Monetary Spillovers in Financial Markets: Policymakers and Premia

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

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1 Introduction

Deepening international linkages between markets have intensified global spillovers of the Federal Reserve’s monetary policy into foreign financial markets, as noted by Rey [2013]. Debates over spillovers 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 establish a novel fact on how currencies and bonds react asymmetrically to the Fed’s announcements, and use it to test different channels of spillovers. 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. These two forms of heterogeneity in how countries receive the Fed’s spillovers, while suggestive when each is studied in isolation, are potent when studied together.

I divide explanations for monetary spillovers into three exhaustive 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 of complete markets, as the majority of international finance models embed complete markets. 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 incomplete markets. I expand the fact to add one new empirical condition for successful models in this class to match. Specifically, I document that the
monetary policies of most other countries do not spill over into global markets. In addition to incorporating heterogeneity in how countries receive Fed spillovers, successful models of spillovers must incorporate heterogeneity in whether central banks generate spillovers.

To illustrate the paper’s argument, 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. 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. 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, and any other taxonomy of the channels of Fed spillovers into foreign currency and bond markets can be mapped into these three. I use these asymmetric asset responses and other evidence to argue against the first two hypotheses, in favor of the third hypothesis.

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

(a) Ten-Year Bonds

(b) Currencies

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

As in this example, I first establish the key empirical fact on 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 histori-
cally (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 illustrates these patterns, as the dollar depreciates more against
the Australian dollar, and Australian yields fall more.

This fact owes its causal interpretation and its precision to four methodological components:
high-frequency returns around Fed announcements, long-maturity instruments, non-announcement
windows, and inferred monetary shocks. The methodology compares currency and long-maturity
bond returns in sixty-minute and daily windows around Fed announcements to returns in windows
outside Fed announcements, and estimates how assets react to latent monetary shocks from the
differences. First, the combination of high frequency returns in announcement windows and returns
in non-announcement windows allows me to isolate the effects of monetary policy. Market variation
driven by monetary policy can be separated from market variation driven by other forms of news and
from idiosyncratic market noise. Next, long-maturity bonds and currencies allows me to capture
expected responses 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 miss them altogether.² I embed these
components within a latent factor model.

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 against these hypotheses, consider the two possible scenarios in Figure 1.
The first scenario is that the Reserve Bank of Australia is expected to ease more than the Bank
of Japan. This explanation predicts that Australian yields should fall more than Japanese yields
and that the Australian dollar should depreciate against the yen. While the observed responses

¹This is not driven by the zero lower bound for interest rates, as I discuss in Section 4.
²As 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.
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 the majority of theories use models with complete markets rather than incomplete markets, as surveyed by Engel [2014]. Again, I illustrate my argument against these hypotheses by considering the two possible scenarios in Figure 1. 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 in complete markets reflect the ratio of stochastic discount factors. While the observed responses of bond yields support this explanation, the observed responses of currencies do not. The second scenario is that the stochastic discount factor of Japanese investors temporarily rose less than that of Australian investors, and this now violates the observed responses of bond yields. More generally, I show that no hypothesized set of risk premia shifts across the nine countries can be consistent with asymmetries in both currency and bond markets concurrently when using models of complete markets, 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. In models with complete markets, currencies respond to both shocks, while long-maturity bonds only respond to temporary shocks. Introducing two sources of hetero-
geneity 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 of such a specification using some commonly used models. The findings suggest that models of complete markets 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. Figure 1 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 market frictions, and future research is needed to evaluate them.

To guide future research on models in this class, I expand my empirical fact to offer a new restriction, on heterogeneity in whether other countries’ monetary policies spill into markets. 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, which is in itself novel to the literature both in scope and in statistical power, requires that successful models go beyond incorporating asymmetries in the reactions of assets. They must incorporate asymmetries among central banks.

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 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%.

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 are ones involving borrowing constraints as in Bruno and Shin [2017] or segmented markets as in Gabaix and Maggiori [2015]. My paper offers evidence from a well-identified setting that financial frictions explain asset returns better than models of frictionless risk-sharing. This informs the debate between complete and incomplete markets frameworks in explaining the overall carry trade.

The paper proceeds as follows. Section 2 reviews the literature on empirical spillover patterns. Section 3 discusses the empirical framework and the data. Section 4 introduces the main fact on asymmetries in currency and bond markets, and uses these asymmetries jointly to argue against two main channels of spillovers: central banks reacting to the Fed and risk premia shifting under complete markets. Section 5 uses the term structures of foreign bond yields to provide further evidence that central banks do not react to the Fed. Section 6 argues further that models of complete markets do not explain spillovers by showing the formal modeling tensions. Section 7 discusses models with incomplete markets, and demonstrates the divergence between spillovers emanating from the Fed and from other central banks. Section 8 concludes.

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.
2 Review of Literature on Spillovers

Beyond contributing to the two more theoretical debates discussed in Section 1, I contribute to
the much broader body of literature on documenting empirical patterns in spillovers. Papers here
have found Fed spillovers in every conceivable asset. The most comprehensive papers include Rey
[2015] and Miranda-Agrippino and Rey [2015], which look at the Fed’s effects on a wide range
of markets. However, there are many papers that examine more specific markets. To review a
handful: Brusa et al. [2017] study equity markets, Fratzscher 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, Binici and Hutchison [2016] for instance find that spillovers
are muted when the recipient country has strong fundamentals. Hausman and Wongswan [2011],
Miyajima et al. [2014], Eichengreen and Gupta [2015], and Aizenman, Chinn and Ito [2016] find
that spillovers are stronger when recipient countries are more financially integrated with the US.

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, Chinn and Ito [2016] briefly study spillovers from the Eurozone, Japan,
and the UK.

3 Empirical Framework and Data

This section introduces the empirical framework and the data used to identify and characterize 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

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4 The one exception that Rey [2015] and Miranda-Agrippino and Rey [2015] find is foreign direct investment.
5 However, Dedola et al. [2017] find no consistent macroeconomic indicators that explain heterogeneity.
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 an-
nouncements, long-maturity assets, non-announcement windows, and inferred (i.e. latent) mon-
etary 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 in-
terpretations, and they provide power. My windows are predominantly sixty-minute windows, al-
though 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 win-
dows 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.
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.

3.1.1 Equation

Equation (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 ($\alpha$), the product of a univariate or multivariate vector of coefficients ($\beta$) and a univariate monetary shock ($m_t$), and a univariate or multivariate error $\epsilon_t$. I primarily test whether assets react differentially to monetary shocks, i.e. test for equality between different elements of the multivariate vector $\beta$. In Section 5, I also test for whether assets react to monetary shocks in the first place ($\beta = 0$).

$$r_t = \alpha + \beta m_t + \epsilon_t$$  \hspace{1cm} (1)

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

In Equation (1), the only observed data are the high-frequency and long-maturity asset returns $r_t$ during announcement windows. However, I let the errors $\epsilon_t$ take on a distribution 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.

3.1.2 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.\footnote{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.}

\footnote{Currency markets are liquid and traded around-the-clock, but bond markets in smaller countries often have poor liquidity or limited hours.}
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. For instance, idiosyncratic noise in currency markets starts to overwhelm monetary shocks in windows beyond twelve to sixteen hours. Bond markets are approximately half as noisy as currency markets, and so daily windows work when necessary.

3.1.3 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 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, 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. In Equation (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). In addition, currencies are driven by the infinite-horizon exchange rate, although I argue both in Section 4.2 and in Appendix C that this does not react to Fed announcements.

\[
\Delta s_j^t/\$ = \sum_{k=1}^{\infty} \Delta i_t^j + \sum_{k=1}^{\infty} \Delta i_t^j + \sum_{k=1}^{\infty} \Delta p_t^j/\$ + \Delta s_j^{\infty}/\$ \tag{2}
\]

In Equation (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

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8Exchange 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.

9Anderson 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.
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).

\[ 10 \Delta y^j_{tt+10} = \sum_{k=1}^{10} \Delta \hat{v}^j_{t+k-1} + \sum_{k=1}^{10} \Delta \gamma^j_{t+k-1} \]  

(3)

Note that throughout this paper, the operator \( \Delta \) 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 bonds
are effectively martingales, but it is useful for the formal models in Section 6. In this paper,
shocks are realized at time \( t \).

\[ \Delta x \equiv E_t x - E_{t-1} x \quad (\text{not } x_t - x_{t-1}) \]  

(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. Given the general noise in financial markets, ascribing movements in interest
rates to announcements several years ago is difficult. Moreover, I estimate a vector autoregression
model in Appendix B following Wright [2012]. This maps closely to my empirical framework on
all other dimensions, as it incorporates non-announcement windows, high-frequency returns, and
inferred shocks, but it utilizes short-maturity rather than long-maturity instruments. I similarly
find no reactions at conventional significance levels, but I take this as evidence of low power rather
than as evidence of the channels behind Fed spillovers.

3.1.4 Non-Announcement Windows

Non-announcement windows, which are used to parameterize error term \( \epsilon_t \) in Equation (1), serve
as reference points for announcement windows and are essential for ensuring that my results do
not ascribe background noise during announcement windows to the Fed’s effects. I identify non-
announcement windows in the same way that I identify my announcement windows, as a combina-
tion 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

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10 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.

11 However, my empirical specifications are exact rather than approximate. For instance, although I observe
\( E_{t-1} x_t - E_{t-1} x_{t-1} \) for empirical specifications, the constant in the specifications formally controls for the (typically negligible)
expected change in the response variable \( E_{t-1} x_t - E_{t-1} x_{t-1} \), leaving \( \Delta x_t \) to be explained by Fed shocks.

12 Formally, time periods are annual and prior expectations of shocks, i.e. \( E_{t-1} x_t \), are assumed to settle just prior
to the Fed’s announcement.
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 2, exactly one week prior to the introduction’s example announcement.

Figure 2: Currency Movements over Announcement and Non-Announcement Windows

(a) Currencies on January 25 (Fed Shock)  (b) Currencies on January 18 (No Event)

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.

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 studies do, it identifies from excess variation over announcement windows. 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.

3.1.5 Inferred Shocks

Inferred shocks or latent shocks, which are represented by $m_t$ in Equation (1), are important for precise and comprehensive identification. Since $m_t$ is estimated, this turns Equation (1) into a

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13 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.
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 emanating from smaller central banks. Conceptually, my specification with inferred shocks closely resembles one with interactive fixed effects.\footnote{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.}

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). 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]).\footnote{However, my inferred shocks correlate reasonably with existing measures of monetary shocks including these, evidence from surveys, and Romer and Romer [2004] shocks, as documented in Appendix B.} Such approaches lose some of the common surprise. 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 reveal information about future rates at horizons longer than two years, especially in the last decade. 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 them 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. Those results are reported in Appendix B.

Second, inferred shocks do not require additional data beyond the response variables $r_t$ in Equation (1), and so they can be constructed at an intraday frequency for all central banks, which I do in Section 7.2. 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 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.

3.2 Methodology

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

3.2.1 Estimation Procedure

The method explains the excess variation in announcement windows over non-announcement windows as a combination of time-varying shocks and asset-varying coefficients. To illustrate the estimation procedure 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) in its multivariate form:

\[
\begin{bmatrix}
\Delta s_{e/\$}^t \\
\Delta s_{\ell/\$}^t \\
\Delta s_{\gamma/\$}^t
\end{bmatrix}
= 
\begin{bmatrix}
\alpha_{e/\$} \\
\alpha_{\ell/\$} \\
\alpha_{\gamma/\$}
\end{bmatrix}
+ 
\begin{bmatrix}
\beta_{e/\$} \\
\beta_{\ell/\$} \\
\beta_{\gamma/\$}
\end{bmatrix}
\begin{bmatrix}
m_t^e \\
m_t^\ell \\
m_t^\gamma
\end{bmatrix}
+ 
\begin{bmatrix}
\epsilon_{e/\$}^t \\
\epsilon_{\ell/\$}^t \\
\epsilon_{\gamma/\$}^t
\end{bmatrix}
\tag{5}
\]

I present the Gaussian likelihood function associated with Equation (5) next. Errors \(\epsilon\) are assumed to have some covariance matrix \(\Sigma\), where \(\Sigma\) 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.\(^{16}\)

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

Since this term involves the product of estimated quantities \(m_t\) and \(\beta\), I cannot analytically solve the system of interlocking first-order conditions.\(^{17}\) 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 iter-

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\(^{16}\)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 \(\Sigma\) in this specification as fixed because of the overwhelming amount of non-announcement data available to estimate it. In Appendix B, I treat \(\Sigma\) 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.
ation. This yields estimates for parameters \((\alpha, \beta, m_t)\). I make one scaling assumption, \(\forall m_t = 1\), since I cannot identify the magnitudes of \(m_t\) and \(\beta\) separately.

In Appendix B, 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 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 \((\alpha, \beta)\) 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 adjustments with missing data. In Appendix B, I discuss alternative approaches to estimating Equation (5) and their limitations, including Identification by Heteroskedasticity by Rigobon [2003].

I next turn to the two ways I concurrently test asymmetries in \(\beta\) in Equation (1).

### 3.2.2 Average Coefficient

A natural way to test asymmetries in \(\beta\) would be a series of pairwise tests. This is reasonable if \(\beta\) has two or three elements, but it is in comprehensible in practice as \(\beta\) 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 \(\beta_e/8 > \beta_y/8\), but I cannot reject \(\beta_e/8 \neq \beta_f/8\) or \(\beta_y/8 \neq \beta_f/8\). At least one of these tests must be wrong.

Instead, I test for asymmetries in a closely related way by comparing each element of \(\beta\) to an average of the other \((n - 1)\) elements. In this example, the test for asymmetries in how the euro

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18Rohde and Cappe [2011] among many others argue that computing the marginal likelihood of continuous latent factors \(m_t\) is intractable, and instead recommend a modified expectation step that uses a variational posterior distribution for \(m_t \sim N(\mu_t, V_t)\). Specifically, I take the expectation of the log-likelihood function with respect to \(m_t | (\mu_t, V_t)\) initially, and also augment the log-likelihood function with the Kullback-Leibler divergence between the variational posterior and the prior \(N(0, 1)\). I then maximize the expression alternately with respect to parameters \((\mu_t, V_t)\) and \((\alpha, \beta)\), which maps to the original two steps. When conducting robustness checks with the shocks themselves, I set the shocks \(m_t\) to be their MAP estimates \(\mu_t\), i.e. the means of their posterior distributions. Further details can be found in Appendix B. Moreover, I check that the results are not sensitive to utilizing variational methods to solve for parameters. In Appendix B, I use MCMC methods instead to solve for parameters in the main specifications. The results, in Figure 17, indicate that both approaches yield the same results.

19Appendix B is more exhaustive, but the main limitation of Identification by Heteroskedasticity is that its solution algorithm, GMM, is not guaranteed to converge in high-dimensional spaces whereas the EM algorithm is. I do successfully implement it when utilizing currency data in Equation (1) and find very similar results to those generated by the EM algorithm. I find convergence to be an issue when utilizing bonds data.

20Regardless, I show the p-values for all thirty-six pairwise tests in Appendix F for my main results.
reacts to Fed announcements, relative to how the yen or pound reacts, becomes:

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

### 3.2.3 Lower-Dimensional Structure

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

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

Figure 3: Currency Reactions to the Fed Easing, January 25, 2012

![Figure 3: Currency Reactions to the Fed Easing, January 25, 2012](image)

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) 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 \( \beta^{\text{L}/\text{S}} \) are much closer to \( \beta^{\text{Y}/\text{S}} \) than to \( \beta^{\text{E}/\text{S}} \), one possible structure that may emerge in

\(^{21}\)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.
Equation (5) is as follows:

\[
\begin{bmatrix}
\Delta s_t^e/s \\
\Delta s_t^e/\£ \\
\Delta s_t^e/Y \\
\end{bmatrix}
= 
\begin{bmatrix}
1 & 0 & 0 & 1 \\
0 & 1 & 0 & 1 \\
0 & 1 & 0 & 1 \\
\end{bmatrix}
\begin{bmatrix}
\alpha/e/s \\
\alpha(e, \£)/s \\
\alpha(e, Y)/s \\
\end{bmatrix}
+ 
\begin{bmatrix}
1 & 0 & 0 & 1 \\
0 & 1 & 0 & 1 \\
0 & 1 & 0 & 1 \\
\end{bmatrix}
\begin{bmatrix}
\beta/e/s \\
\beta(e, \£)/s \\
\beta(e, Y)/s \\
\end{bmatrix}
m_t^s + 
\begin{bmatrix}
\epsilon_t^e/s \\
\epsilon_t^e/\£ \\
\epsilon_t^e/Y \\
\end{bmatrix}
\]

Proposed Lower-Dimensional Structure

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

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.

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.

3.3.1 Monetary Announcements

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

Nakamura and Steinsson [2017] note that monetary announcements may actually be informational announcements, releasing the Fed’s private information about fundamentals. I continue 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 7.2. Collection details can be found in the appendix.

### 3.3.2 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.\(^{22}\)

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.\(^ {23}\) For countries without bonds on a liquid futures exchange, I use an intraday benchmark rate published by Thomson Reuters based on reported transactions. However, some of these returns are illiquid around Fed announcements.

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

In addition to removing observations that risk overlapping with inflation, unemployment, or

\(^{22}\) 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.

\(^{23}\) Examples include the Sydney Futures Exchange for Australian data, the Chicago Mercantile Exchange for American data, the Eurex Exchange for Swiss and German data, the London International Financial Futures and Options Exchange for British data, the Osaka Securities Exchange for Japanese data, etc.
foreign monetary releases, I also prune extreme observations, defined as returns over announcement windows whose distance from mean announcement returns exceed the most extreme 1% threshold. I specifically measure Mahalanobis distance, which generalizes Euclidean distance to multivariate and correlated data.

4 Currency and Bond Asymmetries

I examine asymmetries in currency markets and in bond markets following Fed announcements to show that monetary spillovers from the Fed are neither consistent with shifts in the paths of foreign central banks, nor consistent with transitory shifts in foreign risk premia under complete markets. First, I introduce a new fact 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.

This section has two contributions. First, I identify asymmetries in the responses of currencies and ten-year sovereign bonds to Fed announcements, using high-frequency currency and bond returns. This is a new fact to the literature in its own right, and shows that spillovers are heterogeneous across countries. Asymmetric responses across countries can be subtle, and common methods in the literature that utilize noisier shocks or lower-frequency returns 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 7.2, I apply this methodology to other central banks and find that most central banks, with the exception of the European Central Bank, do not generate asymmetric spillovers. In Appendix F, I show that this finding is robust to different time periods (pre-crisis and post-crisis) and different states (recessionary and expansionary).

The second major contribution of this section is to show that my fact is inconsistent with two major classes of explanations: central banks reacting to the Fed, and risk premia shifting in a complete markets framework. For both of these explanations, results from currency and bond markets suggest opposing stories 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 in a complete markets framework, 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 discount factors of high-rate countries rise.

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 of complete markets. Therefore, if the Fed causes foreign central
banks to adjust their paths of policy rates or if the Fed triggers shifts in foreign risk premia, these
portfolios should be 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.

4.1 Baseline Fact: Asymmetric Asset Responses

Before testing the classes of explanations for spillovers, I first present the baseline fact: Fed an-
nouncements affect currencies and bonds of different countries differentially. This shows that mon-
etary policy spills over into other countries heterogeneously. In currencies, when the Fed tightens
(eases), the dollar appreciates (depreciates) by most against high-rate currencies like the Australian
dollar and least against low-rate currencies like the Japanese yen. In bonds, when the Fed tightens
(eases), ten-year bond yields of high-rate countries rise (fall) more than yields of low-rate countries.

I utilize the Expectation Maximization algorithm to fit the multivariate version of Equation
(1), with country-specific coefficients \((\alpha, \beta)\) 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.

\[
\begin{bmatrix}
\Delta s_{t\epsilon}^\$/S \\
\Delta s_{t\£}^\$/S \\
\Delta s_{tY}^\$/S
\end{bmatrix} = \begin{bmatrix}
\alpha_{\epsilon/\$} \\
\alpha_{\£/\$} \\
\alpha_{Y/\$}
\end{bmatrix} + \begin{bmatrix}
\beta_{\epsilon/\$} \\
\beta_{\£/\$} \\
\beta_{Y/\$}
\end{bmatrix} m_t^\$ + \begin{bmatrix}
\epsilon_{\epsilon/\$} \\
\epsilon_{\£/\$} \\
\epsilon_{Y/\$}
\end{bmatrix}
\]

I estimate Equation (1) separately for currencies and for bonds in this section. In the next
section, I estimate them jointly.

Consider currencies first. I plot the coefficients \(\beta\) in Figure 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 \(\beta\) in Figure 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 2 and (3). Since the US components are common across curren-
cies, differential appreciation and depreciation of the dollar against different currencies points to
Figure 4: Currency Responses to US Monetary Shocks

Notes: The figure depicts by how much the dollar appreciates against a given reference currency when it appreciates by 1% on average, following a Fed tightening. Standard error bars are computed 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.

differential movements by foreign central banks or in currency premia, either under complete or incomplete markets. Differential responses in yields similarly point to differential movements by foreign central banks or in term premia.

\[
\Delta s_j^t/\$ = \sum_{k=1}^{\infty} \Delta i^t_{t+k-1} - \sum_{k=1}^{\infty} \Delta i^t_{t+k-1} + \sum_{k=1}^{\infty} \Delta p^j_{t+k-1} + \Delta s_{CE}^j/\$
\]

\[
10 \Delta y^j_t(t, t+10) = \sum_{k=1}^{10} \Delta i^j_{t+k-1} + \sum_{k=1}^{10} \Delta \gamma^j_{t+k-1}
\]

The Fed’s asymmetric spillovers into currency and bond markets are special among central banks. In Section 7.2, I discuss spillovers from all ten central banks in more detail, but I consider asymmetric spillovers from the Reserve Bank of New Zealand (RBNZ) as an example here. I show that the Reserve Bank of New Zealand has no asymmetric effects on either currencies or bonds in Figure 6, with the slight exception of Australian assets.

Although the empirical specifications are computed on a country-by-country basis, I confirm

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24I address differential movement in the infinite-horizon exchange rate in Appendix C.
that the cross-sectional sorting of asymmetries by countries maps closely to the level of interest rates in those countries. In Figures 4 and 5, the canonical low-rate countries of Japan and Switzerland are found together, while the canonical high-rate countries of Australia, New Zealand, and Norway are found together. I test this formally by interacting the specification in Equation 1 with each country’s pre-shock interest rate spread relative to the US, and by removing country-specific parameters. Equation 6 shows an illustration of the revised equation. I test whether $\beta_1$ in this model is significant, i.e. whether interest rates help predict variation across countries and across time in currency and bond returns. I find that 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 D, I show that this result is robust to dropping any country from the specification.

$$\begin{bmatrix} \Delta s_t^{e/s} \\ \Delta s_t^{e/£} \\ \Delta s_t^{e/¥} \end{bmatrix} = \begin{bmatrix} 1 & i_{t-1}^e - i_{t-1}^s \\ 1 & i_{t-1}^e - i_{t-1}^£ \\ 1 & i_{t-1}^e - i_{t-1}^¥ \end{bmatrix} \begin{bmatrix} \alpha_0 \\ \alpha_1 \end{bmatrix} + \begin{bmatrix} 1 & i_{t-1}^e - i_{t-1}^s \\ 1 & i_{t-1}^e - i_{t-1}^£ \\ 1 & i_{t-1}^e - i_{t-1}^¥ \end{bmatrix} \begin{bmatrix} \beta_0 \\ \beta_1 \end{bmatrix} m_t + \begin{bmatrix} \epsilon_t^{e/s} \\ \epsilon_t^{e/£} \\ \epsilon_t^{e/¥} \end{bmatrix}$$

Interest rates are one of several significant predictors, and I test others in Appendix D: local measures of equity and currency volatility, the cross-currency basis, trade flows, shares of trade flows invoiced in dollars, cross-border bank positions, cross-border portfolio debt positions, cross-
Figure 6: Market Reactions to NZ Monetary Shocks

(a) Currency Responses

(b) Bond Responses

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

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 7, 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.

4.2 Test: Cross-Border Bond Portfolio

While the baseline fact provides suggestive evidence, I formally test and show in this section that Fed spillovers are explained neither by central banks reacting to the Fed nor by risk premia temporarily shifting under complete markets. (I focus on transient shifts in risk premia here, and relax
the assumption in Section 6.\textsuperscript{25}) Conceptually, I exploit the tension between the cross-sectional sorting of countries in currency markets and in bond markets. Methodologically, I construct a portfolio for each country that has no direct exposure to its central bank or to transitory shifts in its risk premia under complete markets. Therefore under these two channels, these portfolios should be symmetrically exposed to the Fed across countries, e.g. the Japanese and Australian portfolios should react indistinguishably to Fed announcements. Empirically, I find that the portfolios continue to have strongly asymmetric exposures to the Fed. This asymmetry suggests that fluctuations in risk premia under incomplete markets best explain spillovers.

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) to Equation (3), such that portfolio returns are expressed in Equation (7). 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 Fed at long horizons.\textsuperscript{26} 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

\textsuperscript{25}The assumption on transitory shocks is viewed differently across fields. In much of the macroeconomics literature, 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 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.

\textsuperscript{26}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.
given sixty-minute window around announcements.27

\[
\frac{\Delta r^j_t}{S} = \frac{\Delta s^j_t}{S} + 10\Delta y^j_t(t, t + 10)
\]

Portfolio Return  Currency Exposure  Foreign Bond Exposure

\[\approx \sum_{k=1}^{\infty} \Delta v^S_{t+k-1} + \sum_{k=1}^{\infty} \Delta p^j_{t+k-1} + \sum_{k=1}^{10} \Delta \gamma^j_{t+k-1}
\]

US Path of Rates  Currency Premia  10Y Term Premia

It is immediately apparent that portfolio returns in Equation (7) 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 under complete markets. Although it is not immediately apparent in Equation (7), 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 (8) and (9) respectively. Under complete markets, 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 derivation in Section 6.28

\[
\frac{\Delta s^j_t}{S} = \Delta m^S_t - \Delta m^j_t
\]

Exchange Rate  US SDF  Foreign SDF

\[10 \Delta y^j_t(t, t + 10) = \Delta m^j_t - \left(\Delta \log \Lambda^j_{t+10} + \Delta L_t \left(\Lambda^j_{t+10}\right)\right)
\]

10Y Yield  SDF  10Y Pricing Kernel

As such, I take the same two steps to construct the portfolio in this revised framework. First, I again add Equation (8) to Equation (9). Second, I assume that properties of the pricing kernel at long horizons do not react differentially to the Fed across countries, which is equivalent to assuming

27I 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 C.

28Equations (2) and (8) are equivalent representations of exchange rates, and Equations (3) and (9) are equivalent representations of bonds. For instance, in Equation (8), the actual realizations of the stochastic discount factors embed all future changes in expectations of rates and risk premia in Equation (2).
that shocks are transitory.\footnote{An alternate concern is that shocks are transitory but have not decayed away in ten years. I show in Appendix F that my results are robust when using thirty-year bonds.}

\[
\begin{align*}
\Delta r^j_t / S & = \Delta s^j_t / S + 10 \Delta y^j_t (t, t + 10) \\
\approx \Delta m^j_t \\
\text{Portfolio Return} & \quad \text{Currency Exposure} & \quad \text{Foreign Bond Exposure} & \quad \text{US SDF}
\end{align*}
\] (10)

I illustrate with the example of power utility and a stationary autoregressive log consumption process (with correlation $\rho$) to make Equations (8) and (9) 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, the foreign consumption basket becomes more dear and its relative price rises. 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.

\[
\begin{align*}
\Delta s^j_t / S & = \left( -\gamma^j \Delta c^j_t \right) - \left( -\gamma^j \Delta c^j_t \right) \\
\text{Exchange Rate} & \quad \text{US SDF} & \quad \text{Foreign SDF}
\end{align*}
\]

\[
10 \Delta y^j_t (t, t + 10) = \left( -\gamma^j \Delta c^j_t \right) + \gamma (\rho^j)^{10} \sigma^j \epsilon^j_t \approx \left( -\gamma^j \Delta c^j_t \right)
\]

10Y Yield \quad SDF \quad 10Y Pricing Kernel

Thus, I have two equivalent representations of this portfolio, and they show the portfolio is not directly exposed to two channels of spillovers. Equation (7) makes clear that these portfolios do not depend directly on foreign monetary policy. Equation (10) makes clear that these portfolios do not depend on foreign stochastic discount factors under a complete markets framework. In short, I should not expect to find any heterogeneity in these portfolios across countries under either explanation 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. Equation (5) becomes in this context:

\[
\begin{bmatrix}
\Delta s^e_t / S + 10 \Delta y^e_t (t, t + 10) \\
\Delta s^L_t / S + 10 \Delta y^L_t (t, t + 10) \\
\Delta s^Y_t / S + 10 \Delta y^Y_t (t, t + 10)
\end{bmatrix} =
\begin{bmatrix}
\alpha^e_t / S \\
\alpha^L_t / S \\
\alpha^Y_t / S
\end{bmatrix} +
\begin{bmatrix}
\beta^e_t / S \\
\beta^L_t / S \\
\beta^Y_t / S
\end{bmatrix} m^S_t +
\begin{bmatrix}
\epsilon^e_t / S \\
\epsilon^L_t / S \\
\epsilon^Y_t / S
\end{bmatrix}
\]

Long-Maturity Portfolio Returns

I plot the coefficients $\beta$ in Figure 7, and 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 7 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 an incomplete markets framework.

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

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

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 7 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 8. Interestingly, the point estimates for the long-maturity portfolio become more asymmetric compared to the short-maturity portfolio.\footnote{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 8 illustrates the argument that asymmetries deepen when currencies and bonds are
Figure 8: Cross-Border Portfolio Responses to US Monetary Shocks

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

Notes: The figures compare the reactions of two types of portfolios to Fed announcements. The left figure shows by how much a portfolio that shorts a given country’s risk-free bond and lends at the US risk-free 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 risk-free 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 4 and 7, except with common axes for comparison purposes.

linked, rather than being offset. 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. 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.

5 Evidence from Bond Term Structures

I offer further evidence from each country’s foreign bond term structure individually, to show that its central bank does not react to the Fed. Section 4 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, or explanations in which central banks partially react to the Fed. This section addresses those concerns by offering evidence for each recipient country on its own, to show that the country’s risk premium rather than its path of short rates reacts to Fed announcements, using data from a rich term structure of bond yields. Since the 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 a country’s yield curve most dominated by its short rates, i.e. short-maturity yields, do not respond to Fed announcements, while the parts most dominated by its term premia, i.e. distant forward yields, do respond. Second, I use a Gaussian affine term structure model to decompose a country’s yields explicitly into the path of short rates and term premia, and show again that term premia do react to Fed announcements while the paths of rates do not.

I take one additional step to mollify concerns that this approach lacks statistical power. Specifically, I show that the paths of short rates in foreign countries respond to announcements from their own central banks, as expected in theory. My methodology is thus sensitive enough to detect other expected responses in the paths of short rates, which should diminish worries that it is too underpowered to detect central banks responding to Fed announcements. Moreover, in doing so, I show 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 Section 7.

I answer a simpler question in this section than in Section 4, as I test whether a given bond yield reacts to the Fed at all rather than whether it reacts differently than other yields. Therefore, I use a simpler empirical methodology, called inference by heteroskedasticity. This maintains the same core empirical principles of high-frequency returns around Fed announcements, long-maturity instruments, non-announcement windows, and inferred shocks. However, it requires fewer assumptions than the previous empirical approach. I introduce this methodology first.

5.1 Methodology: 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.

I again start with Equation (1) (rewritten below) and now focus on univariate returns $r_t$ and monetary shocks $m_t$. 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. The methodology is also transparent, as it gets all its statistical power from a single moment.

$$r_t = \alpha + \beta m_t + \epsilon_t$$

As an example, suppose I want to test whether Australian bond yields react to Fed announce-
ments. I write Equation (1) and its non-announcement counterpart as follows.

Announcement Windows: $\Delta y_t^{AUD} = \alpha + \beta m_t^S + \epsilon_t$

Non-announcement Windows: $\Delta y_t^{AUD} = \epsilon_t$

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.

Announcement Windows: $\mathbb{V}_{t-1} (\Delta y_t^{AUD}) = \beta^2 \mathbb{V}_{t-1} (m_t^S) + \mathbb{V}_{t-1} (\epsilon_t)$

Non-announcement Windows: $\mathbb{V}_{t-1} (\Delta y_t^{AUD}) = \mathbb{V}_{t-1} (\epsilon_t)$

$$\mathbb{V}_{t-1} (\Delta y_t^{AUD}) > \mathbb{V}_{t-1} (\Delta y_t^{AUD}) \implies \beta \neq 0 \quad (11)$$

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. For instance, in Appendix B, I show that the Brown-Forsythe test strongly outperforms the F-test on simulated data with high kurtosis. Moreover, in Appendix F, I 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.

I illustrate this method by testing how all ten central banks affect the ten-year bond yields of all ten countries. The results, in Table 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 consistent with other results in the international finance literature in which the dollar and euro are more dominant than their peer currencies. 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 which component reacts to Fed announcements.

$$10 \Delta y_t^j (t, t + 10) = \sum_{k=1}^{10} \Delta y_t^{j,k-1} + \sum_{k=1}^{10} \Delta y_t^{j,k-1}$$

10Y Yield 10Y Path of Rates 10Y Term Premia
Table 1: Excess Volatility in 10Y Bond Returns

<table>
<thead>
<tr>
<th>Country</th>
<th>AUD</th>
<th>CAD</th>
<th>CHF</th>
<th>EUR</th>
<th>GBP</th>
<th>JPY</th>
<th>NOK</th>
<th>NZD</th>
<th>SEK</th>
<th>USD</th>
</tr>
</thead>
<tbody>
<tr>
<td>Australia</td>
<td>144%</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
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<tr>
<td>Canada</td>
<td>30%</td>
<td>82%</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
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<tr>
<td>Switzerland</td>
<td></td>
<td></td>
<td>124%</td>
<td>96%</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Euro</td>
<td>34%</td>
<td>39%</td>
<td>88%</td>
<td>37%</td>
<td>133%</td>
<td>90%</td>
<td>24%</td>
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<td></td>
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<td></td>
<td></td>
<td>79%</td>
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<td></td>
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<td></td>
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<td></td>
<td></td>
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<td></td>
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<tr>
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<td></td>
<td></td>
<td>26%</td>
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<tr>
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<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td>39%</td>
</tr>
<tr>
<td>United States</td>
<td>207%</td>
<td>144%</td>
<td>40%</td>
<td>216%</td>
<td>46%</td>
<td>54%</td>
<td>25%</td>
<td>82%</td>
<td></td>
<td>233%</td>
</tr>
</tbody>
</table>

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

5.2 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. My results are not driven by weak statistical power, as the short ends of foreign yield curves respond 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.

\[ T_0 \frac{\Delta y_f^j(t, t + T_0)}{T_0 \text{Yield}} = \left( \sum_{k=1}^{T_0} \Delta \gamma_{t+k-1}^j + \frac{T_0}{T_0} \sum_{k=1}^{T_0} \Delta \gamma_{t+k-1}^j \right) \approx \sum_{k=1}^{T_0} \Delta \gamma_{t+k-1}^j \]

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 C, I argue that long-run inflation forecasts are extremely stable over time.

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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.

\[(10 - T_1) \Delta y_j(T_1, t + 10) \approx (10 - T_1) \sum_{k=1}^{10-T_1} \Delta \gamma_j^{T_1+k-1} + \sum_{k=1}^{10-T_1} \Delta \gamma_j^{T_1+k-1} \]

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 2017 to borrow and lend between 2023 to 2027).

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 and 3.

The findings are stark. Table 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, suggesting that weak statistical power does not explain my lack of findings for the Fed. Table 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, rather than the policies of central banks, adjust to announcements by the Fed in each individual country.

The results are robust to different cutoffs. I illustrate with one extreme example: ten-year forward twenty-year rates, or rates that can be locked into in 2017 for borrowing and lending between 2027 and 2047. It is implausible that central banks regularly release guidance at such horizons.

Table 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 F. This confirms that Fed announcements induce strong shifts in term premia for each country.
## Table 2: Excess Volatility in Daily 1Y Bond Returns

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

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

## Table 3: Excess Volatility in Daily 6F4Y Bond Returns

<table>
<thead>
<tr>
<th></th>
<th>AUD</th>
<th>CAD</th>
<th>CHF</th>
<th>EUR</th>
<th>GBP</th>
<th>JPY</th>
<th>NOK</th>
<th>NZD</th>
<th>SEK</th>
<th>USD</th>
</tr>
</thead>
<tbody>
<tr>
<td>Australia</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td>87%</td>
<td>28%</td>
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<tr>
<td>Canada</td>
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<td>Switzerland</td>
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<tr>
<td>Euro</td>
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<td></td>
<td></td>
<td>19%</td>
<td>32%</td>
<td></td>
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<tr>
<td>United Kingdom</td>
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<td>Japan</td>
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<td>Norway</td>
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<td></td>
<td>30%</td>
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<tr>
<td>New Zealand</td>
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<td>Sweden</td>
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<td></td>
<td></td>
</tr>
<tr>
<td>United States</td>
<td></td>
<td>38%</td>
<td>36%</td>
<td>30%</td>
<td>50%</td>
<td>29%</td>
<td></td>
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</tbody>
</table>

Notes: The table tests whether the six-year forward four-year bond of the column country is more volatile around announcements by the row central bank than at other times, using daily returns. If returns are more volatile at the 1% level using the Brown-Forsythe test, the cell records the excess ratio of standard deviations (announcement window to non-announcement window standard deviations, minus 100%). If returns are not statistically more volatile, the cell is left blank. The full table, including ratios significant at the 5% level and insignificant ratios, can be found in the Appendix as Table 14. The Fed affects most other countries’ six-year forward four-year bonds.
### Table 4: Excess Volatility in Daily 10F20Y Bond Returns

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

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

### 5.3 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. As before, the paths of rates respond to their own central banks, suggesting that my methodology does not suffer from weak statistical power.

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 E.

I again utilize the Inference by Heteroskedasticity method to measure how these foreign primitives react to announcements by the Fed and by other central banks at a daily frequency. The results are presented in Tables 5 and 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 5: **Excess Volatility in Daily 10Y Rate Returns**

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

Notes: The table tests whether the model-estimated ten-year path of rates of the column country is more volatile around announcements by the row central bank than at other times, using daily returns. If returns are more volatile at the 1% level using the Brown-Forsythe test, the cell records the excess ratio of standard deviations (announcement window to non-announcement window standard deviations, minus 100%). If returns are not statistically more volatile, the cell is left blank. The full table, including ratios significant at the 5% level and insignificant ratios, can be found in the Appendix as Table 16. The Fed affects few other countries’ estimated paths of rates.
### Table 6: Excess Volatility in Daily 10Y Term Returns

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

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

## 6 Models of Complete Markets

I allow for more complex stochastic discount factors under complete markets, and show that the restrictions that my results require are either impossible or economically implausible for such models to match. Section 4 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 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 4. 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 source 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.\(^{31}\) This gives the stochastic discount factors

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\(^{31}\)While the correlation between currencies and bonds could add a third restriction, I find that its insight largely
enough mathematical freedom to match my results, but economically these two restrictions are
densely 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 use complete markets and one form
of heterogeneity to explain currency markets, and so 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 find 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 of complete markets that may be too great.

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.

6.1 Example 1: Power Utility and Simple Dynamics

I illustrate the tension between currency movements and bond movements using power utility and a
simple process for consumption. This follows closely the example in Section 4.2, although I generate
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. 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
(\(\rho, \beta, \gamma\)) to be common. I relax this assumption later.

\[
\log \bar{C}_t = \rho \log \bar{C}_{t-1} + \sigma^i \epsilon_t^i \quad \text{where} \quad \rho \in [0, 1] \quad \text{and} \quad \epsilon_t^i \sim N(0, 1)
\]

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

duplicates the insights from currencies and bonds separately. Specifically, this third potential restriction is that
\(\rho(\Delta m_J \Delta m_A, \Delta y_J \Delta y_A) < 0\). This is a useful restriction on its own, as shown in Section 4.2, but it does not
meaningfully strengthen the distinct results on currencies and bonds unless strong modeling assumptions are made.
I thus do not consider it in this section.
and the Japanese stochastic discount factor equal one another having adjusted for exchange rates.

\[
\beta \left( \frac{C^A_t}{C^A_{t-1}} \right)^{-\gamma} = \beta \left( \frac{C^J_t}{C^J_{t-1}} \right)^{-\gamma} \frac{S_t}{S_{t-1}}
\]

\(\text{Australian SDF} \quad \text{Japanese SDF}\)

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.\(^{32}\) I use this finding to generate my first restriction.

\[
\mathbb{E}_{t-1} (s_t - s_{t-1}) + r^A_{f,t-1} - r^J_{f,t-1} = \frac{1}{2} \gamma^2 \left( (\sigma^J)^2 - (\sigma^A)^2 \right) > 0 \tag{12}
\]

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.

\[
\mathbb{E}_{t} \left( \beta^k \left( \frac{C^A_{t+k}}{C^A_t} \right)^{-\gamma} \frac{1}{P^A_t(t,t+k)} \right) = 1 \quad \Rightarrow \quad \log P^A_t(t,t+k) = k \log \beta - \gamma \left( \mathbb{E}_{t} c_{t+k} - c_t \right) + \gamma^2 \mathbb{V}_t c_{t+k}
\]

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

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

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

\[
\mathbb{V}_{t-1} \left( \Delta y^J_t(t,t+k) \right) - \mathbb{V}_{t-1} \left( \Delta y^A_t(t,t+k) \right) = \frac{\gamma^2}{k^2} \left( 1 - \rho^k \right)^2 \left( (\sigma^J)^2 - (\sigma^A)^2 \right) < 0 \tag{13}
\]

Equations (12) and (13) are exactly contradictory. The former condition requires Japanese shocks to be larger than Australian shocks, to match the excess returns investors earn for holding

\(^{32}\)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].
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 \((\beta, \gamma)\) are heterogeneous across countries.

The only possible resolution is for \(\rho\) to vary by countries. However, I argue that ten-year bonds approximate infinite-maturity bonds in my sample, such that \(\rho^k \approx 0\). I rely upon two pieces of 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 II, 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.

6.2 Example 2: Epstein-Zin Utility and Complex Dynamics

I show that the tension between currency movements and bond movements persists, using a much richer model with Epstein-Zin utility and with dynamic consumption processes. In this model, a single source of heterogeneity is insufficient. Two sources 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 \(\phi\) and \(\rho\)), 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.

\[
\begin{align*}
c^i_t - c^i_{t-1} &= \mu + \phi x^i_{t-1} + \sigma^i_{t-1} \eta^i_t \\
x^i_t &= \rho x^i_{t-1} + \varphi \sigma^i_{t-1} \epsilon^i_t \\
(\sigma^i_t)^2 &= \sigma^2 + v \left( (\sigma^i_{t-1})^2 - \sigma^2 \right) + \sigma_w w^i_t
\end{align*}
\]

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
components of shocks, are only described in the appendix and suppressed here in the ellipses. Additionally lognormal. Terms that average out across countries over time, namely the idiosyncratic differential loadings $1 + \beta^i_e$ on the global components of shocks. I utilize that modeling innovation, and in fact decompose all shocks $(\eta, e, w)$ into both global and idiosyncratic components. Global components of $(\eta, e)$ have constant global volatility, while idiosyncratic components continue to have idiosyncratic stochastic volatility. Finally, shocks decompose into these two components with weightings $(\alpha_\eta, \alpha_e, \alpha_w)$. I present the updated dynamics.

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

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

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

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.

$$\forall_{t-1} (\Delta m^J_t) > \forall_{t-1} (\Delta m^A_t) \quad \text{and} \quad \forall_{t-1} (\Delta y^J_t) > \forall_{t-1} (\Delta y^A_t) \quad (14)$$

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

$$\forall_{t-1} (\Delta m^J_t) = \alpha_\eta (\gamma \sigma)^2 (1 + \beta^i_\eta)^2 + \alpha_e (1 - \rho)^{-1} (\gamma - 1/\psi) \phi \varphi_e \sigma^2 (1 + \beta^i_e)^2$$

$$+ \alpha_w (1 - v)^{-1} (\gamma - 1/\psi)(1 - \gamma) K_0 \sigma_w)^2 (1 + \beta^i_w)^2 + \ldots$$

$$\forall_{t-1} (\Delta y^J_t) = \alpha_e (1 - \rho)^{-1} (1/\psi) \varphi \varphi_e \sigma^2 (1 + \beta^i_e)^2$$

$$+ \alpha_w (1 - v)^{-1} (1/\psi - \gamma - \gamma/\psi) K_0 \sigma_w)^2 (1 + \beta^i_w)^2 + \ldots$$

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

These expressions make clear the difficulties this model faces in matching my empirical findings. First, they show that a single source 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 (14). However, in this setup Japan and Australia have the same variance in bond yields, violating the bond restrictions in Equation (14). A related argument applies

---

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

---
to $1 + \beta_i^\psi$ 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 sources 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 $1 + \beta_i^\eta_A > 1 + \beta_i^\eta_J$ to match the currency restriction in Equation (14). In turn, this requires $1 + \beta_i^e_J > 1 + \beta_i^e_A$ to match the bond restriction in Equation (14).\(^{34}\)

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 $\eta$ and $e$ load relatively more on the stochastic discount factors more while shocks $w$ load relatively more on bonds. This illustrates the general tension that models of complete markets must confront when matching the asymmetries I document, even when rich heterogeneity is incorporated.

### 6.3 General Restrictions on Stochastic Discount Factors

In this section, I characterize the tension between currency and bond movements in a general preference-free framework with complete markets, allowing for higher order moments and deriving results applicable to both nominal or real stochastic discount factors. To show the tension, I decompose a general stochastic discount factor into transitory and permanent components. The results from currencies require the permanent component of stochastic discount factors to be more volatile through Fed announcements in Japan than in Australia. By contrast, the results from bonds require the transitory components of stochastic discount factors to be more volatile through Fed announcements in 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 $L_{t-1}$.\(^{35}\) Entropy captures higher-order moments, al-

---

\(^{34}\)These conditions are necessary but not sufficient; the actual minimum gap between the two loadings depends on the exact parameterization of the model.

\(^{35}\)This is equivalent to half the conditional variance of the log of a random variable when working with lognormal random variables.
though there remains an open debate over the importance of higher-order moments to currency risk, with estimates ranging from under 20% by Jurek and Xu [2014] to as high as 40% by Chernov et al. [2014]. Second, in keeping with the notation of Alvarez and Jermann [2005], I distinguish between the pricing kernel \( \Lambda \) and the stochastic discount factor \( M \), where \( M_{t+k} \) is the ratio of pricing kernels \( \Lambda_{t+k} \) and \( \Lambda_{t-1} \), 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 infinite-horizon limit.

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

I first turn to my results from currencies. Backus et al. [2001] generalizes the expression in Equation (12) 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.

\[
\mathbb{E}_{t-1} (s_t - s_{t-1}) + r_{J,t-1} - r_{J,t-1} = L_{t-1} \left( \frac{\Lambda_t^j}{\Lambda_{t-1}^j} \right) - L_{t-1} \left( \frac{\Lambda_t^A}{\Lambda_{t-1}^A} \right) > 0
\]

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

\[36\] 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.
expectation of the stochastic discount factor spanning those two periods.

\[ P_t(\infty) = \mathbb{E}_t \left( \frac{\Lambda_{\infty}^t}{\Lambda_{\infty}^i} \right) = \mathbb{E}_t \left( \frac{\Lambda_{\infty}^{i,F} \Lambda_{\infty}^{i,T}}{\Lambda_{\infty}^{i,F} \Lambda_{\infty}^{i,T}} \right) \]

Alvarez and Jermann [2005] argue that at infinite maturities, there is no transitory component of the pricing kernel \( \Lambda_{\infty}^{i,T} \). 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.

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

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.

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

I compute the empirical entropy of gross innovations to yields following Fed announcements in my sample, and confirm in Appendix H 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.\(^{37}\) This yields the second restriction.

\[ \lim_{n \to \infty} \left( L_{t-1} \left( \exp \left( n \Delta y_t^A \right) \right) - L_{t-1} \left( \exp \left( n \Delta y_t^J \right) \right) \right) = L_{t-1} \left( \frac{\Lambda_{t}^{A,T}}{\Lambda_{t-1}^{A,T}} \right) - L_{t-1} \left( \frac{\Lambda_{t}^{J,T}}{\Lambda_{t-1}^{J,T}} \right) > 0 \quad (16) \]

Equations (15) and (16) illustrate this tension again. The transitory component of the stochastic discount factor today must be more volatile in Australia than in Japan, but the overall stochastic discount factor today must be more volatile in Japan than in Australia. As a result, either the permanent component must be more volatile in Japan than in Australia, or the correlation between the Japanese components must be higher than in Australia. These results are preference-free and encompass higher moments. Models with a single source of heterogeneity cannot match these equations simultaneously.

To make this discussion more concrete, I decompose the stochastic discount factors in the power utility and Epstein-Zin examples into permanent and transitory components. First consider the

\[ 37 \text{A long literature discusses the existence of “peso problems” in currency markets, i.e. the possibility that limited historical samples do not include observations of extreme events and thus bias calculations of risk. Burnside et al. [2011] estimate returns from currency portfolios that hedge such extreme events through options, and they find results that are quantitatively smaller but qualitatively consistent with existing work, suggesting that higher order moments do not greatly bias the sample.} \]
power utility example. The permanence of a consumption shock is driven by $\rho$. If $\rho = 1$, shocks are permanent. Agents cannot smooth away a permanent shock, and so both bond yields and the transitory component of the stochastic discount factor remain constant. If $\rho < 1$, 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.

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

Next consider the example with Epstein-Zin utility. I derive these expressions in Appendix H, 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 $\eta$ are permanent, and so they only appear when computing the permanent components of entropy. By contrast, shocks to the trend and volatility of consumption growth have both permanent and transitory elements. Consumption itself follows a random walk, but the trend component of consumption growth is a stationary autoregressive process.

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.

$$\begin{bmatrix}
L_{t-1} \left( \Lambda_i^i / \Lambda_{t-1} \right) \\
L_{t-1} \left( \Lambda_{i,T}^i / \Lambda_{t-1}^T \right) \\
L_{t-1} \left( \Lambda_{i,P}^i / \Lambda_{t-1}^P \right)
\end{bmatrix} = A \begin{bmatrix}
(1 + \beta^\iota_0)^2 \\
(1 + \beta^\iota_1)^2 \\
(1 + \beta^\iota_w)^2
\end{bmatrix}$$

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

7 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 complete markets — do not seem consistent with my empirical results, leaving only models of incomplete markets to explain monetary spillovers. Future research is needed to select a specific model in this class. In this section, I do two things. First, I illustrate with a simple example, using the model of segmented markets by Gabaix and Maggiori [2015]. Second,
I document one additional condition for selecting among models: they must embed heterogeneity across central banks. Specifically, I show that the Fed and ECB uniquely generate global spillovers, while most other central banks do not, as briefly seen in Table 1.

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 — a convenient grouping under the previous two channels — becomes highly suggestive 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 forms of countervailing heterogeneity, but many models of incomplete markets require only one form or two complementary forms of heterogeneity. I do not select a specific model, but I demonstrate with a simple model of segmented markets.

I document one additional constraint for such models: they must make frictions heterogeneously sensitive to different central banks, responding only to the Fed and the ECB. I extend the paper’s methodologies to document spillovers around the announcements of the other nine central banks in my sample. Foreign assets react to the ECB, with assets of non-Eurozone continental European countries reacting especially strongly. However, foreign assets do not globally react to other central banks.

These findings are noteworthy in their own right, both in scope and in statistical power. There is a paucity of literature formally studying spillovers emanating from other central banks, except for the ECB’s effects on European assets. This is likely due to two limitations. First, data on observed monetary shocks can be difficult to collect in smaller countries, although my methodology bypasses this issue by estimating inferred shocks. Second, conventional methods struggle to identify Fed spillovers at reasonable significance levels, and this challenge grows for spillovers from smaller central banks. However, my methodology continues to get its statistical power from high-frequency returns, long-maturity assets, and inferred shocks, and so I can estimate spillovers or the lack thereof with precision.

The results showing the divergence in central banks echo work in other domains of international finance, where the dollar and euro are dominant. For instance, Gopinath [2015] finds that trade is invoiced overwhelmingly in dollars and euros, Bruno and Shin [2015] find cross-border bank loans are most commonly in dollars and euros, and Maggiori et al. [2017] find that the foreign currency exposures of international portfolios are biased toward dollars and euros. This result adds to a growing collection of work on the relative hegemony of American and European institutions.
Moreover, the findings guide the search for specific models to explain spillovers. Models in the international finance literature do not typically link shocks to their monetary or non-monetary origins, such as Gabaix and Maggiori [2015]. Canonical models in the international macroeconomics literature link shocks to symmetric monetary authorities, such as Obstfeld and Rogoff [1995] and Corsetti and Pesenti [2001].\(^{38}\) A comprehensive model of monetary spillovers requires synthesizing insights from both literatures and this paper, linking financial frictions to monetary authorities, and doing so asymmetrically.

### 7.1 Example Model of Segmented Markets

I use a simple model of incomplete markets, following Gabaix and Maggiori [2015], to show that this model can match my empirical results. 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 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 \(\theta\) across a set of assets are defined by three terms: the overall (scalar) constraint \(\Gamma\) of intermediaries, the variance-covariance matrix \(\Sigma\) of asset returns, and the expected returns \(E_t(p_{t+1} - p_t)\) of assets.\(^{39}\)

\[
\theta = \Gamma^{-1} \Sigma^{-1} (E_t p_{t+1} - p_t)
\]

The crucial assumption needed to allow this model to explain spillovers is that the Fed adjusts the constraint \(\Gamma\). There are a variety of explanations for this. 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.

\[
\frac{\partial p_t}{\partial \Gamma} = - \Gamma^{-1} (E_t p_{t+1} - p_t)
\]

Fluctuations in prices are negatively proportional to expected returns, through intermediaries.

\(^{38}\)One notable exception is Benigno [2002], who explicitly considers differential bargaining power between monetary authorities.

\(^{39}\)\(\Gamma\) 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.
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, I test more directly whether actual measures of intermediaries’ positions correlate with asymmetries in currency and bond markets, through Equation (6) (replicated here). In Appendix D, I show evidence for several tests. 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 countries rise more than yields of lender countries. Such results fit with my example model.

\[
\begin{bmatrix}
\Delta s_{e/t} \\
\Delta s_{f/t} \\
\Delta s_{Y/t}
\end{bmatrix} =
\begin{bmatrix}
1 & X_{t-1}^{e} & 1 \\
1 & X_{t-1}^{f} & 1 \\
1 & X_{t-1}^{Y} & 1
\end{bmatrix}
\begin{bmatrix}
\alpha_0 \\
\alpha_1 \\
\beta_0 \\
\beta_1
\end{bmatrix}
\begin{bmatrix}
\mu_t \\
\epsilon_{e/t} \\
\epsilon_{f/t} \\
\epsilon_{Y/t}
\end{bmatrix}
\]

7.2 Spillovers by Other Central Banks

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

I first show results from currency markets in Table 7, to complement results from bond markets in Table 1. The previous pattern holds up: the Fed continues to have strong effects on most currency markets. For instance, the Australian dollar (measured against an equal-weighted basket of currencies, except for the US dollar) is 73% more volatile through Fed announcement windows than non-announcement windows at the 1% level. Moreover, central banks have the power to affect their own currencies.

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

I next employ the latent factor model to each central bank, to identify asymmetries in how
Table 7: Excess Volatility in 60M Currency Returns

<table>
<thead>
<tr>
<th></th>
<th>AUD</th>
<th>CAD</th>
<th>CHF</th>
<th>EUR</th>
<th>GBP</th>
<th>JPY</th>
<th>NOK</th>
<th>NZD</th>
<th>SEK</th>
<th>USD</th>
</tr>
</thead>
<tbody>
<tr>
<td>Australia</td>
<td>293%</td>
<td>185%</td>
<td></td>
<td></td>
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<tr>
<td>Canada</td>
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<td></td>
<td>41%</td>
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<td>16%</td>
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<tr>
<td>Switzerland</td>
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<td>152%</td>
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<tr>
<td>Euro</td>
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<td></td>
<td></td>
<td></td>
<td></td>
<td>37%</td>
<td>137%</td>
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Notes: The table tests whether the currency of the column country is more volatile in the sixty minutes around announcements by the row central bank than in other sixty-minute intervals. (Currencies are measured against an equal-weighted basket of all other currencies except the one whose central bank is making announcements.) If returns are more volatile at the 1% level using the Brown-Forsythe test, the cell records the excess ratio of standard deviations (announcement window to non-announcement window standard deviations, minus 100%). If returns are not statistically more volatile, the cell is left blank. The full table, including ratios significant at the 5% level and insignificant ratios, can be found in the Appendix as Table 18. The Fed and ECB have spillover effects, but most other central banks only affect their own currencies.

Currencies and bonds react to that central bank’s monetary shocks. The most stark results come from the ECB, plotted in Figure 9. When the ECB tightens, the euro appreciates less against other continental European currencies (e.g. the Norwegian krone) than against non-European currencies, and European bond yields rise more than non-European bond yields. This is noteworthy, although the results from currency and bond markets do not rule out central banks reacting to the ECB; nor do they rule out shifts in risk premia under complete markets. For instance, if the Norges Bank follows the ECB, both Norwegian bond yields would rise and the euro would appreciate less against the Norwegian krone when the ECB tightens, as I observe. Similarly, if the Norwegian stochastic discount factor had a positive innovation, again both Norwegian bond yields would rise and the euro would appreciate less against the Norwegian krone, as I observe. Thus, the channel for the ECB’s spillovers remains an open question.

I also test whether assets react asymmetrically to other central banks. Unsurprisingly, I find few strong asymmetries — if assets are not reacting to other central banks as in Tables 1 and 7, they cannot react asymmetrically. For instance, consider the Bank of Japan in Figure 10. When the yen appreciates, it does so symmetrically against all currencies; and when Japanese bond yields move, other bond yields do not move asymmetrically (or at all).

The exceptions are again with neighboring central banks. In Section 4, I illustrated that Australian assets have some exposure to the Reserve Bank of New Zealand. I similarly show that New Zealand assets have some exposure to the Reserve Bank of Australia in Figure 11. When either 40These empirical results are robust to excluding the Swiss franc and Swiss bond yields in the three years during which the Swiss National Bank formally capped the franc’s exchange rate against the euro.
Figure 9: Market Reactions to EU Monetary Shocks

(a) Currency Responses

(b) Bond Responses

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

Figure 10: Market Reactions to JP Monetary Shocks

(a) Currency Responses

(b) Bond Responses

Notes: The figures depict the reactions of currency and bond markets to announcements by the Bank of Japan. The left figure shows by how much the JPY appreciates against a given reference currency when it appreciates by 1% on average; and the right figure shows by how much foreign ten-year yields rise when Japanese ten-year yields rise by 1%. Standard error bars in both pictures are computed versus the average reaction across currencies or foreign bonds; and the shading of the coefficient bars refers to the lower-dimensional structure, whereby assets of the same color react similarly and assets of different colors react dissimilarly following BoJ announcements. The JPY appreciates symmetrically against all currencies and foreign yields do not rise asymmetrically when the BoJ tightens.
central bank tightens, its currency depreciates by slightly less against its counterpart than other currencies. In this example, the Australian dollar appreciates or depreciates less against the New Zealand dollar (and Canadian dollar), and New Zealand bond yields rise and fall with Australian yields. I show the full results for all ten central banks in Appendix G.

Figure 11: 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. 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.

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41 In addition, Lustig and Richmond [2017] have a wider sample, that includes many emerging markets.
8 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 unequivocal negative answers to the leading classes of explanations. The only explanations with which my findings align are models with incomplete markets, which may be broadly informative to academics about the nature of the international financial system.

These results offers guidance to policymakers too. Despite concerns of the trilemma being replaced by a “dilemma,” monetary independence in developed markets survives. Central banks do not yet feel the need to offset the Fed’s actions on their 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.
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Appendix

I present the table of contents for the appendix. New material is being added to the appendices regularly, so please ensure you have the latest version of the paper by clicking [here].

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

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

C. Long-Run Assumptions: I argue that various long-run variables, such as inflation, do not change through Fed announcements, which is important for the paper’s interpretations of the empirical findings.

D. Characterizing Asymmetric Responses: I document the measured predictor variables that best explain asymmetric reactions to the Fed in currency and bond markets, as suggestive evidence of the underlying mechanisms.

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

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

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

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

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

A.1 Australia

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

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

A.2 Canada

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

A.3 Eurozone

European rate announcements primarily target the Main Refinancing Rate (although they concurrently target other rates too) and are made by the European Central Bank. Announcements are made at 1:45 PM CET. Announcements are generally made on Thursday afternoons and occasionally on Wednesday afternoons. Until November 2001, regularly scheduled announcements were made twice a month; from November 2001 - December 2014, regularly scheduled announcements were made monthly; and starting in January 2015, there are eight regularly scheduled announcements per year. This leads to a total of 193 regularly scheduled announcements (21 in 2001, 12 in 2002 - 2014, and 8 in 2015 - 2016). The ECB has made two unscheduled rate announcements (following September 11, 2001, and during the financial crisis), which are excluded from the sample.
European rate announcements discuss only the current rate and do not discuss the future path of monetary policy. 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 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.4 Japan

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

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42 The ECB has changed this as of the middle of 2016, but this only affects a few announcements.
43 Technically, the opening paragraph bids Happy New Year; this is the first paragraph with content.
Since the Bank of Japan does not release announcements at a preset time, market participants might anticipate unusually late announcements as being unusually important. As such, I check that the core results are robust to excluding the fourteen announcements that take place after 2:00 PM JST.

A.5 New Zealand

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

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

A.6 Norway

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

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

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

A.8 Switzerland

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

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

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

Equally unexpectedly, the SNB abandoned the peg on January 15, 2015. While the franc was not officially pegged to the Euro over the intervening forty months, strong appreciation pressures combined with the cap led to far lower volatility. The time series of the Swiss franc-euro exchange rate depicts this, in Figure 12.

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

A.9 United Kingdom

British rate announcements target the Official Bank Rate and are made by the Bank of England. Announcements are always made at 12:00 PM BST. Announcements are almost always made on
Notes: The figure depicts the 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.

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. 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

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44 The Bank of England has changed this as of August 2015, but this only affects a few announcements.
England announcements versus other periods, and this finding is statistically significant via the Brown-Forsythe test.

A.10 United States

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

One regularly scheduled announcement on January 31, 2006 closely overlaps with a Senate confirmation vote for Chairman Bernanke. Since the vote was largely anticipated, the announcement is not excluded outright but I check that the core results are robust to its exclusion. Nakamura and Steinsson [2017] have raised concerns that FOMC announcements signal the Fed’s private information about fundamentals, even though announcements do not officially release new information about fundamentals. If true, monetary announcements should look qualitatively like official announcements about fundamentals: both types of announcements would concurrently release information about the state of the economy and the Fed’s reaction to it. As such, I collect unemployment releases by the Bureau of Labor Statistics, and look at the reaction of currencies and bonds globally to that. I compare such reactions to the reactions following Fed announcements in Figure 13. There is a sharp difference: the patterns that emerge following monetary announcements look little like the patterns that emerge following labor announcements. For instance, whereas the Canadian dollar and yen react similarly to each other following Fed announcements, they react very differently following BLS announcements; and whereas Australian and New Zealand bonds react similarly to each other following Fed announcements, they react very differently following BLS announcements. FOMC announcements are not just announcements about fundamentals.
Figure 13: 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.
B Empirical Framework

This section discusses the empirical framework behind the paper’s core equation. First, it introduces the notation for the fully general equation. Second, it discusses the two leading methodologies used to identify its parameters: a simplified variance test and a maximum likelihood-based approach (i.e. the Expectation Maximization algorithm). The section then provides details of the approach used to find the model’s lower-dimensional mapping. Finally, the section discusses the implementations and 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.

This section is not yet complete, so please check back soon.

B.1 Model

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

\[
r_t(i_t) = X_t^\alpha(i_t)\alpha + X_t^\beta(i_t)\beta m_t + \epsilon_t(i_t) \quad \forall \ t
\]  

Each component of Equation (17) is defined as follows:

- \( r_t \) is a \( C_r \times 1 \) vector of asset returns at time \( t \).
- \( i_t \) is a \( C_i \times 1 \) 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.
- \( \alpha \) is a \( C_{\alpha} \times 1 \) vector of constants.
- \( \beta \) is a \( C_{\beta} \times C_m \) matrix of coefficients (i.e. factor loadings).
- \( X_t^\alpha \) and \( X_t^\beta \) are covariate matrices at time \( t \) with dimensions \( C_r \times C_{\alpha} \) and \( C_r \times C_{\beta} \) respectively.
- \( m_t \) is a \( C_m \times 1 \) vector of factors at time \( t \) with underlying distribution \( m_t \sim N(\bar{\mu}, \Omega) \).
- \( \epsilon_t \) is a \( C_r \times 1 \) vector of residuals at time \( t \) with underlying distribution \( \epsilon_t \sim N(\bar{0}, \Sigma) \). \( \Sigma \) is learned from asset returns through non-event windows (denoted as \( \tilde{r}_t \)). To ensure that the mean of the residuals is zero in this estimation, I subtract the mean asset return through non-event windows from asset returns through event windows (\( r_t \)) before fitting this model to those returns.

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

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\begin{bmatrix}
\tilde{r}_t^{e/s} \\
\tilde{r}_t^{y/s} \\
\tilde{r}_t^{l/s} \\
r_t
\end{bmatrix}
= \begin{bmatrix}
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0 & 1 & 0 & 1 \\
0 & 0 & 1 & 1 \\
0 & 1 & 0 & 0
\end{bmatrix}
\begin{bmatrix}
\tilde{\alpha}_e^b \\
\tilde{\alpha}_y^b \\
\tilde{\alpha}_l^b \\
\tilde{\alpha}_y^f
\end{bmatrix}
+ \begin{bmatrix}
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0 & 1 & 0 & 1 \\
0 & 0 & 1 & 1 \\
0 & 1 & 0 & 0
\end{bmatrix}
\begin{bmatrix}
\tilde{\beta}_e^b \\
\tilde{\beta}_y^b \\
\tilde{\beta}_l^b \\
\tilde{\beta}_y^f
\end{bmatrix}
\times m_t
+ \begin{bmatrix}
\tilde{\epsilon}_t^{e/s} \\
\tilde{\epsilon}_t^{y/s} \\
\tilde{\epsilon}_t^{l/s} \\
\epsilon_t
\end{bmatrix}
\begin{bmatrix}
1 \\
1 \\
1 \\
0
\end{bmatrix}
\]
\[
\begin{bmatrix}
  r_{t}/e^{s} \\
  r_{t}/\xi
\end{bmatrix}
= \begin{bmatrix}
  1 & 0 & 0 \\
  0 & 1 & 0
\end{bmatrix}
\times \begin{bmatrix}
  \alpha^{e} \\
  \alpha^{k}
\end{bmatrix}
+ \begin{bmatrix}
  1 & 0 & 0 \\
  0 & 1 & 0
\end{bmatrix}
\times \begin{bmatrix}
  \beta^{e} \\
  \beta^{k}
\end{bmatrix}
\times m_{t}
+ \begin{bmatrix}
  \epsilon^{e}/s \\
  \epsilon^{k}/s
\end{bmatrix}
\]

\[
\begin{bmatrix}
  r_{t}^{e} \\
  r_{t}^{k}/s
\end{bmatrix}
= \begin{bmatrix}
  1 & 0 \\
  0 & 1
\end{bmatrix}
\times \begin{bmatrix}
  \alpha^{e} \\
  \alpha^{k}
\end{bmatrix}
+ \begin{bmatrix}
  1 & 0 \\
  0 & 1
\end{bmatrix}
\times \begin{bmatrix}
  \beta^{e} \\
  \beta^{k}
\end{bmatrix}
\times m_{t}
+ \begin{bmatrix}
  \epsilon^{e}/s \\
  \epsilon^{k}/s
\end{bmatrix}
\]

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 \( X^{e} \) and \( X^{k} \) 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.

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).

\section*{B.2 Variance Tests}

Consider a single-asset variant of Equation (17) without time-varying covariates (i.e. \( C_{r} = 1 \) and \( X_{t}^{\beta} = X^{\beta} \)) in which I want to test \( H_{0} : \beta = 0 \). This corresponds to testing whether that asset has any exposure to monetary shocks \( m_{t} \) emanating from a given central bank. In this simplified setting, it is overkill to make various structural assumptions, to employ sophisticated algorithms to fit the model, or even to derive point estimates for \( \beta \). Instead, this hypothesis can be tested by a simple variance test. Taking the variance of Equation (17) yields the following, where missing returns are simply dropped:

\[
\text{Var}(r_{t}) = X^{\beta} \Omega \beta^{T} X^{\beta T} + \text{Var}(\epsilon_{t})
\]

Notice that unless \( \beta = 0 \), 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:

\[
\text{Var}(r_{t}) > \text{Var}(\epsilon_{t}) \implies \beta \neq 0
\]

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

\[
\text{Var}(r_{t}) > \text{Var}(\hat{r}_{t}) \implies \beta \neq 0
\]
Although the F-test for equality of variances is the best known variance test, I employ the Brown-Forsythe test for equality of variances instead. This choice is discussed further below.

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

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

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

Third, the choice of the Brown-Forsythe test over the F-test means that asset returns need not be normal. Broadly, the Brown-Forsythe test computes its test statistic via absolute deviations from the median, rather than squared deviations from the mean as the F-test does, and both of these adjustments ensure the test remains robust to fat-tailed data. To further illustrate its relative advantages, I simulate repeated samples of fat-tailed data with unitary variance and varying excess kurtosis, and test a random subset of each sample against another random subset under both the Brown-Forsythe and the F-test. Table 8 shows the percentage of time that these tests reject at the 5% level. When the data has no excess kurtosis (i.e. a standard normal), both tests correctly reject 5% of the time. However, for moderate or high values of excess kurtosis, the F-test rejects far too frequently; whereas the Brown-Forsythe test continues to reject 5% of the time correctly.

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

Notes: The table needs an updated caption.

### B.3 Expectation Maximization Algorithm

To identify the parameters $(\alpha, \beta, m)$ in the general setting of Equation (17), I use the Expectation Maximization algorithm. This approach finds the maximum likelihood estimates of the parameters. The likelihood function is written below, but the maximum likelihood parameter estimates cannot be solved directly, due to the product of $\beta$ and $m$. The Expectation Maximization algorithm, by contrast, updates each of $(\alpha, \beta, m)$ sequentially until the algorithm converges on the optimal estimates.
This specification is a latent Gaussian model (a type of hidden Markov model), and is common in the signal processing literature. Such models are frequently solved in that literature through the Expectation Maximization algorithm.

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

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

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

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

Rearranging the first order condition for $\beta$ presents some complications, as $\beta$ is not easily isolated:

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

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

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

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

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

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

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

### B.3.1 Markov Chain Monte Carlo

### B.4 Lower-Dimensional Mapping

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

\[
\begin{bmatrix}
e_t/\$ \\
e_{\text{Y}}/\$
\end{bmatrix} = \begin{bmatrix} 1 & 0 \\ 0 & 1 \\
\end{bmatrix} \times \begin{bmatrix} \alpha^e \\ \beta^e \end{bmatrix} + \begin{bmatrix} 1 & 0 \\ 0 & 1 \\
\end{bmatrix} \times \begin{bmatrix} \alpha^\text{Y} \\ \beta^\text{Y} \end{bmatrix} \times m_t + \begin{bmatrix} \epsilon_{e_t}/\$ \\ \epsilon_{\text{Y}_t}/\$
\end{bmatrix}
\]

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 (18). 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 \(\alpha\) and \(\beta\) but also the variational posterior parameters for each monetary shock, is a prime candidate for such an approach.
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 4, 5, and 7, 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.
\[ EBIC(\theta) = 2\mathcal{L}(\theta) - \log(n)|\theta| - 2\log(|\theta|)|\theta| \]

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 \( k = 1, \ldots, 9 \). The formula for Stirling numbers of the second kind is:

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

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

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

\[
\begin{bmatrix}
\begin{bmatrix}
  r_{e/s,t} \\
  r_{c,t} \\
  r_{Y,c,t} \\
  r_{£,c,t} \\
  r_{e,b,t} \\
  r_{Y,b,t} \\
  r_{£,b,t}
\end{bmatrix} \\
\end{bmatrix} =
\begin{bmatrix}
  1 & 0 & 0 & 0 \\
  0 & 1 & 0 & 0 \\
  0 & 0 & 1 & 0 \\
  0 & 0 & 0 & 1
\end{bmatrix}
\times
\begin{bmatrix}
\begin{bmatrix}
  \alpha_{e,c} \\
  \alpha_{c,c} \\
  \alpha_{Y,c} \\
  \alpha_{£,c} \\
  \alpha_{e,b} \\
  \alpha_{Y,b} \\
  \alpha_{£,b}
\end{bmatrix} \\
\end{bmatrix} +
\begin{bmatrix}
\begin{bmatrix}
  \beta_{e,c} \\
  \beta_{c,c} \\
  \beta_{Y,c} \\
  \beta_{£,c} \\
  \beta_{e,b} \\
  \beta_{Y,b} \\
  \beta_{£,b}
\end{bmatrix} \\
\end{bmatrix}
\times
m_t +
\begin{bmatrix}
\begin{bmatrix}
  \epsilon_{e/s,t} \\
  \epsilon_{c,t} \\
  \epsilon_{Y,c,t} \\
  \epsilon_{£,c,t} \\
  \epsilon_{e,b,t} \\
  \epsilon_{Y,b,t} \\
  \epsilon_{£,b,t}
\end{bmatrix}
\end{bmatrix}
\]

### B.5 Identification by Heteroskedasticity

An alternate approach to identify the parameter \( \beta \) for Equation (17) is to use Identification by Heteroskedasticity. This approach is a generalized methods of moments (GMM) estimator. While this approach gives equivalent results to the Expectation Maximization algorithm when it converges, it has poor convergence properties in the first place due to the severe non-linearities and high dimensionality of the estimator. As such, it is not utilized except as a robustness check. This section discusses both the implementation and the issues.

To understand the principle behind Identification by Heteroskedasticity in this context, I first present an extremely simplified version of Equation (17) to build intuition. In this example, asset returns are demeaned, monetary shocks \( m_t \) are univariate, there is no missing data, there is one coefficient per asset (\( C_r = C_\beta \)), and \( X_t^\beta = I_{C_r} \). In this simplified framework, asset returns are measured around the set of event windows \( E \) and the set of non-event windows \( N \):

\[
r_t = \beta m_t + \epsilon_t \quad \forall \ t \in E
\]
\[
\tilde{r}_t = \epsilon_t \quad \forall \ t \in N
\]
Taking the second moments of these systems yields the following:

\[ E(\mathbf{r}_t^T \mathbf{r}_t) = \mathbf{\beta} \mathbf{\beta}^T \mathbf{m}_t + \mathbf{\Sigma} \]

\[ E(\tilde{\mathbf{r}}_t^T \tilde{\mathbf{r}}_t) = \mathbf{\Sigma} \]

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

\[ E(\mathbf{r}_t^T \mathbf{r}_t) - E(\tilde{\mathbf{r}}_t^T \tilde{\mathbf{r}}_t) = \mathbf{\beta} \mathbf{\beta}^T \]

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

\[ r_t(i_t) = \mathbf{X}^{\mathbf{\beta}}(i_t) \mathbf{\beta} \mathbf{m}_t + \epsilon_t(i_t) \quad \forall \ t \in E \]

\[ \tilde{r}_t(i_t) = \epsilon_t(i_t) \quad \forall \ t \in N \]

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.

\[
\frac{1}{\sum_{t \in E,(j,k) \in i_t} 1} \sum_{t \in E,(j,k) \in i_t} r_t(j)r_t(k) = \frac{1}{\sum_{t \in E,(j,k) \in i_t} 1} \sum_{t \in E,(j,k) \in i_t} \mathbf{X}^{\mathbf{\beta}}(j)\mathbf{\beta}^T \mathbf{X}^{\mathbf{\beta}}(k)^T \mathbf{m}_t^2 + \sigma(j,k) \\
\frac{1}{\sum_{t \in N,(j,k) \in i_t} 1} \sum_{t \in N,(j,k) \in i_t} \tilde{r}_t(j)\tilde{r}_t(k) = \sigma(j,k) \\
\frac{1}{\sum_{t \in E,(j,k) \in i_t} 1} \sum_{t \in E,(j,k) \in i_t} r_t(j)r_t(k) - \frac{1}{\sum_{t \in N,(j,k) \in i_t} 1} \sum_{t \in N,(j,k) \in i_t} \tilde{r}_t(j)\tilde{r}_t(k) = \mathbf{X}^{\mathbf{\beta}}(j)\mathbf{\beta}^T \mathbf{X}^{\mathbf{\beta}}(k)^T
\]

There are \( \frac{1}{2} \times C_r \times (C_r + 1) \) unique equations to estimate \( C_\mathbf{\beta} \) parameters. This is an over-determined system, and so the Generalized Method of Moments is utilized to find the best fit. The sample moment condition for asset returns at time \( t \) and variance-covariance entry \((j,k)\) is defined as follows:

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

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 \( W \):
\[ g(\beta) = \left[ \sum_t g_t, (j=1, k=1)(\beta) \ldots \sum_t g_t, (j=1, k=C_r)(\beta) \ldots \sum_t g_t, (j=C_r, k=C_r)(\beta) \right]^T \]

\[ \hat{\beta} = \arg \min_{\beta} g(\beta)^T W g(\beta) \]

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

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

### B.6 Fixed Effects

An alternate approach to identify the parameter \( \beta \) for Equation (17) would be to use fixed effects for coefficients, and ignore variation in monetary shocks over time. Again, consider a simplified version of Equation (17) in which asset returns are demeaned, monetary shocks \( m_t \) are univariate, and there is no missing data, there is one coefficient per asset (\( C_r = C_\beta \), and \( X_t^\beta = I_{C_r} \):

\[
\begin{bmatrix}
\epsilon / S \\
r_t \\
r_L^t / S \\
r_Y^t / S \\
\end{bmatrix}
= 
\begin{bmatrix}
\beta \epsilon / S \\
\beta \ell / S \\
\beta \eta / S \\
\beta \xi / S \\
\end{bmatrix}
\]

\[
\begin{bmatrix}
\epsilon / S \\
r_t \\
r_L^t / S \\
r_Y^t / S \\
\end{bmatrix}
= 
\begin{bmatrix}
\epsilon / S \\
\ell / S \\
\eta / S \\
\xi / S \\
\end{bmatrix}
\]

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

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

Figure 15: Fixed Effects Estimator

(a) Small Shocks  (b) Large Shocks

Notes: The figures need an updated caption.

B.7 Vector Autoregression

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

B.8 Measured Shocks

B.8.1 Traditional Measures

An alternate approach is to use traditional measures of monetary shocks, rather than estimating or
inferring them. There are five popular ones: movements in short-term rates (e.g. Fed Funds futures)
around announcements, movements in medium-term rates (e.g. the two-year Treasury) around
announcements, differences between actual policy and surveyed expectations, shocks constructed
through the narrative method (i.e. Romer and Romer shocks), and policy deviations from the
Taylor Rule.

Four of these are immediately problematic; the exception is medium-term rates. Short-term
measures and surveys focus only on the surprises delivered to the short end of the yield curve; but
especially in the past decade, monetary surprises are most commonly delivered to the medium end
of the yield curve. For instance, I graph how often surveys corrected anticipated the monetary
announcement over my sample below. They were accurate 80+% of the time, and so a measure of
surprises constructed from survey data would involve retaining 20% of the data at most.

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

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

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

B.8.2 Average Return

An alternate approach to identify the parameter \( \beta \) for Equation (17) would be to use take the average of \( r_t \) as a measure of the monetary shock at time \( t \). This is equivalent to regressing equity returns on the market return (i.e. their average return), as is common in the finance literature.

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

To show each of these biases in more depth, I simulate 500 samples and nine assets as before, and estimate coefficients using an oracle estimator (that has access to the true underlying shocks), the Expectation Maximization algorithm, and the Average Return method. In each sample, coefficients take a randomly chosen value \( \in \{0.75, 1.00, 1.25\} \). I first explore the coefficient bias by restricting to the estimates where the underlying true coefficient is 0.75 or 1.25. I plot the density of coefficients obtained by the oracle estimator against the Expectation Maximization, and against the Average Return. There is a clear bias towards one in the case of the latter.

Second, I focus on the coefficients that are actually one, and plot the density of t-statistics for each method’s estimates. If standard errors are computed correctly, these should follow a standard normal distribution (e.g. 5% of the t-statistics should make a Type I error). For the Expectation Maximization algorithm, that happens. However, for the Average Return method, that does not happen – the standard errors are too small, and so the Type I errors made are far too numerous.
Figure 17: Average Return

(a) Density (1)  (b) Density (2)

(c) Density (3)

Notes: The figures need an updated caption.
Figure 18: 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).
Figure 19: Market Reactions to Monetary Shocks (EM vs GMM)

(a) Currency Responses to US Shocks (EM)

(b) Currency Responses to US Shocks (GMM)

(c) Currency Responses to EU Shocks (EM)

(d) Currency Responses to EU Shocks (GMM)

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

C.1 Overview

In this section, I test the validity of the approximations in Equation (7). The cross-border bond portfolio in Equation (7) is not directly affected by foreign monetary policy if the following assumption holds.

\[ \sum_{k=11}^{\infty} \Delta^{j}_{t+k-1} - \Delta s^{j/\infty}_{\infty} \approx 0 \]

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).

\[ \left( \sum_{k=11}^{\infty} \Delta^{r}_{t+k-1} - \Delta q^{j/\infty}_{\infty} \right) + \left( \sum_{k=11}^{\infty} \Delta^{\pi}_{t+k-1} - \Delta \hat{p}^{j}_{\infty} \right) \approx 0 \]

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 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.

C.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.

C.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.

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

C.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.

C.4.1 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 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.

C.4.2 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

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.
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 4) and the six-year forward four-year maturity (for Section 5).

First, I offer evidence from returns in the Treasury inflation-protected securities (TIPS) market. Table 10 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.

Table 10: Excess Volatility in US Inflation Estimates

<table>
<thead>
<tr>
<th></th>
<th>Nominal</th>
<th>Real (TIPS)</th>
<th>Inflation</th>
</tr>
</thead>
<tbody>
<tr>
<td>7F3Y</td>
<td>42**</td>
<td>24**</td>
<td>18</td>
</tr>
<tr>
<td>10F20Y</td>
<td>51**</td>
<td>33**</td>
<td>−9</td>
</tr>
</tbody>
</table>

Notes: The table tests whether six types of bonds are more volatile around announcements by the Fed than at other times, using daily returns. The six bonds are the nominal yield, the TIPS yield, and the difference between the two, for both the seven-year forward three-year bond and the ten-year forward twenty-year bond. 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 11 shows that the Fed does not affect estimates of inflation over the six-year forward four-year and ten-year forward twenty-year maturities.

Table 11: Excess Volatility in Foreign Inflation Estimates

<table>
<thead>
<tr>
<th></th>
<th>E.U.</th>
<th>U.K.</th>
</tr>
</thead>
<tbody>
<tr>
<td>6F4Y</td>
<td>11</td>
<td>7</td>
</tr>
<tr>
<td>10F20Y</td>
<td>13</td>
<td>12</td>
</tr>
</tbody>
</table>

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.

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

47 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.
D 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.

D.1 Methodology

The methodology takes the empirical framework discussed in Section 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 (6), here as Equation (19). I estimate $\beta_1$ for each measured pre-announcement predictor $X_{t-1}$ in Equation (1).

$$
\begin{bmatrix}
\Delta s_e/\$ \\
\Delta s_£/\$ \\
\Delta s_Y/\$
\end{bmatrix}
= 
\begin{bmatrix}
1 & X_{t-1}^{e,\$} \\
1 & X_{t-1}^{£,\$} \\
1 & X_{t-1}^{Y,\$}
\end{bmatrix}
\begin{bmatrix}
\alpha_0 \\
\alpha_1
\end{bmatrix}
+ 
\begin{bmatrix}
1 & X_{t-1}^{e,\$} \\
1 & X_{t-1}^{£,\$} \\
1 & X_{t-1}^{Y,\$}
\end{bmatrix}
\begin{bmatrix}
\beta_0 \\
\beta_1
\end{bmatrix}
m_t + 
\begin{bmatrix}
\epsilon_e/\$ \\
\epsilon_£/\$ \\
\epsilon_Y/\$
\end{bmatrix}
$$

A statistically significant $\beta_1$ means that the candidate predictor $X_{t-1}$ is meaningful for predicting variation in responses to Fed monetary announcements both across countries and across time. I estimate Equation (19), 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.

D.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.

D.2.1 Interest Rates

I let $X_{t-1}$ 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.

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48 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.
D.2.2 Volatility

I let $X_{i,t-1}^i$ 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.\footnote{The VIX uses implied volatility, whereas I compute historical volatility.} 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).

D.2.3 Limits to Arbitrage

I let $X_{i,t-1}^i$ 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.

D.2.4 Trade Flows

I let $X_{i,t-1}^i$ 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.

\[
X_{i,t-1}^i = \frac{\text{Exports}^{i \rightarrow \text{US}} + \text{Imports}^{i \rightarrow \text{US}}}{\sum_c (\text{Exports}^{i \rightarrow c} + \text{Imports}^{i \rightarrow c})}
\]

\[
X_{i,t-1}^i = \frac{\text{Exports}^{i \rightarrow \text{US}} - \text{Imports}^{i \rightarrow \text{US}}}{\text{Exports}^{i \rightarrow \text{US}} + \text{Imports}^{i \rightarrow \text{US}}}
\]

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.
D.2.5 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.

D.2.6 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.

$$X_{i,t-1} = \frac{\text{Assets}^{i \to \text{US}} + \text{Liabilities}^{i \to \text{US}}}{\sum_c (\text{Assets}^{i \to c} + \text{Liabilities}^{i \to c})}$$

$$X_{i,t-1}' = \frac{\text{Assets}^{i \to \text{US}} - \text{Liabilities}^{i \to \text{US}}}{\text{Assets}^{i \to \text{US}} + \text{Liabilities}^{i \to \text{US}}}$$

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.

D.2.7 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.
D.2.8 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.

D.2.9 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 (19) 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 C. 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.
E 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 ($\mu$, $\Phi$, $\Sigma$, $\sigma^2$).

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

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

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 risk premia 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 $\mu$ and $\Phi$. 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 $40 \times 40$ covariance matrix for the cross-section of yields, whereas my modification uses 800 observations.

3. Finally, $\Sigma$ and $\sigma^2$ are defined from the residuals of various regressions. $\Sigma$ is defined as the variance-covariance matrix from which errors $v_{t+1}$ in the state variable regression are drawn. $\sigma^2$ 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 $\sigma^2$, although I follow ACM’s methodology exactly nonetheless.

I present the results for the decomposition of the ten yield curves in Figure 20, 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 20: Decomposition of International Yield Curves

(a) Australia

(b) Canada

(c) Eurozone

(d) Japan

(e) New Zealand

(f) Norway
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.
## Robustness Checks

### Table 12: Excess Volatility in 10Y Bond Returns

<table>
<thead>
<tr>
<th></th>
<th>AUD</th>
<th>CAD</th>
<th>CHF</th>
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<td>25**</td>
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</tbody>
</table>

Notes: The table, which supports Table 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 13: Excess Volatility in Daily 1Y Bond Returns

<table>
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<tr>
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<th>AUD</th>
<th>CAD</th>
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Notes: The table, which supports Table 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 14: Excess Volatility in Daily 6F4Y Bond Returns

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</table>

Notes: The table, which supports Table 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 2023 to 2027, in 2017) are more volatile around announcements by the row central bank than at other times. The cell shows the excess ratio of standard deviations for that asset (announcement window standard deviation over non-announcement window standard deviation, minus 100%). Significance is assessed at the 1% (**) and 5% (*) level by the Brown-Forsythe test.

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

<table>
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Notes: The table, which supports Table 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 16: Excess Volatility in Daily 10Y Rate Returns

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Notes: The table, which supports Table 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 17: Excess Volatility in Daily 10Y Term Returns

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<td>17**</td>
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</table>

Notes: The table, which supports Table 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.
### Table 18: Excess Volatility in 60M Currency Returns

<table>
<thead>
<tr>
<th></th>
<th>AUD</th>
<th>CAD</th>
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<th>GBP</th>
<th>JPY</th>
<th>NOK</th>
<th>NZD</th>
<th>SEK</th>
<th>USD</th>
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</thead>
<tbody>
<tr>
<td>Australia</td>
<td>293**</td>
<td>13*</td>
<td>1</td>
<td>5</td>
<td>6</td>
<td>23*</td>
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<td>41**</td>
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</tr>
<tr>
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<td>−5</td>
<td>185**</td>
<td>5</td>
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<td>152**</td>
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<td>48**</td>
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<td>120**</td>
<td>23**</td>
<td>60**</td>
<td>70**</td>
<td>317**</td>
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</tbody>
</table>

Notes: The table, which supports Table 7, tests whether returns of the column country’s currency are more volatile around announcements by the row central bank than at other times. The cell shows the excess ratio of standard deviations for that asset (announcement window standard deviation over non-announcement window standard deviation, minus 100%). Significance is assessed at the 1% (**) and 5% (*) level by the Brown-Forsythe test. Currencies are measured against an equal-weighted basket of all other currencies, but the one whose central bank is making announcements.

### Table 19: Excess Volatility in 60M Currency Returns (F-test)

<table>
<thead>
<tr>
<th></th>
<th>AUD</th>
<th>CAD</th>
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<th>NOK</th>
<th>NZD</th>
<th>SEK</th>
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<td>23**</td>
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<td>125**</td>
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<td>37**</td>
<td>30**</td>
<td>17**</td>
<td>35**</td>
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<td>120**</td>
<td>23**</td>
<td>60**</td>
<td>70**</td>
<td>317**</td>
</tr>
</tbody>
</table>

Notes: The table tests whether returns of the column country’s currency are statistically different around announcements by the row central bank than at other times. The cell shows the excess ratio of standard deviations for that asset (announcement window standard deviation over non-announcement window standard deviation, minus 100%). Significance is assessed at the 1% (**) and 5% (*) level by the F-test. Currencies are measured against an equal-weighted basket of all other currencies, but the one whose central bank is making announcements.
Table 20: Excess Volatility in 60M Currency Returns (Kolmogorov-Smirnov test)

<table>
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<td>120**</td>
<td>23*</td>
<td>60**</td>
<td>70**</td>
<td>317**</td>
</tr>
</tbody>
</table>

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

Table 21: Pairwise Comparisons on Currency Responses to US Monetary Shocks

<table>
<thead>
<tr>
<th></th>
<th>AUD</th>
<th>CAD</th>
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<th>GBP</th>
<th>JPY</th>
<th>NOK</th>
<th>NZD</th>
<th>SEK</th>
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<td>0.132</td>
<td>0.000</td>
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<td>0.680</td>
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<tr>
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<td>0.664</td>
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<td>0.000</td>
<td>0.000</td>
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<tr>
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<td>0.013</td>
<td>0.132</td>
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<td>0.004</td>
<td>0.065</td>
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<td>0.007</td>
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<td>0.065</td>
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<td>0.001</td>
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<td>0.000</td>
<td>0.001</td>
<td>0.885</td>
<td>0.389</td>
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</tbody>
</table>

The table supports Figure 4 by implementing pairwise comparisons among the coefficients associated with each currency. Figure 4 depicts by how much the dollar appreciates against a given reference currency when it appreciates by 1% on average, following a Fed tightening. This table shows the p-values that emerge from a two-sided two-sample t-test between the relative appreciation for two reference currencies.
Figure 21: 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 22: 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 23: 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 risk-free 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 is virtually nonexistent outside the US, the right figure modifies the methodology to offset the loss of power. Specifically, it both uses US thirty-year bonds as an additional variable in the regression to offer some high-frequency identification, and it uses the regular version of the Bayesian information criterion to identify the lower-dimensional structure.

Table 22: Pairwise Comparisons on Bond Responses to US Monetary Shocks

<table>
<thead>
<tr>
<th></th>
<th>AUD</th>
<th>CAD</th>
<th>CHF</th>
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</tbody>
</table>

The table supports Figure 6 by implementing pairwise comparisons among the coefficients associated with each country’s bond yield. Figure 6 depicts by how much yields of a given country’s ten-year bonds rise when US ten-year yields rise by 1%, following a Fed tightening. This table shows the p-values that emerge from a two-sided two-sample t-test between the relative rise for the bonds of two countries.
Table 23: Pairwise Comparisons on Portfolio Responses to US Monetary Shocks

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

The table supports Figure 7 by implementing pairwise comparisons among the coefficients associated with portfolios for each country. Figure 7 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.

G Other Central Banks

Figure 24: 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 11.
Figure 25: 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 26: 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 27: 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 9.

Figure 28: 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 29: 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 10.

Figure 30: 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 asymmetrically when the Norges Bank tightens, with the possible exception of Swedish assets.
Figure 31: 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. This is a duplicate of Figure 6.

Figure 32: 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 33: Market Reactions to American Monetary Shocks

(a) Currencies

(b) Bonds

Notes: The figures depict the reactions of currency and bond markets to announcements by the Federal Reserve. The left figure shows by how much the USD appreciates against a given reference currency when it appreciates by 1% on average; and the right figure shows by how much foreign ten-year yields rise when American ten-year yields rise by 1%. Standard error bars in both pictures are computed 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 4 and 5.
H Models of Complete Markets

This appendix supports Section 6 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.

H.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.

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

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.

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

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.

\[
L_{t-1} \left( \exp \left( n\Delta y_i^t \right) \right) = L_{t-1} \left( \frac{\Lambda_{i,T}^t}{\Lambda_{i,T}^{t-1}} \right)
\]

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.
H.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.

H.2.1 Model Setup

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

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

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:

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

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

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 $e_t^g$ and an idiosyncratic component $e_t^i$. To incorporate structured heterogeneity, different countries $i$ have differential loadings $1 + \beta^g_i$ on the global components of shocks. For simplicity, $1 + \beta^g_i \geq 0$ 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 $\alpha$, yielding the updated utility function and dynamics:

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

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

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

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

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 where \( g_t \) 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 \( j \):

\[
E_{t-1} \left( \theta \log \frac{C_{it}^{t}}{C_{i,t-1}^{t}} \right)^{\theta/\psi} R_{a,t}(i)^{(1-\theta)} R_{j,t} = 1
\]

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

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

which makes the log SDF:

\[
m_{i}^{a} = \theta \log \delta - \frac{\theta}{\psi} (c_{i}^{t} - c_{i-1}^{t}) + (\theta - 1)r_{a,t}^{i} + 1/2 \psi_{t}^{i} = \theta \log \delta - \frac{\theta}{\psi} (c_{i}^{t} - c_{i-1}^{t}) + (\theta - 1)r_{a,t}^{i}
\]

H.2.2 The Consumption Asset

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

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

Second, I conjecture that the price-dividend ratio \( z_{a,t} \) is a linear function of a country’s state variables \( x_{i}^{t} \) and \( (\sigma_{i}^{t})^{2} \), as in [Bansal and Yaron 2004]:

\[ z_{a,t}^{i} = A_{0} + A_{1}x_{i}^{t} + A_{2}(\sigma_{i}^{t})^{2} \]

To solve the coefficients, I use Equation (20), which is the log Euler equation, and price the consumption asset itself \( j = a \):

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

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.
where:
\[
r_{a,t} \approx \kappa_{a,0} + \chi(z_a) \left( A_0 + A_1 \left( \rho x_{t-1} + \varphi_e \left( \sqrt{\sigma_e} \sigma(1 + \beta_i^e) e_t^z + \sqrt{1 - \alpha_e} \sigma_{t-1}^i e_t^i \right) \right) \right. \\
+ A_2 \left( \sigma^2 + v \left( (\sigma_{t-1}^i)^2 - \sigma^2 \right) \right) + \sigma_w \left( \sqrt{\sigma_w} \left( 1 + \beta_i^w \right) w_t^z + \sqrt{1 - \alpha_w} \sigma_{t-1}^i w_t^i \right) \right) \\
- \left( A_0 + A_1 x_{t-1} + A_2 \left( \sigma_{t-1}^i \right)^2 \right) + \left( \mu + \phi x_{t-1} + \left( \sqrt{\alpha_{\eta}} \sigma \left( 1 + \beta_i^\eta \right) \eta_t^z + \sqrt{1 - \alpha_{\eta}} \sigma_{t-1}^i \eta_t^i \right) \right) \right]
\]

(23)

Using Equation (23), I expand the Euler equation (22) into:
\[
\theta \log \delta + (1 - \gamma) \left( \mu + \phi x_{t-1}^i \right) + \theta \left( \kappa_{a,0} + \chi(z_a) \left( A_0 + A_1 \rho x_{t-1} + A_2 \left( \sigma^2 + v \left( (\sigma_{t-1}^i)^2 - \sigma^2 \right) \right) \right) \right. \\
- \left( A_0 + A_1 x_{t-1} + A_2 \left( \sigma_{t-1}^i \right)^2 \right) + \frac{1}{2} \theta^2 \chi(z_a)^2 A^2 \varphi_e^2 \left( \alpha_e \sigma^2 \left( 1 + \beta_i^e \right)^2 + (1 - \alpha_e) \left( \sigma_{t-1}^i \right)^2 \right) \\
+ \frac{1}{2} \theta^2 \chi(z_a)^2 A^2 \sigma_w^2 \left( \alpha_w \left( 1 + \beta_i^w \right)^2 + (1 - \alpha_w) \right) + \frac{1}{2} \left( 1 - \gamma \right) \left( \alpha_{\eta} \sigma^2 \left( 1 + \beta_i^\eta \right)^2 + (1 - \alpha_{\eta}) \left( \sigma_{t-1}^i \right)^2 \right) = 0
\]

This expression must hold for any arbitrary value of the state variables \( x_{t-1}^i \) and \( (\sigma_{t-1}^i)^2 \), 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.

\[
x_{t-1}^i : \quad \phi (1 - \gamma) + \theta \chi(z_a) A_1 \rho - \theta A_1 = 0
\]

\[
(\sigma_{t-1}^i)^2 : \quad \theta \chi(z_a) A_2 v - \theta A_2 + \frac{1}{2} \theta^2 \left( 1 - \frac{1}{\psi} \right)^2 (1 - \alpha_{\eta}) + \frac{1}{2} \theta^2 \chi(z_a)^2 A^2 \varphi_e^2 (1 - \alpha_e) = 0
\]

This yields the following solutions for the coefficients, using the approximation that \( \chi(z_a) = 1 \) over short windows:
\[
A_1 = \phi (1 - \rho)^{-1} \left( 1 - \frac{1}{\psi} \right)
\]
\[
A_2 = \frac{1}{2} (1 - \gamma) \left( 1 - \frac{1}{\psi} \right) \left( 1 - \alpha_{\eta} \right) + \left( 1 - \alpha_e \right) \phi^2 \left( \frac{\varphi_e}{1 - \rho} \right)^2
\]

With the coefficients, I can now return to Equation (23) to simplify it. I group together all constants as \( K_a \).
\[
r_{a,t}^i = K_a + \frac{1}{\psi} x_{t-1}^i - \left( \sigma_{t-1}^i \right)^2 A_2 (1 - v) + A_1 \varphi_e \left( \sqrt{\alpha_e} \sigma \left( 1 + \beta_i^e \right) e_t^z + \sqrt{1 - \alpha_e} \sigma_{t-1}^i e_t^i \right) \\
+ A_2 \sigma_w \left( \sqrt{\sigma_w} \left( 1 + \beta_i^w \right) w_t^z + \sqrt{1 - \alpha_w} \sigma_{t-1}^i w_t^i \right) + \sqrt{\alpha_{\eta}} \sigma \left( 1 + \beta_i^\eta \right) \eta_t^z + \sqrt{1 - \alpha_{\eta}} \sigma_{t-1}^i \eta_t^i
\]

(24)

H.2.3 Stochastic Discount Factor

With the expression for the consumption asset, I return to Equation (21) to generate an expression for the stochastic discount factor. That expression is written below:
\[
m_t^i = \theta \log \delta - \frac{\theta}{\psi} \left( c_t^i - c_{t-1}^i \right) + (\theta - 1) r_{a,t}^i
\]

I plug Equation (24) (along with the dynamics for consumption) to get an expression relating
the stochastic discount factor to underlying shocks:

\[ m_t = \theta \log \delta - \frac{\theta}{\psi} \left( \mu + \phi x_{t-1}^i + (\sqrt{\alpha_0} \sigma (1 + \beta_\eta^i) \eta_t^i + \sqrt{1 - \alpha_0 \sigma_{t-1}^i \eta_t}) \right) \]

\[ + (\theta - 1) \left( K_\theta + \phi \frac{1}{\psi} x_{t-1}^i - (\sigma_{t-1}^i)^2 A_1 (1 - v) + A_1 \varphi e \left( \sqrt{\alpha_e \sigma (1 + \beta_\eta^i) e_t^i + \sqrt{1 - \alpha_e \sigma_{t-1}^i e_t^i}} \right) + A_2 \sigma_w \left( \sqrt{\alpha_w (1 + \beta_w^i) w_t^i + \sqrt{1 - \alpha_w w_{t-1}^i}} \right) + \sqrt{\alpha_\eta \sigma (1 + \beta_\eta^i) \eta_t^i + \sqrt{1 - \alpha_\eta \sigma_{t-1}^i \eta_t^i}} \right) \]

This expression can be simplified, as follows.

\[ m_{t}^i = K_m - \phi \frac{1}{\psi} x_{t-1}^i + (\gamma - 1/\psi)(1 - \gamma)K_0 (\sigma_{t-1}^i)^2 \]

\[ - (1 - \rho)^{-1} (\gamma - 1/\psi) \phi \varphi_e \left( \sqrt{\alpha_e \sigma (1 + \beta_\eta^i) e_t^i + \sqrt{1 - \alpha_e \sigma_{t-1}^i e_t^i}} \right) \]

\[ - (1 - v)^{-1} (\gamma - 1/\psi)(1 - \gamma)K_0 \sigma_w \left( \sqrt{\alpha_w (1 + \beta_w^i) w_t^i + \sqrt{1 - \alpha_w w_{t-1}^i}} \right) \]

\[ - \gamma \left( \sqrt{\alpha_\eta \sigma (1 + \beta_\eta^i) \eta_t^i + \sqrt{1 - \alpha_\eta \sigma_{t-1}^i \eta_t^i}} \right) \]

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

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

### H.2.4 Long-Maturity Bonds

Long-maturity bonds can be priced similarly to the consumption asset. The major difference is that the dividend process is not a function of shocks; in this case, I set it to be a constant \( \mu_b^i \). 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 \( (\sigma_t^i)^2 \):

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

\[ z_{b,t}^i = B_0 + B_1 x_{t-1}^i + B_2 (\sigma_{t-1}^i)^2 \]

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

\[ r_{b,t}^i \approx \kappa_{b,0} + \chi(z_b) \left( B_0 + B_1 (\rho x_{t-1}^i + \varphi_e \left( \sqrt{\alpha_e \sigma (1 + \beta_\eta^i) e_t^i + \sqrt{1 - \alpha_e \sigma_{t-1}^i e_t^i}} \right) \right) \]

\[ + B_2 \left( \sigma^2 + v \left( (\sigma_{t-1}^i)^2 - \sigma^2 \right) + \sigma_w \left( \sqrt{\alpha_w (1 + \beta_w^i) w_t^i + \sqrt{1 - \alpha_w w_{t-1}^i}} \right) \right) \]

\[ - \left( B_0 + B_1 x_{t-1}^i + B_2 (\sigma_{t-1}^i)^2 \right) + \mu_b^i \]

As before, I use the Euler equation, Equation (20), to identify the coefficients \( B_1 \) and \( B_2 \):

\[ \mathbb{E}_{t-1} m_t^i + \mathbb{E}_{t-1} r_{b,t}^i + \frac{1}{2} \nabla_t-1 (m_t^i + r_{b,t}^i) = 0 \]
which expands to the following:

\[
\kappa_{b,0} + \chi(z_b) \left( B_0 + B_1 \rho x_{t-1}^i + B_2 \left( \sigma^2 + v \left( (\sigma_{t-1}^i)^2 - \sigma^2 \right) \right) \right) - \left( B_0 + B_1 x_{t-1}^i + B_2 (\sigma_{t-1}^i)^2 \right) + \mu_b^i + K_m \\
- \frac{\phi}{\psi} x_{t-1} + (\gamma - 1/\psi)(1 - \gamma) K_0 (\sigma_{t-1}^i)^2 + \frac{1}{2} \left( \alpha_e \left( \chi(z_b) B_1 - (1 - \rho)^{-1}(\gamma - 1/\psi) \phi \right)^2 \varphi_{x}^e \sigma^2 (1 + \beta_e^i)^2 \\
+ (1 - \alpha_e) \left( \chi(z_b) B_1 - (1 - \rho)^{-1}(\gamma - 1/\psi) \phi \right)^2 \varphi_{e}^x (\sigma_{t-1}^i)^2 + \alpha_{\phi} \gamma^2 \sigma^2 (1 + \beta_{\phi}^i)^2 + (1 - \alpha_{\phi}) \gamma^2 (\sigma_{t-1}^i)^2 \\
+ \alpha_w \left( \chi(z_b) B_2 - (1 - v)^{-1}(\gamma - 1/\psi)(1 - \gamma) K_0 \right) (\sigma_{t-1}^i)^2 + (1 - \alpha_{\psi}) \left( \chi(z_b) B_2 - (1 - v)^{-1}(\gamma - 1/\psi)(1 - \gamma) K_0 \right) (\sigma_{t-1}^i)^2 \right) = 0
\]

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

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

\[
(\sigma_{t-1}^i)^2 : \quad \chi(z_b) B_2 v - B_2 + (\gamma - 1/\psi)(1 - \gamma) K_0 \\
+ \frac{1}{2} (1 - \alpha_e) \left( \chi(z_b) B_1 - (1 - \rho)^{-1}(\gamma - 1/\psi) \phi \right)^2 \varphi_{x}^e + \frac{1}{2} (1 - \alpha_{\phi}) \gamma^2 = 0
\]

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

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).

\[
\tilde{r}_{b,t}^i = K_b + \frac{\phi}{\psi} x_{t-1}^i - (\gamma - 1/\psi + \gamma/\psi) K_0 (\sigma_{t-1}^i)^2 \\
- (1 - \rho)^{-1}(1/\psi) \phi \varphi_e \left( \sqrt{\alpha_e} \sigma (1 + \beta_e^i) e_t^i + \sqrt{1 - \alpha_e} \sigma_{t-1}^i \epsilon_t^i \right) \\
- (1 - v)^{-1}(1/\psi - \gamma - \gamma/\psi) K_0 \sigma_w \left( \sqrt{\alpha_w} (1 + \beta_w^i) w_t^i + \sqrt{1 - \alpha_w} w_{t-1}^i \right)
\]