters (β,γ) are heterogeneous across countries.

The only possible resolution is for ρ 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

[12]This pattern holds up in my sample with varying significance too, but I have a shorter sample and thus less power than Mueller *et al.* (2017).

evidence. 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 B.5, I find that that *thirty*-year bonds in Australia and other high-rate countries are more volatile than those in Japan and other low-rate countries, around Fed announcements.

2.4.2 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 dynamic consumption processes. In this model, a single form of heterogeneity is insufficient. Two forms of heterogeneity are mathematically sufficient, but they are economically implausible together. 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. International models do not typically make such nuanced distinctions, as countries are either more or less exposed to the US overall in such frameworks.

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 modeling setups. 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 others (e.g. high ϕ and ρ), this is the model of long-run risk by Bansal and Yaron (2004). I present the consumption dynamics for country i, although I modify the consumption dynamics to incorporate heterogeneity next.

$$c_t^i - c_{t-1}^i = \mu + \phi x_{t-1}^i + \sigma_{t-1}^i \eta_t^i$$

$$x_t^i = \rho x_{t-1}^i + \varphi_e \sigma_{t-1}^i e_t^i$$

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

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 into two independent components: a global component e^z_t and an idiosyncratic component e^i_t. Different countries i have differential loadings $1+\beta^i_e$ on the global components of shocks.[13] I utilize that modeling innovation, and in fact decompose all shocks (η, e, w) into both global and idiosyncratic components. Global components of (η, e) have constant global volatility, while idiosyncratic components continue to have idiosyncratic stochastic volatility. Finally, shocks decompose into these two components with weightings $(\alpha_\eta, \alpha_e, \alpha_w)$. I present the updated dynamics.

$$c^i_t - c^i_{t-1} = \mu + \phi x^i_{t-1} + \left(\sqrt{\alpha_\eta}\sigma\left(1+\beta^i_\eta\right)\eta^z_t + \sqrt{1-\alpha_\eta}\sigma^i_{t-1}\eta^i_t\right)$$

$$x^i_t = \rho x^i_{t-1} + \varphi_e\left(\sqrt{\alpha_e}\sigma\left(1+\beta^i_e\right)e^z_t + \sqrt{1-\alpha_e}\sigma^i_{t-1}e^i_t\right)$$

$$\left(\sigma^i_t\right)^2 = \sigma^2 + v\left(\left(\sigma^i_{t-1}\right)^2 - \sigma^2\right) + \sigma_w\left(\sqrt{\alpha_w}\left(1+\beta^i_w\right)w^z_t + \sqrt{1-\alpha_w}w^i_t\right)$$

As before, the empirical results in currencies require that returns in the Japanese stochastic discount factor m^J_t be more volatile than returns in the Australian one m^A_t; and as before, the empirical results in bonds require changes in Australian yields to be more volatile than changes in Japanese yields.

$$\mathbb{V}_{t-1}\left(\Delta m^J_t\right) > \mathbb{V}_{t-1}\left(\Delta m^A_t\right) \quad \text{and} \quad \mathbb{V}_{t-1}\left(\Delta y^A_t\right) > \mathbb{V}_{t-1}\left(\Delta y^J_t\right) \tag{2.11}$$

In Appendix B.5, I derive the expressions for the variances in innovations for stochastic discount factors m^i_t and for infinite-maturity bond yields y^i_t 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.

$$\begin{aligned}\mathbb{V}_{t-1}\left(\Delta m^i_t\right) = \alpha_\eta\left(\gamma\sigma\right)^2\left(1+\beta^i_\eta\right)^2 \quad &+ \quad \alpha_e\left(\left(1-\rho\right)^{-1}\left(\gamma - 1/\psi\right)\phi\varphi_e\sigma\right)^2\left(1+\beta^i_e\right)^2 \\ &+ \quad \alpha_w\left(\left(1-v\right)^{-1}\left(\gamma-1/\psi\right)\left(1-\gamma\right)K_0\sigma_w\right)^2\left(1+\beta^i_w\right)^2 \quad + \quad \ldots\end{aligned}$$

$$\begin{aligned}\mathbb{V}_{t-1}\left(\Delta y^i_t\right) = \alpha_e&\left(\left(1-\rho\right)^{-1}\left(1/\psi\right)\phi\varphi_e\sigma\right)^2\left(1+\beta^i_e\right)^2 \\ &+ \quad \alpha_w\left(\left(1-v\right)^{-1}\left(1/\psi - \gamma - \gamma/\psi\right)K_0\sigma_w\right)^2\left(1+\beta^i_w\right)^2 \quad + \quad \ldots\end{aligned}$$

[13] Without loss of generality, I restrict $(1+\beta^i) \geq 0$.

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

These expressions make clear the difficulties this model faces in matching my empirical findings. First, they show that a single form of heterogeneity is insufficient. For instance, suppose only the global loading $1+\beta^i_\eta$ 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 (2.11). However, in this setup Japan and Australia have the same variance in bond yields, violating the bond restrictions in Equation (2.11). A related argument applies to $1+\beta^i_e$ or $1+\beta^i_w$: 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 forms of heterogeneity are mathematically sufficient, but they are economically 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 (2.11). In turn, this requires $1+\beta^A_e > 1+\beta^J_e$ to match the bond restriction in Equation (2.11).[14]

This is economically implausible, 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 easily generate these two results. Countries that are relatively more exposed to the Fed through trade flows, bank linkages, and other common channels are likely be relatively more exposed in all dimensions of consumption. 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 η and e load relatively more on the stochastic discount factors more while shocks w load relatively more on bonds. This

[14]These conditions are necessary but not sufficient; the actual minimum gap between the two loadings depends on the exact parameterization of the model.

illustrates the general tension that models of complete markets must confront when matching the asymmetries I document, even when rich heterogeneity is incorporated.

2.4.3 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 market completeness, 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 Australia 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-1}.[15] 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).[16] Second, in keeping with the notation of Alvarez and Jermann (2005), I distinguish between the pricing kernel Λ and the stochastic discount factor M, where M_{t+k} is the ratio of pricing kernels Λ_{t+k} and Λ_{t-1}, i.e. the growth rate of pricing kernels between future period $t+k$ and the pre-announcement period $t-1$. Third, I assume that each pricing kernel is the product of two components: a martingale permanent component, and a residual transitory component. 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

[15]This is equivalent to half the conditional variance of the log of a random variable when working with lognormal random variables.

[16]Gavazzoni *et al.* (2013) note a similar tension between currency returns and bond returns for unconditional asset returns, but conclude that higher-order moments are responsible. I show that higher-order moments do not resolve the tension around Fed announcements.

infinite-horizon limit.

$$M_t = \frac{\Lambda_t^i}{\Lambda_{t-1}^i} = \frac{\Lambda_t^{i,\mathbb{P}} \Lambda_t^{i,\mathbb{T}}}{\Lambda_{t-1}^{i,\mathbb{P}} \Lambda_{t-1}^{i,\mathbb{T}}}$$

I first turn to my results from currencies. Backus *et al.* (2001) generalizes the expression in Equation (2.9) from variance to entropy. Specifically, the excess currency return between two countries is equal to the differences in entropy of stochastic discount factors 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.

$$\begin{aligned} \underbrace{\mathbb{E}_{t-1}\left(s_t - s_{t-1}\right) + r_{f,t-1}^A - r_{f,t-1}^J}_{\text{Excess Currency Return}} &= L_{t-1}\left(\frac{\Lambda_t^J}{\Lambda_{t-1}^J}\right) - L_{t-1}\left(\frac{\Lambda_t^A}{\Lambda_{t-1}^A}\right) > 0 \qquad (2.12) \\ &= L_{t-1}\left(\frac{\Lambda_t^{J,\mathbb{P}}}{\Lambda_{t-1}^{J,\mathbb{P}}} \frac{\Lambda_t^{J,\mathbb{T}}}{\Lambda_{t-1}^{J,\mathbb{T}}}\right) - L_{t-1}\left(\frac{\Lambda_t^{A,\mathbb{P}}}{\Lambda_{t-1}^{A,\mathbb{P}}} \frac{\Lambda_t^{A,\mathbb{T}}}{\Lambda_{t-1}^{A,\mathbb{T}}}\right) > 0 \end{aligned}$$

This is the generalized restriction that much of the international asset pricing literature matches with various models. 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. As a third example, Ready *et al.* (2017) argues 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, rationalizing the unconditional carry trade. However, this prediction must be more nuanced to align with bond markets.

In bond markets, I continue to assume that my results for ten-year and thirty-year bonds 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.

$$P_t^i(\infty) = \mathbb{E}_t\left(\frac{\Lambda_\infty^i}{\Lambda_t^i}\right) = \mathbb{E}_t\left(\frac{\Lambda_\infty^{i,\mathbb{P}} \Lambda_\infty^{i,\mathbb{T}}}{\Lambda_t^{i,\mathbb{P}} \Lambda_t^{i,\mathbb{T}}}\right)$$

Alvarez and Jermann (2005) argue that at infinite maturities, there is no transitory component of the pricing kernel $\Lambda_\infty^{i,\mathbb{T}}$. 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 and for innovations in log yields.

$$P_t^i(\infty) = \frac{1}{\Lambda_t^{i,\mathbb{T}}} \quad \Rightarrow \quad \lim_{n\to\infty} n\Delta y_t^i = \Delta\lambda_t^{i,\mathbb{T}} = \log\left(\frac{\Lambda_t^{i,\mathbb{T}}}{\Lambda_{t-1}^{i,\mathbb{T}}}\right) - \mathbb{E}_{t-1}\log\left(\frac{\Lambda_t^{i,\mathbb{T}}}{\Lambda_{t-1}^{i,\mathbb{T}}}\right)$$

I take the entropy of exponentiated innovations in yields. Since entropy is invariant with respect to constant addition or multiplication, e.g. $L(a+cx) = L(x)$, this relates the entropy of gross bond yield innovations to the entropy of the transitory components of the stochastic discount factor.

$$\lim_{n\to\infty} L_{t-1}\left(e^{n\Delta y_t^i}\right) = L_{t-1}\left(\frac{\Lambda_t^{i,\mathbb{T}}}{\Lambda_{t-1}^{i,\mathbb{T}}}\right)$$

I compute the empirical entropy of gross innovations to yields following Fed announcements in my sample, and confirm in Appendix B.5 that it is statistically higher in high-rate countries at both ten-year and thirty-year maturities, as would be expected without large higher-order moments.[17] This yields the second restriction.

$$\underbrace{\lim_{n\to\infty} L_{t-1}\left(\exp\left(n\Delta y_t^A\right)\right) - L_{t-1}\left(\exp\left(n\Delta y_t^J\right)\right)}_{\text{Difference in (Gross) Bond Yield Entropy}} = L_{t-1}\left(\frac{\Lambda_t^{A,\mathbb{T}}}{\Lambda_{t-1}^{A,\mathbb{T}}}\right) - L_{t-1}\left(\frac{\Lambda_t^{J,\mathbb{T}}}{\Lambda_{t-1}^{J,\mathbb{T}}}\right) > 0 \qquad (2.13)$$

Equations (2.12) and (2.13) 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 form of heterogeneity cannot match these equations simultaneously.

To make this discussion more concrete, 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 ρ. If $\rho = 1$, shocks are permanent. Agents cannot smooth away a permanent shock, and so both bond yields and

[17] 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 results that are quantitatively smaller but qualitatively consistent with existing work, suggesting that higher order moments do not greatly bias the sample.

the transitory component of the stochastic discount factor remain constant. If $\rho < 1$, shocks are 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.

$$\begin{bmatrix} L_{t-1}\left(\Lambda_t^i/\Lambda_{t-1}^i\right) \\ L_{t-1}\left(\Lambda_t^{i,\mathbb{T}}/\Lambda_{t-1}^{i,\mathbb{T}}\right) \\ L_{t-1}\left(\Lambda_t^{i,\mathbb{P}}/\Lambda_{t-1}^{i,\mathbb{P}}\right) \end{bmatrix} = \begin{bmatrix} \gamma^2\left(\sigma^i\right)^2 \\ 0 \\ \gamma^2\left(\sigma^i\right)^2 \end{bmatrix} \text{ if } \rho = 1; \quad \begin{bmatrix} L_{t-1}\left(\Lambda_t^i/\Lambda_{t-1}^i\right) \\ L_{t-1}\left(\Lambda_t^{i,\mathbb{T}}/\Lambda_{t-1}^{i,\mathbb{T}}\right) \\ L_{t-1}\left(\Lambda_t^{i,\mathbb{P}}/\Lambda_{t-1}^{i,\mathbb{P}}\right) \end{bmatrix} = \begin{bmatrix} \gamma^2\left(\sigma^i\right)^2 \\ \gamma^2\left(\sigma^i\right)^2 \\ 0 \end{bmatrix} \text{ if } \rho < 1$$

Next consider the example with Epstein-Zin utility. I derive these expressions in Appendix B.5, using the approach of Hansen and Scheinkman (2009), and I again abstract from the components of entropy that are common across all countries. Each shock has a different loading. Level shocks to consumption η 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.

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.

$$\begin{bmatrix} L_{t-1}\left(\Lambda_t^i/\Lambda_{t-1}^i\right) \\ L_{t-1}\left(\Lambda_t^{i,\mathbb{T}}/\Lambda_{t-1}^{i,\mathbb{T}}\right) \\ L_{t-1}\left(\Lambda_t^{i,\mathbb{P}}/\Lambda_{t-1}^{i,\mathbb{P}}\right) \end{bmatrix} = A \begin{bmatrix} (1+\beta_\eta^i)^2 \\ (1+\beta_e^i)^2 \\ (1+\beta_w^i)^2 \end{bmatrix}$$

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

2.5 Models of Incomplete Markets

Two classes of explanations — ones in which central banks react to the Fed, and ones in which risk premia shift per models with full risk-sharing — do not seem consistent with my empirical results, leaving only models with incomplete markets to explain monetary spillovers. Future research is needed to develop a specific model in this class. In this section, I illustrate with a simple model of segmented markets à la Gabaix and Maggiori (2015).

Currency and bond markets sort countries in contradictory ways under the first two classes of explanations, but they sort countries in complementary ways under many models of incomplete markets. The distinction between high-rate and low-rate countries — coincidental under the previous two channels — becomes important under this channel. For instance, in models with segmented markets such as Gabaix and Maggiori (2015) or Alvarez *et al.* (2009), profit-seeking intermediaries are the marginal investors in bonds and currencies. When financial constraints tighten, intermediaries offload high-rate assets, causing both high-rate currencies to depreciate and high-rate bond yields to rise versus low-rate currencies and yields. Alternatively, in models with leverage constraints such as Maggiori (2017) or Bruno and Shin (2015), bonds and exchange rates are exposed to two related forms of heterogeneity. As an example, in Maggiori (2017), currencies are exposed to trade financing costs and bonds to leverage constraints, and these assets fall together when all financial constraints tighten concurrently. Models of complete markets require two types of countervailing heterogeneity, but many models of incomplete markets require only one form or two complementary forms of heterogeneity. I demonstrate with a model of segmented markets.

Specifically, consider Gabaix and Maggiori (2015). In this setting, assets in segmented markets are priced by constrained intermediaries. Currencies and bonds in which intermediaries have positive positions, e.g. those in high-rate countries, depreciate together versus those in which intermediaries have negative positions, when constraints on intermediaries tighten. This setup — which focuses on frictions rather than incomplete spanning — matches my empirical results, and I offer further evidence directly from observed measures of intermediaries' positions.

My example follows the multi-asset generalization version of Gabaix and Maggiori (2015). In this setting, intermediaries' positions θ across a set of assets are defined by three terms: the overall (scalar) constraint Γ of intermediaries, the variance-covariance matrix Σ of asset returns, and the expected returns $\mathbb{E}_t p_{t+1} - p_t$ of assets.[18]

$$\underbrace{\theta}_{\text{Positions}} = \Gamma^{-1} \underbrace{\Sigma^{-1} \left(\mathbb{E}_t p_{t+1} - p_t\right)}_{\text{Risk-Adjusted Profit}}$$

The crucial assumption for this model is that the Fed adjusts the constraint Γ. There are several potential explanations. Drechsler *et al.* (2017) argue that the cost of leverage is tied to the Fed's nominal rate, and so Fed actions shift intermediaries' abilities to borrow. Bruno and Shin (2015) and Banerjee *et al.* (2016) argue that Fed actions shift banks' net worth (due to unhedged balance sheet exposures), and so they shift intermediaries' investing capabilities. Empirical work by Miranda-Agrippino and Rey (2015) finds strong correlations between Fed actions and "market fear," which may reflect direct shifts in the risk aversion of intermediaries. I make further simplifying assumptions that the external supply, terminal payoffs, and riskiness of assets remain constant through Fed announcements, although this is not necessary for the model.

$$\underbrace{\frac{\partial p_t}{\partial \Gamma}}_{\text{Change in Prices}} = -\underbrace{\Gamma^{-1} \left(\mathbb{E}_t p_{t+1} - p_t\right)}_{\text{Risk-Neutral Positions}}$$

Fluctuations in prices are negatively proportional to expected returns, through intermediaries. First, intermediaries hold positive positions in assets with high returns, e.g. high-rate Australian bonds and the Australian dollar. Second, when constraints tighten, their prices fall the most to incentivize intermediaries to continue holding them. The opposite logic holds for assets with low returns, e.g. low-rate Japanese bonds and the Japanese yen. This stylized example matches the empirical results from currency and bond markets.

Although the level of interest rates is indirectly revealing about intermediaries' positions, Chapter 1 shows that actual measures of intermediaries' positions correlate with asymmetries in currency and bond markets. Measures of cross-border bank positions from the BIS and cross-border equity positions from the IMF correlate with asymmetries. For instance, the dollar appreciates most against currencies of borrower countries, and appreciates least against currencies of lender countries, when the Fed tightens. Similarly, the long-maturity bond yields of borrower

[18] Γ represents a reduced-form constraint that prevents intermediaries from otherwise taking infinite positions. For instance, Gabaix and Maggiori (2015) offer interpretations ranging from contracting frictions to risk aversion.

countries rise more than yields of lender countries. Such results fit with my example model.

While this model matches the relative movements of bonds and currencies, it requires further work. Theoretically, the model also predicts — in absolute terms — that Australian yields rise and Japanese yields *fall* when the Fed tightens, while empirically I find that Australian yields rise by a lot and Japanese yields rise by little when the Fed tightens. Without resorting to exogenous differences between Australian and Japanese assets, this requires further complexity within the intermediary sector. Quantitatively, Lustig and Verdelhan (2016) find that frictions cannot explain exchange rate movements, as the magnitude of frictions needed to explain some properties of exchange rates make it unable to match other properties. In short, this stylized model is qualitatively consistent with my fact, but further research is needed to account for these complications.

2.6 Conclusion

The asymmetries in currency and bond markets around the globe following Fed announcements are illuminating when considered together. 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 negative answers to the leading classes of explanations. The only explanations with which my findings align are models with incomplete markets, which illustrates the importance of frictions and heterogeneity in the international financial system.

These results offer guidance to policymakers, on debates over both global policy and capital controls. First, despite concerns of the trilemma being replaced by a “dilemma," monetary independence in developed markets survives; and this weakens the case for global monetary coordination. Jaime Caruana, former head of the Bank of International Settlements, has argued that since countries must follow the Fed *ex post*, they should be allowed to advise its policy *ex ante*. My results suggest that this rationale may be overstated for developed countries. Second, policymakers have discussed imposing capital controls to insulate their economies from monetary spillovers. Under complete markets, capital controls prevent agents from hedging risk optimally and decrease welfare; but under incomplete markets, they have the potential to mitigate frictions and improve welfare.

More generally, my results suggest that central banks do not yet feel the need to offset the Fed's actions on their real markets. On the other hand, central banks face challenges from the Fed over influence on their *financial* markets. As these markets grow in size and complexity, policymakers may want to consider preemptively wresting power over local investors back from the Fed. Otherwise, the Fed's financial spillovers today may spark the global real crises of tomorrow.

Chapter 3

Foreign Dollar Reserves and Financial Stability

3.1 Introduction

The foreign reserve portfolios of central banks vary greatly in size and in composition. For instance, Hong Kong's reserves exceed its GDP while Australian reserves are twenty times smaller than its GDP, and Peruvian reserves are overwhelmingly dollars while Romanian reserves are biased towards euros. The literature has posited both theoretical and empirical explanations to understand why countries hold foreign reserves, and papers have relied on the one piece of available information — the sizes of foreign reserves — to test theories. This paper generates a second piece of information — their *compositions* by currency — to explain foreign reserves.

In this paper, I estimate the currency composition of countries' foreign reserves, and show that the currency shares of countries' foreign reserves can be explained by the foreign currency shares of their financial systems' liabilities. First, I estimate currency shares of foreign reserves using a Bayesian dynamic linear model, in which I project changes in the size of a country's foreign reserves onto the returns of major reserve currencies. The results are novel, and show that dollars are far more prevalent than believed in foreign reserve portfolios. Second, I show that the dollar shares of countries' portfolios are correlated in the cross-section with the dollar shares of their financial systems' liabilities, particularly when controlling for swap lines — a form of emergency central bank-to-central bank lending that substitute for official foreign reserves. Third, I explain these findings using a model in which central banks use foreign reserves to hedge liquidity shocks to their financial sectors, which have foreign currency liabilities — particularly in the presence of foreign exchange transactions costs. Separately, the empirical findings also provide suggestive evidence against other explanations for foreign reserves. For instance, they are inconsistent with explanations in which foreign reserves provide fiscal space to governments during recessions, as these dollar-heavy portfolios are too lightly diversified across other reserve assets that appreciate during recessions.

The paper's main empirical finding is that the dollar shares of foreign reserve portfolios are massive: across the foreign reserves of seventy-seven emerging and developed countries, the average dollar share is 80-85%, and this number is fairly stable over time. The euro represents a small share for most countries; but for a handful of countries (e.g. Romania and Morocco), its share is comparable or larger than that of the dollar. These results are novel to the literature, as few countries report the composition of their foreign reserves publicly, and even fewer do so at any reasonable frequency. The only comprehensive public source of information — the IMF's currency composition of official foreign exchange reserves (COFER) data — only publishes currency shares aggregated across reporting countries, and does not even release the names of the countries who report to preserve their confidentiality. As such, the aggregated estimates are both skewed by large countries and are incomplete due to non-reporting countries. For instance, at the start of 2015, the IMF reported that they had recorded the currency composition for 55-60% of all foreign reserves, and 65% of those verified reserves were held in dollars.

The paper's second finding is that these dollar and euro shares correlate with the dollar and euro shares of external financial liabilities for emerging markets, although not for developed markets. There are two plausible reasons for this divergence. First, developed countries have been historically more willing than emerging countries to seek assistance from the International Monetary Fund, as Bird and Mandilaras (2011) find, and thus do not rely upon foreign reserves to stabilize their financial systems. This weakens the link between the currency shares of foreign reserves and financial liabilities for developed countries. Second and more recently, developed countries have largely benefited from "swap lines" to the Federal Reserve — a form of direct lending between central banks — which provide emergency dollar funding. The ECB has also established euro swap lines, although these have not been used widely yet. Swap lines generate dollar and euro reserves on demand, and thus substitute for actual dollar and euro reserves during liquidity crises. Indeed, the correlation between the foreign currency shares of reserves and banking liabilities holds when restricting to the countries that historically did not receive dollar or euro swap lines during crises.

To estimate the currency compositions of foreign reserves, I develop a dynamic linear model that projects returns in a country's aggregate portfolio value onto returns in reserve currencies over time. While the underlying portfolio shares of currencies are confidential, this is feasible because most countries report the values of their *total* portfolios at monthly frequencies. To illustrate the core insight, suppose a portfolio (with unknown dollar and euro shares) grows by 10% (in dollar terms) over a period in which the euro appreciates by 20% against the dollar. This portfolio must be 50% dollars and 50% euros. In practice, the methodology is more complicated for a few reasons. Portfolio weights may change over time, portfolios may grow or shrink due to inflows and outflows, and the set of potential reserve currencies is potentially large. Thus, the paper augments the dynamic linear model to allow for these complications. Finally, limited public data on currency shares are available for a handful of countries and in the aggregate, and so the paper further augments the dynamic linear model to incorporate this information, and uses Bayesian algorithms to find solutions. The methodology is novel, and can be applied to other cases in which the compositions of liquid portfolios are unavailable.

These two facts — the prevalence of dollar reserves and the correlation between dollar (and euro) shares in foreign reserves and financial liabilities — seem most consistent with models of financial stability, and I develop one. Specifically, I build a model in which a central bank hedges liquidity shocks for its banks that have borrowed cheaply and excessively in dollars, due to limited liability frictions. The central bank has one of three choices to mitigate liquidity shocks: it can print local currency and exchange it for dollars, it can hold diversified foreign reserves and exchange those for dollars, or it can hold dollars directly. However, the central bank faces foreign exchange costs for the first two strategies, and so it hold dollar reserves *ex ante*.[1] This model generates a causal link between dollar shares of foreign reserves and financial liabilities.

[1] Although currency markets are liquid, foreign exchange costs should be conceptualized broadly, e.g. the market panic that may follow large and sudden foreign exchange transactions by a central bank.

The model does not formally incorporate swap lines, but these are natural substitutes in practice for foreign reserves as they allow central banks to produce dollars on demand.[2]

In addition to supporting an explanation grounded in financial stability, the empirical findings on the currency shares can also provide new evidence against competing explanations for the purpose of foreign reserves. The literature has used the sizes of foreign reserves to argue against several possible explanations, by showing that foreign reserves are too large to smooth exchange rate fluctuations arising from current account imbalances, to be byproducts of sterilized exchange rate intervention, or to hedge rollover risk for short-term government debt. But size data cannot easily distinguish an explanation focused on financial stability from an alternate explanation: foreign reserves provide fiscal space to governments during recessions, as discussed by Fernandez-Arias and Montiel (2009), or more generally foreign reserves smooth consumption shocks, as discussed by Dominguez (2010). However, my findings on the currency compositions of foreign reserves are more consistent with an explanation based on financial stability than one based on fiscal space. Under the fiscal space explanation, reserve portfolios — which are similar to sovereign wealth funds — should be broadly diversified, with meaningful positions in Swiss francs and yen (currencies that appreciate strongly during global recessions). The overwhelming dominance of dollars in foreign reserve portfolios contradicts this alternate explanation, and supports the financial stability explanation, in which foreign reserves mitigate shocks to dollarized financial systems.[3]

The paper proceeds as follows. Section 3.2 reviews the literature on foreign reserves. Section 3.3 discusses the empirical framework and the data used to estimate the currency shares of foreign reserves. Section 3.4 documents the results and tests them against financial liabilities data. Section 3.5 develops a model that links foreign reserves and financial liabilities. Section 3.6 concludes.

[2] In the baseline model, liquidity shocks are exogenous and banks are financed in dollars. In a refinement of the model, liquidity shocks are endogenous, and banks could be financed in local currency — but crucially by investors whose stochastic discount factors are dollar-based. Even in this setting, foreign reserves are necessary to mitigate liquidity shocks. Since investors care about dollar returns, they may refuse to rollover funding to a solvent and liquid country if they anticipate a currency depreciation. Foreign reserves allow a central bank to stabilize the exchange rate, which stops emergent liquidity shocks. The model's predictions are still being developed, and so this model is not presented in this version.

[3] In addition, the popularity of swap lines is further suggestive evidence in favor of a financial stability explanation. Swap lines are a source of temporary funding that entail no nominal capital gains. Thus, they should only be valuable to mitigate liquidity shocks, and not to increase a country's budget.

3.2 Literature Review

This paper contributes to two groups of literature: on the motivations for holding foreign reserves, and on the dollar's dominance in the international financial system. With respect to the first strand of literature, reserves are strongly associated with lower risks of crises, as Catao and Milesi-Ferretti (2014) find empirically, and policymakers are broadly advised to follow the Greenspan-Guidotti rule-of-thumb, which advises foreign reserves in excess of external short-term debt.

However, the specific reason that countries hold foreign reserves remains contested. Consider the five most popular (partially overlapping) hypotheses: (i) foreign reserves are the byproduct of sterilized exchange rate interventions, (ii) foreign reserves stabilize exchange rates during current account fluctuations, (iii) foreign reserves hedge rollover risk for short-term government debt, (iv) foreign reserves stabilize consumption during recessions, and (v) foreign reserves ensure financial stability. Dominguez (2010) argues against the first explanation by noting that foreign reserve buildups are too large to be accidental. The second and third explanations have generated more serious debates. For instance, Aizenman and Sun (2012) argues against the second explanation, noting that countries were unwilling to spend reserves to stabilize their currencies during the financial crisis. On the other hand, Dominguez (2012) argues for this explanation, arguing that countries did actively manage their currencies during the crisis; and Dominguez (2014) discusses examples from non-Eurozone European countries. The third explanation originates in the literature on "original sin" (a country's inability to issue local-currency debt), and recent work includes Bianchi *et al.* (2013), who build and calibrate a model to generate policy recommendations. However, Obstfeld *et al.* (2010) argue that these two explanations seem largely inconsistent with the sizes of foreign reserve portfolios. They use back-of-the-envelope calculations to argue that these explanations would require foreign reserves worth 0.1-0.5% of GDP weekly, whereas in practice foreign reserves vastly dwarf that estimate, even under extreme assumptions on the length of crises.

The fourth and fifth explanations are more consistent with large foreign reserve portfolios, and have been supported alternately by Fernandez-Arias and Montiel (2009) and Dominguez (2010) and by Obstfeld *et al.* (2010) respectively, among others. Other papers include Kim and Ryou (2011), who argue that foreign reserve portfolios largely fail mean-variance efficiency tests, suggesting that they have purposes beyond stores of value. None of these papers utilize portfolio compositions to test hypotheses, however; and that is where my paper can contribute to this literature. Finally, Eichengreen *et al.* (2017) has floated non-economic reasons that countries hold foreign reserves (e.g. to strengthen geopolitical alliances), but that is beyond the scope of this paper.

Moreover, a set of papers indirectly argues that foreign reserves are useful for financial stability by noting that countries are reluctant to turn to the IMF, and foreign reserves substitute for IMF programs. Bird and Mandilaras (2011) and Fernandez-Arias and Levy-Yeyati (2012) argue that the IMF is unpopular for several reasons: programs take a long time to negotiate and often involve conditionality. Furthermore, Joyce and Razo-Garcia (2011) put forward a model to show why reserves are favored over IMF programs, and note that the programs are often too small in practice. (They were both insufficient for the Mexican and East Asian crises of the 1990s, and despite the recent quota expansions, emerging markets often hold reserves many times larger than their IMF allocations.) Finally, while regional funds have emerged, e.g. the Latin American Reserve Fund, Rosero (2014) note that their advantages over foreign reserves are still unproven.

Separately, this paper contributes to the literature on the dollar's dominance in the international financial system, by establishing its dominance in most countries' reserve portfolios. The dollar's ubiquity has been well-established in other domains, e.g. Gopinath (2015) in trade invoicing and Bruno and Shin (2015) in bank lending. However, there are concerns that foreign reserve portfolios are adjusting away from the dollar. Truman and Wong (2006), Wong (2007), and Wooldridge (2006) discuss the composition of foreign reserves from limited public data, and note cautiously that these fears may be overstated. My results validate this claim more forcefully, showing that the dollar remains as prevalent as ever across a wider set of countries.

3.3 Empirical Framework

The paper's core contribution is to estimate the currency composition of foreign reserves for individual countries, as most countries do not report their own breakdown, and the IMF only reports an aggregated and incomplete breakdown to preserve confidentiality. (In fact, the IMF does not release the names of the countries who contribute to the series.) I estimated a modified dynamic linear model using Bayesian algorithms.

The paper uses two main pieces of data: the sizes of reserves holdings at the country level and reserve currency returns, and projects the former onto the latter. To illustrate the key insight, consider the following toy scenario: a country who passively holds only dollars and euros reports that its foreign reserves have risen (in dollar terms) by 20% over a month. If the euro appreciated against the dollar by by 40% during that month, the portfolio must be 50% dollars and 50% euros.

Of course, this insight does not generalize to multiple assets without further structure. For instance, consider a third asset: the Japanese yen. If that appreciates versus the dollar by, say, 20%, the portfolio holdings are indeterminate. More observations alone are insufficient, as countries may change their portfolio weights continuously. However, under some reasonable assumptions about portfolio share stickiness and under some prior beliefs on overall portfolio shares (released either in the aggregate by the IMF, or by a few specific countries), solutions can be found. As such, I augment a dynamic linear model with a Bayesian prior, and use Bayesian methods (Markov Chain Monte Carlo) to find the solution.

In addition, this approach does not account for flows: portfolios can grow or shrink outside of currency fluctuations, if foreign reserves are actively added. If inflows are correlated with currency returns, this can potentially bias the estimated currency shares. Finding a suitable instrument or bias-free sub-sample is difficult, and so I use various structural assumptions to control the bias.

This methodology of uncovering portfolios is novel, although the general insight has been used to estimate confidential baskets to which countries peg their currencies. Specifically, Fidrmuc (2010), Frankel and Wei (2008), and Frankel and Xie (2010) similarly decompose local currency movements into various reserve currencies, to find the *de facto* peg. However, in these papers, portfolio weights are much stickier, and the methodologies do not worry about incorporating prior data or adjusting for flows.

3.3.1 Data

All but a handful of central banks do not disclose the composition of their foreign reserves. But the IMF, on behalf of central banks, reports two key pieces of information: the total value of individual central bank reserves on a monthly basis (e.g. Japan holds $1.212 trillion as of August

2014), and the quarterly composition of central bank portfolios in aggregate (e.g. Euros composed 24.7% of known central bank portfolios in 2004, worldwide). The paper relies on these two pieces of information, along with monthly currency returns for large reserve currencies (obtained from the Federal Reserve).

The first piece of data — the total value of individual central bank reserves, on a monthly basis — are collected by the IMF for seventy-five central banks, and by the Federal Reserve for two more central banks. (Supranational entities like the European Central Bank are omitted.) The seventy-seven countries are broken down by region: twenty-five in Western Europe, fifteen in Eastern Europe, eleven in Middle East and North Africa, twelve in East Asia, and fourteen in the Americas. Coverage naturally gets better over time, with approximately forty countries reporting their total values as early as 2000 and almost all countries reporting by 2010. Each country is analyzed from when they start reporting reserves to the IMF (at the earliest, in 2000) until 2013.

The second piece of data, formally known as the Currency Composition of Official Foreign Exchange Reserves (COFER) database, is particularly tantalizing. In its disaggregated form, it is precisely what this paper seeks. Yet the disaggregated version is inaccessible; the IMF states:

> [Composition] data for individual countries are kept strictly confidential given the sensitive nature of the data. Access to individual country data is limited to only four IMF staff on a need-to-know basis.

No modern papers have bypassed this restriction; to my knowledge, the only exception is Eichengreen and Mathieson (2000), which accessed the underlying data two decades ago. Thus, I use the public and aggregated version, which is used to construct prior beliefs about portfolio composition. In addition, a handful of countries report their currency shares publicly, as documented by Wong (2007) and Truman and Wong (2006). I do not explicitly use this information in my model, but I do use it as an after-the-fact check of the results.

Finally, I use monthly currency returns, collected by the US Federal Reserve. Seven currencies correspond to approximately 97% of global reserves, and so the set of reserve currencies is defined as: the dollar, the euro, the pound, the yen, the Swiss franc, the Canadian dollar, and the Australian dollar. While the Chinese yuan has been discussed for several years as an emerging reserve currency, this has failed to appear in the data, and so it is not included. The IMF only broke out the yuan in the December 2016 COFER update, and it constituted a mere 1% of portfolio holdings.

Since the IMF data breaks out gold from currencies, I also do not consider gold. Gold may be formally incorporated in a future version, but most countries have small gold reserves (with a few notable examples, e.g. Switzerland).

Finally, foreign reserves are typically invested in riskfree bonds in that currency than just in the currency itself. However, the monthly variation in riskfree returns is minuscule compared to the monthly variation in exchange rates, and Wooldridge (2006) similarly notes that foreign reserve managers are largely concerned about currency rather than interest rate risk. In my specification, constant differences in riskfree rates across countries will be captured by a constant.

3.3.2 Methodology

This section develops the dynamic linear model for estimating the currency shares. The portfolio size at time $t+1$ reflects two components: the portfolio size at time t (times the gross return on that portfolio) plus any inflows or outflows:

$$P_{t+1} = P_t(1 + r_{t+1}) + F_{t+1}$$

Therefore, the overall growth rate of the portfolio can be decomposed into the net return and (scaled) inflows and outflows:

$$g_{t+1} = \frac{P_{t+1} - P_t}{P_t} = r_{t+1} + f_{t+1}$$

Finally, I decompose the portfolio net return into the weighted return by currency, where weights can also adjust at every point in time. This formulation — in which coefficients can change over time — is known as a dynamic linear model, and it can be written as:

$$g_{t+1} = \sum_{k=1}^{K} w_t^k r_{t+1}^k + f_{t+1} \tag{3.1}$$

In addition, I impose the restrictions that weights sum to one and are non-negative, as central banks do not meaningfully short currencies in their foreign reserves portfolios.

$$w_t^k \in [0, 1] \qquad \sum_{k=1}^{K} w_t^k = 1$$

There are three small adjustments. First, I estimate Equation (3.1) in logs. Second, I embed the restriction that weights sum to one by using relative returns rather than absolute returns. Third, I allow weights to adjust at annual rather than the monthly level, to keep the parameter space relatively compact. In the revised formulation, Equation (3.2), $T(t)$ is defined as a function that converts a given month t into the corresponding year.

$$g_{t+1} - r_{t+1}^K = \sum_{k=1}^{K-1} w_{T(t)}^k (r_{t+1}^k - r_{t+1}^K) + f_{t+1} \tag{3.2}$$

Flows

In Equation (3.2), the estimation strategy must contend with f_{t+1}, which represents active inflows and outflows from the foreign reserves. Flows are likely correlated with currency movements, and they are almost always unobserved, although Dominguez *et al.* (2012) notes that the IMF has recently started requesting more nuanced information that may help in estimating these components. Depending on the coverage of the results, this may be incorporated in a future iteration of the estimation strategy.

This classic omitted variables problem leads to misleading portfolio estimates. For instance, consider a defensive central bank that always increases its US dollar holdings when negative economic shocks hit the world. Since the dollar also tends to appreciate during such periods, estimating Equation (3.2) naively will overestimate the portfolio weight on dollars. Reserves increase precisely when dollars are performing well, and I will give too much credit to existing dollar holdings.

There are two broad classes of solutions: an instrumental variables approach, or a structural approach. The former is virtually impossible: it requires a variable that drives currency returns (e.g. the dollar-euro exchange rate) but does *not* affect how a central bank actively responds. All the classic macroeconomic drivers of exchange rates (e.g. interest rates, current account imbalances, etc) would not pass the exclusion restriction.

The second solution is more feasible, by using a flexible structural model to model central bank choices. The key assumption is that central banks respond disproportionately to larger moves in currency markets, i.e. small shocks trigger no policy change, whereas large shocks do. This can be embedded through three steps. First, I incorporate a cubic term for the local exchange rate against a basket of SDRs (a mix of dollars, euros, yen, and pounds). This allows large movements in the exchange rate to absorb variation in portfolio returns.

$$g_{t+1} - r_{t+1}^{K} = \sum_{k=1}^{K-1} w_{T(t)}^{k} (r_{t+1}^{k} - r_{t+1}^{K}) + \sum_{m=0}^{3} \beta_{T(t)}^{m} e_{t}^{m} \tag{3.3}$$

Second, I detrend the portfolio returns using a standard Hodrick-Prescott filter, allowing for slow-changing trends in foreign reserve accumulation. Finally, I weight the residuals inversely to the size of local currency returns, allowing small movements (representing calm and stable times) to drive the model's estimation.

Weights

The dynamic linear model in Equation (3.3) requires further structure, particularly on time-varying coeffcients w_t and β_t. First, I impose a Markovian assumption on how these coefficients change through time; and second, I parameterize that stochastic process.

First, a dynamic linear model imposes the Markovian assumption that the values of these coefficients at a given point in time are only directly related to the previous and subsequent values. This assumption seems reasonable: conditional on knowing the portfolio share of dollars in 2004, I can assume that the portfolio shares of dollars in 2003 and 2005 are independent. This allows a simplification of the standard likelihood function into a more tractable representation, where D_t represents data and θ all other parameters:

$$\mathbb{P}(\{w_j, \beta_j, \theta\} | \{D_t\}) \propto \mathbb{P}(\{D_t\} | \{w_j, \beta_j, \theta\}) \mathbb{P}(\{w_j, \beta_j, \theta\})$$

$$\begin{aligned} \mathbb{P}(\{w_j, \beta_j, \theta\} | \{D_t\}) \propto \quad & \left[\prod_{t=1}^{T} \mathbb{P}(D_t | w_j, \beta_j, \theta)\right] \left[\prod_{j=2}^{T(t)} \mathbb{P}(w_j | w_{j-1})\right] \times \\ & \left[\prod_{j=2}^{T(t)} \mathbb{P}(\beta_j | \beta_{j-1})\right] \mathbb{P}(w_1) \mathbb{P}(\beta_1) \mathbb{P}(\theta) \end{aligned} \tag{3.4}$$

Second, I select the functional forms for the probability distributions of weights at $t+1$, conditional on weights at t. Specifically, weights w_{t+1} are distributed according to a product prior of the trapezoid and normal distribution. The normal distribution comes from IMF's COFER data, which holds the aggregate portfolio weights across many countries. The trapezoid distribution relates w_{t+1} to w_t, and it is calibrated to reflect some portfolio stickiness across time. (The trapezoid distribution obeys the bounds $[0, 1]$ on portfolio weights, unlike the normal distribution, and so it is very similar to a truncated normal distribution.) Moreover, β_{t+1} is distributed normally, centered around β_t, which suggests that response functions of central banks to local currency volatility are sticky across time. These assumptions of portfolio stickiness seem reasonable, as Lim (2007) shows that portfolio weights are relatively stable in the aggregate data.

Markov Chain Monte Carlo

Finally, I estimate the distribution of parameters in Equation (3.4) using the Metropolis-Hastings algorithm. This algorithm is applied to each country in isolation, to find each country's set of parameters.

The Metropolis-Hastings algorithm stochastically searches the parameter space, moving to regions with higher density and away from regions with lower density, using a "jumping distribution." While calibrating the jumping distribution can be difficult in practice, I use the adaptive MCMC approach given by Roberts and Rosenthal (2009), in which the distribution is calibrated automatically to yield reasonable search processes. Moreover, the exact choice of a jumping distribution does not affect the long-run convergence properties, as long as it searches the parameter space adequately.

I run the algorithm 50 million times per country, and discard the first 49 million as the "burn-in" period. For almost all countries, the parameter draws appear stationary and convergence seems reasonable.

3.4 Results

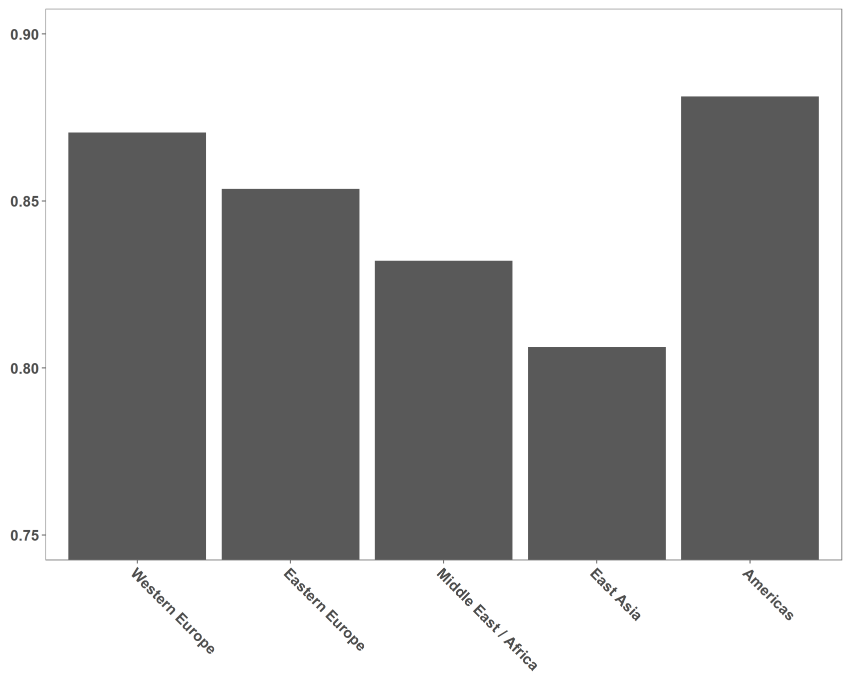
This section depicts the results from the simulation. First, I show the summary statistics that emerge from the dynamic linear model, which are economically interesting on their own. Dollar shares in foreign reserves are large, and this is relatively stable across countries and across time. Euro shares are smaller, and most other currencies have negligible shares. Second, I show that the dollar and euro shares of foreign reserves can be explained by the shares of external financial liabilities denominated in those currencies. This is suggestive of a model of financial stability, which I develop further in Section 3.5.

3.4.1 Summary Statistics

In this section, I first show the average dollar and euro shares across countries, taking the simple average across different groups of countries. I also focus on the dollar, euro, pound, and yen; although I also generate results for the Swiss franc, the Australian dollar, and the Canadian dollar.

Figure 3.1 shows the average dollar share by region, and these are uniformly high. Unsurprisingly, the shares are especially high in the Americas. More surprisingly, they are especially high in Europe; but this may be because holding euros for Eurozone countries is inefficient, leaving the dollar as the main reserve asset.

Figure 3.1: *Dollar Shares of Foreign Reserves*

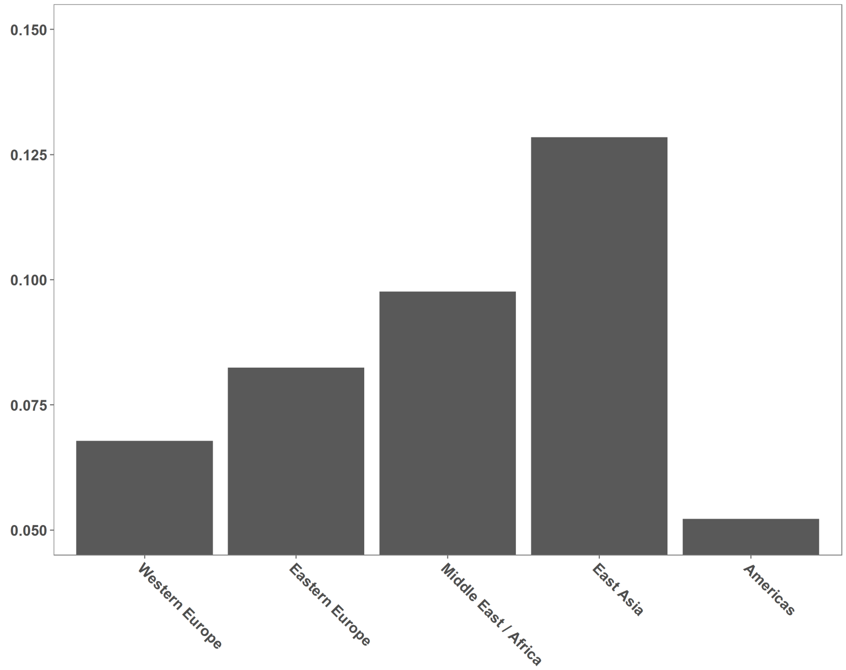


Notes: The figure depicts the average dollar share across the foreign currency reserves of countries in a given region, where dollar shares are averaged first across time within country and second across countries. Dollar shares are extremely high and somewhat heterogeneous, with countries in Europe and in the Americas having particularly high dollar shares.

Figure 3.2 shows the average euro share by region, and these are generally low, although the graph masks some large outliers. For instance, a handful of non-Eurozone eastern European countries and Middle East / North Africa countries (e.g. Romania and Morocco) have enormous euro shares.

Figure 3.3 shows the average shares for the pound and yen. These are uniformly low, but the contrast is informative. In the official COFER statistics, both have comparable shares — but in the

Figure 3.2: *Euro Shares of Foreign Reserves*



Notes: The figure depicts the average euro share across the foreign currency reserves of countries in a given region, where euro shares are averaged first across time within country and second across countries. Euro shares are low and somewhat heterogeneous, with countries in the Middle East, Africa, and Asia having particularly high euro shares.

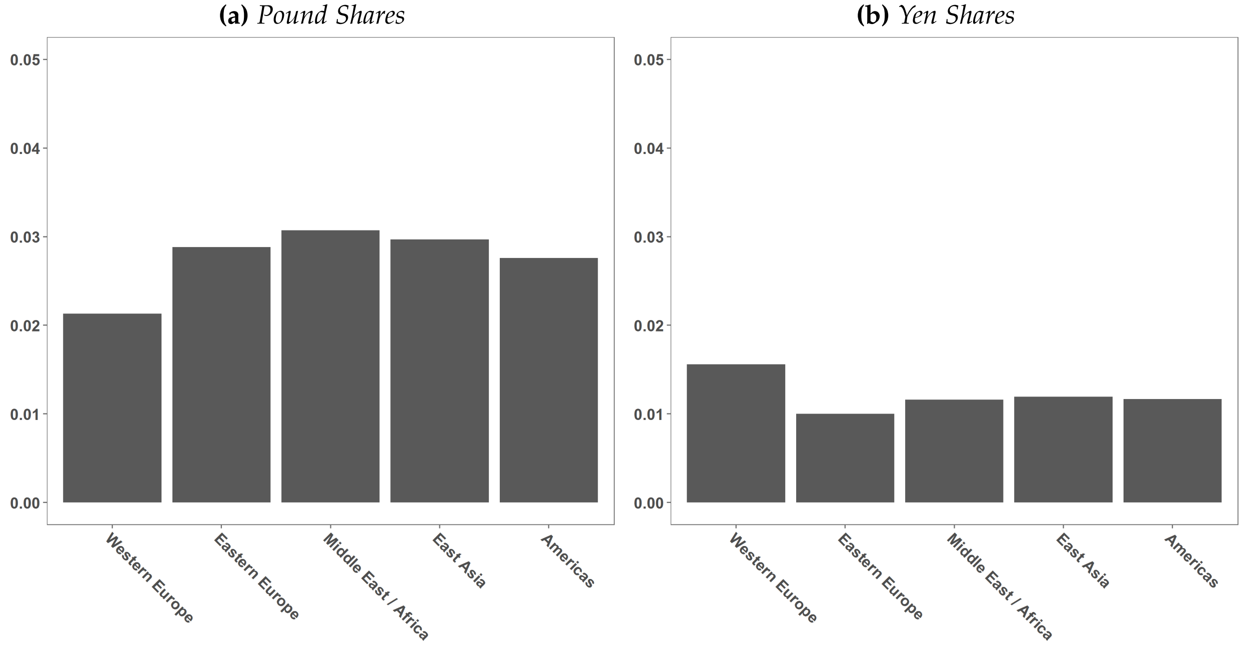
disaggregated data, the pound appears twice as dominant as the yen. This is suggestive evidence against reserves as a source of fiscal space during recessions, as the yen (not the pound) is the ultimate safe-haven asset.

Finally, Figure 3.4 shows changes in dollar and euro shares over time, benchmarked to their shares at the start of the sample (2004). The euro has gained some share at the dollar's expense, but the magnitudes are economically small, at approximately one percentage point. However, these small magnitudes may be a function of overly tight priors; and so in a future version, I will check that these findings are robust to looser priors.

Finally, I compare my estimates to the limited public data available, as reported by Wong (2007). Specifically, I focus on the sixteen countries (which excludes countries in the Eurozone and the United States) that report their currency compositions at some frequency, and compute the cross-sectional correlation between my estimated shares and the reported shares, for each of the dollar and euro. The results are qualitatively promising: both the dollar and euro shares are positively correlated at the 90% significance level.[4] However, the magnitudes are quite different,

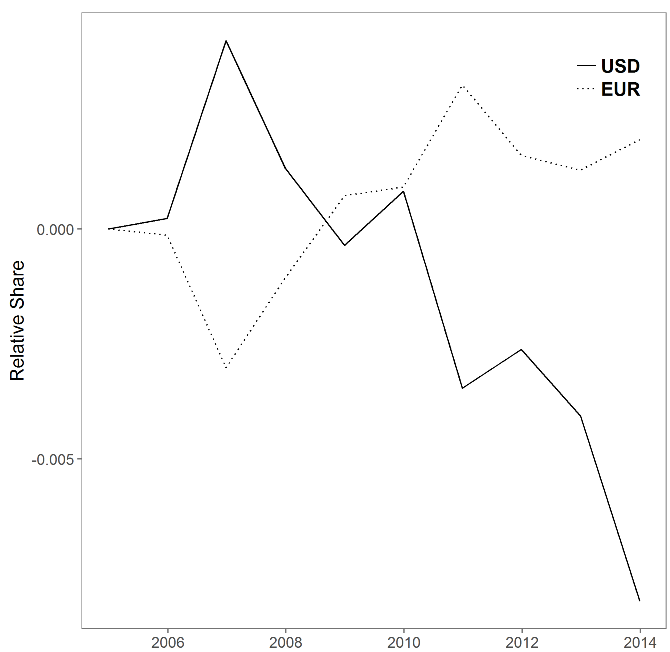
[4]I find $\rho \approx 0.15$ for both the dollar and euro estimates. I compute standard errors by bootstrap rather than

Figure 3.3: *Pound and Yen Shares of Foreign Reserves*



Notes: The figures depict the average pound and yen shares across the foreign currency reserves of countries in a given region, where shares are averaged first across time within country and second across countries. Pound shares and yen shares are both low and largely homogeneous, although pound shares are surprisingly higher than yen shares across the globe.

Figure 3.4: *Dollar and Euro Shares over Time*



Notes: The figure depicts the change, relative to 2004, in the average dollar and euro shares across global foreign currency reserves, where currency shares are averaged across countries for each year. While the euro share has grown at the expense of the dollar share, the growth is small, and the dollar's dominance in foreign reserve portfolios is economically stable.

for two reasons. First, as mentioned, my model priors are tight. Second, the countries that report publicly are the ones who have small reserve holdings, where this methodology is expected to work least well.

3.4.2 Bank Liabilities

In this section, I explain the heterogeneity in currency shares across countries using the heterogeneity in the currency shares of their financial liabilities. Specifically, I regress the dollar and euro shares in foreign reserves on the dollar and euro shares in banking liabilities, and show this is significant once swap lines — a form of direct lending between central banks that substitutes for official reserves — are taken into account.

Broadly, I regress the the dollar and euro shares of foreign reserves across countries on the dollar and euro shares of their banking systems' external liabilities, as calculated from the Bank of International Settlement's Locational Banking Statistics. The BIS data have some known limitations that I address. First, not all countries report their banking systems' positions. However, fifty countries do, and this includes most developed countries (e.g. the US, UK, Germany, and Japan) and many financial hubs (e.g. Luxembourg, the Cayman Islands, and Jersey) — and the number of reporters has grown steadily over time. Second, a country's coverage of its own financial system has grown steadily over time, making earlier reports less representative of the financial system at that time than later reports. Thus, I first infer countries' external liabilities as equal to the assets held on them by reporting countries, to bypass the limited coverage issue. Second, I focus on the latest year the data are available to estimate shares, as this has the widest coverage both in terms of number of reporting countries and number of reporting financial institutions within each country.

This regression on its own yields few significant results, and is not reported. In Table 3.1, however, I include an indicator for emerging countries, and the interactive term is highly significant. This means that, for emerging markets, higher fractions of banking liabilities in dollars and euros predicts higher dollar and euro shares for foreign reserves.

The distinction between developed and emerging countries may seem arbitrary, but one

parametrically, given the low sample sizes.

plausible and consistent explanation involves swap lines, which were extended to many developed countries and to few emerging countries. Specifically, swap lines are channels by which one central bank can temporarily lend its local currency directly to another central bank for liquidity management, and these plausibly crowd out foreign reserves as they function as temporary and “on-demand" foreign reserves. For instance, Aizenman *et al.* (2011) and Morelli *et al.* (2015) argue that swap lines empirically and theoretically, respectively, substitute for foreign reserves; and Allen and Moessner (2011) argues that foreign currency banking liabilities directly predict swap lines. (Moreover, Bordo *et al.* (2014) — who provide a comprehensive history on swap lines — note that William Poole, president of the St. Louis Federal Reserve Bank, objected on extending swap lines to central banks with large dollar reserves on the grounds of redundancy.) Swap lines were particularly popular during the financial crisis, when countries' financial systems suffered dollar shortages, and McGuire and von Peter (2009), Fleming and Klagge (2010), and Rose and Spiegel (2012) argue that swap lines alleviated these shortages both in the time series and in the cross-section.

As such, in Table 3.1, I include an indicator for receiving dollar and euro swap lines during and after the financial crisis. The interactive term remains significant, although with wider confidence levels. Of course, countries that have received swap lines historically may not predict countries that will receive swap lines going forward, and that may partially explain the smaller confidence intervals.

A second explanation directly addresses the divide between emerging and developed markets in the context of supranational entities like the International Monetary Fund or European Central Bank. Developed countries everywhere are more willing to turn to the IMF for historical reasons; and Eurozone countries are of course far more able to turn to the European Central Bank for direct assistance during financial crises. In other words, countries without swap lines or without access to supranational entities have to hold reserves explicitly; countries with swap lines or with access to supranational entities hold reserves implicitly.

Table 3.1: *Explaining Reserve Portfolios by Bank Liabilities*

	Specification *Dependent Variable: Reserve Share*			
	Dollar		*Euro*	
	(1)	(2)	(3)	(4)
Bank Share	-0.130	-0.085	-0.068	-0.016
	(0.083)	(0.078)	(0.048)	(0.060)
Emerging (Indicator)	-0.111^{**}		-0.063^{*}	
	(0.046)		(0.034)	
Bank Share x Emerging	0.243^{**}		0.219^{***}	
	(0.099)		(0.079)	
No Swap Lines (Indicator)		-0.097^{**}		-0.009
		(0.045)		(0.042)
Bank Share x No Swap Lines		0.183^{*}		0.145^{*}
		(0.094)		(0.085)
Constant	0.895^{***}	0.889^{***}	0.116^{***}	0.076^{*}
	(0.029)	(0.030)	(0.029)	(0.039)
Observations	77	77	77	77
R^2	0.085	0.065	0.097	0.075

Notes: The table regresses the dollar and euro shares of countries' foreign reserves on the dollar and euro shares of their banking systems' liabilities, across the set of seventy-seven countries. Columns (1) and (2) focus on dollars, and Columns (3) and (4) on euros. All specifications include a dummy that is interacted with the banking liabilities variable: for Columns (1) and (3), this dummy is an indicator for emerging countries; for Column (2), this dummy is an indicator for countries that did not receive dollar swap lines; and for Column (4), this dummy is an indicator for countries that did not receive euro swap lines. Significance is assessed at the 10% (*), 5% (**), and 1% (***) level. For all specifications, the interacted term is significant, suggesting that for emerging countries and for countries without swap lines, the currency shares of the financial systems' liabilities correlate with the currency shares' of their foreign reserve portfolios in the cross-section.

3.5 Model

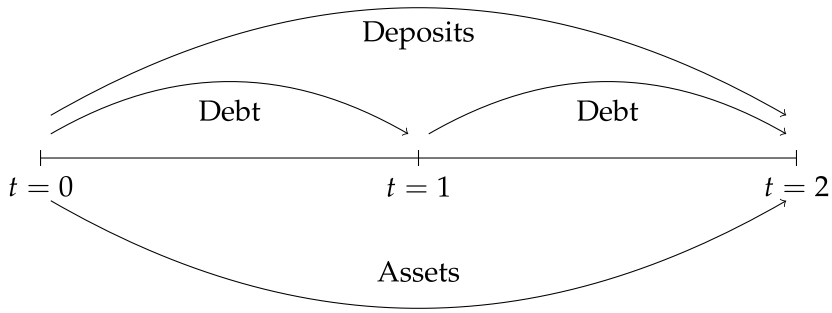
In this section, I develop a model to explain the empirical patterns — the prevalence of dollar reserves and the correlation between currency shares in reserves and financial liabilities. In the model, a country's central bank provides liquidity to its private banking system who have taken out excessive dollar debt, in order to protect domestic depositors. During liquidity crises, central banks need to lend dollars, but crucially it faces foreign exchange transaction costs during crises (in addition to general costs of inflation). Thus, printing local currency or holding other currencies and swapping those into dollars during crises is more costly than holding dollars directly *ex ante*. Although this model does not incorporate swap lines, this is equivalent to allowing the central bank to generate dollars during crises directly, removing the need for dollar reserves. Gopinath and Stein (2018) offer a similar model in which *ex ante* reserves mitigate the need to raise emergency funding.

Of course, nominal transaction costs in foreign exchange markets are tiny. For instance, I examine high-frequency exchange rate data for the dollar versus the New Zealand dollar (the least liquid of the major currencies) in August 2011 — a month that included a 7% drop in the S&P 500, a 20% fall in France's CAC 40, and a downgrade in the US credit rating — and find that the mean and median transaction costs are 2-3 basis points; and even when transaction costs rise, they dissipate within a minute or so. However, I conceptualize implicit foreign exchange transaction costs during crises to be large, particularly when central banks need to swap billions in short intervals. For instance, this might signal bad news to markets, and end up generating market panic and deeper liquidity crunches prematurely.

3.5.1 Private Banks

The agents in the model are banks, who generate liquidity mismatches with dollar debt. As in the standard Diamond and Dybvig (1983) model, banks borrow and lend across borders to fund long-term assets with short-term liabilities and deposits. Developed markets offer lower-yielding projects and cheaper funding compared to emerging markets, and this assumption — coupled with limited liability on behalf of banks — leads banks to under-hedge dollar debt. Central banks will be introduced later, as trying to protect domestic depositors who have invested in these banks.

The liquidity mismatch is standard. At $t = 0$, banks issue short-term debt (expiring at $t = 1$) and are given long-term deposits (expiring at $t = 2$) to fund long-term projects (paying off at $t = 2$). At $t = 1$, banks issue new short-term debt (expiring at $t = 2$) to pay off expiring short-term debt. At $t = 2$, banks cash in assets and settle all remaining liabilities.



Specifically, there are two countries (the US and Thailand), and each country c offers a continuum of projects to banks with gross riskless return $\sim U[1, R^c]$. If projects are liquidated early (at $t = 1$), they yield zero. As such, a bank that invests x^c in a country gets at $t = 2$:

$$\int_{R^c - x^c}^{R^c} r dr = R^c x^c - \frac{1}{2}(x^c)^2$$

Banks also raise financing from risk-neutral lenders in each country, to supplement exogenous deposits D. But financial markets have supply constraints, where each marginal dollar borrowed in each market has increasing costs. The costs are governed by a country-specific and time-varying α_t^c for each country c:

$$\int_0^X (1 + \alpha^c x) dx = X \left(1 + \frac{\alpha^c}{2} X\right)$$

Note that in this model, each country only offers projects and funding in its local currency, and so raising financing from US investors is equivalent to raising dollar funding.

Finally, exchange rates are assumed to be pegged to unity, and there are no impediments to moving funding across borders — in other words, foreign exchange markets are frictionless. As such, the bank's optimization problem is:

$$\max_{\{x\}_c, \{X_1\}_c} \sum_c \left(R^c x^c - \frac{(x^c)^2}{2} \right) - \sum_c X_1^c \left(1 + \frac{\alpha^c}{2} X_1^c \right) - D$$

subject to budget constraints at $t = 0$ (in which long-term investments are made) and $t = 1$ (in which short-term debt is refinanced), and various non-negativity constraints:

$$\sum_c X_0^c + D = \sum_c x^c$$

$$\sum_c X_1^c = \sum_c X_0^c \left(1 + \frac{\alpha^c}{2} X_0^c\right)$$

$$x^c, X_0^c, X_1^c \geq 0 \quad \forall\, c$$

Now, I introduce a second state of the world at $t = 1$: the *crisis* state. Several things happen during a crisis. Most importantly, foreign exchange markets develop convex frictions, paramterized by f^c for country c. Suppose a bank wants to move y dollars into a given country c; during a crisis, the bank will only receive the following:

$$\int_0^y (1 - f^c y) dy = y \left(1 - \frac{f^c}{2} y\right)$$

In addition, during a crisis, short-term borrowing costs shift and exchange rates temporarily deviate from unity. These are not important for the results qualitatively, but they both seem empirically valid and make the quantitative results more stark.

Banks have limited liability, and can default (which changes the cost of borrowing by risk-neutral lenders *ex ante*, denoted by κ). Finally, banks can save through s, although given the limited liability constraint, they will not do so in practice. Thus, the bank's optimization problem can be revised.

$$\max_{s,x,X,y} \mathbb{E} \left[\max \left\{ 0, \sum_c \left(R^c x^c - \frac{(x^c)^2}{2} \right) - \sum_c X_1^{c,s} \left(1 + \frac{\alpha_1^{c,s}}{2} X_1^c\right) - D \right\} \right]$$

subject to budget constraints at $t = 0$ (in which long-term investments are made), $t = 1$ (non-crisis state), and $t = 1$ (crisis state), as well as a constraint for transferring wealth in either state and various non-negativity constraints:

$$\sum_c X_0^c + D = \sum_c s^c + \sum_c x^c$$

$$s^c + X_1^c + y^c E^c = X_0^c \left(1 + \frac{\alpha_0^c}{2} X_0^c\right) \frac{1}{\kappa} \quad \forall c$$

$$s^c + X_1^c + y^c \left(1 - \frac{f^c}{2} y^c\right) E^c = X_0^c \left(1 + \frac{\alpha_0^c}{2} X_0^c\right) \frac{1}{\kappa} \quad \forall c$$

$$\sum_c y^c = 0$$

$$s^c, x^c, X_0^c, X_1^c \geq 0 \quad \forall\, c$$

To illustrate with a specific parameterization, consider a bank that has license to raise funds and invest in both countries (the US and Thailand). Compared to the US, Thailand has more profitable investments but less developed financial markets. If there is a crisis at $t = 1$ (with $p = 0.01$), exchange rate frictions appear. In this scenario, the costs of funding in the US and Thailand switch and there are no exchange rate changes. While I can get more dramatic results when the costs of funding both rise and the exchange rate fluctuates — which is empirically more consistent — this is done to illustrate the importance of foreign exchange frictions, as the overall funding menu offered to the bank is kept the same.

At $t = 0$, the bank has two choices. It can take a safe plan, in which it keeps debt low and remains solvent during the crisis; or it can take out a risky plan, in which it takes out high debt and defaults during the crisis. Table 3.2 depicts the options. In this example, the risky plan is more profitable, and so a risk-neutral bank will implement this.

Table 3.2: *Private Bank Optimization*

Name	State	Country	Variable	Safe Plan	Risky Plan
Debt	$t = 0$	US	X_0^1	1.60	1.74
		Thailand	X_0^2	0.08	0.07
Investment	$t = 0$	US	x^1	0.00	0.00
		Thailand	x^2	1.98	2.11
Debt	No Crisis ($t = 1$)	US	$X_1^{1,1}$	4.24	4.80
		Thailand	$X_1^{2,1}$	0.17	0.19
Inflows	No Crisis ($t = 1$)	US	$y_1^{1,1}$	-0.08	0.00
		Thailand	$y_1^{2,1}$	0.08	0.00
Debt	Crisis ($t = 1$)	US	$X_1^{1,2}$	4.06	N/A
		Thailand	$X_1^{2,2}$	0.54	N/A
Inflows	Crisis ($t = 1$)	US	$X_1^{1,2}$	0.20	N/A
		Thailand	$X_1^{2,2}$	-0.20	N/A
Expected Profit				69.79	73.09

Notes: The table shows the optimal scenarios under the assumption that a private bank wishes to remain solvent or default in a crisis. Both plans involve taking heavy US debt and investing in Thailand, although the plan under the default assumption involves more leverage. The plan that defaults in a crisis is more profitable, leading to an overly leveraged banking sector in the aggregate.

3.5.2 Central Banks

When private banks default, the central bank does not sit idle: it wishes to protect depositors, either by bailing out distressed banks or printing the deposits itself. Suppose the cheaper option is to bail out distressed banks. If banks' short-term liabilities are in foreign currencies, domestic currency is useless — and so the central bank can either incur foreign exchange transaction costs to gather foreign currency, or it can lend out of previously accumulated foreign currency reserves. In this model, central banks accumulate precautionary savings in foreign currencies before turmoil.

Specifically, if a bank is on the verge of defaulting in a crisis, the central bank can choose to bail out the bank by extending loans L^c at $t = 1$ in each currency c. These are not gifts, as banks must repay at $t = 2$. However, banks can borrow these loans without encountering the convex financing costs or foreign exchange costs. Naturally, central banks want to minimize the funds transferred, as long as banks pay off their short-term obligations, and so it does not lend wastefully. I assume that bailouts are unexpected for banks when banks are optimizing. As such, at $t = 1$ in the crisis state, the central bank solves the following optimization problem.

$$\min_{X_1, y, L} \quad \frac{1}{2} \sum_c \beta_c (L^c)^2$$

subject to a solvency constraint for the bank to which it lends at $t = 2$, a constraint to ensure that the bank can remain afloat at $t = 1$, and the usual constraints on internal transfers and non-negativity.

$$\sum_c \left(R^c x^c - \frac{(x^c)^2}{2} \right) \geq \sum_c X_1^c \left(1 + \frac{\alpha_1^c}{2} X_1^c \right) + D + \sum_c L^c$$

$$s^c + L^c + X_1^c + y^c \left(1 - \frac{f^c}{2} y^c \right) E^c = X_0^c \left(1 + \frac{\alpha_0^c}{2} X_0^c \right) \frac{1}{\kappa} \quad \forall c$$

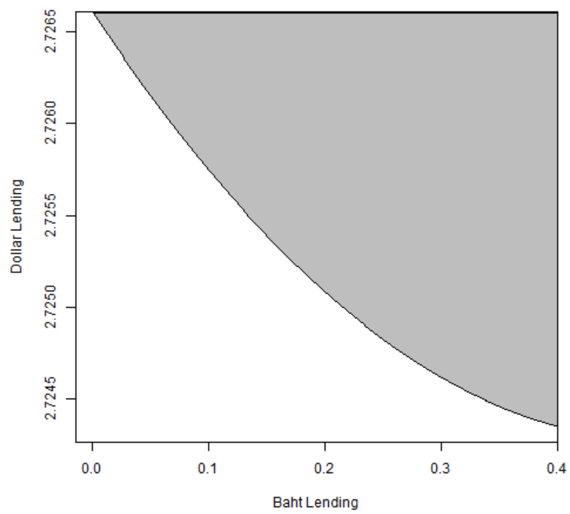
$$\sum_c y^c = 0$$

$$X_1^c \geq 0$$

This optimization yields a "frontier" of lending packages across the two currencies — dollars and Thai baht — depicted in Figure 3.5. A central bank picks a loan package from the line, as that reflects the minimum transfer needed to bail out the bank.

The next step is determining which loan package to pick (and whether a central bank even wishes to bail out the bank in the first place). The one power that it has is the printing press, and

Figure 3.5: *Central Bank Lending*



Notes: The figure depicts the lending frontier for a central bank that chooses to bail out the risky bank in Table 3.2 during a crisis, across a mixture of dollar and Thai baht lending. The line depicts efficient lending mixtures. A loan package from the upper-right portion of the graph is gratuitous, and a loan package from the lower-left portion is insufficient to bail out a bank.

it can print domestic currency — although at the cost of inflation π.

Now, consider a central bank that is determining whether to let a failing bank actually fail or not. If the private bank fails, the central bank has to print money to cover depositors. If the bank lends to it, it must hit the support frontier identified earlier, using existing reserves and freshly printed money. The central bank can also use foreign exchange markets, although it is subject to the same frictions during a crisis.

Thus, the central bank solves *ex ante*:

$$\min_{\pi} \mathbb{E}\left[\pi_0^2 + \pi_1^2\right]$$

subject to budget constraints at $t = 0$ and $t = 1$ across domestic and foreign currencies, and a bailout function $B(\cdot)$ that checks whether the lending frontier has been reached during a bailout, and returns zero if so.

$$\pi_0 = \sum_c s^c$$

$$\pi_1 + s^H + y^H \left(1 - \frac{f^H}{2} y^H\right) E^H = \sum_i L_i^H + \sum_j D_j$$

$$s^c + y^c \left(1 - \frac{f^c}{2} y^c\right) E^c = \sum_i L_i^c \quad \forall c \neq \text{Home}$$

$$B\left(L_i^H, L_i^c\right) = 0 \quad \forall i$$

For instance, consider the same example. In Table 3.3, central banks find it optimal to bail out their risky banks, versus bailing out depositors directly. Note, however, that the domicile of the bank matters greatly — while Thai and American banks perform the same strategies, their supervisors find it differentially difficult to mitigate liquidity shocks. The Federal Reserve, which can print dollars on demand, need not store reserves and can bear inflationary costs if a crisis emerges. The Bank of Thailand, which cannot, stores dollar reserves ahead of time to hedge the possibility of a crisis.

Table 3.3: *Central Bank Optimization*

Name	State	Variable	Federal Reserve	Bank of Thailand
Inflation	$t = 0$	π_0	0.03	2.62
	No Crisis ($t = 1$)	π_1^1	0.00	0.00
	Crisis ($t = 1$)	π_1^1	2.70	0.66
Dollar Savings	$t = 0$	s^1	0.03	2.62
Baht Savings	$t = 0$	s^2	0.00	0.00
Dollar Lending	Crisis ($t = 1$)	L^1	2.73	2.72
Baht Lending	Crisis ($t = 1$)	L^2	0.00	0.36
Expected Loss			-0.07	-6.89

Notes: The table shows the optimal plans for each central bank if the defaulting private bank is their responsibility. Both central banks choose to bail out the defaulting bank, although the Federal Reserve can print dollars as needed and thus stores no reserves. The Bank of Thailand cannot, and thus holds large dollar reserves *ex ante*.

This model thus sheds light on why foreign central banks hold large dollar reserves when their financial systems are heavily dollarized. Failing to do so would be costly for the central bank, which would either have to incur large transaction costs to generate dollars for its financial system or would have to let its financial system collapse.

3.6 Conclusion

This paper uses a Bayesian dynamic linear model to document a novel empirical fact: dollars are widely prevalent in the foreign reserve portfolios of most countries. This pattern can be explained, both empirically and theoretically, from the dollarization of countries' financial liabilities. Pragmatic central banks that wish to bypass foreign exchange markets during liquidity crises (and are either unable or unwilling to turn to the IMF or the Fed directly) must hold dollar reserves prior to crises. Alternate explanations, such as ones in which foreign reserves provide fiscal space to governments, seem less consistent with the fact.

Policymakers have recently focused on the global ramifications of countries' foreign reserves, as Steiner (2014) explains. For instance, the East Asian savings glut — largely by East Asian governments — is blamed for fueling the American mortgage bubble during the 2000s, and the geopolitical consequences of foreign countries holding large quantities of US debt has drawn attention in recent years. This paper both provides new data and plausible explanations to such debates.

In addition, this paper shows another arena in which dollar hegemony reigns. Even as the US's presence in goods markets shrinks, its presence in financial markets remains enduringly dominant — whether in bank loans, trade invoicing patterns, monetary policy, or now sovereign reserves.

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Appendix A

Appendix to Chapter 1

A.1 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.1.1 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.1.2 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.1.3 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.[1] 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[2] 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.

A.1.4 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

[1] The ECB has changed this as of the middle of 2016, but this only affects a few announcements.

[2] Technically, the opening paragraph bids Happy New Year; this is the first paragraph with content.

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 Tokohu 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.

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.1.5 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.1.6 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.1.7 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.1.8 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 Swiss franc-euro exchange rate depicts this, in Figure A.1.

Figure A.1: *Swiss Franc / Euro Exchange Rate*

Notes: The figure depicts the Swiss franc-euro exchange rate from 2010 - 2016. The noteworthy event in this time series is that cap imposed by the Swiss National Bank from September 2011 until January 2015, which prevented the franc from appreciating to below 1.20 francs per euro. Volatility in the exchange rate was substantially lower in this period than outside this period, but it did not drop to zero. In other words, the cap was not a peg.

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.1.9 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

Thursdays, with only 2% of announcements being made on Mondays or Wednesdays. Until late 2016, regularly scheduled announcements were made twelve times annually; they are now made eight times 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.[3] 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.

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.

[3] The Bank of England has changed this as of August 2015, but this only affects a few announcements.

A.1.10 United States

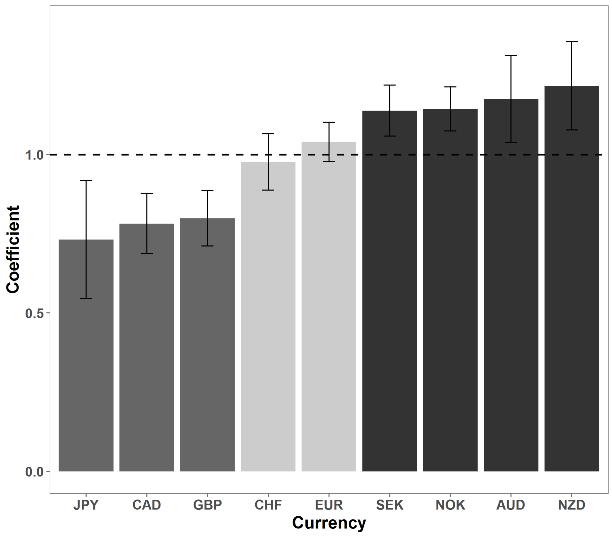
American rate announcements (known as FOMC announcements) target the Federal Funds Rate and are made by the Federal Reserve. Until early 2013, announcements were made at 2:15 PM EST; and starting in January 2013, announcements are made at 2:00 PM EST. Some announcements from 2011 - 2012 are made at 12:30 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.

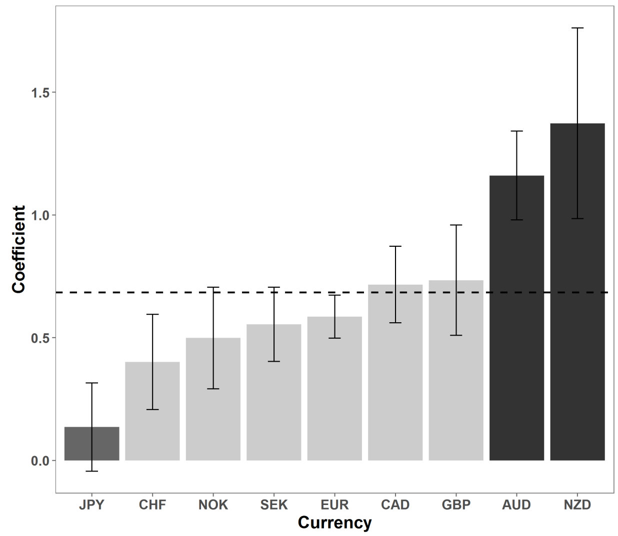
Nakamura and Steinsson (2017) have raised concerns that FOMC announcements signal the Fed's private information about fundamentals, even though announcements do not officially release new information about fundamentals. If true, monetary announcements should look qualitatively like official announcements 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 A.2. 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. FOMC announcements are not just announcements about fundamentals.

Figure A.2: *Market Reactions to Various US Shocks*

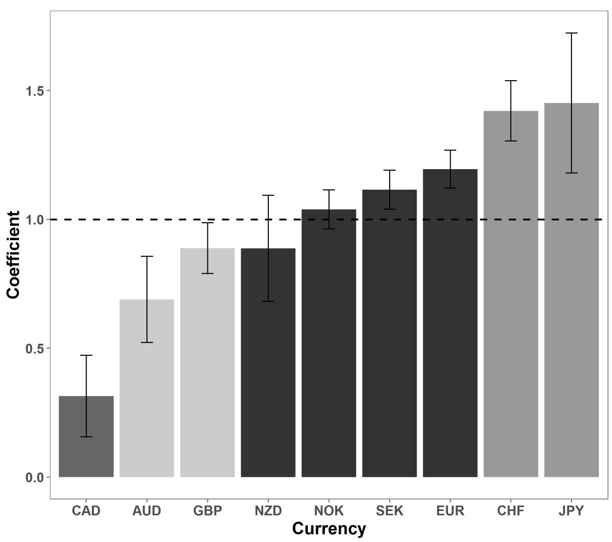
(a) *Currency Responses to Monetary Shocks*



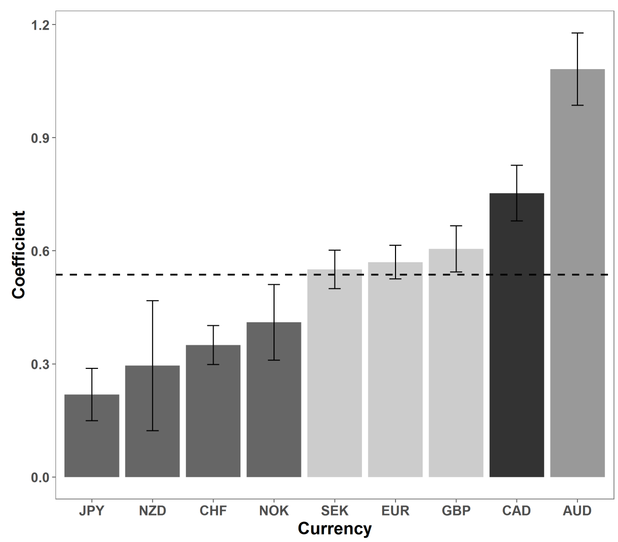
(b) *Bond Responses to Monetary Shocks*



(c) *Currency Responses to Fundamentals Shocks*



(d) *Bond Responses to Fundamentals Shocks*



Notes: The figures depict the reactions of currency and bond markets to monetary and fundamentals announcements in the US. The two top figures are reactions to Fed announcements (as in the main paper), 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 against 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. The responses in asset markets to FOMC announcements differ from responses to BLS announcements, suggesting that FOMC announcements do not deliver news about fundamentals and lending credence to their interpretation as monetary shocks.

A.2 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 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.

A.2.1 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 (α, β, m) in the following specification:

$$r_t(i_t) = X_t^{\alpha}(i_t)\alpha + X_t^{\beta}(i_t)\beta m_t + \epsilon_t(i_t) \quad \forall\ t \tag{A.1}$$

Each component of Equation (A.1) 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 $X_t^{\alpha}(i_t)$, $X_t^{\beta}(i_t)$, 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.
- α is a $C_\alpha \times 1$ vector of constants.
- β is a $C_\beta \times C_m$ matrix of coefficients (i.e. factor loadings).
- X_t^{α} and X_t^{β} 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 $m_t \sim N(\vec{0}, \Omega)$.

- ϵ_t is a $C_r \times 1$ vector of residuals at time t with underlying distribution $\epsilon_t \sim N(\vec{0}, \Sigma)$. Σ 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. Note that because ϵ_t is multivariate normal, any subset $\epsilon_t(i_t)$ is also multivariate normal.

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, $X_t^\alpha = X_t^\beta$.

$$\begin{bmatrix} r_t^{€/\$} \\ r_t^{¥/\$} \\ r_t^{£/\$} \end{bmatrix} \left(\begin{bmatrix} 1 \\ 1 \\ 0 \end{bmatrix} \right) = \begin{bmatrix} 1 & 0 & 0 & y_t^{€} \\ 0 & 1 & 0 & y_t^{¥} \\ 0 & 0 & 1 & y_t^{£} \end{bmatrix} \left(\begin{bmatrix} 1 \\ 1 \\ 0 \end{bmatrix} \right) \times \begin{bmatrix} \alpha^{€} \\ \alpha^{¥} \\ \alpha^{£} \\ \alpha^{y} \end{bmatrix} + \begin{bmatrix} 1 & 0 & 0 & y_t^{€} \\ 0 & 1 & 0 & y_t^{¥} \\ 0 & 0 & 1 & y_t^{£} \end{bmatrix} \left(\begin{bmatrix} 1 \\ 1 \\ 0 \end{bmatrix} \right) \times \begin{bmatrix} \beta^{€} \\ \beta^{¥} \\ \beta^{£} \\ \beta^{y} \end{bmatrix} \times m_t + \begin{bmatrix} \epsilon_t^{€/\$} \\ \epsilon_t^{¥/\$} \\ \epsilon_t^{£/\$} \end{bmatrix} \left(\begin{bmatrix} 1 \\ 1 \\ 0 \end{bmatrix} \right)$$

$$\begin{bmatrix} r_t^{€/\$} \\ r_t^{¥/\$} \end{bmatrix} = \begin{bmatrix} 1 & 0 & 0 & y_t^{€} \\ 0 & 1 & 0 & y_t^{¥} \end{bmatrix} \times \begin{bmatrix} \alpha^{€} \\ \alpha^{¥} \\ \alpha^{£} \\ \alpha^{y} \end{bmatrix} + \begin{bmatrix} 1 & 0 & 0 & y_t^{€} \\ 0 & 1 & 0 & y_t^{¥} \end{bmatrix} \times \begin{bmatrix} \beta^{€} \\ \beta^{¥} \\ \beta^{£} \\ \beta^{y} \end{bmatrix} \times m_t + \begin{bmatrix} \epsilon_t^{€/\$} \\ \epsilon_t^{¥/\$} \end{bmatrix}$$

$$\begin{aligned} r_t^{€/\$} &= \alpha^{€} + \alpha^y y_t^{€} + \left(\beta^{€} + \beta^y y_t^{€}\right) m_t + \epsilon_t^{€/\$} \\ r_t^{¥/\$} &= \alpha^{¥} + \alpha^y y_t^{¥} + \left(\beta^{¥} + \beta^y y_t^{¥}\right) m_t + \epsilon_t^{¥/\$} \end{aligned}$$

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_t^α and X_t^β 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.

$$\begin{bmatrix} r_t^{€/\$} \\ r_t^{¥/\$,1} \\ r_t^{¥/\$,2} \end{bmatrix} = \begin{bmatrix} 1 & 0 \\ 0 & 1 \\ 0 & 1 \end{bmatrix} \times \begin{bmatrix} \alpha^{€} \\ \alpha^{¥} \end{bmatrix} + \begin{bmatrix} 1 & 0 \\ 0 & 1 \\ 0 & 1 \end{bmatrix} \times \begin{bmatrix} \beta^{€} \\ \beta^{¥} \end{bmatrix} \times m_t + \begin{bmatrix} \epsilon_t^{€/\$} \\ \epsilon_t^{¥/\$} \\ \epsilon_t^{¥/\$,2} \end{bmatrix}$$

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).

A.2.2 Expectation Maximization Algorithm

To identify the parameters (α, β, m) in the general setting of Equation (A.1), 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 β and m. The Expectation Maximization algorithm, by contrast, updates each of (α, β, m) sequentially until the algorithm converges on the optimal estimates.

$$\max_{\alpha,\beta,\{m_t\}_{t=1}^T} -\frac{1}{2T}\sum_{t=1}^{T}\left[\left(r_t(i_t) - X_t^{\alpha}(i_t)\alpha - X_t^{\beta}(i_t)\beta m_t\right)^T \Sigma(i_t,i_t)^{-1}\left(r_t(i_t) - X_t^{\alpha}(i_t)\alpha - X_t^{\beta}(i_t)\beta m_t\right)\right]$$

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. In general, the Expectation Maximization algorithm alternately takes the expectation of log-likelihood function with respect to the latent monetary shocks m_t and then maximizes the expression with respect to the parameters (α, β). In this context, the marginal likelihood of continuous latent factors m_t is intractable, and so instead I implement a commonly-used variation of the approach for such latent factor models.

Specifically, I first 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 $N(\mu_t, V_t)$ and the prior $N(0, I)$. I impose a scaling assumption that the ex ante variance of monetary shocks is set at I, which is standard in factor models.

$$\max_{\alpha,\beta,\{\mu_t,V_t\}_{t=1}^T} -\frac{1}{T}\sum_{t=1}^T\left(\frac{1}{2}\mathbb{E}_{m|\mu,V}\left[\left(r_t(i_t)-X_t^\alpha(i_t)\alpha-X_t^\beta(i_t)\beta m_t\right)^T\Sigma(i_t,i_t)^{-1}\right.\right.$$
$$\left.\left.\left(r_t(i_t)-X_t^\alpha(i_t)\alpha-X_t^\beta(i_t)\beta m_t\right)\right]+\frac{1}{2}\log|V_t|-\frac{1}{2}\mathrm{Tr}(V_t)-\frac{1}{2}\mu_t^T\mu_t\right)$$

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

I then maximize the expression alternately with respect to parameters (μ_t, V_t) and (α, β), which maps to the original two steps of taking the expectation and maximizing. In other words, I take first order conditions with respect to the underlying parameters, and update them iteratively until the algorithm converges. The rearranged conditions for (V_t, μ_t, α) are presented first.

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

Rearranging the first order condition for β is trickier, as β is not easily isolated.

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

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 $\left(\mu_t\mu_t^T+V_t\right)$ is a scalar.

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

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 β 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 β can be isolated.

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

Standard errors for α and β are computed by bootstrap, sampling vectors of asset returns at time $t = 1, ..., T$ with replacement. There are no analytic expressions for standard errors, due to missing data concerns. When conducting robustness checks with the shocks themselves, I set the shocks m_t to be their MAP estimates μ_t, i.e. the means of their posterior distributions.

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.

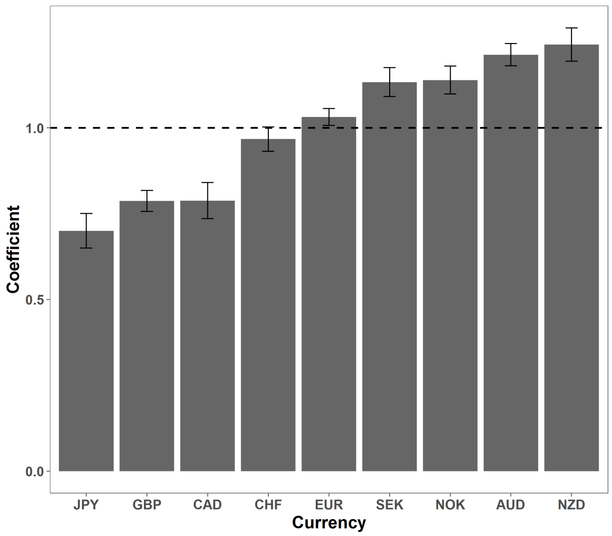
In addition, I check that this solution technique is robust by re-solving the three core specifications (with currencies, bonds, and cross-border portfolios) using Markov Chain Monte Carlo to optimize the log-likelihood expression directly. Figure A.3 shows that the results are virtually identical to those generated by the EM algorithm, in the main paper. However, MCMC is a computationally cumbersome solution technique in general. By contrast, the EM algorithm is much quicker, which is especially useful when finding the optimal lower-dimensional structure of the model.

A.2.3 Lower-Dimensional Mapping

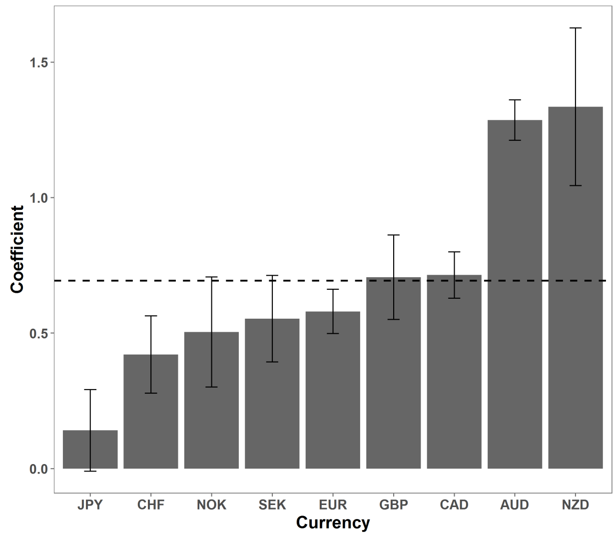
This paper routinely maps the factor model to the lowest-dimensional structure possible, in which the assets of partner countries have similar reactions to monetary shocks, and thus share coefficients (e.g. Norwegian and Swedish assets often react similarly). An example lower-

Figure A.3: *Market Reactions to US Monetary Shocks (MCMC Algorithm)*

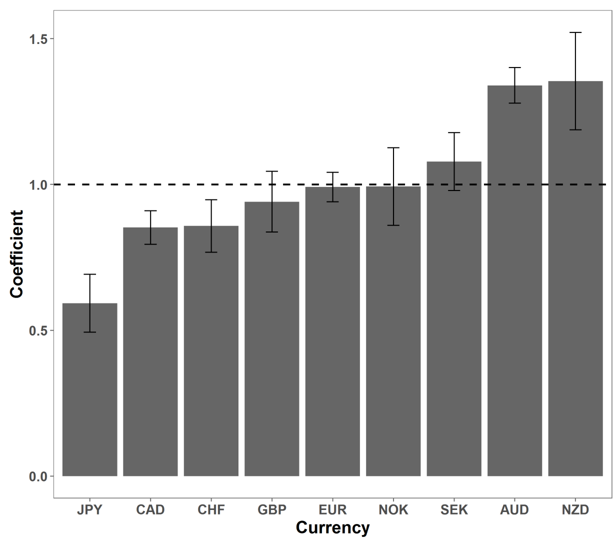
(a) *Currencies*



(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, but the parameters are estimated using MCMC rather than variational methods. 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 against the mean reaction across all currencies, bonds, or portfolios. These results mirror Figures 1.4 and 1.5, indicating that both this method and variational methods yield the same parameter estimates. However, I largely utilize variational approaches throughout this paper, as those are much faster than MCMC approaches; and this is importantly when finding the optimal lower-dimensional structure.

dimensional structure is below, in which the yen and pound share a coefficient.

$$\begin{bmatrix} r_t^{€/\$} \\ r_t^{¥/\$} \\ r_t^{£/\$} \end{bmatrix} = \begin{bmatrix} 1 & 0 \\ 0 & 1 \\ 0 & 1 \end{bmatrix} \times \begin{bmatrix} \alpha^{€} \\ \alpha^{¥,£} \end{bmatrix} + \begin{bmatrix} 1 & 0 \\ 0 & 1 \\ 0 & 1 \end{bmatrix} \times \begin{bmatrix} \beta^{€} \\ \beta^{¥,£} \end{bmatrix} \times m_t + \begin{bmatrix} \epsilon_t^{€/\$} \\ \epsilon_t^{¥/\$} \\ \epsilon_t^{£/\$} \end{bmatrix}$$

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 (A.2). Third, I find the best model fit using the (extended) Bayesian Information Criterion, which trades off the model log likelihood $\mathcal{L}(\theta)$ 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 α and β but also the variational posterior parameters for each monetary shock, is a prime candidate for such an approach.

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

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 $k = 1, ..., 9$. 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:

$$\begin{bmatrix} r_{c,t}^{€/\$} \\ r_{c,t}^{¥/\$} \\ r_{c,t}^{£/\$} \\ r_{b,t}^{€} \\ r_{b,t}^{¥} \\ r_{b,t}^{£} \end{bmatrix} = \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \\ 0 & 0 & 0 & 1 \end{bmatrix} \times \begin{bmatrix} \alpha_c^{€} \\ \alpha_c^{¥,£} \\ \alpha_b^{€} \\ \alpha_b^{¥,£} \end{bmatrix} + \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & 1 \\ 0 & 0 & 0 & 1 \end{bmatrix} \times \begin{bmatrix} \beta_c^{€} \\ \beta_c^{¥,£} \\ \beta_b^{€} \\ \beta_b^{¥,£} \end{bmatrix} \times m_t + \begin{bmatrix} \epsilon_{c,t}^{€/\$} \\ \epsilon_{c,t}^{¥/\$} \\ \epsilon_{c,t}^{£/\$} \\ \epsilon_{b,t}^{€} \\ \epsilon_{b,t}^{¥} \\ \epsilon_{b,t}^{£} \end{bmatrix}$$

A.2.4 Identification by Heteroskedasticity

An alternate approach to identify the parameter β for Equation (A.1) is to use Identification by Heteroskedasticity, which uses the generalized method of moments (GMM). While this approach gives equivalent results to the Expectation Maximization algorithm when it converges, it has poor convergence properties in my setting due to the severe non-linearities and high dimensionality of my specifications. As such, it is not utilized except as a robustness check. This section discusses both the implementation, augmented to handle missing data and to incorporate a lower-dimensional model structure, and the practical issues.

To understand the principle behind Identification by Heteroskedasticity in this context, I first present an extremely simplified version of Equation (A.1) to build intuition. In this example, asset returns are demeaned, monetary shocks m_t are univariate, there are 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:

$$\mathbb{E} r_t r_t^T = \beta\beta^T \mathbb{E} m_t^2 + \Sigma$$

$$\mathbb{E} \tilde{r}_t \tilde{r}_t^T = \Sigma$$

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

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

I now present the more general framework and the exact operational steps to fit β via GMM. This framework allows for missing data, shared coefficients between series, and non-trivial covariate matrices X^β; 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.

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

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

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

There are $\frac{1}{2} \times C_r \times (C_r + 1)$ unique equations to estimate C_β parameters. This is an overdetermined 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:

$$g_{t,(j,k)}(\beta) = \begin{cases} \left(\sum_{t \in E,(j,k) \in i_t} 1\right)^{-1} r_t(j) r_t(k) - X^\beta(j) \beta \beta^T X^\beta(k)^T \text{ if } t \in E, (j,k) \in i_t \\ -\left(\sum_{t \in N,(j,k) \in i_t} 1\right)^{-1} \tilde{r}_t(j) \tilde{r}_t(k) - X^\beta(j) \beta \beta^T X^\beta(k)^T \text{ if } t \in N, (j,k) \in i_t \\ 0 \text{ otherwise} \end{cases}$$

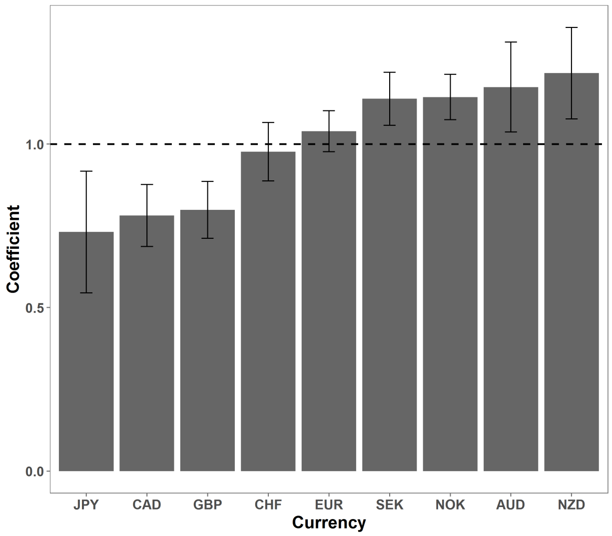
The sample moment conditions are aggregated across time and stacked via the usual definitions; and parameter estimates are solved through the standard framework, which estimate deviations from zero under some weighting matrix W:

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

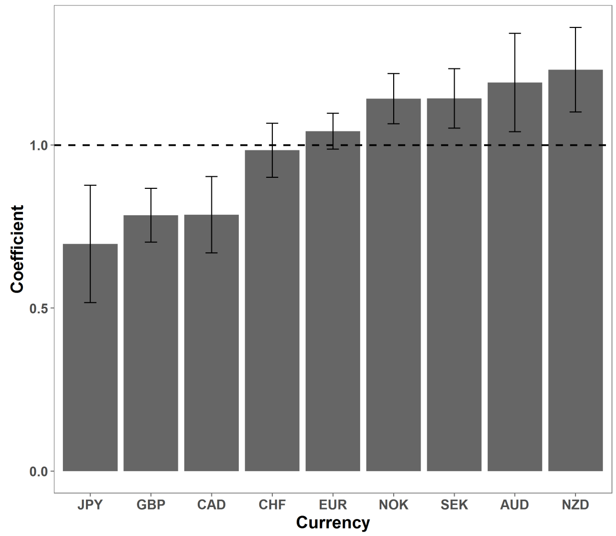
There are large computational issues that make this methodology ill-suited for the paper. In general, this is a very high-dimensional and a very non-linear problem. With respect to the former concern of dimensionality: for my benchmark specification with nine currencies, I have 45 moments; and for my 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. These results are presented in Figure A.4. In my specifications with bonds (100+ moments) or with the optimal weighting matrix, it fails to converge.

Figure A.4: *Market Reactions to Monetary Shocks (EM vs GMM)*

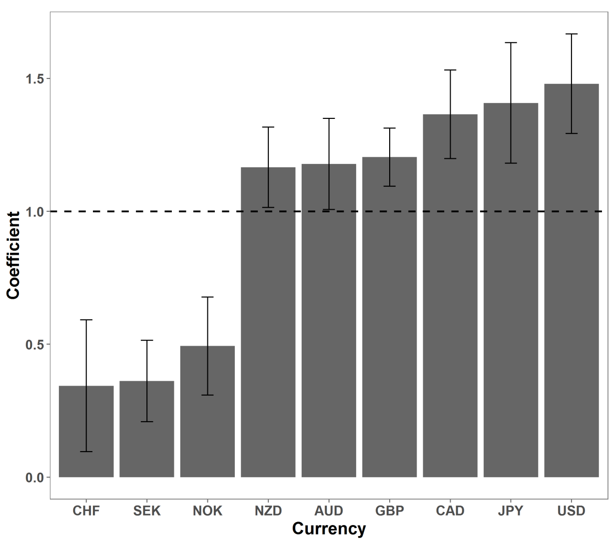
(a) *Currency Responses to US Shocks (EM)*



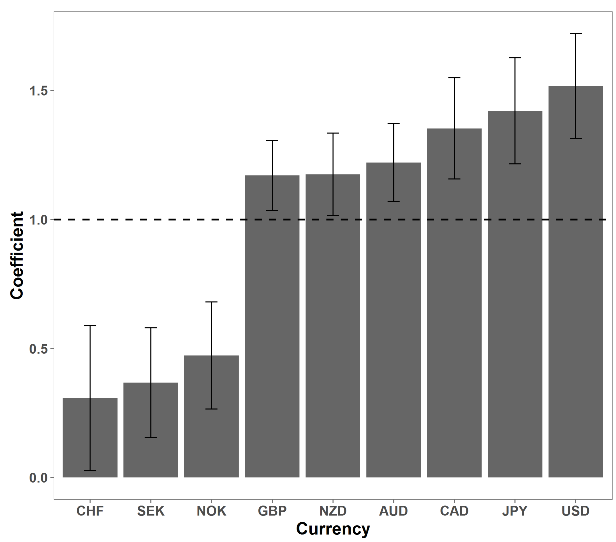
(b) *Currency Responses to US Shocks (GMM)*



(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 reference currency when it appreciates by 1% on average following announcements by the Fed (ECB). Standard error bars in all pictures are computed against the mean reaction across all currencies. The two approaches yield virtually identical estimates for both sets of data.

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.

A.2.5 Fixed Effects

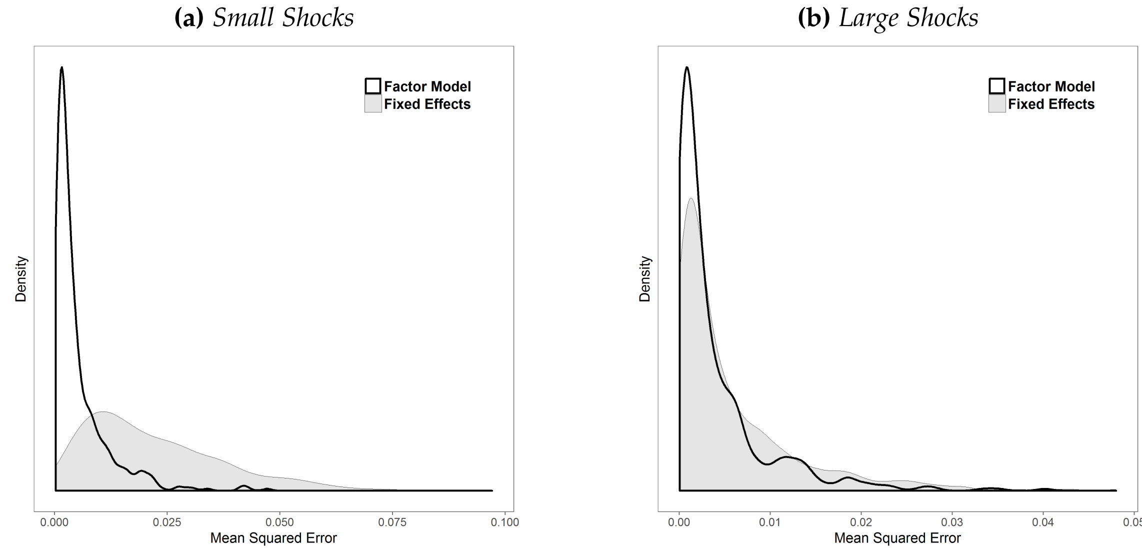
An alternate approach to identify the parameter β for Equation (A.1) would be to use fixed effects for coefficients, and ignore variation in monetary shocks over time. This is feasible and computationally quick, but it lacks statistical power by ignoring variation in monetary shocks.

Again, consider a simplified version of Equation (A.1) in which asset returns are demeaned (and so α can be ignored), monetary shocks m_t are univariate, there are no missing data, there is one coefficient per asset ($C_r = C_\beta$), and $X_t^\beta = I_{C_r}$. Practically, since monetary shocks have zero mean, this approach must regress the *absolute value* of asset returns on a set of currency fixed effects. As before, the variance-covariance matrix is parameterized with information from non-announcement windows. Assets that systematically respond more to monetary shocks will have larger absolute movements, and therefore will have larger coefficients.

$$\left| \begin{bmatrix} r_t^{€/\$} \\ r_t^{£/\$} \\ r_t^{¥/\$} \end{bmatrix} \right| = \begin{bmatrix} \beta^{€/\$} \\ \beta^{£/\$} \\ \beta^{¥/\$} \end{bmatrix} + \begin{bmatrix} \epsilon_t^{€/\$} \\ \epsilon_t^{£/\$} \\ \epsilon_t^{¥/\$} \end{bmatrix}$$

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 has far fewer moments and suffers from weak statistical power. This is easy to show by simulation. I simulate asset returns driven by monetary shocks and by other noise, and compare the mean squared error for the coefficients estimated under the Expectation Maximization algorithm to those estimated under the Fixed Effects approach. The Expectation Maximization algorithm always outperforms, particularly when the shocks are small relative to the background noise. The mean squared error density plots are presented in Figure A.5 for relatively small and relatively large shocks.

Figure A.5: *Fixed Effects Estimator*



Notes: The figure plots the distribution of mean squared errors for estimating the parameters, for two approaches on simulated data: the Expectation Maximization algorithm and the Fixed Effects estimator. The simulated data either embed “small" shocks (where shocks are twice as large as the background noise) or “large shocks" (four times as large). The Expectation Maximization algorithm always outperforms the Fixed Effects estimator. For large shocks, the performance gap is small, but the performance gap is very wide for small shocks, justifying the Expectation Maximization algorithm as the paper’s preferred approach.

A.2.6 Observed Shocks

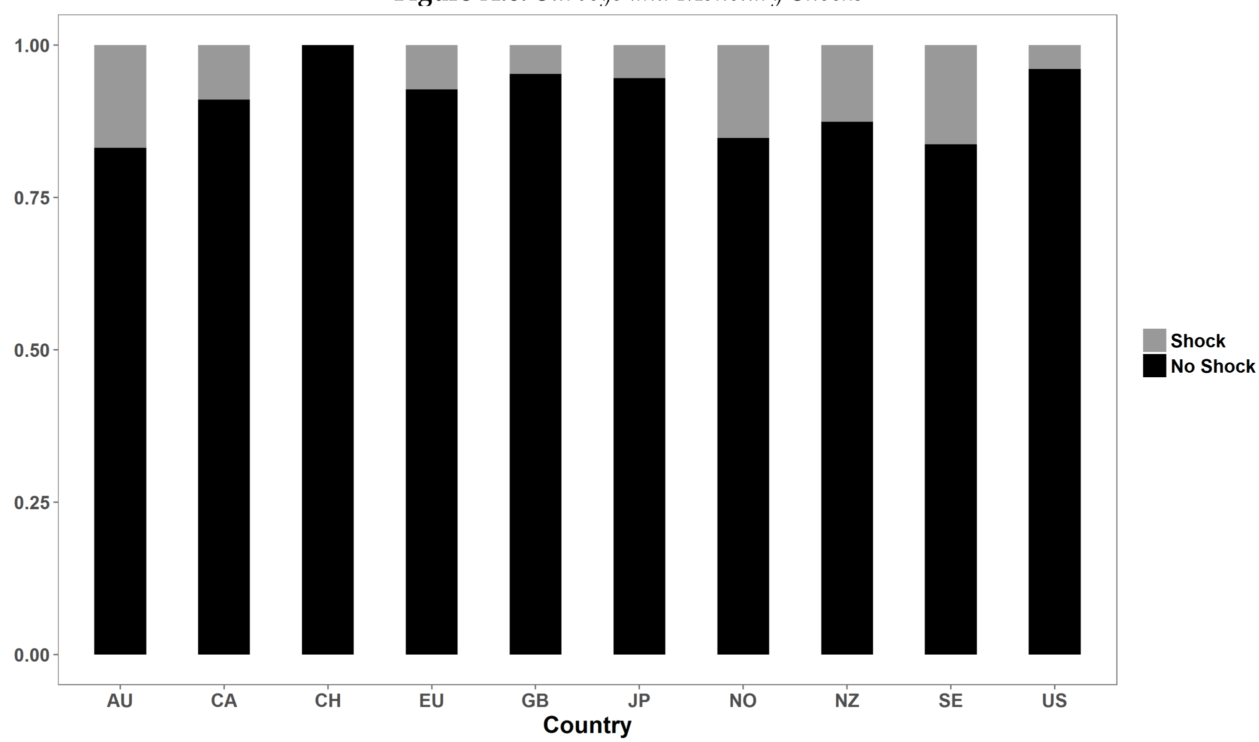
An alternate approach is to use observed measures of monetary shocks, rather than estimating or inferring them. There are broadly two classes of observed shocks. The first uses traditional measures of monetary shocks, such as the difference between actual policy and surveyed expectations, movements in yields, etc. Most of these measures are inappropriate for high-frequency usage. The second sets the monetary shock to be the average currency or bond return over that window. This introduces a bias in the standard errors. Regardless, my core results are qualitatively unchanged with most of these measures, and my latent shocks correlate well with these measured shocks too.

Traditional Measures

The first class of observed shocks are traditional measures of monetary shocks. There are five popular ones: movements in short-term rates (e.g. Fed Funds futures) around announcements, differences between actual policy and surveyed expectations, shocks constructed through the narrative method (i.e. Romer and Romer shocks), policy deviations from the Taylor Rule, and movements in medium-term rates (e.g. the two-year Treasury) around announcements.

The first four are immediately problematic. 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, Figure A.6 show that surveys correctly anticipated the monetary announcement for each of the ten central banks over 80% of the time, which means that a measure of shocks constructed from survey data would only identify shocks over 20% of my sample. Moreover, the Romer and Romer (2004) and Taylor Rule shocks are sensitive to the methodology and model used. For instance, central banks have strongly deviated from traditional Taylor Rule forecasts in the past decade, making this approach unreliable.

Figure A.6: *Surveys and Monetary Shocks*



Notes: The figure plots the percentage of time that survey forecasters correctly versus incorrectly anticipated the announced policy rate, for announcements by ten central banks from 2001 - 2016. Surveyed expectations come from Bloomberg, and they are constructed as the median forecast made by analysts at most major Wall Street broker-dealers for each announcement. Forecasters correctly anticipated 80+% of announcements, making this a statistically weak measure of shocks.

While I avoid these four measures of monetary shocks, I still check that they are correlated with my estimated latent shocks. In Table A.1, I correlate my shocks against those constructed from short-term rates and surveys. In the case of the United States, I also correlate my shocks against those constructed via Fed Funds futures data, Nakamura and Steinsson (2017), and Romer and Romer (2004). The correlations are almost all positive and significant.

Table A.1: *Comparing Monetary Shocks*

Country	Survey	01M	01Y	FFR	N-S	R-R
Australia	0.47^{**}	0.32^{**}	0.56^{**}			
Canada	0.33^{**}	0.07	0.50^{**}			
Switzerland		0.17	0.31^{*}			
Euro	0.19^{*}	-0.06	0.38^{**}			
United Kingdom	0.26^{**}	0.04	0.29^{**}			
Japan	-0.03	-0.04	0.22^{**}			
Norway	0.44^{**}	0.12	0.55^{**}			
New Zealand	0.29^{**}	0.27^{**}	0.73^{**}			
Sweden	0.43^{**}	-0.02	0.46^{**}			
United States	0.22^{*}	0.37^{**}	0.53^{**}	0.26^{*}	0.65^{**}	0.07

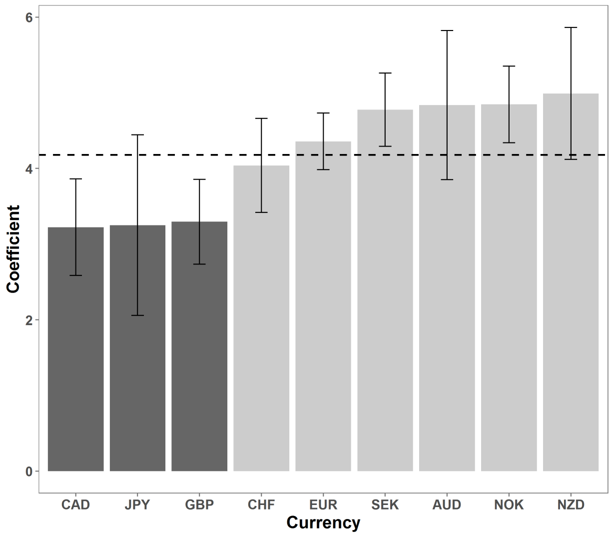
This table correlates the paper's latent monetary shocks against common observed measures of monetary shocks, using the Spearman correlation. All countries have shocks constructed from one-month and one-year yields; and all countries except Switzerland have shocks constructed from survey data (in Switzerland, all announcements were correctly anticipated so there are no surprises). For the United States, I also utilize Fed Funds futures shocks, Nakamura and Steinsson (2017) shocks, and Romer and Romer (2004) shocks (updated by Coibion *et al.* (2017)). Significance is assessed at the 1% (**) and 5% (*) level. Broadly, my latent shock correlates with most observed measures of shocks.

The most promising measure of observed shocks are medium-term yields (e.g. two-year rates), which solves obvious maturity or methodology concerns. However, there are still subtler reasons to prefer inferred shocks to these measured shocks. First, medium-term rates may still be insufficient to capture the entire path of shocks along the entire yield curve, as Boyarchenko *et al.* (2017) have argued. Second, medium-term domestic rates only capture components of Fed shocks that affect domestic medium-term assets, and potentially miss components of Fed shocks that affect foreign or long-term assets. Latent shocks, which are inferred directly from long-term foreign assets, do not suffer from this problem. While this makes the shocks harder to interpret, this paper is about identifying asymmetries in β rather than identifying operational measures of m_t.

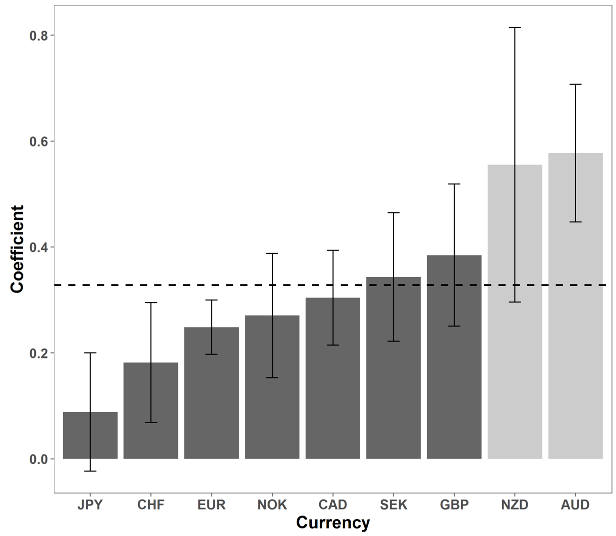
Regardless, I still check my results are robust to utilizing movements in the two-year Treasury yield as shocks, as in Hanson and Stein (2015). Figure A.7 presents the three core specifications (with currencies, bonds, and cross-border portfolios) using this measure of shocks. The results are qualitatively equivalent to those generated using latent shocks, in the main paper, although the standard error bars are wider due to the reasons outlined.

Figure A.7: *Market Reactions to US Monetary Shocks (2Y Treasury)*

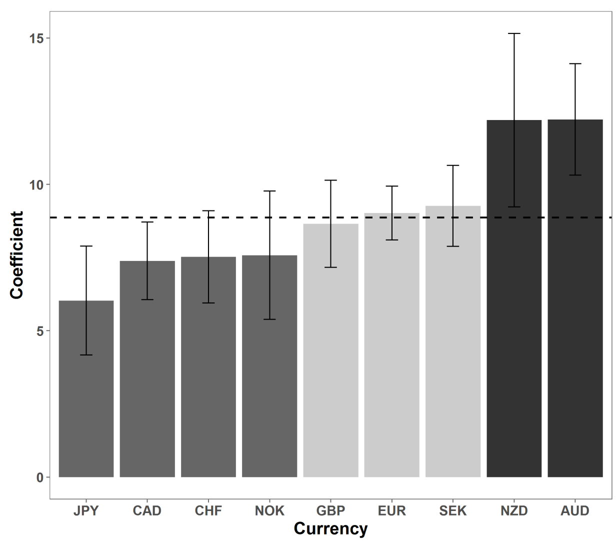
(a) *Currencies*



(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 against 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 results qualitatively mirror Figures 1.4 and 1.5, which use latent shocks, suggesting that the results are not economically sensitive to the choice of shocks.

Average Return

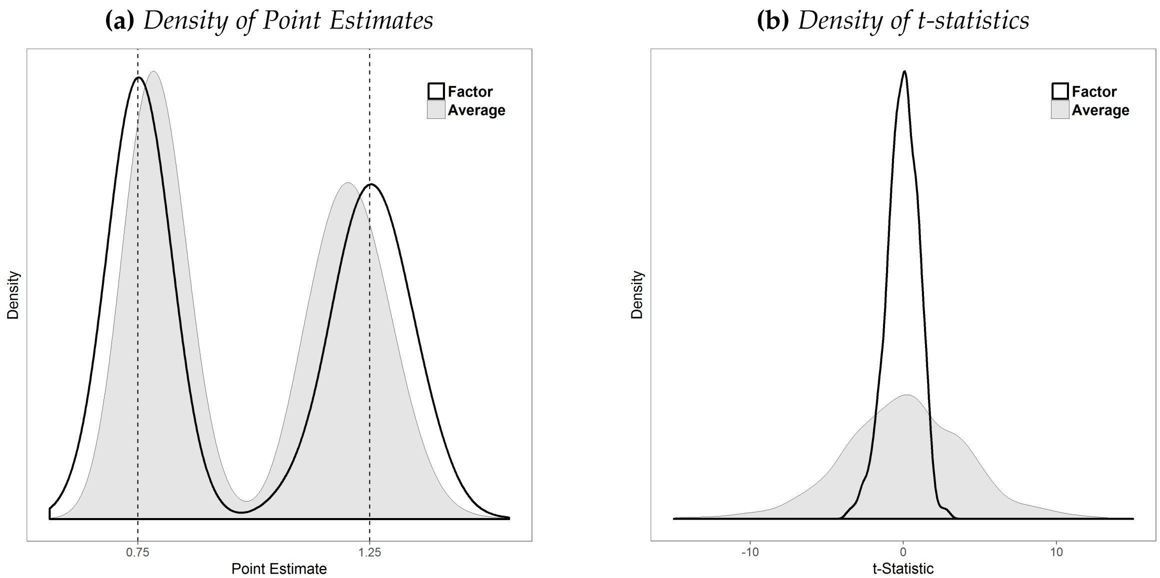
The second class of observed shocks are ones in which shocks are the average asset return r_t at time t. For instance, in the finance literature, individual equity returns are frequently regressed on the market return (i.e. their average return).

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 to zero – dissipate but do not vanish. 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.

Since it is an empirical question of whether nine assets are sufficient to eliminate the bias, I conduct a simulation exercise and find that the bias persists, particularly with standard errors. Specifically, I simulate 500 samples in which coefficients are drawn independently and uniformly $\in \{0.75, 1.00, 1.25\}$, and shocks and errors are drawn using data from announcement and non-announcement windows. I then compare the bias in coefficients and standard errors when computed by the Factor Model (solved by the Expectation Maximization algorithm) and the Average Return method.

Figure A.8 shows the results. The first panel show coefficient bias by focusing on the point estimates of coefficients whose true underlying value $\in \{0.75, 1.25\}$. The Factor Model has no bias, whereas the Average Return method shows a slight bias towards one. This is problematic but not highly so. The second panel shows the bias in standard errors, which is highly problematic. I focus on t-statistics for coefficients whose true underlying value $= 1.00$. While the distribution of t-statistics for the Factor Model is approximately normal, it is highly fat-tailed for the Average Return method.

Figure A.8: *Checking for Bias in Average Return Method*



Notes: The figures depict the results of a simulation exercise, in which I simulate 500 samples and compare the Factor Model against the Average Return method. In each sample, coefficients are drawn independently and uniformly $\in \{0.75, 1.00, 1.25\}$, and shocks and errors are drawn from distributions parameterized using announcement and non-announcement windows. The first panel show coefficient bias by focusing on the density of point estimates of coefficients whose true underlying value $\in \{0.75, 1.25\}$, and they show a slight bias towards one for the Average Return (but not the Factor Model). The second panel shows the standard error bias by focusing on the density of the t-statistics for coefficients whose true underlying value $= 1.00$, and the distribution generated by the Average Return method is too fat-tailed suggesting severe standard error biases.

A.3 Characterizing Asymmetric Responses

This section discusses different variables that successfully predict the asymmetric responses in currencies, bonds, and cross-border bond portfolios. In the paper, I focus on the level of interest rates as determining the cross-section of responses. In this section, I both test that characterization more rigorously, and I evaluate other characterizations. The results are suggestive of the underlying mechanism behind Fed spillovers, and should guide future work in identifying the exact channels. I first introduce the methodology used, and then discuss the results for various candidate predictors.

A.3.1 Methodology

The methodology takes the empirical framework discussed in Section 1.3, and replaces currency-specific coefficients with coefficients interacted with predictor variables. I generalize the three-currency example with interest rates in the paper, Equation (1.6), here as Equation (A.3). I estimate

β_1 for each measured pre-announcement predictor X^i_{t-1} in Equation (A.3).

$$\begin{bmatrix} \Delta s_t^{€/\$} \\ \Delta s_t^{£/\$} \\ \Delta s_t^{¥/\$} \end{bmatrix} = \begin{bmatrix} 1 & X_{t-1}^{€,\$} \\ 1 & X_{t-1}^{£,\$} \\ 1 & X_{t-1}^{¥,\$} \end{bmatrix} \begin{bmatrix} \alpha_0 \\ \alpha_1 \end{bmatrix} + \begin{bmatrix} 1 & X_{t-1}^{€,\$} \\ 1 & X_{t-1}^{£,\$} \\ 1 & X_{t-1}^{¥,\$} \end{bmatrix} \begin{bmatrix} \beta_0 \\ \beta_1 \end{bmatrix} m_t^{\$} + \begin{bmatrix} \epsilon_t^{€/\$} \\ \epsilon_t^{£/\$} \\ \epsilon_t^{¥/\$} \end{bmatrix} \tag{A.3}$$

A statistically significant β_1 means that the candidate predictor X^i_{t-1} is meaningful for predicting variation in responses to Fed monetary announcements both across countries and across time. I estimate Equation (A.3), using different candidate predictors, for the three main specifications: currencies, bonds, and cross-border bond portfolios (that combine a short position in a currency with a short position in its bond). I consider a candidate predictor to be important if it is statistically significant at the 5% level in each of these three specifications.

A.3.2 Results

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

Interest Rates

I let X^i_{t-1} measure an interest rate differential against the US for country i just prior to 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 maturities are significant in all specifications. During Fed announcements, currencies of high-rate countries move against currencies of low-rate countries, long-maturity yields of high-rate countries move more than those of low-rate countries, and the returns on cross-border bond portfolios of high-rate countries are more volatile than those of low-rate countries.

I subject this specific candidate predictor to additional robustness checks, by dropping each country one-by-one and testing whether the level of interest rates is still predictive. Across all maturities and all specifications, the coefficient remains highly significant.[4]

Volatility

I let X^i_{t-1} reflect four different volatility metrics for country i, benchmarked to the US, just prior to time t. The first metric is historical 30-day local equity market volatility (minus US volatility) using equity index data from Datastream, and it is closely related to the VIX.[5] The second, third, and fourth metrics extract different moments of currency volatility, using Bloomberg data on currency options against the dollar. I extract implied volatility from the 25-delta call, the 50-delta call, and the 75-delta call. One metric is the implied volatility from the 50-delta call, which measures expected volatility. Another metric is the difference between implied volatility from the 25-delta and 75-delta call, which measures skew. 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 measures kurtosis.

Of these four candidate predictors, volatility skew is significant in all specifications. During Fed announcements, currencies with high skew move against currencies with low skew, long-maturity yields of high-skew currencies move more than those of low-skew currencies, and the returns on cross-border bond portfolios of high-skew currencies are more volatile than those of low-skew currencies. In particular, Australia, New Zealand, and Norway have high skew while Japan and Switzerland have low skew, implying that the former are risky currencies (with strong downside volatility) and the latter are safe-haven currencies (with limited downside volatility).

Limits to Arbitrage

I let X^i_{t-1} measure the ten-year cross-currency basis for that currency versus the dollar just prior to time t, sourced from Bloomberg. The cross-currency basis quantifies the deviation from covered interest parity, and so should proxy for limits to arbitrage in a given market. However, this measure is insignificant in some specifications, and so it is not pursued further. Moreover, my core findings are present in both the pre-crisis and post-crisis eras, whereas the cross-currency basis was virtually zero prior to the financial crisis.

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[4]There is one minor exception: for the specifications using currencies and for interest rates at one-month maturities, I find significance at the 10% level on occasion.

[5]The VIX uses implied volatility, whereas I compute historical volatility.

Trade Flows

I let X^i_{t-1} measure trade flows against the US in the quarter prior to the announcement, sourced from the IMF's Direction of Trade Statistics (DOTS) database. I construct two measures for a given country: a measure of the US's trade importance, and a measure of the trade balance versus the US. For a given country i, the first and second measures are constructed as follows.

$$X^i_{t-1} = \frac{\text{Exports}^{i \to \text{US}} + \text{Imports}^{i \to \text{US}}}{\sum_c \left(\text{Exports}^{i \to \text{c}} + \text{Imports}^{i \to \text{c}} \right)}$$

$$X^i_{t-1} = \frac{\text{Exports}^{i \to \text{US}} - \text{Imports}^{i \to \text{US}}}{\text{Exports}^{i \to \text{US}} + \text{Imports}^{i \to \text{US}}}$$

The second measure is significant in all specifications. During Fed announcements, currencies whose countries have high bilateral imports versus the US move against currencies whose countries have high bilateral exports, long-maturity yields of high-import countries move more than those of high-export countries, and the returns on cross-border bond portfolios of high-import countries are more volatile than those of high-export countries. In particular, Australia and New Zealand import heavily from the US, while most other countries export on net to the US.

Dollar Invoicing

I let X^i_{t-1} measure the fraction of a given country i's trade invoiced in dollars, based on data published by Gopinath (2015). Since time variation in this data is limited and inconsistent, I focus only on cross-sectional variation and compute three different metrics: dollar invoicing in exports, dollar invoicing in imports, and total dollar invoicing.

Each metric is insignificant in at least one of the three specifications. Broadly, countries with high dollar shares include the Pacific countries of Australia, Canada, and Japan, along with Norway (which has substantial trade in oil), while European countries have low dollar shares. However, there is substantial heterogeneity in how assets of the high-dollar countries react. For instance, the dollar appreciates or depreciates most against the Australian dollar and Norwegian krone, and least against the Canadian dollar and Japanese yen. Assets of European countries are typically in between these two extremes. This creates a non-monotonic pattern relating dollar invoicing to asset heterogeneity, making it a poor predictor variable.

Zhang (2017) argues that dollar invoicing actually drives the underlying heterogeneity in spillovers, and finds that dollar invoicing both predicts currency and bond variation. Our divergent empirical findings can be explained in three ways empirically: his sample includes emerging and developed markets whereas I focus on developed markets, he computes dollar invoicing as a fraction of the overall consumption basket whereas I compute it as a fraction of the traded basket, and he uses short-maturity bonds whereas I use long-maturity bonds.

Bank Positions

I let X^i_{t-1} measure external banking positions against the US in the quarter prior to the announcement, sourced from the BIS's Locational Banking Statistics (LBS) database. I compute two measures that parallel the trade measures for a given country: a measure of the US's banking importance, and a measure of the banking asset-liability balance versus the US. For a given country i, the first and second measures are constructed as follows.

$$X^i_{t-1} = \frac{\text{Assets}^{i \to \text{US}} + \text{Liabilities}^{i \to \text{US}}}{\sum_c \left(\text{Assets}^{i \to \text{c}} + \text{Liabilities}^{i \to \text{c}} \right)}$$

$$X^i_{t-1} = \frac{\text{Assets}^{i \to \text{US}} - \text{Liabilities}^{i \to \text{US}}}{\text{Assets}^{i \to \text{US}} + \text{Liabilities}^{i \to \text{US}}}$$

The second measure is significant in all specifications. During Fed announcements, currencies whose countries have high liability positions versus the US move against currencies whose countries have high asset positions, long-maturity yields of high-liability countries move more than those of high-asset countries, and the returns on cross-border bond portfolios of high-liability countries are more volatile than those of high-asset countries. In particular, Australia and New Zealand have high liability positions while Japan has high asset positions, which is consistent with the former having high interest rates and the latter low interest rates.

Portfolio Debt Positions

I let X^i_{t-1} measure portfolio debt positions against the US in the quarter prior to the announcement, sourced from the IMF's Coordinated Portfolio Investment Survey (CPIS) database. I compute two measures that parallel the trade and banking measures for a given country: a measure of the US's portfolio debt importance, and a measure of the portfolio debt asset-liability balance versus the US.

Each of the two measures is statistically insignificant in at least one of the specifications. For example, both Australia and Canada have similarly large liability positions in portfolio debt versus the US, but the Australian dollar and Canadian dollar have strongly different reactions to Fed announcements.

Portfolio Equity Positions

I let X^i_{t-1} measure portfolio equity positions against the US in the quarter prior to the announcement, sourced from the IMF's Coordinated Portfolio Investment Survey (CPIS) database. I compute two measures that parallel the trade, banking, and portfolio equity measures for a given country: a measure of the US's portfolio equity importance, and a measure of the portfolio equity asset-liability balance versus the US.

The second measure is significant in all specifications. During Fed announcements, currencies whose countries have high asset positions versus the US move against currencies whose countries have high liabilities positions, long-maturity yields of high-asset countries move more than those of high-liability countries, and the returns on cross-border bond portfolios of high-asset countries are more volatile than those of high-liability countries. Surprisingly, many high-rate countries like Norway and New Zealand actually have large asset positions in US equities, while low-rate countries like Switzerland and Japan have large liability positions. This may be because the countries with high interest rates are smaller and have underdeveloped equity markets locally.

Distance

Finally, I let X^i_{t-1} measure the distance between the US and a given country, based on conceptual work by Lustig and Richmond (2017) and based on distance data by Mayer and Zignago (2011). This is a statistically significant predictor in all specifications. During Fed announcements, currencies of distant countries move against currencies of close countries, long-maturity yields of distant countries move more than those of close countries, and the returns on cross-border bond portfolios of distant countries are more volatile than those of close countries. Canada and the

UK, whose assets react similarly to each other, are the closest countries; and Australia and New Zealand, whose assets react similarly to each other, are the furthest countries. (Japan is an outlier.)

Lustig and Richmond (2017) find distance to predict asymmetric responses in currencies for *all* countries. I thus extend this test to monetary announcements from all other central banks and to specifications of Equation (A.3) with currencies. The results here are statistically mixed. For monetary announcements from the Eurozone, Australia, and New Zealand, distance does predict asymmetries. The currencies of close countries — respectively, European countries, New Zealand, and Australia — appreciate or depreciate less than the currencies of distant countries, versus the home currency. For monetary announcements from Canada and the United Kingdom, distance predicts asymmetries negatively. The Australian and New Zealand dollars appreciate or depreciate less both against the pound and Canadian dollar, despite the substantial distance between those countries. Finally, distance does not predict asymmetries in currency markets following monetary announcements from the remaining four countries of Japan, Norway, Sweden, and Switzerland.

More generally, the asymmetries are economically small for shocks generated by most central banks, as documented further in Appendix A.5. Outside of the Federal Reserve and European Central Bank, monetary announcements by most central banks pass into their currencies largely symmetrically, and I estimate this with precision. As such, the effects that Lustig and Richmond (2017) discuss may be better explained by shocks to fundamentals or by other global shocks, rather than by identified monetary shocks.

A.4 Robustness Checks

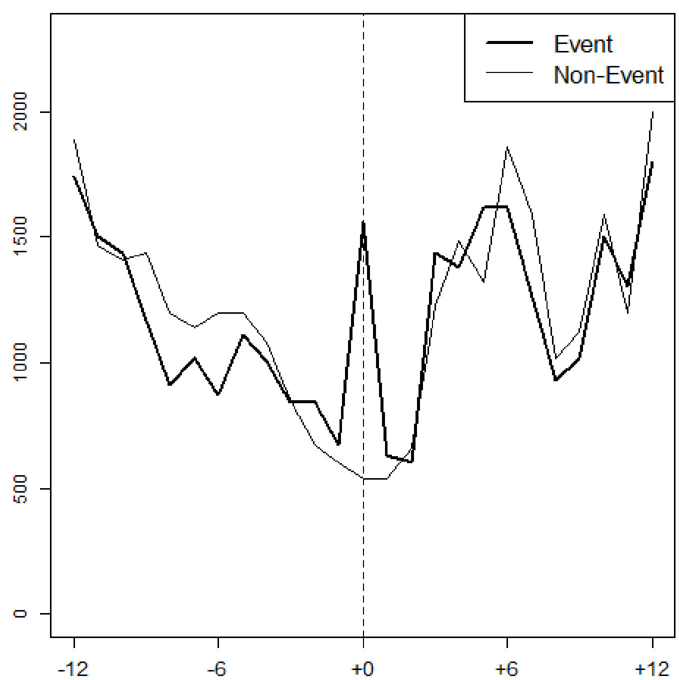
This section describes the various supporting tables and figures for the paper. Table A.2 shows that currency returns over sixty-minute windows do not revert in the subsequent hours, and Figure A.9 illustrates that market liquidity remains high during announcements. Tables A.3 and A.4 support Figures 1.4 and 1.5 respectively by showing the pairwise standard errors for asset asymmetries. Figures A.10, A.11, and A.12 respectively show that the asset asymmetries are robust to splitting the sample over time (pre-crisis and post-crisis), over state (recession and expansion), and over shock type (tightening and easing).

Table A.2: *Testing for Reversion of Currency Returns in High-Frequency Windows*

	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK
3h Windows	0.14	0.10	0.16	0.17	0.17	0.06	0.13	0.16	0.03
7h Windows	0.12	0.11	0.12	0.17	0.18	0.14	0.15	0.19	0.07
11h Windows	0.09	0.13	0.10	0.18	0.19	0.15	0.11	0.20*	0.07
15h Windows	0.14	0.13	0.08	0.05	0.21*	0.13	−0.02	0.16	−0.04
19h Windows	0.10	0.05	0.14	0.10	0.19*	0.17	0.10	0.10	0.06
23h Windows	0.09	0.04	0.14	0.12	0.20*	0.17	0.11	0.09	0.09

This table looks for evidence of reversion of currency returns, measured against the dollar, beyond sixty-minute windows around Fed announcements. I compute the Spearman correlation of currency returns measured over windows of a given length (that start forty-five minutes after a Fed announcement) to those returns measured over sixty-minute windows (fifteen minutes before to forty-five minutes after the announcement). Significance is assessed at the 1% (**) and 5% (*) level. There is virtually no evidence of reversion of currency returns; and in a few cases the returns are positively correlated, although this is not significant in the aggregate.

Figure A.9: *Australian Bond Volume (Intraday)*



Notes: The figure depicts intraday volume for the Australian ten-year bond, to illustrate how markets remain liquid through announcements even in extreme cases of timezone mismatch. I divide announcement and non-announcement days into hourly blocks, and plot the total traded volume (where I compute the median volume across hours on announcement or non-announcement days). Values are plotted with respect to the time of the announcement. In absolute terms, futures liquidity remains high throughout the day, and in unreported results, the median bid-ask spread similarly remains constant throughout the day. In relative terms, while market liquidity drops at night (as Fed announcements take place at approximately 4:00 or 6:00 AM in Australia), liquidity rises specifically during announcements and exhibits volume comparable to volume traded during Australian business hours.

Table A.3: *Pairwise Comparisons on Currency Responses to US Monetary Shocks*

	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK
AUD		0.000	0.047	0.132	0.000	0.003	0.736	0.329	0.680
CAD	0.000		0.013	0.000	0.812	0.664	0.000	0.000	0.000
CHF	0.047	0.013		0.132	0.007	0.004	0.006	0.021	0.010
EUR	0.132	0.000	0.132		0.000	0.003	0.004	0.065	0.001
GBP	0.000	0.812	0.007	0.000		0.498	0.000	0.000	0.000
JPY	0.003	0.664	0.004	0.003	0.498		0.000	0.001	0.001
NOK	0.736	0.000	0.006	0.004	0.000	0.000		0.405	0.885
NZD	0.329	0.000	0.021	0.065	0.000	0.001	0.405		0.389
SEK	0.680	0.000	0.010	0.001	0.000	0.001	0.885	0.389	

The table supports Figure 1.4 by implementing pairwise comparisons among the coefficients associated with each currency. Figure 1.4 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.

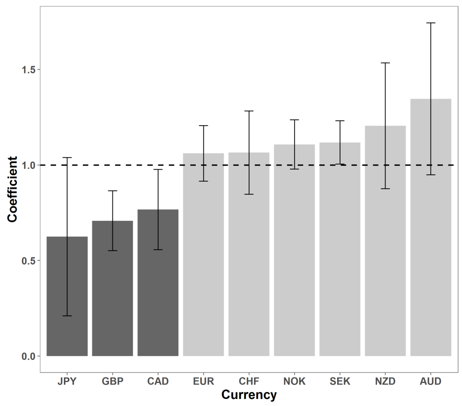
Table A.4: *Pairwise Comparisons on Bond Responses to US Monetary Shocks*

	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK
AUD		0.000	0.000	0.000	0.008	0.000	0.000	0.356	0.000
CAD	0.000		0.048	0.032	0.887	0.000	0.144	0.005	0.194
CHF	0.000	0.048		0.076	0.247	0.000	0.439	0.000	0.721
EUR	0.000	0.032	0.076		0.059	0.066	0.430	0.000	0.175
GBP	0.008	0.887	0.247	0.059		0.000	0.181	0.012	0.229
JPY	0.000	0.000	0.000	0.066	0.000		0.028	0.000	0.001
NOK	0.000	0.144	0.439	0.430	0.181	0.028		0.000	0.653
NZD	0.356	0.005	0.000	0.000	0.012	0.000	0.000		0.001
SEK	0.000	0.194	0.721	0.175	0.229	0.001	0.653	0.001	

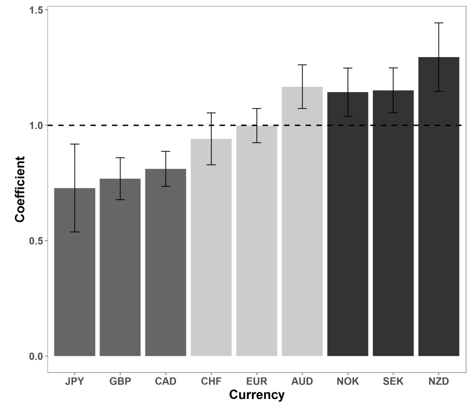
The table supports Figure 1.5 by implementing pairwise comparisons among the coefficients associated with each country's bond yield. Figure 1.5 depicts by how much yields of a given country's ten-year bonds rise when US ten-year yields rise by 1%, following a Fed tightening. This table shows the p-values that emerge from a two-sided two-sample t-test between the relative rise for the bonds of two countries.

Figure A.10: *Market Reactions to US Monetary Shocks, across Time*

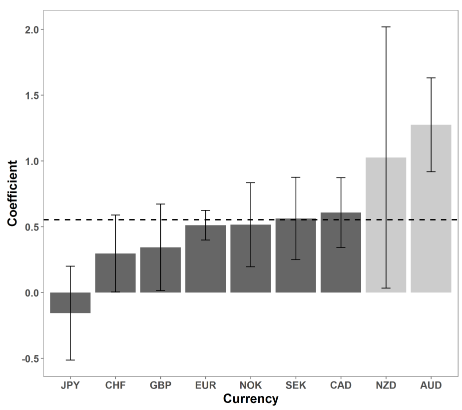
(a) *Currencies, pre-Crisis*



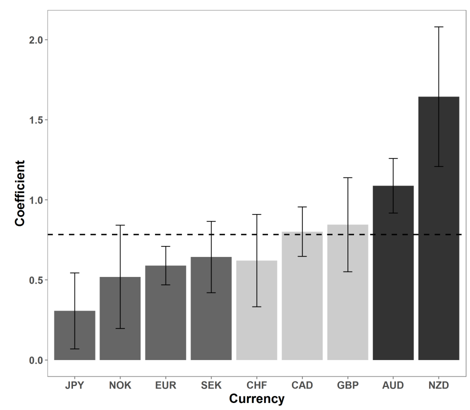
(b) *Currencies, post-Crisis*



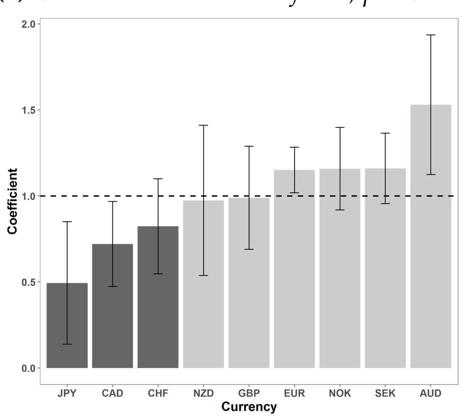
(c) *Bonds, pre-Crisis*



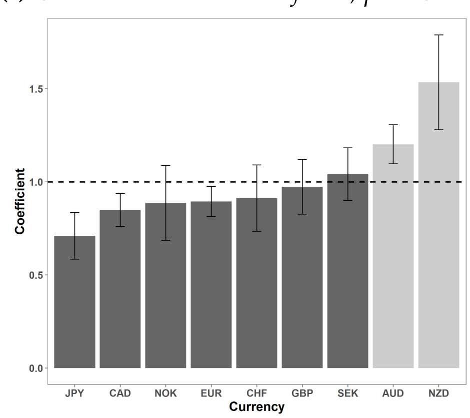
(d) *Bonds, post-Crisis*



(e) *Cross-Border Bond Portfolios, pre-Crisis*



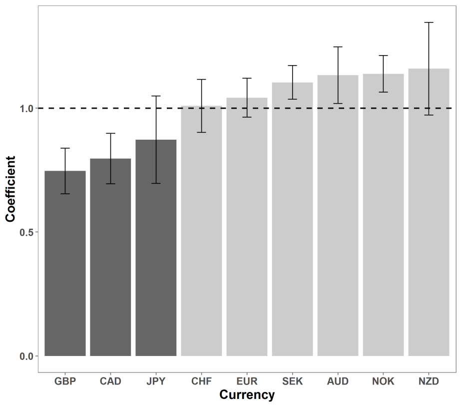
(f) *Cross-Border Bond Portfolios, post-Crisis*



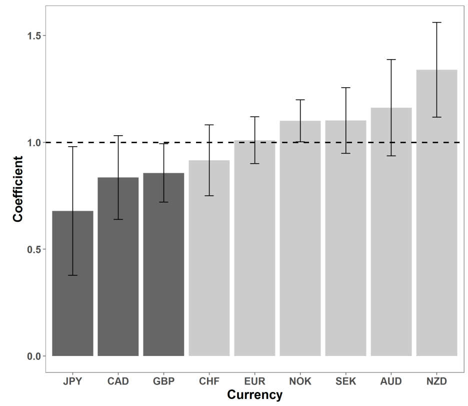
Notes: The figures depict the reactions of currencies, bonds, and cross-border bond portfolios to monetary announcements in the US in two different periods: the pre-crisis sample (2001 until mid-2008) on the left, and the post-crisis sample (mid-2009 until 2016) on the right. The currency figures show 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 against 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 A.11: *Market Reactions to US Monetary Shocks, across States*

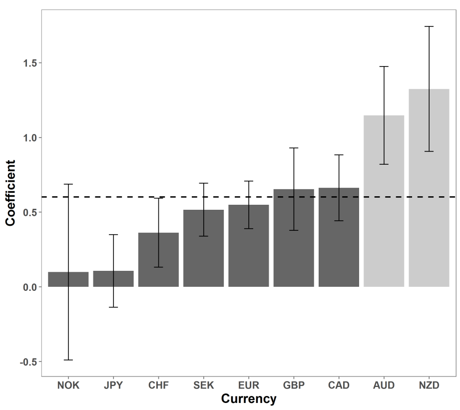
(a) *Currencies, Expansions*



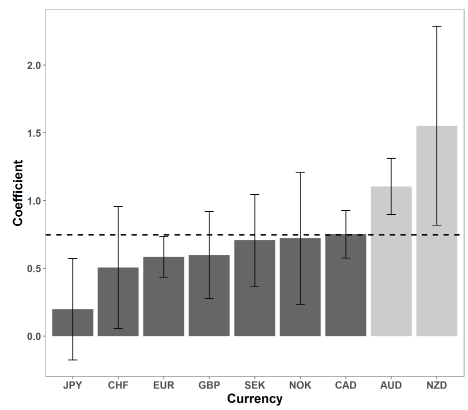
(b) *Currencies, Recessions*



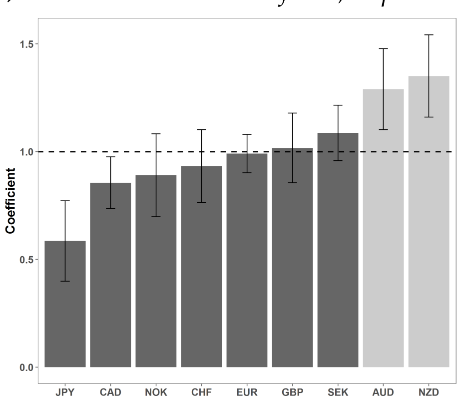
(c) *Bonds, Expansions*



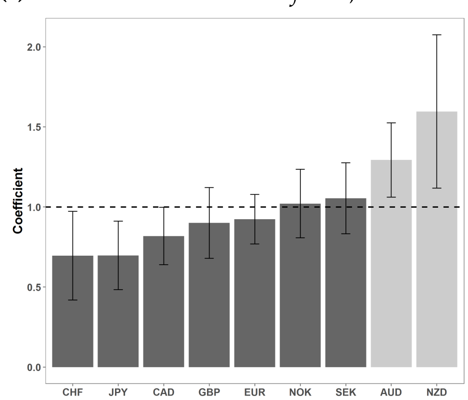
(d) *Bonds, Recessions*



(e) *Cross-Border Bond Portfolios, Expansions*



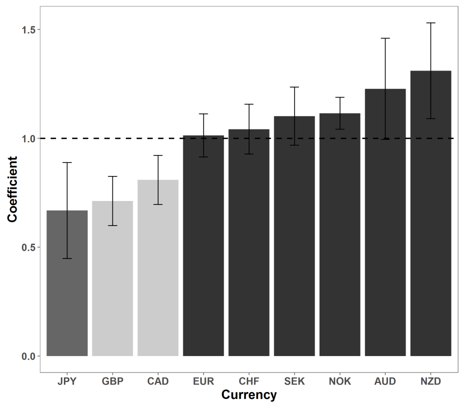
(f) *Cross-Border Bond Portfolios, Recessions*



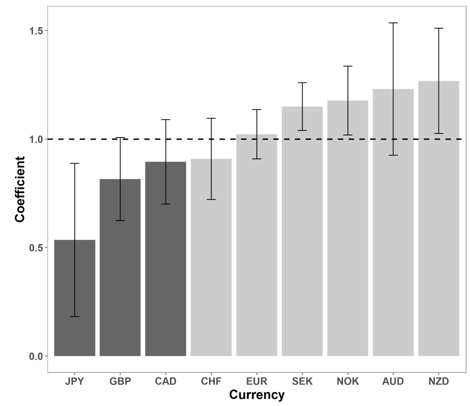
Notes: The figures depict the reactions of currencies, bonds, and cross-border bond portfolios to monetary announcements in the US in two different states: expansionary states (with above-average GDP growth) on the left, and recessionary states (below-average GDP growth) on the right. The currency figures 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 riskfree rate appreciates when the average portfolio appreciates by 1%. Standard error bars in all pictures are computed against 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 A.12: *Market Reactions to US Monetary Shocks, across Shock Types*

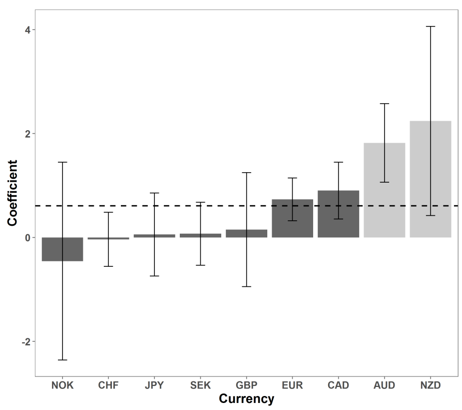
(a) *Currencies, Tightening*



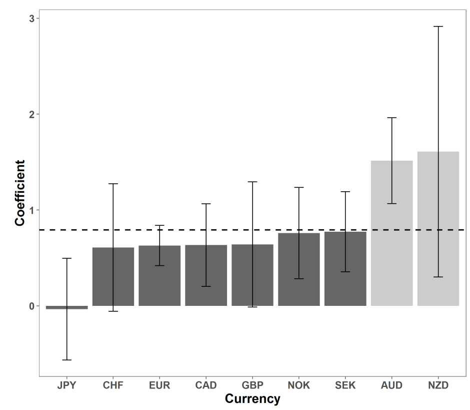
(b) *Currencies, Easing*



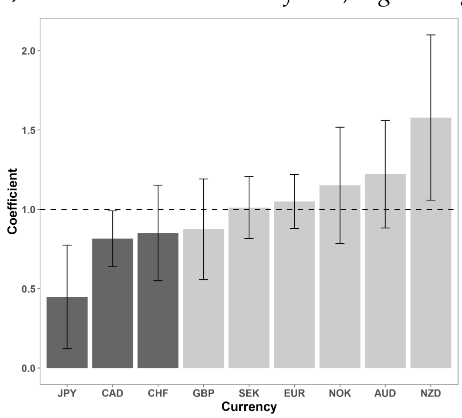
(c) *Bonds, Tightening*



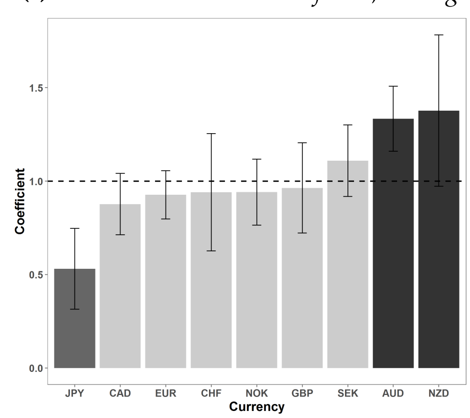
(d) *Bonds, Easing*



(e) *Cross-Border Bond Portfolios, Tightening*



(f) *Cross-Border Bond Portfolios, Easing*



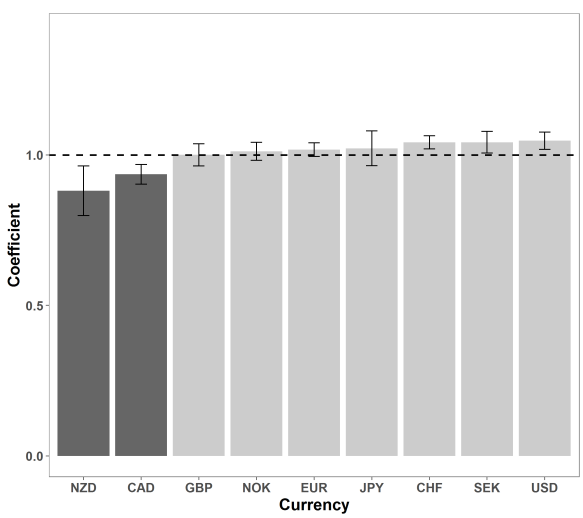
Notes: The figures depict the reactions of currencies, bonds, and cross-border bond portfolios to monetary announcements in the US of two different types: tightening (defined as positive US ten-year yield movements in sixty-minute windows around Fed announcements) on the left, and easing (negative movements) 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 against 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).

A.5 Other Central Banks

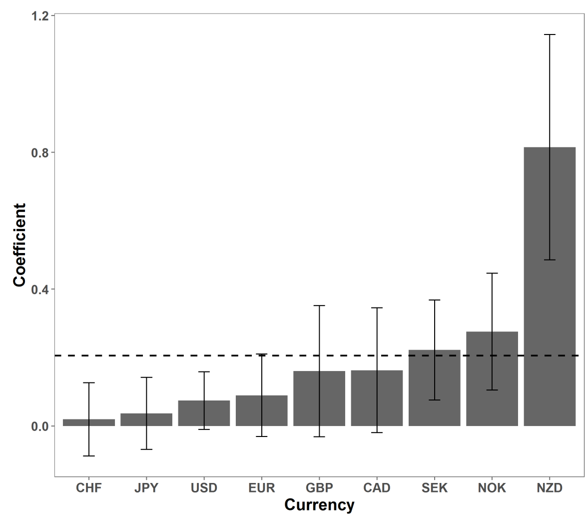
This section presents the reactions of bond and currency markets to announcements emanating from all ten central banks, using the factor model. The paper shows market responses to the Federal Reserve, European Central Bank, Reserve Bank of Australia, Bank of Japan, and Reserve Bank of New Zealand. This section duplicates those figures, and also includes market responses to the Bank of Canada, the Swiss National Bank, the Bank of England, the Norges Bank, and the Swedish Riksbank. The Fed and ECB are the only central banks that generate strong and asymmetric market responses. All other central banks generate negligible and symmetric responses.

Figure A.13: *Market Reactions to Australian Monetary Shocks*

(a) *Currencies*



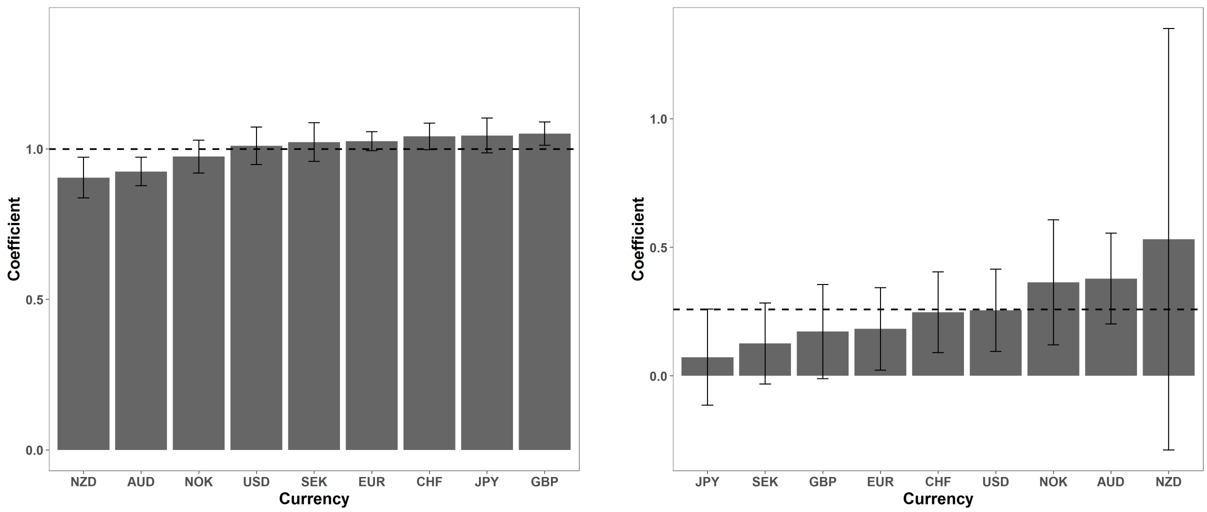
(b) *Bonds*



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 against 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 1.8.

Figure A.14: *Market Reactions to Canadian Monetary Shocks*

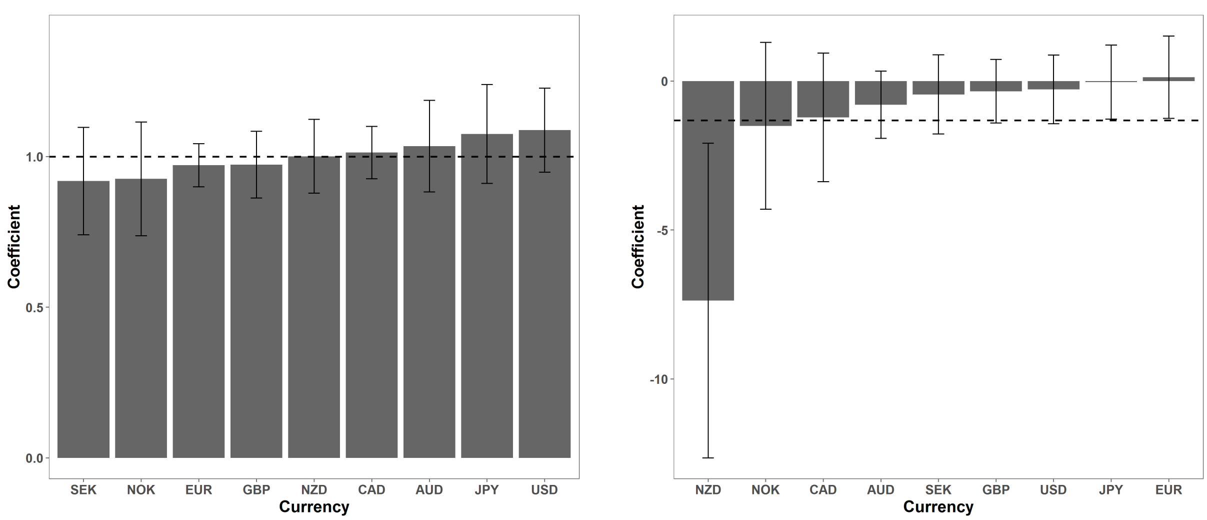
(a) *Currencies* **(b)** *Bonds*



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 against 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 A.15: *Market Reactions to Swiss Monetary Shocks*

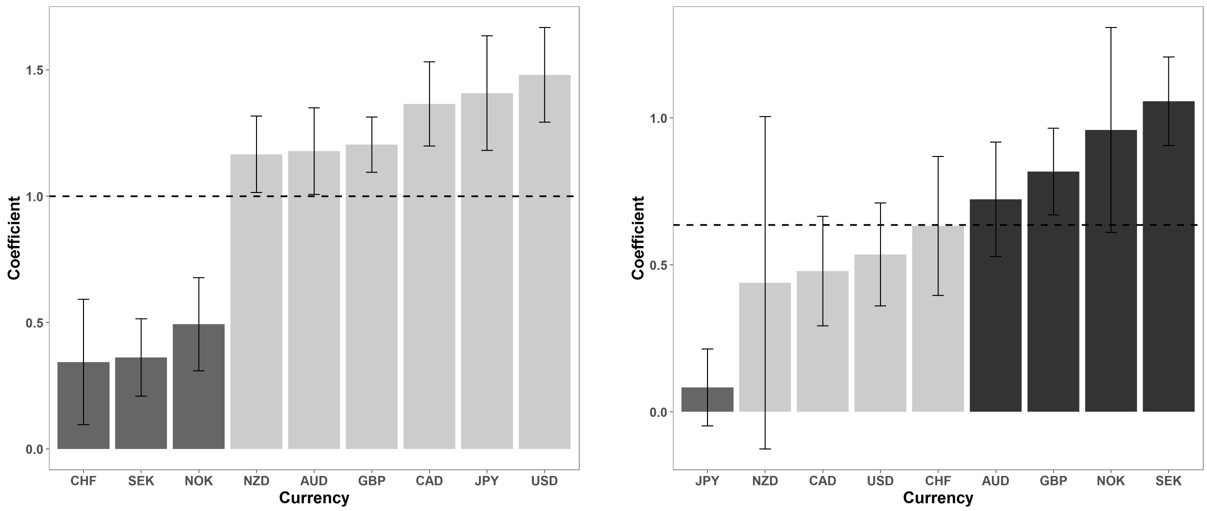
(a) *Currencies* **(b)** *Bonds*



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 against 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 A.16: *Market Reactions to European Monetary Shocks*

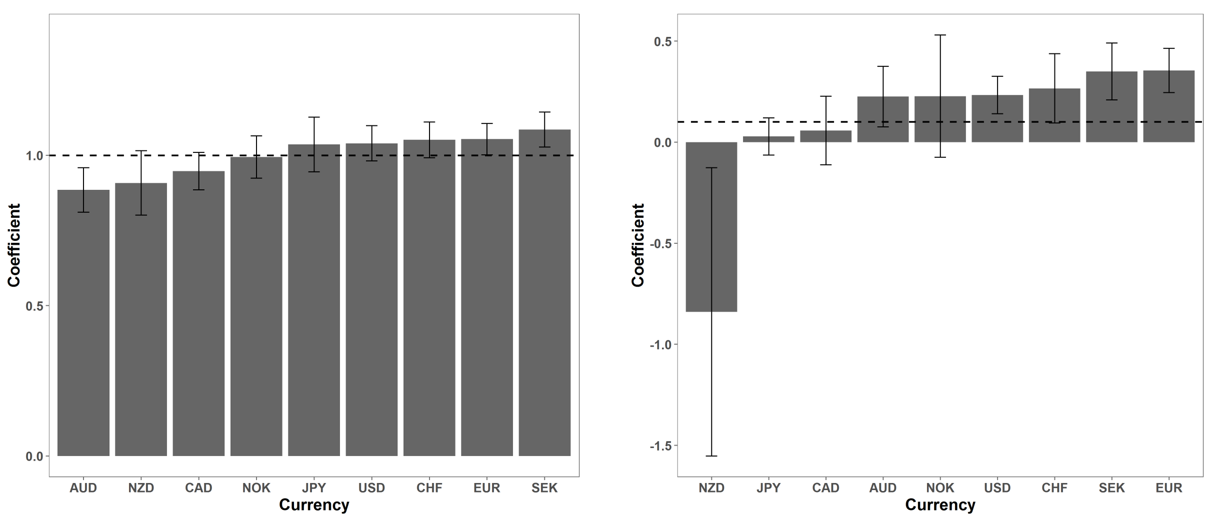
(a) *Currencies* **(b)** *Bonds*



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 against 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 1.6.

Figure A.17: *Market Reactions to British Monetary Shocks*

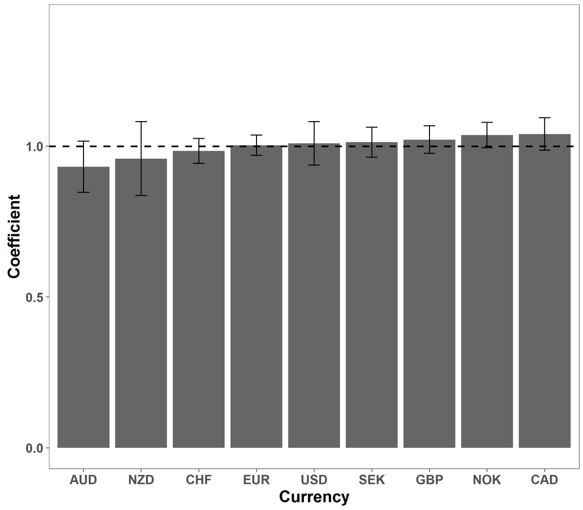
(a) *Currencies* **(b)** *Bonds*



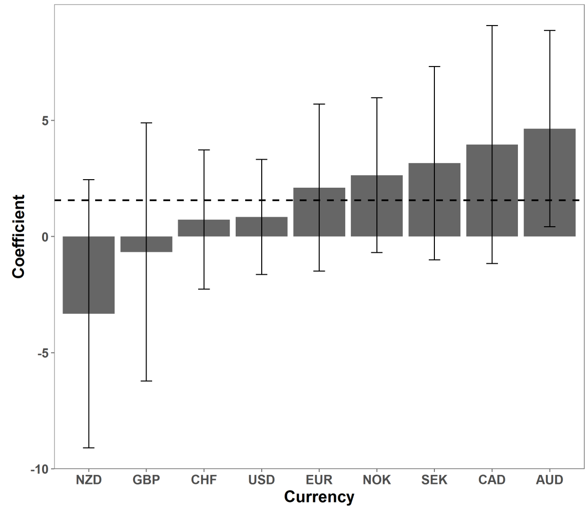
Notes: The figures depict the reactions of currency and bond markets to announcements by the Bank of England. The left figure shows by how much the GBP appreciates against a given reference currency when it appreciates by 1% on average; and the right figure shows by how much foreign ten-year yields rise when British ten-year yields rise by 1%. Standard error bars in both pictures are computed against 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 A.18: *Market Reactions to Japanese Monetary Shocks*

(a) *Currencies*



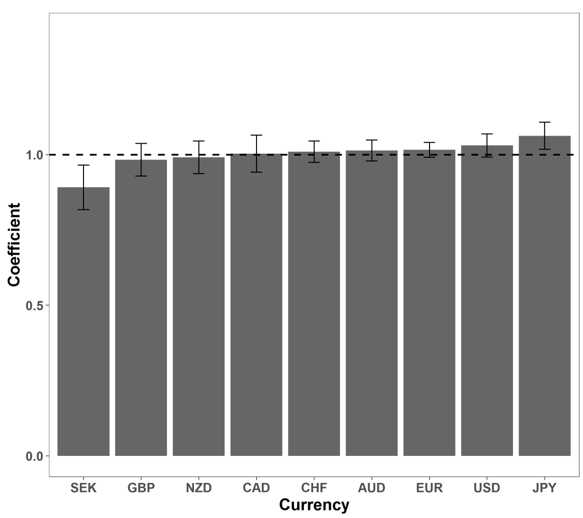
(b) *Bonds*



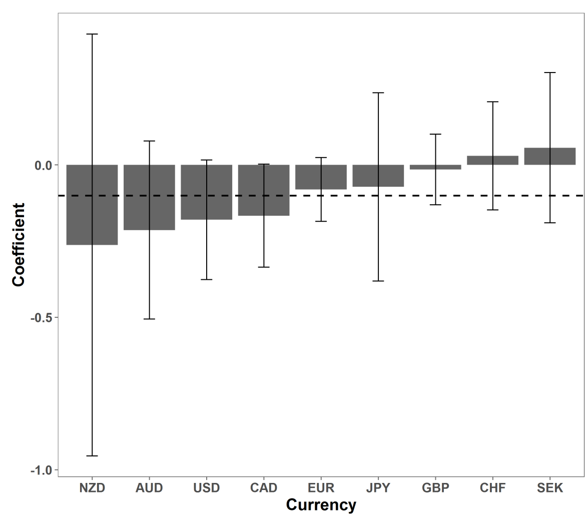
Notes: The figures depict the reactions of currency and bond markets to announcements by the Bank of Japan. The left figure shows by how much the JPY appreciates against a given reference currency when it appreciates by 1% on average; and the right figure shows by how much foreign ten-year yields rise when Japanese ten-year yields rise by 1%. Standard error bars in both pictures are computed against 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 1.7.

Figure A.19: *Market Reactions to Norwegian Monetary Shocks*

(a) *Currencies*



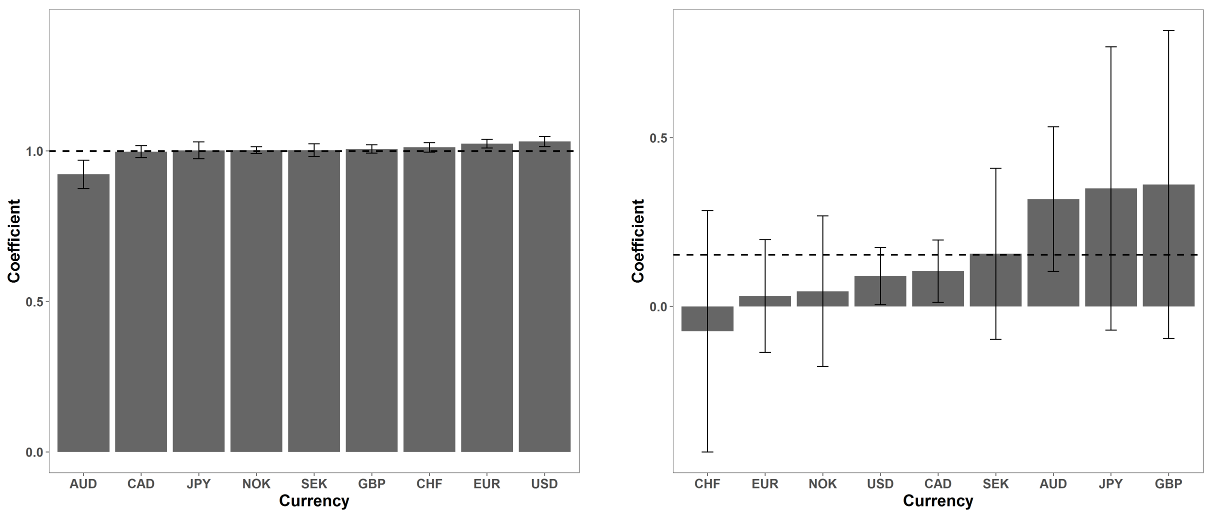
(b) *Bonds*



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 against 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 when the Norges Bank tightens, with the possible exception of Swedish assets.

Figure A.20: *Market Reactions to New Zealand Monetary Shocks*

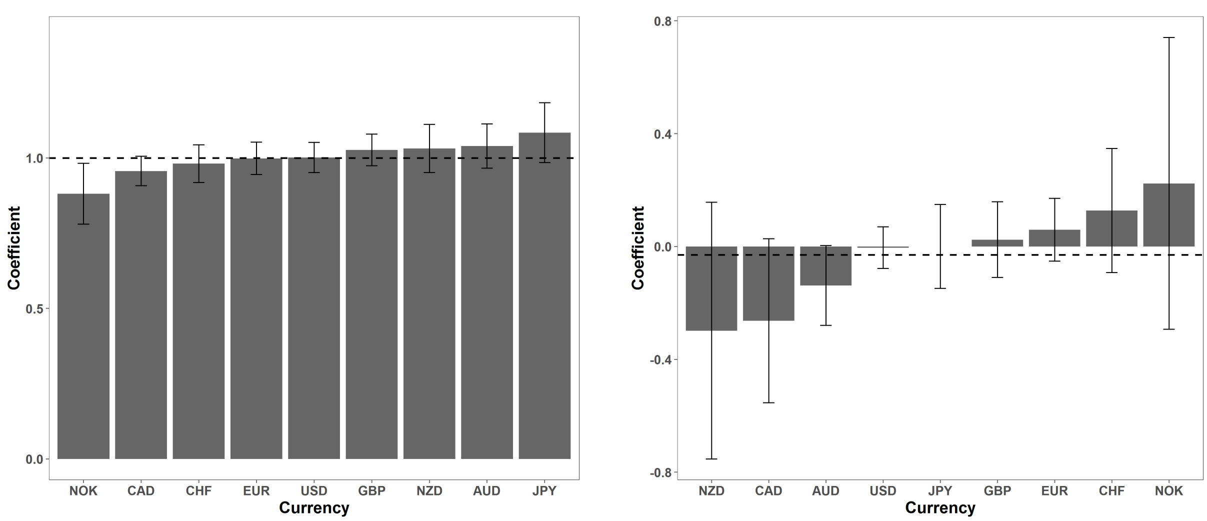
(a) *Currencies* **(b)** *Bonds*



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 against 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.

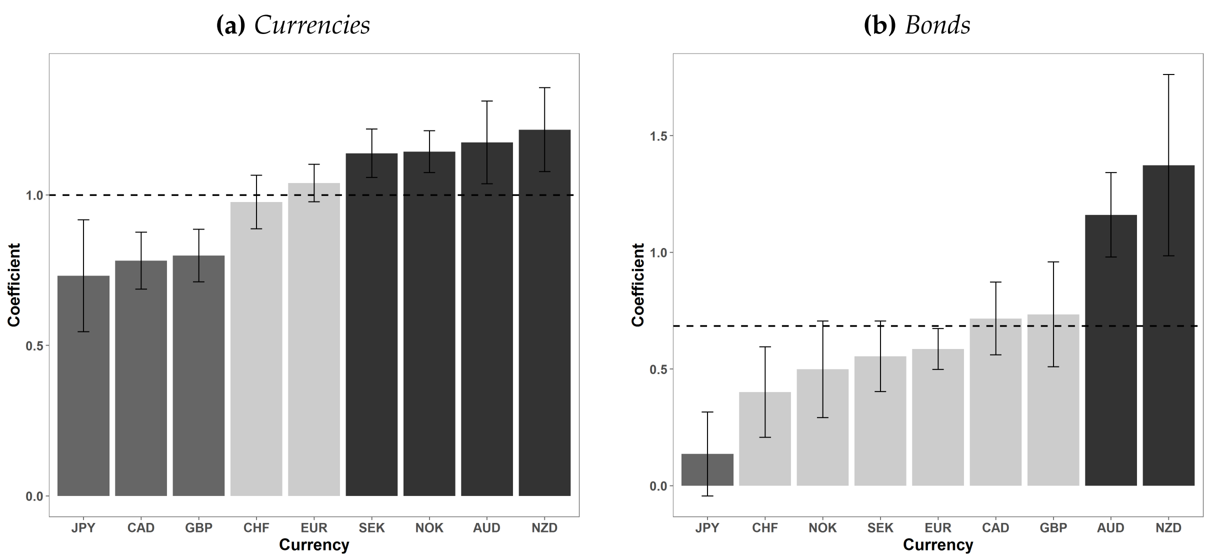
Figure A.21: *Market Reactions to Swedish Monetary Shocks*

(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 against 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 A.22: *Market Reactions to American Monetary Shocks*



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 against 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 1.4 and 1.5.

Appendix B

Appendix to Chapter 2

B.1 Variance Test

Take the following equation, utilized heavily in Chapter 1:

$$r_t = X_t^{\alpha}\alpha + X_t^{\beta}\beta m_t + \epsilon_t \quad \forall\ t \tag{B.1}$$

Now, consider a single-asset variant of Equation (B.1) without time-varying covariates (i.e. $C_r = 1$ and $X_t^{\beta} = X^{\beta}$) 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 β. Instead, this hypothesis can be tested by a simple variance test. Taking the variance of Equation (B.1) yields the following, where missing returns are simply dropped:

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

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:

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

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

$$\mathbb{V}\left(r_t\right) > \mathbb{V}\left(\tilde{r}_t\right) \implies \beta \neq \vec{0}$$

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 requires only one meaningful assumption: 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.

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

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 its complement under both the Brown-Forsythe and the F-test. Table B.1 shows the percentage of time that these tests reject at the 5% level. As kurtosis increases, the F-test performs poorly but the Brown-Forsythe test correctly rejects only 5% of the time.

Table B.1: *Variance Test Performances on Simulated Data*

Kurtosis	BF Test	F-Test
0	0.05	0.05
4	0.04	0.23
8	0.05	0.39

Notes: The table tests whether the Brown-Forsythe test and F-test correctly fail to reject or incorrectly reject equality of variances, for two identically distributed series with increasing kurtosis, and depicts the percentage of simulations that these tests reject at the 5% level. When the data have 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. This illustrates the importance of using the Brown-Forsythe test over the F-test for testing variances.

B.2 Long-Run Assumptions

B.2.1 Overview

In this section, I test the validity of the approximations in Equation (2.3). The cross-border bond portfolio in Equation (2.3) is not directly affected by foreign monetary policy if the following assumption holds.

$$\underbrace{\sum_{k=11}^{\infty} \Delta i^{j}_{t+k-1}}_{\text{Foreign Future Nominal Rates}} - \underbrace{\Delta s^{i/\$}_{\infty}}_{\text{Long-Run Exchange Rate}} \approx 0$$

This assumption states that the conditional expectations of long-run foreign variables, namely long-run foreign nominal interest rates and infinite-horizon exchange rates, do not shift on net through Fed announcements. In turn, this expression can be decomposed into foreign real factors (real rates and the real infinite-horizon exchange rate) and foreign price factors (inflation and the infinite-horizon price level).[1]

$$\left(\underbrace{\sum_{k=11}^{\infty} \Delta r^{j}_{t+k-1}}_{\text{Foreign Real Rates}} - \underbrace{\Delta q^{j/\$}_{\infty}}_{\text{Long-Run Real Exchange Rate}} \right) + \left(\underbrace{\sum_{k=11}^{\infty} \Delta \pi^{j}_{t+k-1}}_{\text{Foreign Inflation}} - \underbrace{\Delta \hat{p}^{i}_{\infty}}_{\text{Long-Run Foreign Price Level}} \right) \approx 0$$

I argue that these additional foreign terms are individually unlikely to react to the Fed. First, I confirm that foreign central banks do not explicitly set nominal interest rates at horizons longer than a few years. Since foreign real or price factors could change implicitly, I next invoke monetary

[1]The infinite-horizon US price level is common to all portfolios, and thus can be excluded.

neutrality to argue against foreign real factors reacting to the Fed at long horizons. Finally, I show direct evidence against foreign price factors reacting at long horizons.

B.2.2 Nominal Interest Rates

I first examine the statements of foreign central banks during their *own* scheduled announcements, and verify that they do not explicitly set nominal interest rates at horizons longer than a few years. Three foreign central banks provide future rate guidance in their own announcements, using only suggestive phrases: "extended period of time" by the European Central Bank, "over the next few years" by the Bank of England, "for the time being" by the Bank of Japan. These phrases suggest that foreign central banks explicitly guide nominal rates over horizons of a few years, at best.

In fact, the only central bank in my sample that gives explicit calendar-based guidance is the Fed. Specifically, 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. These examples similarly suggest that the Fed explicitly sets rates a few years in advance at most.

B.2.3 Real Factors

While it is important to check for explicit central bank guidance over long horizons, foreign real or price factors could adjust implicitly at long horizons. I first focus on shifts in long-horizon foreign real factors: distant foreign real rates (ten years from now) and the infinite-horizon real exchange rate. For both, I invoke monetary neutrality to argue against responses to Fed announcements. Long-horizon real variables are driven by fundamentals (e.g. demographic shifts, technological improvements, etc) and not by monetary news.

Consider real rates first. Domestic macroeconomic models commonly show monetary neutrality to real variables within a few years, once prices adjust. For instance, Uhlig (2005) finds monetary neutrality with respect to real rates is restored at the two-year horizon. Even Nakamura and Steinsson (2017), who argue for monetary non-neutrality at unusually long horizons, still note that neutrality is restored at the ten-year horizon.

Consider the infinite-horizon real exchange rate next. Most papers similarly discuss this as being driven by real determinants, such as productivity costs and trade costs as in Bordo *et al.*

(2017). Moreover, Carvalho *et al.* (2017) use a calibrated model to argue that monetary shocks to the real exchange rate are offset at horizons of two to five years, implying that the infinite-horizon real exchange rate should be unchanged by monetary news. More generally, Chong *et al.* (2012) argue that the real exchange rate converges rapidly to its long-run value, regardless of the shock.

B.2.4 Price Factors

Finally, I focus on shifts in long-horizon foreign price factors, and I show empirical evidence to establish that these do not respond to the Fed. I make the argument in two ways. First, I look at long-run inflation forecasts, and show that these vary too little to explain my results. Second, I extract expectations of inflation from inflation-linked securities, and show that these do not respond at long horizons to the Fed.[2]

Inflation Forecasts

I first examine inflation forecasts at long horizons, and show that they do not vary enough to explain my results. Specifically, these forecasts change a few basis points per year, whereas my results find that assets move a few basis points *per announcement*.

Nominal bond yields move 1.8 annualized basis points on average through Fed announcements, or just over five basis points annually. Moreover, 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). Finally, inflation in a foreign country is also exposed to its own shocks. Thus, long-run inflation forecasts should vary substantially more than five basis points per year for inflation to plausibly explain my results.

In fact, long-run inflation forecasts vary at most around five basis points annually. I show this in several ways. First, I consider the IMF's World Economic Outlook, which makes five-year inflation forecasts for the ten countries in my sample. I look at the median absolute revision in the forecast as the five-year ahead forecast in one year becomes the four-year ahead forecast in the next year. The revision is three basis points. Second, I look at the Fed's Survey of Professional

[2] Although I focus on inflation, note that the infinite-horizon price level enters the expression too. However, it enters with the *opposite* sign as inflation, and so it only serves to dampen the effects of inflation.

Forecasters and find that their median absolute revision for the five-year ahead forecast is six basis points. Finally, I consider the European Central Bank's forecast, and their median absolute revision for the five-year ahead forecast is zero basis points.

Inflation-Linked Securities

I next examine estimates of inflation from inflation-linked securities, and show that they do not react to Fed announcements. The relative advantage of this approach is that it yields high-frequency measures of inflation, while the relative disadvantage is that it identifies inflation expectations and inflation risk premia. I study expected inflation over two maturities: maturities beyond ten years (for Section 2.2) and the six-year forward four-year maturity (for Section 2.3).

First, I offer evidence from returns in the Treasury inflation-protected securities (TIPS) market. Table B.2 shows that Fed announcements drive nominal US yields and real US yields for seven-year forward three-year and ten-year forward twenty-year maturities, but fail to drive the difference, which reflects expectations of inflation and inflation premia.[3]

Table B.2: *Excess Volatility in US Inflation Estimates*

	Nominal	Real (TIPS)	Inflation
7F3Y	42^{**}	24^{**}	18
10F20Y	51^{**}	33^{**}	-9

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. 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. 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.

Second, I offer evidence from inflation swaps for the two foreign countries with liquid markets: Germany and the United Kingdom. Inflation is estimated directly from inflation swaps, which renders nominal yields unnecessary. Table B.3 shows that the Fed does not affect estimates of inflation over the six-year forward four-year and ten-year forward twenty-year maturities.

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

[3]The 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 B.3: *Excess Volatility in Foreign Inflation Estimates*

	E.U.	U.K.
6F4Y	11	7
10F20Y	13	12

Notes: The table tests whether estimates of inflation, derived from inflation swaps, are more volatile around announcements by the Fed than at other times, using daily returns. The four tests use German and British data over six-year forward four-year and ten-year forward twenty-year maturities. 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. The Fed does not affect inflation estimates at any forward maturity or in any country.

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*.

B.3 Structural Decomposition of International Yield Curves

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 Gaussian affine term structure model of Adrian *et al.* (2013), hereafter ACM. ACM apply this model to US data. However, I am the first paper to systematically apply it to the yield curves of other countries, apart from Jennison (2017), who applies the model to Australian data. I sketch the steps (noting one modification from the original paper) and document the results below.

1. The key variables needed to decompose yield curves are the risk neutral bond pricing parameters A_n^{RF} and B_n^{RF}. For instance, ACM note that the time average of future short rates over the next n months is defined as $-n^{-1}(A_n^{RF} + B_n^{RF'} X_t)$, where X_t refers to the state variables at t. Moreover, the difference between yields and these rates is the average of term premia over the next n periods. In turn, A_n^{RF} and B_n^{RF} are built recursively, using other model parameters $(\mu, \Phi, \Sigma, \sigma^2)$.

$$A_n^{RF} = A_{n-1}^{RF} + B_{n-1}^{RF'}\mu + \frac{1}{2}\left(B_{n-1}^{RF'}\Sigma B_{n-1}^{RF} + \sigma^2\right) - A_1^{RF}, n = 2, ..., 120$$

$$B_n^{RF'} = B_{n-1}^{RF'}\Phi - B_1^{RF'}, n = 2, ..., 120$$

   Finally, to initialize the sequence with A_1^{RF} and B_1^{RF}, ACM propose regressing the one-month Treasury bill on pricing factors X_t, and defining the constant and coefficient matrix as A_1^{RF}

and B_1^{RF} respectively. Implicitly, this assumes away a risk premium at the shortest end of the maturity curve, and so a regression of observed yields directly on the state variable will uncover the risk-neutral pricing parameters.

2. Consider μ and Φ. These are the parameters in a VAR regression of state variables X_t:

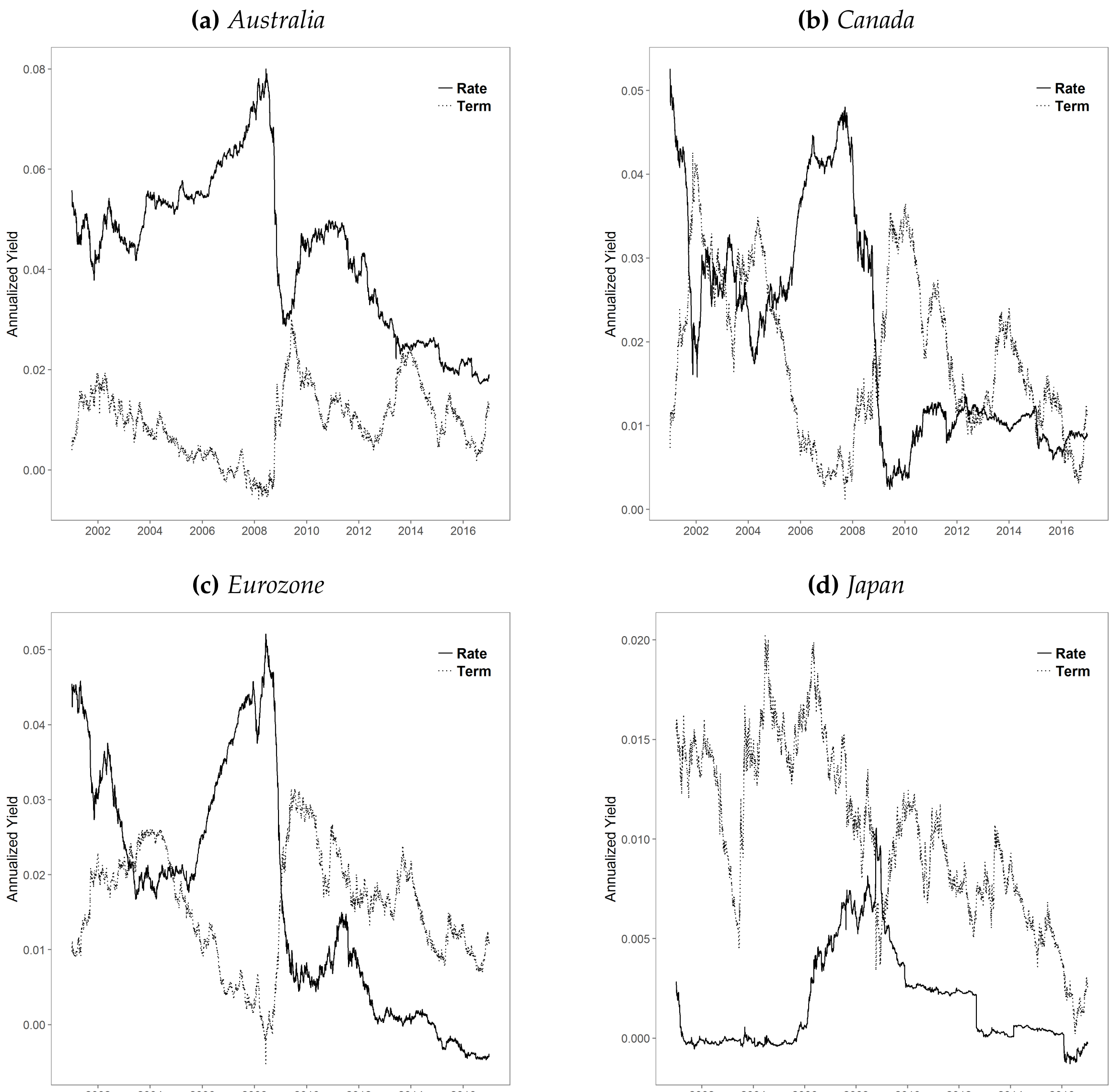
$$X_{t+1} = \mu + \Phi X_t + v_{t+1}$$

   In turn, state variables X_t are the first five factors from the cross-section of yields at maturities every three months. Operationally, ACM measure this cross-section of yields at the monthly frequency to get the principal components, and then apply these weights (i.e. eigenvectors) to the cross-section of yields measured at the daily frequency to get daily factors. I make one small difference, because I have less data than ACM: I extract eigenvectors using yields measured at the weekly (rather than monthly) frequency. For instance, consider the case of New Zealand, which has the largest data limitations — the original methodology would use 200 observations to estimate a 40×40 covariance matrix for the cross-section of yields, whereas my modification uses 800 observations.

3. Finally, Σ and σ^2 are defined from the residuals of various regressions. Σ is defined as the variance-covariance matrix from which errors v_{t+1} in the state variable regression are drawn. σ^2 is defined through a more complex process, defined in the paper, that regresses excess returns on state variables. Practically, the results are largely insensitive to σ^2, although I follow ACM's methodology exactly nonetheless.

I present the results for the decomposition of the ten yield curves in Figure B.1, over a ten-year horizon. Each country has its own idiosyncrasies, but all countries show large drops in the expected path of rates during the financial crisis.

Figure B.1: *Decomposition of International Yield Curves*



(a) *Australia* **(b)** *Canada* **(c)** *Eurozone* **(d)** *Japan*

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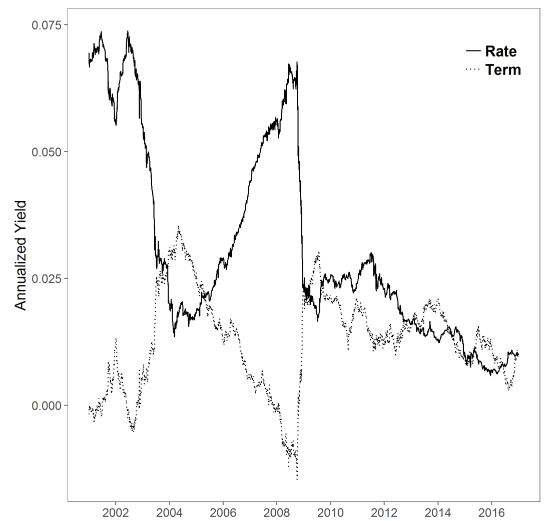
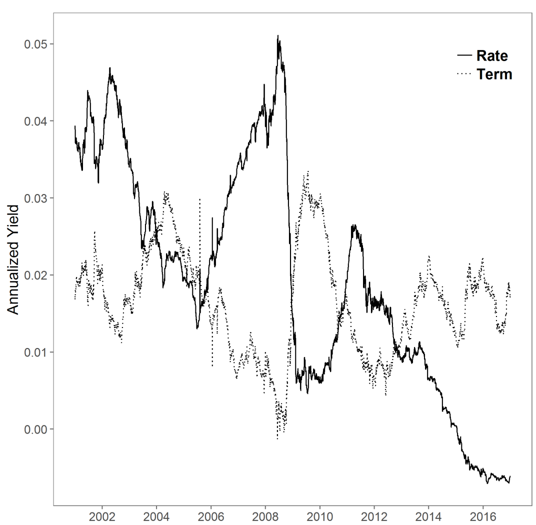
Figure B.1: *(Continued) Decomposition of International Yield Curves*

(e) *New Zealand*

(f) *Norway*



(g) *Sweden*



(h) *Switzerland*

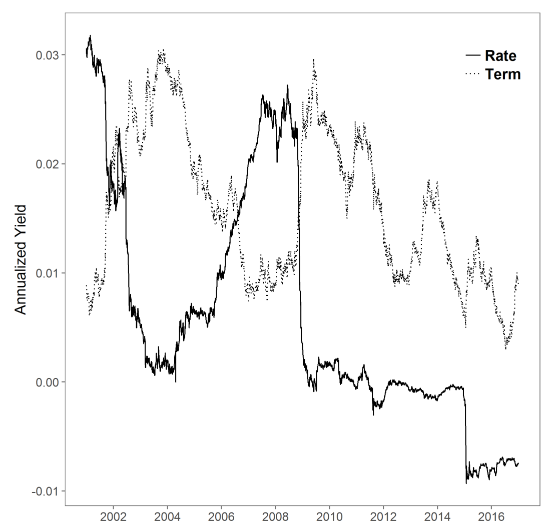


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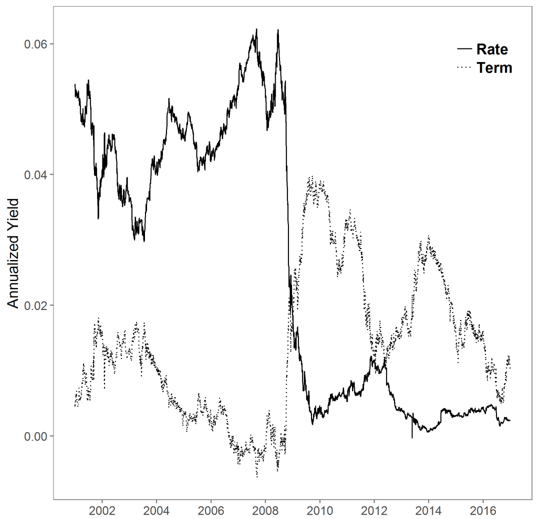
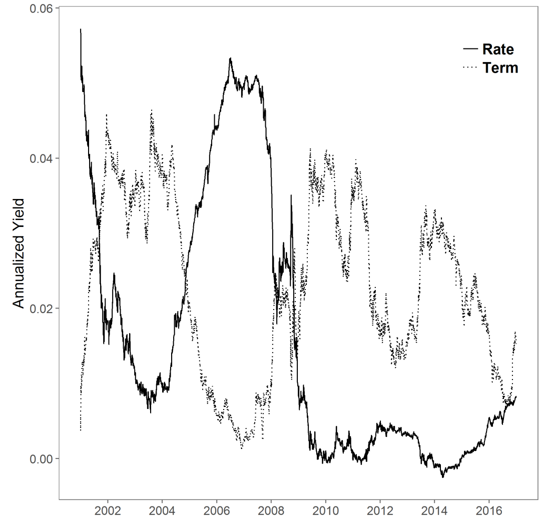
Figure B.1: *(Continued) Decomposition of International Yield Curves*

(i) *United Kingdom*



(j) *United States*



Notes: The figures depict the decomposition of international yield curves into the expected paths of rates and term premia at a ten-year horizon, per the model of Adrian *et al.* (2013), over 2001 - 2016 for most countries and over 2005 - 2016 for New Zealand. I apply the model exactly as given, with only one small modification to handle limited data more robustly. Values are expressed in annualized yields. All countries show a strong drop in the expected paths of rates around the financial crisis, as expected. This model computes the decomposition at the daily frequency, and thus can be used to estimate responses in the paths of rates and term premia to the Fed at high frequencies.

B.4 Robustness Checks

This section describes the various supporting tables for the paper. Table B.4 supports Figure 2.2 by showing the pairwise standard errors for asset asymmetries. Figure B.2 shows that the results are robust to using thirty-year instead of ten-year bonds. Tables B.5, B.6, B.7, B.8, B.10, and B.11 support Tables 2.1, 2.2, 2.3, 2.4, 2.5 and 2.6 respectively by showing the full results for the various inference by heteroskedasticity tests. Table B.9 shows that the results on forward yields are robust to the exact cutoffs utilized.

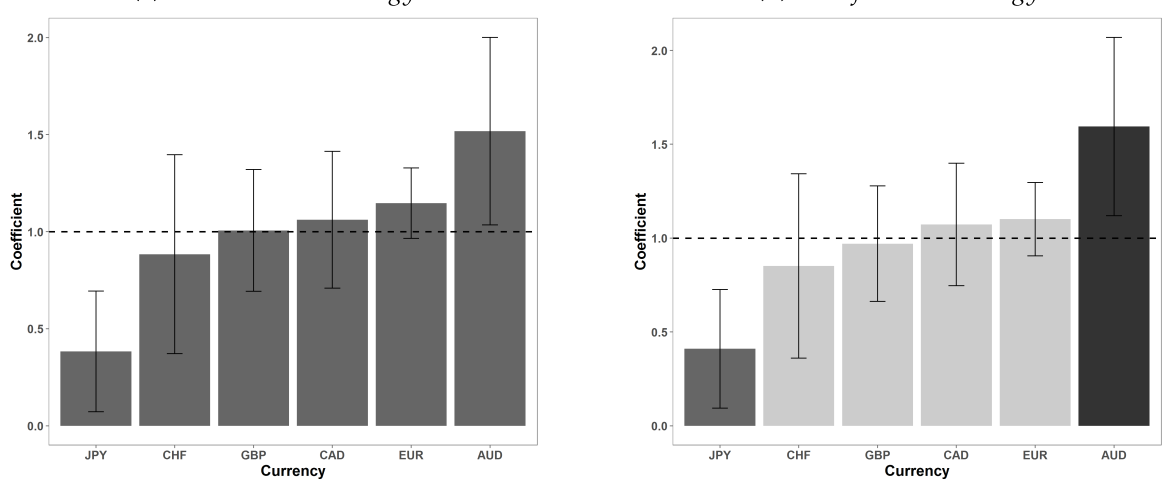
Table B.4: *Pairwise Comparisons on Portfolio Responses to US Monetary Shocks*

	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK
AUD		0.000	0.000	0.000	0.001	0.000	0.011	0.555	0.028
CAD	0.000		0.022	0.883	0.203	0.002	0.157	0.000	0.011
CHF	0.000	0.022		0.090	0.587	0.000	0.968	0.010	0.148
EUR	0.000	0.883	0.090		0.486	0.016	0.178	0.002	0.012
GBP	0.001	0.203	0.587	0.486		0.000	0.681	0.004	0.186
JPY	0.000	0.002	0.000	0.016	0.000		0.001	0.000	0.000
NOK	0.011	0.157	0.968	0.178	0.681	0.001		0.019	0.303
NZD	0.555	0.000	0.010	0.002	0.004	0.000	0.019		0.079
SEK	0.028	0.011	0.148	0.012	0.186	0.000	0.303	0.079	

The table supports Figure 2.2 by implementing pairwise comparisons among the coefficients associated with portfolios for each country. Figure 2.2 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 B.2: *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 against 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 are 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 B.5: *Excess Volatility in 10Y Bond Returns*

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

Notes: The table, which supports Table 2.1, 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 B.6: *Excess Volatility in Daily 1Y Bond Returns*

	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK	USD
Australia	86**	−20	−2	−16	−10	1	−10	18	−18	−13
Canada	−12	67**	5	−1	−6	18	11	13	−11	−5
Switzerland	14	−17	108**	−13	13	11	67	36**	−13	12
Euro	4	−17	45**	53**	−4	−27	12*	18	22*	12
United Kingdom	28**	−23	29**	−16	22	−11	21*	9	13	2
Japan	0	−3	−8	5	−2	104**	−3	1	−2	2
Norway	12	−16	12	1	−3	9	131**	40	−18	50*
New Zealand	53**	−3	−5	4	9	4	−33	92**	−25	−4
Sweden	14	6	−22	−1	7	25	−8	−1	133**	1
United States	4	−6	11	20*	−2	−1	12	−20	10	79**

Notes: The table, which supports Table 2.2, 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 B.7: *Excess Volatility in Daily 6F4Y Bond Returns*

	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK	USD
Australia	10	−7	3	−5	4	−1	3	14*	0	−12
Canada	1	14	10	−4	−3	0	−18	14	−19	−8
Switzerland	30*	30*	18	−9	17	8	15	87**	28**	41*
Euro	8	3	19**	32**	3	0	5	−4	20	5
United Kingdom	12	−1	5	1	1	−17	−12	13	−12	−6
Japan	0	4	−1	3	0	−2	−6	1	4	11*
Norway	9	30**	−8	13	14	−5	15*	38**	6	−6
New Zealand	−7	−5	6	2	1	13	−2	7	−24	−5
Sweden	5	10	3	0	5	−15	−2	15*	19	18
United States	38**	36**	30**	50**	29**	7	18	57**	29**	51**

Notes: The table, which supports Table 2.3, tests whether daily returns of the column country's six-year forward four-year sovereign bonds (e.g. the rate one can guarantee from 2024 to 2028, in 2018) 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 B.8: *Excess Volatility in Daily 10F20Y Bond Returns*

	AUD	CAD	CHF	EUR	GBP	JPY	USD
Australia	−15	1	1	7	13	20	−7
Canada	−4	−14	−1	−14	−24	2	−7
Switzerland	54**	6	−8	−12	18	−1	30*
Euro	6	8	11*	21*	6	10	6
United Kingdom	11*	−1	−6	−10	2	−7	−4
Japan	4	3	5	4	−7	−6	9*
United States	18*	32**	36**	32**	25*	−9	51**

Notes: The table, which supports Table 2.4, 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 B.9: *Excess Volatility in Daily Bond Returns around Fed Announcements*

Forward Yield	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK	USD
4F6Y	38**	32**	35**	47**	35**	1	21*	41**	25**	51**
5F5Y	38**	32**	24**	46**	30**	5	15	53**	26**	50**
6F4Y	38**	36**	30**	50**	29**	7	18	57**	29**	51**
7F3Y	31**	23*	14	42**	18*	4	23*	51**	19**	42**
8F2Y	30**	34**	27**	47**	25**	2	20*	53**	19**	43**

This table tests whether daily returns of the column country's forward bonds (maturity denoted by the row) are more volatile around announcements by the Fed 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. The Fed affects all forward maturities across almost all countries.

Table B.10: *Excess Volatility in Daily 10Y Rate Returns*

	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK	USD
Australia	97**	−11	0	−8	−20	−14	−17	7	−15	−5
Canada	0	82**	−18	0	−7	−2	3	19	3	−12
Switzerland	28**	10	91**	25	22*	40*	92**	19	24*	30
Euro	3	−1	17*	38**	−1	−5	11*	−7	12	13
United Kingdom	44**	13	18*	−7	9	−4	30**	−5	14*	2
Japan	2	0	−4	−7	−4	92**	4	−19	−1	21**
Norway	18*	−10	21	20	−3	−13	131**	5	3	9
New Zealand	35**	−9	8	−17	−6	28*	−20	102**	−29	8
Sweden	11	0	15	2	8	8	−10	12	127**	1
United States	11*	40**	22	30**	−1	29	10	3	14*	97**

Notes: The table, which supports Table 2.5, 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 B.11: *Excess Volatility in Daily 10Y Term Returns*

	AUD	CAD	CHF	EUR	GBP	JPY	NOK	NZD	SEK	USD
Australia	29^{**}	-3	-2	2	0	1	-1	18	-2	-4
Canada	2	19^{*}	-6	-18	-14	11	-27	21	-18	-7
Switzerland	30	38^{**}	58^{**}	-10	7	19^{*}	12	26^{*}	14	61^{**}
Euro	12^{*}	1	10	28^{**}	-15	-6	0	-1	5	21^{**}
United Kingdom	15	12	7	0	10	-12	-6	7	-8	-1
Japan	10	0	-7	12^{*}	4	7	3	-3	13^{*}	22^{**}
Norway	11	25^{*}	7	21^{*}	15	-6	46^{**}	20^{**}	6	-27
New Zealand	10	6	5	5	10	19	-3	20^{*}	-27	11
Sweden	19	7	1	11	2	-6	11	22^{*}	59^{**}	25^{*}
United States	27^{**}	46^{**}	25^{**}	35^{**}	32^{**}	3	14	30^{**}	17^{**}	68^{**}

Notes: The table, which supports Table 2.6, 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.

B.5 Models of Complete Markets

This appendix supports Section 2.4, on models with complete markets, in two ways. First, it formally relates higher bond yield entropy in high-rate countries (e.g. Australia) relative to low-rate countries (e.g. Japan) to higher entropy in transitory stochastic discount factors. Second, it provides derivations for the model that uses Epstein-Zin utility and complex dynamics, relating innovations in stochastic discount factors and bond yields to underlying economic shocks. This section will decompose shocks into permanent and transitory components for this model too, but that has not yet been written; please check back soon.

B.5.1 Bond Entropy

In this section, I translate bond yield entropy into entropy of the transitory components of stochastic discount factors, and argue that this is higher in high-rate countries using three steps. First, I derive expressions relating movements in bond yields over announcement and non-announcement windows to Fed-driven entropy in the transitory components of stochastic discount factors. Second, I estimate that term empirically. Third, I correlate that term with the level of interest rates. I find that the correlation is high and statistically significant, across both ten-year and thirty-year bonds.

I first break the transitory components of the stochastic discount factor into Fed-driven components and idiosyncratic components.

$$\left(\frac{\Lambda_t^{i,\mathbb{T}}}{\Lambda_{t-1}^{i,\mathbb{T}}}\right)_{\text{Total}} = \left(\frac{\Lambda_t^{i,\mathbb{T}}}{\Lambda_{t-1}^{i,\mathbb{T}}}\right)_{\text{Fed}} \times \left(\frac{\Lambda_t^{i,\mathbb{T}}}{\Lambda_{t-1}^{i,\mathbb{T}}}\right)_{\text{Other}}$$

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 be isolated as the difference in total and idiosyncratic. Intuitively, I ascribe the *excess* entropy in the transitory components on announcements days to the Fed.

$$L_{t-1}\left(\left(\frac{\Lambda_t^{i,\mathbb{T}}}{\Lambda_{t-1}^{i,\mathbb{T}}}\right)_{\text{Fed}}\right) = L_{t-1}\left(\left(\frac{\Lambda_t^{i,\mathbb{T}}}{\Lambda_{t-1}^{i,\mathbb{T}}}\right)_{\text{Total}}\right) - L_{t-1}\left(\left(\frac{\Lambda_t^{i,\mathbb{T}}}{\Lambda_{t-1}^{i,\mathbb{T}}}\right)_{\text{Other}}\right)$$

In turn, I estimate these terms using exponentiated innovations in yields, using movements in yields during announcement windows for total entropy and movements in yields during non-announcement windows for idiosyncratic entropy.

$$L_{t-1}\left(\exp\left(n\Delta y_t^i\right)\right) = L_{t-1}\left(\frac{\Lambda_t^{i,\mathbb{T}}}{\Lambda_{t-1}^{i,\mathbb{T}}}\right)$$

Once I generate the Fed-driven entropy for a given country's transitory component entropy, 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. The analysis is conducted with two specifications: ten-year bonds for nine countries, and thirty-year bonds for the six countries that issue them (Australia, Canada, Germany, Japan, Switzerland, and the United Kingdom). The point estimates for the correlation range between 0.7 to 0.8; and the ten-year estimates and thirty-year estimates are statistically significant at the 1% and 5% level respectively. This confirms that the transitory component is more volatile in high-rate countries than in low-rate countries.

B.5.2 Solving Epstein-Zin Utility and Complex Dynamics

This section documents the steps needed to derive expressions for currency and bond yield innovations in a model of complete markets with Epstein-Zin utility and complex consumption dynamics. 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 innovations in the stochastic discount factors and in bond yields, as functions of underlying economic shocks. It is important to stress that shocks are realized at time t, rather than $t+1$ as is common in the literature.

Model Setup

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

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

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

$$c_t - c_{t-1} = \mu + \phi x_{t-1} + \sigma_{t-1} \eta_t$$

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

$$x_t = \rho x_{t-1} + \varphi_e \sigma_{t-1} e_t$$

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

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 into two components: a global component e_t^z and an idiosyncratic component e_t^i. To incorporate structured heterogeneity, different countries i have differential loadings $1+\beta_e^i$ on the global components of shocks. For simplicity, $1+\beta_e^i \geq 0$ 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

α, yielding the updated utility function and dynamics:

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

$$c_t^i - c_{t-1}^i = \mu + \phi x_{t-1}^i + \left(\sqrt{\alpha_\eta}\sigma\left(1+\beta_\eta^i\right)\eta_t^z + \sqrt{1-\alpha_\eta}\sigma_{t-1}^i\eta_t^i\right)$$

$$x_t^i = \rho x_{t-1}^i + \varphi_e\left(\sqrt{\alpha_e}\sigma\left(1+\beta_e^i\right)e_t^z + \sqrt{1-\alpha_e}\sigma_{t-1}^i e_t^i\right)$$

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

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:

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

where $z_t = p_t - d_t$, i.e. the log price-to-dividend ratio, and whereg_t is the log growth rate in dividends. It is worth noting that the coefficient on z_t, 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:

$$\mathbb{E}_{t-1}\left(\delta^\theta\left(\frac{C_t^i}{C_{t-1}^i}\right)^{-\theta/\psi} R_{a,t}(i)^{-(1-\theta)}R_{j,t}\right) = 1$$

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:

$$\mathbb{E}_{t-1}\exp\left(\theta\log\delta - \frac{\theta}{\psi}\left(c_t^i - c_{t-1}^i\right) + (\theta-1)r_{a,t}^i + r_{j,t}\right) = 1 \tag{B.2}$$

which makes the log SDF:

$$m_t^i = \theta \log \delta - \frac{\theta}{\psi}\left(c_t^i - c_{t-1}^i\right) + (\theta - 1) r_{a,t}^i \tag{B.3}$$

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:

$$r_{a,t}^i \approx \kappa_{a,0} + \chi(z_a) z_{a,t}^i - z_{a,t-1}^i + \left(c_t^i - c_{t-1}^i\right)$$

Second, I conjecture that the price-dividend ratio $z_{a,t}$ is a linear function of a country's state variables x_t^i and $\left(\sigma_t^i\right)^2$, as in Bansal and Yaron (2004).[4]

$$z_{a,t}^i = A_0 + A_1 x_t^i + A_2 \left(\sigma_t^i\right)^2$$

To solve the coefficients, I use Equation (B.2), which is the log Euler equation, and price the consumption asset itself $j = a$:

$$\theta \log \delta - \frac{\theta}{\psi}\mathbb{E}_{t-1}\left(c_t^i - c_{t-1}^i\right) + \theta \mathbb{E}_{t-1} r_{a,t} + \frac{1}{2}\mathbb{V}_{t-1}\left(-\frac{\theta}{\psi}\left(c_t^i - c_{t-1}^i\right) + \theta r_{a,t}\right) = 0 \tag{B.4}$$

where:

$$\begin{aligned} r_{a,t} \approx & \kappa_{a,0} + \chi(z_a)\left(A_0 + A_1\left(\rho x_{t-1} + \varphi_e\left(\sqrt{\alpha_e}\sigma(1+\beta_e^i)e_t^z + \sqrt{1-\alpha_e}\sigma_{t-1}^i e_t^i\right)\right)\right. \\ & \left. + A_2\left(\sigma^2 + v\left(\left(\sigma_{t-1}^i\right)^2 - \sigma^2\right) + \sigma_w\left(\sqrt{\alpha_w}\left(1+\beta_w^i\right)w_t^z + \sqrt{1-\alpha_w}w_t^i\right)\right)\right) \\ & - \left(A_0 + A_1 x_{t-1}^i + A_2\left(\sigma_{t-1}^i\right)^2\right) + \left(\mu + \phi x_{t-1}^i + \left(\sqrt{\alpha_\eta}\sigma\left(1+\beta_\eta^i\right)\eta_t^z + \sqrt{1-\alpha_\eta}\sigma_{t-1}^i \eta_t^i\right)\right) \end{aligned} \tag{B.5}$$

[4]The state variables of other countries do not enter this expression, since foreign state variables do not add information on the margin relative to domestic state variables for a country's consumption dynamics.

Using Equation (B.5), I expand the Euler equation (B.4) into:

$$\theta \log \delta + (1-\gamma)\left(\mu + \phi x^i_{t-1}\right) + \theta\left(\kappa_{a,0} + \chi(z_a)\left(A_0 + A_1\rho x^i_{t-1} + A_2\left(\sigma^2 + v\left(\left(\sigma^i_{t-1}\right)^2 - \sigma^2\right)\right)\right)\right.$$
$$\left. - \left(A_0 + A_1 x^i_{t-1} + A_2\left(\sigma^i_{t-1}\right)^2\right)\right) + \frac{1}{2}\theta^2\chi(z_a)^2 A_1^2\varphi_e^2\left(\alpha_e\sigma^2\left(1+\beta^i_e\right)^2 + (1-\alpha_e)\left(\sigma^i_{t-1}\right)^2\right)$$
$$+\frac{1}{2}\theta^2\chi(z_a)^2 A_2^2\sigma_w^2\left(\alpha_w\left(1+\beta^i_w\right)^2 + (1-\alpha_w)\right) + \frac{1}{2}(1-\gamma)^2\left(\alpha_\eta\sigma^2\left(1+\beta^i_\eta\right)^2 + (1-\alpha_\eta)\left(\sigma^i_{t-1}\right)^2\right) = 0$$

This expression must hold for any arbitrary value of the state variables x^i_{t-1} and $\left(\sigma^i_{t-1}\right)^2$, 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.) This is also known as the method of undetermined coefficients. As such, I group all terms involving each state variable, and impose the restriction that their coefficients must equal zero.

$$x^i_{t-1}: \quad \phi(1-\gamma) + \theta\chi(z_a)A_1\rho - \theta A_1 = 0$$
$$\left(\sigma^i_{t-1}\right)^2: \quad \theta\chi(z_a)A_2 v - \theta A_2 + \frac{1}{2}\theta^2\left(1-\frac{1}{\psi}\right)^2(1-\alpha_\eta) + \frac{1}{2}\theta^2\chi(z_a)^2A_1^2\varphi_e^2(1-\alpha_e) = 0$$

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

$$A_1 = \phi\left(1-\rho\right)^{-1}\left(1-\frac{1}{\psi}\right)$$
$$A_2 = \frac{1}{2}(1-\gamma)\left(1-\frac{1}{\psi}\right)\frac{(1-\alpha_\eta) + (1-\alpha_e)\phi^2\left(\frac{\varphi_e}{1-\rho}\right)^2}{1-v}$$

With the coefficients, I can now return to Equation (B.5) to simplify it. I group together all constants as K_a.

$$\begin{aligned} r^i_{a,t} =& K_a + \phi\frac{1}{\psi}x^i_{t-1} - \left(\sigma^i_{t-1}\right)^2 A_2\left(1-v\right) + A_1\varphi_e\left(\sqrt{\alpha_e}\sigma(1+\beta^i_e)e^z_t + \sqrt{1-\alpha_e}\sigma^i_{t-1}e^i_t\right) \\ &+ A_2\sigma_w\left(\sqrt{\alpha_w}\left(1+\beta^i_w\right)w^z_t + \sqrt{1-\alpha_w}w^i_t\right) + \sqrt{\alpha_\eta}\sigma\left(1+\beta^i_\eta\right)\eta^z_t + \sqrt{1-\alpha_\eta}\sigma^i_{t-1}\eta^i_t \end{aligned} \tag{B.6}$$

Stochastic Discount Factor

With the expression for the consumption asset, I return to Equation (B.3) to generate an expression for the stochastic discount factor. That expression is written below:

$$m^i_t = \theta\log\delta - \frac{\theta}{\psi}\left(c^i_t - c^i_{t-1}\right) + (\theta-1)r^i_{a,t}$$

I plug Equation (B.6) (along with the dynamics for consumption) to get an expression relating the stochastic discount factor to underlying shocks:

$$
\begin{aligned}
m_t =& \theta \log \delta - \frac{\theta}{\psi}\left(\mu + \phi x_{t-1}^i + \left(\sqrt{\alpha_\eta}\sigma\left(1+\beta_\eta^i\right)\eta_t^z + \sqrt{1-\alpha_\eta}\sigma_{t-1}^i \eta_t^i\right)\right) \\
&+ (\theta - 1)\left(K_a + \phi\frac{1}{\psi}x_{t-1}^i - \left(\sigma_{t-1}^i\right)^2 A_2(1-v) + A_1\varphi_e\left(\sqrt{\alpha_e}\sigma(1+\beta_e^i)e_t^z + \sqrt{1-\alpha_e}\sigma_{t-1}^i e_t^i\right)\right. \\
&\left. + A_2\sigma_w\left(\sqrt{\alpha_w}\left(1+\beta_w^i\right)w_t^z + \sqrt{1-\alpha_w}w_t^i\right) + \sqrt{\alpha_\eta}\sigma\left(1+\beta_\eta^i\right)\eta_t^z + \sqrt{1-\alpha_\eta}\sigma_{t-1}^i\eta_t^i\right)
\end{aligned}
$$

This expression can be simplified, as follows.

$$
\begin{aligned}
m_t^i =& K_m - \phi\frac{1}{\psi}x_{t-1}^i + (\gamma - 1/\psi)(1-\gamma)K_0\left(\sigma_{t-1}^i\right)^2 \\
&-(1-\rho)^{-1}(\gamma-1/\psi)\phi\varphi_e\left(\sqrt{\alpha_e}\sigma(1+\beta_e^i)e_t^z + \sqrt{1-\alpha_e}\sigma_{t-1}^i e_t^i\right) \\
&-(1-v)^{-1}(\gamma-1/\psi)(1-\gamma)K_0\sigma_w\left(\sqrt{\alpha_w}(1+\beta_w^i)w_t^z + \sqrt{1-\alpha_w}w_t^i\right) \\
&-\gamma\left(\sqrt{\alpha_\eta}\sigma(1+\beta_\eta^i)\eta_t^z + \sqrt{1-\alpha_\eta}\sigma_{t-1}^i\eta_t^i\right)
\end{aligned}
$$

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.

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

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 μ_b^i. 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_t^i and $\left(\sigma_t^i\right)^2$:

$$
r_{b,t}^i \approx \kappa_{b,0} + \chi(z_b)z_{b,t}^i - z_{b,t-1}^i + \mu_b^i
$$

$$
z_{b,t}^i = B_0 + B_1 x_t^i + B_2\left(\sigma_t^i\right)^2
$$

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

$$r^i_{b,t} \approx \kappa_{b,0} + \chi(z_b)\left(B_0 + B_1\left(\rho x^i_{t-1} + \varphi_e\left(\sqrt{\alpha_e}\sigma\left(1+\beta^i_e\right)e^z_t + \sqrt{1-\alpha_e}\sigma^i_{t-1}e^i_t\right)\right)\right.$$
$$\left. + B_2\left(\sigma^2 + v\left(\left(\sigma^i_{t-1}\right)^2 - \sigma^2\right) + \sigma_w\left(\sqrt{\alpha_w}\left(1+\beta^i_w\right)w^z_t + \sqrt{1-\alpha_w}w^i_t\right)\right)\right)$$
$$-\left(B_0 + B_1 x^i_{t-1} + B_2\left(\sigma^i_{t-1}\right)^2\right) + \mu^i_b$$

As before, I use the Euler equation, Equation (B.2), to identify the coefficients B_1 and B_2:

$$\mathbb{E}_{t-1}m^i_t + \mathbb{E}_{t-1}r^i_{b,t} + \frac{1}{2}\mathbb{V}_{t-1}\left(m^i_t + r^i_{b,t}\right) = 0$$

which expands to the following:

$$\kappa_{b,0} + \chi(z_b)\left(B_0 + B_1\rho x^i_{t-1} + B_2\left(\sigma^2 + v\left(\left(\sigma^i_{t-1}\right)^2 - \sigma^2\right)\right)\right) - \left(B_0 + B_1 x^i_{t-1} + B_2\left(\sigma^i_{t-1}\right)^2\right) + \mu^i_b + K_m$$
$$-\frac{\phi}{\psi}x_{t-1} + (\gamma - 1/\psi)(1-\gamma)K_0\left(\sigma^i_{t-1}\right)^2 + \frac{1}{2}\left(\alpha_e\left(\chi(z_b)B_1 - (1-\rho)^{-1}(\gamma-1/\psi)\phi\right)^2\varphi^2_e\sigma^2(1+\beta^i_e)^2\right.$$
$$+(1-\alpha_e)\left(\chi(z_b)B_1 - (1-\rho)^{-1}(\gamma-1/\psi)\phi\right)^2\varphi^2_e\left(\sigma^i_{t-1}\right)^2 + \alpha_\eta\gamma^2\sigma^2(1+\beta^i_\eta)^2 + (1-\alpha_\eta)\gamma^2\left(\sigma^i_{t-1}\right)^2$$
$$+\alpha_w\left(\chi(z_b)B_2 - (1-v)^{-1}(\gamma-1/\psi)(1-\gamma)K_0\right)^2\sigma^2_w(1+\beta^i_w)^2$$
$$\left.+(1-\alpha_w)\left(\chi(z_b)B_2 - (1-v)^{-1}(\gamma-1/\psi)(1-\gamma)K_0\right)^2\sigma^2_w\right)\Bigg) = 0$$

As before, since this must hold regardless of x^i_{t-1} and $\left(\sigma^i_{t-1}\right)^2$, I group the respective coefficients and equate them to zero. As before, I also impose $\chi(z_b) = 1$:

$$x^i_{t-1}: \quad \chi(z_b)B_1\rho - B_1 - \phi\frac{1}{\psi} = 0 \quad \Rightarrow \quad B_1 = -\phi\left(1-\rho\right)^{-1}\frac{1}{\psi}$$

$$\left(\sigma^i_{t-1}\right)^2: \quad \chi(z_b)B_2 v - B_2 + (\gamma - 1/\psi)(1-\gamma)K_0$$
$$+\frac{1}{2}(1-\alpha_e)\left(\chi(z_b)B_1 - (1-\rho)^{-1}(\gamma-1/\psi)\phi\right)^2\varphi^2_e + \frac{1}{2}(1-\alpha_\eta)\gamma^2 = 0$$

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

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).

$$\begin{aligned}
r_{b,t}^i =& K_b + \frac{\phi}{\psi} x_{t-1}^i - (\gamma - 1/\psi + \gamma/\psi) K_0 \left(\sigma_{t-1}^i\right)^2 \\
&- (1-\rho)^{-1}(1/\psi)\phi\varphi_e \left(\sqrt{\alpha_e}\sigma(1+\beta_e^i)e_t^z + \sqrt{1-\alpha_e}\sigma_{t-1}^i e_t^i\right) \\
&- (1-v)^{-1}(1/\psi - \gamma - \gamma/\psi)K_0\sigma_w \left(\sqrt{\alpha_w}(1+\beta_w^i)w_t^z + \sqrt{1-\alpha_w}w_t^i\right)
\end{aligned}$$