Risk, Monetary Policy and Asset Prices in a Global World*

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Abstract

We study how monetary policy and risk shocks affect major asset prices (short-term interest rates, stocks, long-term bonds) in three large economies: the US, the euro area, and Japan. Using a high-frequency framework, we fail to find evidence in favor of monetary policy driving asset price cycles through a risk channel. Instead, there is a strong global common component in risk shocks which is not driven by monetary policy. Comparing the impact of monetary policy and risk shocks on asset prices across countries, we document that monetary policy spillovers are economically relatively more (less) important for interest rates and bond prices (stock prices) than risk shocks. The US generates relatively important monetary policy spillovers, but information shocks emanating from the euro area produce relatively the strongest effects on international stock and bond markets. We provide suggestive evidence that monetary policy effects on asset prices may reflect a persistent direct interest rate effect rather than a risk premium effect.

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1 Introduction

Since the global financial crisis, there has been renewed interest in understanding how monetary policy shocks transmit across countries through financial markets and capital flows. The increased synchronization of financial cycles across countries in recent decades (Jordà, Schularick, Taylor, and Ward (2019)) generates the specter of a "hegemon" country, such as the US, whose monetary policy drives risk appetite and thus asset prices globally (Miranda-Agrippino and Rey (2020)). It is therefore not surprising that the Fed Chairman Jerome Powell recently devoted a speech to the topic, arguing that: "... while global factors play an important role in influencing domestic financial conditions, the role of US monetary policy is often exaggerated."[1] After all, in globally integrated capital markets, financial risk conditions and therefore asset returns may naturally comove strongly.

Our analysis takes a less US-centric perspective and assesses the transmission of monetary policy as well as risk shocks to asset prices across three advanced economies, the US, euro area, and Japan, using high-frequency (daily) data over the 2000-2015 period. In this multi-country, multi-shock framework, we fail to find evidence in favor of monetary policy driving asset price cycles through a risk channel. Instead, we find a strong global common component in risk shocks which is not driven by monetary policy. Different asset classes exhibit differential responses to shocks, with monetary policy spillovers being economically more important for interest rates and bond prices, and risk shocks being relatively more important for stock prices. While the US generates relatively important monetary policy spillovers through short- and long-term interest rates, the strongest spillover effects on international stock and bond markets emanate from euro area information shocks which reveal central bank information about the economy. In all, our results lend support to Mr. Powell's conclusion.

Our analysis faces two key measurement issues. In terms of monetary policy shocks, we use the framework of Jarociński and Karadi (2020) to separate measures of policy shocks into pure information shocks and a residual, representing actual monetary policy shocks. In terms of risk variables, we recognize that risk can vary because of changes in the risk aversion of market participants, or changes in the amount of risk (uncertainty). It has become more widely accepted in finance and economics that aggregate risk aversion changes through time; habit models (see e.g. Campbell and Cochrane (1999); Fuhrer (2000)), featuring stochastic risk aversion, are now prevalent. While it is impossible to define risk aversion in a model-free

[1]Speech by Chairman Jerome Powell on "Monetary Policy Influences on Global Financial Conditions and International Capital Flows," at the Eighth High-Level Conference on the International Monetary System sponsored by the International Monetary Fund and Swiss National Bank, Zurich, Switzerland, May 8, 2018.

way, recent research in finance suggests that equity options markets harbor much market-based information on risk aversion. Martin (2017) shows that an option-implied volatility index constitutes a lower bound for the equity premium. Bekaert, Engstrom, and Xu (2020) estimate a measure of aggregate risk aversion, pricing equities and corporate bonds, and find it to be highly correlated with the variance risk premium, essentially the premium earned for selling a claim on realized equity variances. To separate the two risk concepts at the country level, we decompose option-implied variances of country-level stock returns into a conditional variance component and a variance risk premium component, building on Bekaert and Hoerova (2014). For each country, the conditional variance of stock returns (a proxy for uncertainty) is computed using a novel non-linear forecasting model. The variance risk premium component, computed as option-implied variance minus the uncertainty measure, is our proxy for risk aversion. Of course, variation in this measure can also reflect other sources of a wedge between "physical" and "risk-neutral" volatility, such as Knightian uncertainty (Drechsler (2013)).

We then examine how monetary policy (MP) shocks in the US, the euro area, and Japan affect stock market risk aversion and uncertainty across countries on a daily basis, while controlling for macroeconomic shocks. By considering both domestic and foreign monetary policy shocks, and by using data at the daily frequency, we complement the work by Miranda-Agrippino and Rey (2020) who focus on the effect of US monetary policy shocks on global risk and domestic business cycles at the quarterly frequency.[2] They find that monetary policy in the US, the centre country of the international monetary system, has large spillovers to the rest of the world by driving the "Global Financial Cycle."[3] The global financial cycle is reflected in strong comovements of financial asset prices across countries as driven by aggregate risk appetite in international financial markets. More generally, our findings here contribute to the question how central the US is to international financial spillovers. Cerutti, Claessens, and Rose (2019), for example, show that common shocks (such as those emanating from a central country like the US) drive little of the variation in global capital flows. However, Brusa, Savor, and Wilson (2020) claim that US monetary policy is unique in generating global equity announcement premiums.[4]

[2]Ca'Zorzi, Dedola, Georgiadis, Jarociński, Stracca, and Strasser (2020) use monthly data to compare the international transmission of monetary policy of the Fed and the ECB. They document a relatively larger impact of Fed monetary policy on euro area financial markets and real activity as well as on real and financial variables in the rest of the world.

[3]Jiang, Krishnamurthy, and Lustig (2020) build a model to rationalize an outsized impact of U.S. monetary policy shocks on the world economy based on the special demand for dollar safe assets.

[4]We focus on spillovers across advanced economies (US, euro area, Japan). Recent literature documents that monetary policies in advanced economies may have important effects on emerging markets; see, e.g., Kalemli-Özcan (2019) and the references therein. Ehrmann and Fratzscher (2009) study the transmission of US monetary policy shocks to global equity markets, documenting that the degree of global integration of countries is a key determinant for the transmission process.

We further compare the effects of monetary policy shocks and and non-monetary policy-driven "cleansed" risk and uncertainty shocks on asset prices across the three major economies. In financially integrated markets, changes in global risk aversion should affect all financial markets. Moreover, the relative importance of monetary policy versus risk aversion shocks may differ across equities, bonds and other asset classes. Such analysis is missing in the recent papers by Miranda-Agrippino and Rey (2020) and Passari and Rey (2015). For example, Xu (2019) documents that a US-based risk aversion index is a major source of international stock return comovements, but explains less of bond return comovements.

Our main results are as follows. First, we do not find a strong effect of monetary policy shocks on either risk aversion or uncertainty for the three countries we examine. We do not find evidence for cross-country effects either. In contrast, risk aversion and uncertainty comove strongly across countries.

Second, while monetary policy shocks have their usual effect on domestic short-term interest rates, with strong and statistically significant pass-through, there are also economically important international spillover effects. It is not just US monetary policy affecting interest rates in other countries (this is mostly true for the euro area), but also other countries affecting the US. For example, euro area information shocks have strong spillover effects to US and Japanese interest rates. Non-monetary policy-driven uncertainty and risk aversion shocks affect interest rates negatively, but these effects are economically small. Together with our first point above, we argue that the monetary policy effects on asset prices may well reflect a persistent pure interest rate effect, consistent with Binsbergen (2020) who argues against an important role for equity risk premiums in stock returns.

Third, we do not find a special role for US monetary policy when it comes to its effect on equity prices either. US monetary policy very significantly affects US stock returns, but the effect is only marginally economically more important than the effect observed in the euro area. Spillover effects for both pure monetary policy and information shocks are stronger from the euro area to the US than from the US to the euro area. The effects from the cleansed risk aversion and uncertainty shocks on stock prices are of a larger economic magnitude than are pure monetary policy effects, but of the same order of magnitude than are information shock effects.

Fourth, the monetary policy effects on bond returns mostly mirror those of interest rates, with information shocks being less potent than pure monetary policy shocks. The effects of uncertainty and risk aversion shocks on bond markets are in general not very substantial, and less important than monetary policy shocks.

Fifth, the exchange rate effects of monetary policy are consistent with the predictions of uncovered interest parity, whereas the uncertainty and risk aversion effects largely confirm the safe haven status of the Japanese yen. For commodity prices, we find that uncertainty and risk aversion shocks everywhere drive commodity prices down while monetary policy shocks have no effect. Gold prices cleansed of general commodity price changes are negatively affected by US and euro area monetary policy tightening, which leads to higher real interest rates. This is consistent with the idea that gold competes with other safe assets (such as short term money market deposits and perhaps bonds). Gold prices also tend to increase following uncertainty shocks.

Finally, we estimate a daily "global" risk aversion measure that is cleansed of monetary policy effects and macroeconomic news and controls for different time zones. We correlate this cleansed global risk aversion measure with the "global financial cycle" measure constructed by Miranda-Agrippino and Rey (2020). Interestingly, we find the correlation between the two global risk measures to be very high at around 0.7, suggesting that both measures are driven by common factors other than US monetary policy.

Our research relates to a voluminous empirical literature on international spillovers of monetary policies to financial asset prices. Many contributions focus on the spillovers of US monetary policy (e.g., Kim (2001), Faust, Rogers, Swanson, and Wright (2003), Ehrmann and Fratzscher (2005, 2009), Faust, Rogers, Wang, and Wright (2007), Ammer, Vega, and Wongswan (2010), Hausman and Wongswan (2011) among many others). Some papers also consider spillovers to the US, following monetary policy actions of other central banks; e.g., Ehrmann and Fratzscher (2005) analyze ECB's actions, while Craine and Martin (2008) consider Australian monetary surprises. Kearns, Schrimpf, and Xia (2020) examine the interest rate spillovers from seven advanced economy central banks to the rest of the world.

Another literature analyzes channels of international transmission of financial shocks and the role that US monetary policy plays in such transmission. Bruno and Shin (2015a,b) document that a contractionary shock to US monetary policy leads to a decrease in cross-border banking capital flows and a decline in the leverage of international banks. Such a decrease in bank capital flows is associated with an appreciation of the US dollar. Cetorelli and Goldberg (2012) and Buch, Bussiere, Goldberg, and Hills (2019) focus on the bank lending channel, showing, inter alia, that global banks can partially insulate from monetary policy shocks through internal capital markets (see Morais, Peydró, Roldán-Peña, and Ruiz-Ortega (2019); Schmidt, Caccavaio, Carpinelli, and Marinelli (2018), for additional contributions). Durdu, Martin, and Zer (2019) show that a contractionary shock to US monetary policy can lead to capital outflows in other countries due to search-for-yield incentives, and may increase the probability of a banking crisis. Jordà, Schularick, Taylor, and Ward (2019) document that the comovement in credit, house prices, and equity prices across 17 advanced economies has reached historical highs in the past three decades. They highlight the role of equity risk premia in driving the equity market synchronization. Relating to all these papers, we study spillovers across three major advanced economies and assess the relative importance of monetary policy shocks and non-MP-driven risk shocks in driving asset prices.

It is noteworthy that our sample period also covers the period of the Global Financial Crisis and its aftermath, when conventional monetary policy across the world operated at the effective lower bound and central banks employed a range of unconventional monetary policies to support the economy. To analyze the spillovers of unconventional policies across our three advanced economies, we analyze the period from 2008 onwards separately, and employ the policy shock measures developed by Rogers, Scotti, and Wright (2018).

The remainder of the paper is organized as follows. Section 2 describes the conceptual and empirical framework. Section 3 describes the estimation of risk aversion and uncertainty across the three economies and examines the direct (within country) and spillover (cross country) effects of monetary policy shocks on country risk aversion and uncertainty. Section 4 distinguishes between the effects of monetary policy shocks and non-monetary policy-driven risk shocks on interest rates, and stock and bond returns. Section 5 provides additional results on exchange rates, gold prices, and commodity prices, and considers the effects of the post-2008 unconventional monetary policies. Section 6 constructs a high-frequency global risk aversion measure from the country by country non-MP-driven risk aversion shocks, cleansed of monetary policy and macro news influences. The estimation uses a parsimonious global factor model where the global factor news is realized during the day as stock markets open and close in different time zones. Section 7 concludes.

2 Conceptual and Empirical Frameworks

In this section, we first provide a simple conceptual framework in which to interpret our empirical work. We then present the econometric framework we use to gauge the effects of monetary policy and risk shocks.

2.1 Conceptual framework

2.1.1 Domestic policy effects

To provide some conceptual background to our empirical work, Appendix A considers a simple habit model with stochastic risk aversion building on Bekaert, Engstrom, and Xing (2009) (BEX henceforth). In this model, the short-term real interest rate, rf_t, is given by:

$$rf_t = \underbrace{\phi_g g_t + \phi_{RA} RA_t + \phi_{UC} UC_t}_{rf_t^*} + \phi_{MP} MP_t, \quad (1)$$

where g_t represents expected consumption growth; RA_t is a state variable measuring stochastic risk aversion and UC_t is a state variable measuring the time-varying uncertainty of aggregate consumption growth; MP_t is a monetary policy shock. The first three terms represent the equilibrium real interest rate, y_t^*, with the effects well understood. Better growth prospects increase the interest rate ($\phi_g > 0$), precautionary savings effects imply that uncertainty lowers interest rates ($\phi_{UC} < 0$), but risk aversion may increase or decrease the interest rate depending on whether intertemporal smoothing or precautionary savings effects dominate (see also Wachter (2006)). We do not endogenize monetary policy in the model, but monetary policy can affect the short-term interest rate in four ways. It can work through a risk channel by affecting RA_t and UC_t (see Borio and Zhu (2012) for a survey of various economic mechanisms leading to such a link). For example, low interest rates may boost asset and collateral values, thus relaxing VaR constraints of banks (Miranda-Agrippino and Rey (2020)); lower banks' cost of taking leverage, resulting in lower risk premia and higher asset prices (Drechsler, Savov, and Schnabl (2018b)); and even cause banks to relax lending standards (Dell'Ariccia and Marquez (2006), Dell'Ariccia, Laeven, and Suarez (2017)). Rajan (2006) discusses how lax monetary policy may prompt asset managers to "search for yield". Central bank interest rate changes may also reveal the bank's reaction function to market stress (or lack thereof) and consequently change market behavior. Monetary policy can also affect growth expectations g_t when it releases new information, see Gürkaynak, Sack, and Swanson (2005), Nakamura and Steinsson (2018), Jarociński and Karadi (2020) and Miranda-Agrippino and Ricco (forthcoming). Finally, there could be a direct pass-through effect which we model through the MP_t state variable. The indirect effects through g_t, RA_t and UC_t imply that the ϕ_{MP} coefficient does not necessarily measure the full extent of interest rate pass-through.[5]

[5]See, e.g., Pflueger and Rinaldi (2020) for a model with stochastic risk aversion and monetary policy, estimated to match quarterly macroeconomic moments.

The model also features closed form solutions for bond prices and equity prices (more precisely, price-dividend ratios). Long term bond prices are fully determined by long-term interest rates and to a first order approximation, the bond return equals duration times the change in long term interest rates. The long-term interest rate can be decomposed in an Expectation Hypothesis (EH) term, reflecting expectations of future short-term interest rates, and a term premium. Because of mean reversion, a standard monetary policy effect should pass through less strongly to long-term interest rates via the EH effect. However, there can also be indirect effects, through the term premium, which is given by:

$$tp_t \equiv \eta_{RA} RA_t + \eta_{UC} UC_t. \tag{2}$$

Unfortunately, neither η_{RA} nor η_{UC} can be signed. In BEX's application to US interest rates, both coefficients are positive.

For equity returns, monetary policy can produce a discount rate effect directly through its effect on interest rates. With interest rates now at all-time lows, Binsbergen (2020) claims that US equity returns in the last 20 years have not outperformed long-duration fixed income portfolios, suggesting that direct interest rate effects may now be stronger. The standard interpretation of monetary policy effects on stock returns is that they operate through the risk premium (see Bernanke and Kuttner (2005)). In the model, the equity premium is entirely driven by RA_t and UC_t, such that:

$$ep_t \approx \kappa_{RA} RA_t + \kappa_{UC} UC_t. \tag{3}$$

While it is theoretically not guaranteed, BEX find κ_{RA} and κ_{UC} to be positive. Thus, monetary policy can affect stock returns through a direct interest rate effect, through indirect interest rate effects (via g_t, RA_t or UC_t) or through a risk premium effect. However, all these discount rate effects move stock prices in the same direction. Lastly, stocks also react to cash flow news. If we assume that cash flows are directly related to expected growth g_t, there is yet another effect of monetary policy, through the (new) information it releases about the economy. Information shocks have the opposite effect on stock prices than do pure monetary policy shocks, as an increased interest rate here signals positive news about the economy, which should increase stock prices. As is always true, the discount rate effects naturally imply mean-reverting behavior in returns, whereas the cash flow effects ought to be permanent.

2.1.2 International spillovers and asset return comovements

The model above outlines monetary policy and risk effects within a closed economy environment. Of course, the US, euro area and Japan operate in a world economy, where shocks may spill over to other countries. As standard monetary policy seeks to affect short-term interest rates, a discussion of monetary policy spillovers always starts with how interest rate spillovers may occur, viewed from the perspective of standard trilemma theory.

The trilemma states that economies cannot simultaneously control monetary policy, the exchange rate and capital flows (see, e.g., Obstfeld, Shambaugh, and Taylor (2005); Klein and Shambaugh (2015); Aizenman, Chinn, and Ito (2016); Bekaert and Mehl (2019); Jordà, Schularick, and Taylor (2020)). Given that the exchange rates between our three countries are flexible, and capital is mobile, the standard theory implies that monetary authorities should be able to achieve autonomy and no interest rate spillover must happen. However, a variety of alternative economic channels can still lead to short-term interest rate spillovers.[6] For example, monetary policy can reveal information about economic conditions (information about g_t in our model) or affect financial conditions (e.g., uncertainty driving precautionary savings effects, UC_t in our model), see Kearns, Schrimpf, and Xia (2020) for a survey. Such monetary policy effects operating through interest rates obviously may have repercussions for international asset prices, including exchange rates.

However, recent literature argues that the classic trilemma may have morphed into a dilemma between financial openness and monetary policy autonomy. Rey (2015), Bruno and Shin (2015a,b), and Passari and Rey (2015) stress the critical role played by the US dollar and US monetary policy in setting global liquidity and credit conditions (see also Obstfeld (2015) for a discussion). They suggest that non-US central banks have lost their ability to influence domestic interest rates, even in the presence of flexible exchange rates, due to the existence of "US-driven" global financial cycles in liquidity and credit. The main claim in these articles is that monetary policy spillover happens through a risk channel; in particular, Miranda-Agrippino and Rey (2020) trace the effects of their measure of risk aversion on a common component in international asset returns (prices).

Of course, in a financially integrated world, asset prices around the world should comove more or less strongly depending on the global nature of the shocks we outlined in Section 2.1.1. Global shocks to growth prospects g_t, risk RA_t and uncertainty UC_t should

[6] Jotikasthira, Le, and Lundblad (2015) and Bekaert and Ermolov (2019) in fact show that nominal interest rates are highly correlated across countries

induce comovements and such comovements may be different across asset classes. For stock returns, comovements have increased substantially in recent times (see Bekaert and Mehl (2019); Christoffersen, Errunza, Jacobs, and Langlois (2012); Jordà, Schularick, Taylor, and Ward (2019)), but Xu (2019) shows that government bond returns across countries show much less correlation than do stock returns. Baele, Bekaert, and Inghelbrecht (2010) show that the correlation between stock and bond returns in the US exhibits dramatic time variation (from very negative to substantially positive) and has decreased over time. Their results extend to other countries (see Baele, Bekaert, Inghelbrecht, and Wei (2020)). Chaieb, Errunza, and Gibson Brandon (2020) show that bond market integration is far from complete, confirming earlier results by Barr and Priestley (2004). We therefore deem it important to investigate various asset returns separately.

In our empirical framework, presented in the next subsection, we test the effects of monetary policy and risk shocks on asset prices domestically, how they spill over to other countries. Note that in a fully internationally integrated world, CAPM intuition would indicate that the US *should be* the hegemon country. For equity markets, for example, the US represents about 40% of the world's market capitalization. Therefore, any shock affecting the US equity market should spill over strongly to other countries through simple "beta" effects. With Japan and the euro area each representing less than 10% of world market capitalization, the corresponding reverse effects ought to be small. Whereas these are partial equilibrium relations, we nonetheless use them to help interpret our empirical work.

2.2 Empirical framework and hypotheses

Monetary policy shocks are best identified using high frequency data. Since our interest is in the impact on asset prices – which move fast in response to shocks – we conduct our tests mostly using daily data, considering longer term effects briefly in Section 4.

We first test the risk channel of monetary policy with the following regression:

$$\Delta X_{j,t} = \alpha_j + \sum_{i=US,EA,JP} \beta_j^{MP,i} MP_t^i + \sum_{i \neq j} \beta_j^{X,i} \overline{x_t^i} + \sum_{i=US,EA,JP} \gamma_j^i D_{t,i} + \sum_{k=Macro} \delta_j^k Macro_t^k + \sum_{k=Macro} r_j^k \theta_t^k + \varepsilon_{j,t}, \tag{4}$$

where $\Delta X_{j,t}$ represents either first-differenced RA or UC for countries $j = US, EA, JP$. We devote Section 3.1 to computing proxies for UC and RA. MP_t^i stands for the monetary policy shock series in country i on day t, either representing the standard monetary policy shock (with results reported in the Online Appendix), or a vector of a central bank information shock and a "pure" MP shock purged of the effects of information releases. $Macro_t^k$ represent 21 macroeconomic news series to control for variation in g_t and for other macro news at the daily level.

D_t^i represents dummies equal to 1 for MP event days in $i = US$, EA, JP. Analyses using high frequency identification of monetary policy shocks are often run only on the event dates. However, because we contrast the effects of monetary policy induced with non-MP induced shocks here (represented by $\overline{x_t}$), we choose to use all of the data. Including the monetary policy day dummies ensures that the results we obtain using these daily regressions are very similar to "event-only" regressions. Moreover, it is conceivable that the mere release of information, irrespective of the sign or the magnitude of the shock affects uncertainty as information is released to the markets. Brusa, Savor, and Wilson (2020) claim that, following US monetary policy shocks, global stock market returns increase while uncertainty decreases worldwide; but this does not happen on policy days for other countries. Similarly, θ_t^k represent 21 dummies for the macroeconomic news release days.

The main coefficients of interest are the $\beta_j^{MP,i}$s which measure the domestic and spillover effects of monetary policy shocks on X: domestic effects through $\beta_j^{MP,j}$, and spillover effects through $\beta_j^{MP,i}$, $i \neq j$. The second set of coefficients of interest are the $\beta_j^{X,i}$ coefficients. The $\overline{x_t}$ variables (for x=ra or uc) represent "cleansed" risk aversion or uncertainty shocks, that is: $\overline{x_t^i} = \Delta X_{i,t} - E[\Delta X_{i,t}|\boldsymbol{z_t}]$, where the set of $\boldsymbol{z_t}$ instruments include monetary policy and macro shocks, event day dummies and macroeconomic announcement dummies. The expectation is evaluated by a linear projection. For example, for risk aversion, this procedure cleanses risk aversion changes from any monetary policy influences, but it also removes the effects of the extensive set of macroeconomic announcements occurring around the world on risk aversion shocks. As a result, this residual is the non-MP-driven, cleansed risk aversion shock (denoted by $\overline{ra^i}$, $i = US$, EA, JP,[7] and labeled as "non-MP RA shocks" in tables). The procedure is analogous for uncertainty.

One last challenge our analysis must overcome, given its high-frequency nature, is the non-synchronous trading schedules of the three parts of the world economy. The rule of thumb is that subscript t is adjusted to reflect the information set of X_t. In particular, for the US, all US and foreign MP and macroeconomic shocks enter contemporaneously, except for those shocks that are released after the US market closes (those only enter the information set on the next trading day). For the euro area, JP and EA shocks that materialize before or during the European opening hours enter contemporaneously while the other shocks as well as the

[7]Lower case ra is used to differentiate this shock variable with the level variable denoted using upper case RA.

US shocks enter the information set on the next trading day. For Japan, JP shocks that materialize while Japanese financial markets are open enter contemporaneously, while the EA and US shocks dated on the same day enter the information set on the next trading day.

The presence of the cleansed risk aversion or uncertainty shocks from other countries aids the identification of monetary policy shock effects on risk. Imagine a typical US monetary policy announcement on day t, which tends to happen in the early afternoon, US time (GMT-5). The daily US risk aversion change may be influenced by events earlier in the day, during European or Japanese market hours. The presence of $\overline{ra_t^{EA}}$ and $\overline{ra_t^{JP}}$ (or their uc counterparts) controls for these events. At the same time, it reveals how global risk aversion and uncertainty travel across time zones, which we further exploit in Section 6.

The remainder of the analysis focuses on the effect of risk shocks (risk aversion and uncertainty), monetary policy shocks (split up in pure MP and information shocks) and macro-economic announcements on interest changes and several asset returns with the following regression set-up:

$$\begin{aligned} Y_{j,t} = \alpha_j + \sum_{i=US,EA,JP} \beta_j^{MP,i} MP_t^i + \sum_{i=US,EA,JP} \beta_j^{RA,i} \overline{ra_t^i} + \sum_{i=US,EA,JP} \beta_j^{UC,i} \overline{uc_t^i} \\ + \sum_{i=US,EA,JP} \gamma_j^i D_{t,i} + \sum_{k=Macro} \delta_j^k Macro_t^k + \sum_{k=Macro} r_j^k \theta_t^k + \varepsilon_{j,t} \end{aligned} \tag{5}$$

where $Y_{j,t}$ is a financial variable (e.g., changes in interest rates, stock returns etc.) in country j. The main coefficients of interest are the $\beta_j^{MP,i}$, $\beta_j^{RA,i}$ and $\beta_j^{UC,i}$ coefficients so that we can contrast the effects of monetary policy shocks with the effects of non-monetary policy-driven risk aversion and uncertainty shocks.

3 Monetary policy, risk aversion and uncertainty

In this section, we first propose and estimate measures of financial risk aversion and uncertainty, building on the work by Bekaert and Hoerova (2014). They use a decomposition of the squared VIX index, derived from US S&P500 options prices, into the conditional variance of stock returns (a measure of variance forecast or stock market uncertainty, or "UC") and a variance risk premium (a measure of stock market risk aversion, or "RA").[8] Our measures embed a new feature designed to capture parameter (persistence) changes when variance levels increase dramatically (as during the Global Financial Crisis). We then conduct an extensive

[8]We opt to measure uncertainty for the stock market rather than from data on real activity as in the model set out in Section 2, because, obviously, such uncertainty is not observed at high frequencies. Of course, in equilibrium, stock market uncertainty depends on both fundamental uncertainty and on risk aversion.

model selection exercise to obtain the best RA and UC measures for each country. We describe the empirical model to determine UC_t in Section 3.1, and discuss the empirical results in Section 3.2.

In Section 3.3, we then estimate how monetary policy affects risk and uncertainty, using Equation (4). Our analysis provides a direct test of the first building block of a US monetary policy-induced global financial cycle: Does US monetary policy affect stock market risk aversion in the US and other countries? By considering both domestic and foreign monetary policy shocks, and by using data at the daily frequency, we complement the work by Miranda-Agrippino and Rey (2020) who focus on the role of US monetary policy shocks for global financial variables and domestic business cycles at the quarterly frequency. In addition, we contrast the effects of monetary policy on risk aversion versus uncertainty.

3.1 Empirical model for uncertainty and risk aversion

Our starting point are option-implied or "risk-neutral" volatility indices, which can be inferred from option prices (see Britten-Jones and Neuberger (2000) and Bakshi, Kapadia, and Madan (2003)). For example, the VIX index calculation uses a weighted average of European-style S&P500 call and put option prices that straddle a 30-day maturity (22 trading days) and cover a wide range of strikes (see CBOE (2004) for more details). Importantly, this estimate is model-free and does not rely on an option pricing model (see e.g. Bakshi and Madan (2000)). The implied volatility indices for the euro area and Japan are constructed using the same methodology.

While the implied volatility index obviously reflects stock market uncertainty (the expected "amount of risk" over the next 22 trading days), it conceptually must also harbor information about risk aversion. To isolate this component more correlated with risk aversion, we compute the variance premium; the difference between the squared risk-neutral volatility index and an estimate of the "physical" conditional variance (see, e.g., Carr and Wu (2009)). With $IV_{i,t}$ the option-implied variance of the stock market index in country i for contracts, at time t, with a maturity of one month, we therefore posit that:

$$IV_{i,t} = UC_{i,t} + RA_{i,t}, \tag{6}$$

where $UC_{i,t}$ is the conditional physical expectation of realized stock market variance in country i over the next month (22 trading days); and $RA_{i,t}$ constitutes our measure of stock market risk aversion.

To estimate $UC_{i,t}$, we project future realized monthly variances onto a set of current instruments. One benchmark model is proposed in Bekaert and Hoerova (2014), who conduct a horserace between a number of forecasting models:

$$UC_{i,t} \equiv E_t \left[RV_{i,t+22}^{(22)} \right] = \alpha_i + \beta_i^m RV_{i,t}^{(22)} + \beta_i^w RV_{i,t}^{(5)} + \beta_i^d RV_{i,t} + \gamma_i IV_{i,t}. \tag{7}$$

The country indicator i is omitted henceforth for simplicity. In the regression equation above, $RV_{t+22}^{(22)}$ denotes the monthly (22 trading days) stock market realized variance from day $t+1$ to day $t+22$; $RV_t^{(22)}$ denotes the past 22-day (monthly) realized variance from day $t-21$ to day t, $RV_t^{(5)}$ the past 5-day (weekly) realized variance from day $t-4$ to day t, and RV_t the past daily realized variance from day $t-1$ to day t. The use of realized variances at the monthly, weekly and daily frequencies was first proposed by Corsi (2009). The last variable IV_t is expressed in monthly variance units to be of comparable magnitude to RV_t. For example, with the VIX_t expressed in annualized percent, $IV_t = \frac{VIX_t^2}{120000}$. The 5- and 1-day variance are scaled up to be comparable with $RV_t^{(22)}$.

In addition, we consider a new variant of this baseline model that allows the coefficients to depend on the levels of the realized variables:

$$\begin{aligned} UC_t \equiv E_t \left[RV_{t+22}^{(22)} \right] &= \alpha + \frac{\exp(\beta_0^m - \beta_1^m Z_t^m)}{\exp(\beta_0^m - \beta_1^m Z_t^m) + 1} RV_t^{(22)} + \frac{\exp(\beta_0^w - \beta_1^w Z_t^w)}{\exp(\beta_0^w - \beta_1^w Z_t^w) + 1} RV_t^{(5)} \\ &+ \frac{\exp(\beta_0^d - \beta_1^d Z_t^d)}{\exp(\beta_0^d - \beta_1^d Z_t^d) + 1} RV_t + \frac{\exp(\gamma_0 - \gamma_1 Z_t^{iv})}{\exp(\gamma_0 - \gamma_1 Z_t^{iv}) + 1} IV_t, \end{aligned} \tag{8}$$

where the instrument Z_t^m is $RV_t^{(22)}$, Z_t^w is $RV_t^{(5)}$, Z_t^d is RV_t, and Z_t^{iv} is IV_t; as before, parameters and variances are all country-specific. Economically, such non-linear coefficients (a logistic function here) help capture sudden changes in mean reversion in crisis times. For example, when a particular month witnesses tremendous volatility, resulting in high realized variances, it is quite likely that such high variance realization does not persist in the same fashion as does a moderate realization. The estimation uses the standard procedure that minimizes the sum of squared residuals and the longest daily sample possible for each country.

3.2 Uncertainty and risk aversion estimates

We focus on three countries / economic areas: United States (US), euro area (EA), and Japan (JP). Daily option-implied volatilities for the stock market indices in the three countries are obtained from DataStream. The availability of the implied volatility series determines the starting points of our samples across the various countries. Specifically, the data start on November 1, 1989 for Japan; on January 2, 1990 for the United States; on January 4, 1999 for the euro area. To construct the longest possible daily realized variance series, we use a proprietary dataset of high-frequency 5-minute stock market return data for the United States (available to us for the 1990-2010 period); we use daily aggregate market return data from DataStream for 11 years (1989-1999) for Japan and for one year (1999) for the euro area, and we use daily realized variance data representing the sum of squared 5-minute returns from Oxford-Man Institute post-2000 for Japan and the euro area and post-2010 for the United States. All daily variance series are in monthly (squared) units. The data covers the period until May 31, 2018.

The Appendix contains detailed estimation and model selection results; here, we provide a short summary. For the linear models, we consider 5 models: the martingale model of Bollerslev, Tauchen, and Zhou (2009) (no parameters to be estimated), an autoregressive model (2 coefficients including the constant), a model that uses the past monthly realized variance and the past implied variance (3 coefficients including the constant), the Corsi (2009) model (4 coefficients including the constant), and finally the full model with 5 coefficients.

Using the BIC criterion, the full model is the best linear model. In Table 1, we show the parameter estimates for the best linear model for each country. We find that, for the US, all coefficients are statistically significant and positive, which confirms the results in Bekaert and Hoerova (2014). The past monthly and weekly realized variances have similar and strong predictive power, given the similar coefficients. For the euro area, the past weekly realized variance exhibits stronger predictive power than past realized variances at other horizons; the past implied variance, however, is clearly the most important variance predictor. For Japan, only the three realized variance coefficients are statistically different from zero.

For the non-linear model, we consider and estimate all 15 possible models, reflecting the various combinations of time-varying and constant coefficients. The best non-linear model for each country is chosen based on the BIC criterion as well. Our analysis here extends analyses on international variance premiums by Londono (2015) and Bollerslev, Marrone, Xu, and Zhou (2014), who use daily data to construct realized variances and only estimate linear models.

The chosen non-linear models outperform the chosen linear models for all three countries. In Table 2, we show the parameter estimates for the best non-linear model for each country. The complete 9-parameter model never wins. For the US and Japan, assigning a non-linear coefficient to three of the four past variance variables provides the best fit while, for the euro area, there are only two non-linear coefficients.

To provide some economic intuition on how the logistic function works in practice, Table 3 reports the actual parameter values at various values of the instruments, focusing mostly on the right tail of the distribution of the instruments (Zs). Note that the coefficient estimates are higher at the median than at the mean, reflecting the positive skewness in instruments' distributions and the (monotonically) declining nature of the logistic function. While the non-linear model clearly dominates according to the BIC criterion, the non-linearity is economically rather mild. The largest coefficient drop occurs for Japan's $RV^{(22)}$ variable, where the coefficient drops from 0.44 at the median to 0.35 at the 95th percentile.

Tables 1 and 2 also show some statistical properties for our international uncertainty and risk aversion estimates derived from the model above. First, as shown in the last two rows of Table 2, the correlation of the risk variable estimates between the best linear and non-linear models within a country all appear very high, exceeding 0.94, for all three countries. Then, rows "UC-RA corr" in Tables 1 and 2 show the correlation between UC and RA using the linear and non-linear models, respectively. For all countries, the best non-linear estimates (from Table 2) exhibit smaller UC-RA correlations than the best linear estimates (from Table 1). In particular, the UC-RA correlations are fairly low for Japan (0.32) and the US (0.54), but relatively high for the euro area (0.87).

Finally, Table 4 presents more statistical properties of the best non-linear measures, which we use in our analysis henceforth. From Panel A, both UC and RA are lowest on average in the US, relative to Japan and the euro area. However, uncertainty is more skewed in the US than in the other countries. Panels B and C characterize the cross-country correlations of the risk aversion and uncertainty measures. Given that the three countries are in different time zones, there is essentially no overlap in trading. This implies that we can investigate "contemporaneous" correlations from three different information sets or investor perspectives, illustrated in the note to the Table. Panel B produces correlations from all three perspectives (using the overlapping 1/4/1999 to 5/31/2018 sample). Both RA and UC are moderately highly correlated across countries, with the correlations in the 0.7–0.9 range. Note that the correlations are quite similar across time zone perspectives, revealing that there is an important global component in risk and uncertainty. In Panel C, we show that the first principal component of risk aversion (uncertainty) explains 85 to 91% (83.5 to 84%) of the total variation of the risk variables. For risk aversion, the first principal component loads most strongly on euro area risk aversion; for uncertainty, it loads strongly on US uncertainty.

Figure 1 plots the time series of UC and RA for the overlapping sample starting in 2000. For this plot and in all tables onwards, we scale up the measures by a factor of 10000 to put them in squared monthly percentages (comparable to $VIX^2/12$). Note that both UC and RA peak worldwide during the early 2000's, the 2008-09 Global Financial Crisis, and the 2010-2012 euro area sovereign debt crisis, and the 2015 stock market selloff.

3.3 Monetary policy and risk

With the risk measures in hand, we now describe our empirical shock measures (Section 3.3.1), and show the estimation results of Equation (4) for risk aversion (Section 3.3.2; Table 6, Panel A) and uncertainty shocks (Section 3.3.3, Table 6, Panel B). Apart from testing the domestic and foreign risk channel effect of monetary policy, we also examine how non-monetary policy-driven risk shocks are directly correlated across countries. While we sometimes refer to these effects as "risk spillovers," they could simply follow from a global risk shock travelling across time zones. Recall that these risk shocks are cleansed from the effects of monetary policy shocks and macro announcements.

3.3.1 Shocks

Monetary policy shocks Our monetary policy (MP) shocks are derived from high-frequency data. To account for the information content of monetary policy announcements, we use the measures developed by Jarociński and Karadi (2020) for the US and the euro area. Their paper disentangles monetary policy shocks from a contemporaneous information shock by analyzing the high-frequency comovement of interest rates (US: 3-month Federal funds futures rate; EA: 3-month Eonia (Euro Overnight Index Average) interest rate swap rates) and stock prices in a narrow window around the policy announcement (starting 10 minutes before and ending 20 minutes after the announcement). They argue that a pure monetary policy tightening should unambiguously lower stock market valuations through a discount rate effect (higher real interest rates and rising risk premia) and a cash flow effect (expected payoffs declining with the deteriorating outlook caused by the policy tightening). Therefore, they identify a monetary policy shock through a negative high-frequency comovement between interest rate and stock price changes. In contrast, stock markets and interest rates comoving positively is interpreted as an indication for the presence of an accompanying information shock. They document that a "pure" monetary policy tightening leads to a significant tightening of financial conditions (and a contraction in output). Whereas the central bank "information" shock (with a positive shock signalling good news about the economy) leads to improving financial conditions and persistently higher short-term interest rates (as the central bank tightens its policy to counteract the impact on the macroeconomy).[9]

[9] An advantage of using the Jarociński and Karadi (2020) decomposition is that it gives us a consistent decomposition for both the US and the euro area, for our entire sample period. Several recent papers similarly

For Japan, we use a measure from Rogers, Scotti, and Wright (2014) who compute changes in 10-year Japanese government bond futures yields from 15 minutes before to 15 minutes after the announcement. We use changes in 10-year government bond futures as short-term rates in Japan were constrained by the zero lower bound on interest rates over our sample period.

We analyze monetary policy shocks and central bank information shocks for the overlapping sample for the three countries, January 2000 – December 2015. Table 5 provides summary statistics for these measures (all quoted in basis points). Over this time period, we have 137 monetary policy shocks for the US, 263 for the euro area, and 247 for Japan. For monetary policy shocks, a positive (negative) shock indicates monetary policy tightening (easing). For the central bank (CB) information shocks, a positive value indicates good news about the economy and vice versa. All measures are quoted in basis points. Note that the standard deviation of both the pure monetary policy and information shocks is comparable at 6.5 basis points. In the Online Appendix, we replicate our results using baseline monetary policy shock measures based on high-frequency data: 3-month Fed futures for the US (Gürkaynak, Sack, and Swanson (2005)) and 3-month swap rates for the euro area (Jarociński and Karadi (2020)).[10]

Macroeconomic news In addition to the monetary policy shocks, we collect data on macroeconomic news releases and the corresponding survey expectations prior to the news release. As is standard in the literature, we define a macroeconomic news shock as the actual realization minus the survey expectation, divided by the sample standard deviation. For each of the three countries, we construct shocks for the following seven macroeconomic series: GDP growth quarter-on-quarter, industrial production growth, unemployment rate, CPI inflation, current account balance, consumer confidence and manufacturing confidence. We assume that these shocks span new information about changes in g_t in Equation (1).

Risk shocks We project daily changes (first differences) in country risk measures (as constructed in Section 3.1) onto domestic and foreign monetary policy, macroeconomic shocks and their dummies. We call the residual term the country "non-MP" RA or UC shock.

propose measures of monetary policy shocks which control for central bank information effects, e.g., Miranda-Agrippino and Ricco (forthcoming) and Nakamura and Steinsson (2018). Cieslak and Schrimpf (2019) who decompose news conveyed by central bank communication into monetary policy as well as non-monetary news, such as news about economic growth and news affecting financial risk premia.

[10]We are very grateful to Refet Gürkaynak, Marek Jarociński, and John Rogers for sharing their data with us.

3.3.2 Monetary policy and risk aversion

The dependent variables are daily changes in our risk aversion measures in the three countries. To conserve space, we only report the coefficients related to the monetary policy shocks, $\beta_j^{MP,i}$, or to direct risk spillovers, $\beta_j^{RA,i}$.[11] While Equation (4) is run at the country level, we summarize information from all three country regressions by organizing results according to the economic nature of the coefficients (policy or risk aversion effects emanating from the US, the euro area, and Japan). For example, Columns (1), (5) and (8) of Table 6 come from one regression with the left-hand-side variable being the first-differenced US RA and the right-hand-side variables including the 3-month pure MP shocks and CB information shocks from US, EA, and JP[12] as well as non-MP-driven foreign RA shocks.

We start by discussing the monetary policy effects. In the baseline regression, we find that the monetary policy shocks have no significant effects on risk aversion, neither domestically nor across countries. The exception is a positive spillover from the euro area to Japan. When we use three-month shocks (with the results reported in the Online appendix), there are no significant effects either. This appears inconsistent with the assumptions underlying the work in Miranda-Agrippino and Rey (2020). It is also inconsistent, at first glance, with the original findings in Bekaert, Hoerova, and Lo Duca (2013), who find a causal effect of monetary policy shocks on risk aversion in the US. However, these authors focused on a sample ending in 2007, before the Great Recession ushered in an era of unconventional monetary policy. Moreover, they use one-month futures. Unreported results show that one-month MP shocks exhibit a positive but insignificant effect on changes in risk aversion in our sample.

While the MP shock coefficients are mostly positive (but statistically insignificant), information shocks show mostly negative domestic and spillover effects; the coefficients are significant for the domestic effect in the euro area, and the spillover from the euro area to the US. If such shocks reflect positive growth prospects, it is conceivable that they entail lower uncertainty (with results reported below) and lower risk aversion, consistent with the habit model in Section 2.

Overall, we conclude that monetary policy has surprisingly weak effects on risk aversion, from the high-frequency perspective we focus on. The last line in each block reports the $\beta_j^{RA,i}$ coefficients on non-MP-driven risk aversion shocks. The results show strong evidence of international risk aversion spillovers. US risk aversion shocks transmit to both Japanese

[11]In terms of the other unreported coefficients, we find that risk aversion across countries is statistically significantly lower on US monetary policy event days and, additionally, risk aversion in the euro area is lower on EA monetary policy event days.

[12]The central bank information shock is not available for Japan.

and euro area risk aversion, with the former effect economically and statistically the strongest. euro area risk aversion transmits to US risk aversion but the positive effect on Japanese risk aversion is statistically insignificant. Japanese risk aversion only has a statistically significant effect on euro area risk aversion. These results are robust using various measures of MP shocks (see the Online Appendix).

These risk aversion spillovers are potentially consistent with a strong global factor structure in risk aversion whereby, over the course of a day, information about global risk aversion is first released in Japan, then in Europe and the US and spillovers happen as markets open. We attempt to infer such a global risk variable in Section 6.

3.3.3 Monetary policy and uncertainty

While the link between monetary policy and uncertainty has received less attention in the literature, there has been a proliferation of work in macroeconomics on the adverse effects of uncertainty shocks (Bloom (2009) and a voluminous follow-up literature). It is therefore of interest to examine what effects monetary policy exerts on financial uncertainty, if any. There is not much theoretical guidance on the issue but, intuitively, it is conceivable that monetary policy actions help resolve uncertainty about the monetary policy reaction function and therefore lower financial uncertainty. However, it is not clear whether this effect is more prevalent for tightening or easing monetary policy. Without an asymmetry, it should be picked up by the event dummies. Information shocks, however, may have a directional effect: if higher interest rates reflect good news about economic prospects, they may lower financial volatility and vice versa.

From Table 6, Panel B, while the effects of monetary policy shocks on uncertainty are indeed mostly negative, the coefficients are mostly statistically insignificantly different from zero. Euro area monetary policy shocks significantly lower US uncertainty, with the effect surprisingly stronger for "pure" than for "information" shocks. The domestic effect of euro area pure monetary policy shocks on uncertainty is positive and significant. This effect is not significant when simple three-month shocks are considered (see the Online Appendix).

As with risk aversion, we find mostly strong uncertainty spillover effects. The spillovers are strongest from Japan to the euro area and from the euro area to the US. The spillovers are an order of magnitude smaller from the US to Japan and the euro area to Japan. The spillovers from the US to the euro area and from Japan to the US are not significant.

4 Monetary Policy, Risk, and Asset Returns

Our previous results do not support two tenets of the global financial cycle story, as put forward in Miranda-Agrippino and Rey (2020) and Passari and Rey (2015). First, monetary policy does not seem to affect risk aversion significantly. Second, there is no special role for the US "hegemon" (as they put it), in that US monetary policy shocks do not transmit to risk variables across countries. This need not mean that risk spillovers do not play an important role in globalized financial markets; in fact, we show strong international financial spillovers of changes in risk aversion and uncertainty that are not driven by monetary policy; only, we do not find evidence of such spillovers being induced by US monetary policy.

In this section, we first trace out the effects of monetary policy on interest rates, and contrast them with the effects of non-MP-driven uncertainty and risk aversion shocks (Section 4.1). We then examine how monetary policy and risk shocks transmit to stock and 10-year government bond returns domestically and abroad (Section 4.2). Finally, we summarize our empirical evidence of contemporaneous effects (Section 4.3), and explore dynamic effects in these three major asset classes (Section 4.4).

4.1 Monetary policy and interest rates

Table 7 reports our baseline regression with daily changes in 3-month interest rates as the dependent variable. Specifically, we use three-month Treasury interest rates for the US and three-month government yields for the euro area and Germany; for Japan, we use 10-year government bond yields as short-term interest rates barely moved throughout the sample period. As with most financial data used in this article, they are downloaded from DataStream. Interest rates and MP shocks are in basis points.

The regression now includes both cleansed risk aversion and uncertainty shocks, and domestic and foreign risk shocks, as described in Equation (5). Note that risk aversion and uncertainty shocks are slightly positively correlated in the euro area and Japan, but negatively correlated in the US. As before, the columns present the key coefficients (domestic and spillover effects on interest rates) in the three country-specific regressions.

The first goal of this table is to verify that monetary policy does indeed pass through to interest rates as expected. The effect of a 10 basis points tightening of US monetary policy (the MP shocks purged from CB information) is a 3.2-basis point increase in US Treasury rates, or a 32% pass-through. The pass-through is 22% in the euro area and 28% in Germany. In Japan, the pass-through is 63%. All these coefficients are highly statistically significant.[13]

Our second goal is to verify whether there is direct monetary policy spillover. We first confirm that indeed US monetary policy shocks spill over to Germany and the euro area, with the pass-through around 15% (coefficients around 0.15).[14] However, there is no effect on Japanese interest rates. Japanese monetary policy shocks have a negative (but insignificant) effect on the US interest rate, and a significant positive effect on German interest rates. There are no significant spillover effects from the euro area.

The information shocks also pass through for the US and Germany, domestically, in a significant fashion (with coefficients around 0.6 and 0.5, respectively), but the effect is insignificant for the euro area; there are no data for Japan. Information shocks significantly spill over from Europe to the US and Japan, but there are no significant spillover from US information shocks.

In general, our results suggest that there are substantial interest rate spillovers across countries, but they are country and shock dependent. These results seem inconsistent with the findings in Kearns, Schrimpf, and Xia (2020), who show weak evidence of interest spillovers for short-term interest rates.

Finally, we examine the interest rate effects of changes in uncertainty and risk aversion, as predicted by the model outlined in Section 2. We find overwhelmingly negative coefficients for uncertainty and risk aversion shocks. For the US, uncertainty shocks significantly lower US interest rates, consistent with precautionary savings effects, and also lower euro area interest rates. For the euro area, there is a significant negative risk aversion effect on Japanese interest rates. In Japan, risk aversion shocks have negative interest rate effects while the uncertainty shock effects are insignificant. Japanese uncertainty shocks negatively affect US interest rates, but Japanese risk aversion shocks positively affect US interest rates (with the latter effect significant at the 10% level).

4.2 Monetary policy, risk, and stock and bond returns

In this section, we report results for two major asset classes, namely equities (Table 8, Panel A) and long-term (10-year) government bonds (Table 8, Panel B). All returns are measured in

[13] The euro area coefficient is statistically significant at the 10% level; the somewhat weaker results for the euro area compared to Germany are influenced by the sovereign debt crisis period during which interest rates across stressed versus non-stressed euro area countries did not move in tandem. This is also the reason why, in this section, we explicitly look at the German interest rates in addition to the euro area (composite) rates.

[14] Ehrmann and Fratzscher (2005) also document strong reactions of interest rates in the euro area to monetary policy and macroeconomic news in the US.

percent (log first-differences of total return indices multiplied by 100) and local currency, and are sourced from DataStream.

4.2.1 Effects on stock returns

First, US monetary policy tightening leads to negative stock returns in the US. Economically, a 10 basis points 3-month pure MP shock leads to a 85 basis point drop in the stock market, confirming the large effects documented in the seminal Bernanke and Kuttner (2005) article. The domestic MP effect is even slightly stronger in Europe but statistically insignificant in Japan. The information shock effect is, as expected, robustly positive, with the effect in the euro area (0.14) three times as large as in the US (0.04). This renders the overall MP coefficient positive in the euro area (see the Online Appendix).

While Bernanke and Kuttner (2005) argue the effect of MP shocks on stock returns is mostly a risk premium effect, this channel does not square well with our lack of evidence for a risk channel for monetary policy (Table 6). In general, permanent discount rate effects should generate, when viewed as permanent, very large immediate price effects (in a Gordon model the return multiplier is proportional to the price dividend ratio). Taken together, our evidence suggests that the real interest rate effects induced by monetary policy may be viewed as much more persistent than risk effects. Martin (2017) and Bekaert, Engstrom, and Xu (2020) show that equity risk premiums and risk aversion levels, respectively, do not show strong persistence. Binsbergen (2020) also argues for strong pure interest rate effects on stock market returns.

We also observe large international spillover effects. The US pure MP shocks have a negative effect on the euro area stock market, about 30% the magnitude of the own market effect. This effect is not at all surprising. The beta of the euro area with respect to world equity returns is 1.02 (measured over a 1970-2019 sample). Thus, with the US market being 40-45% of world market capitalization, a "direct" CAPM prediction would be a 40% move or more. The US information shock has also a positive and significant effect on the euro area stock market, which is, surprisingly, stronger than the domestic information effect. However, the strongest spillover effects come from the euro area, with both the pure MP and information shocks affecting US and Japanese stock market returns in the expected direction. The effects are economically large, with stock markets (a) dropping around 0.5% in response to a 10 basis point pure MP shock in the euro area and (b) increasing by 70 basis points (Japan) to 1.20% (US) in response to a 10 basis point information shock in the euro area. These effects are much larger in magnitude than simple CAPM predictions would suggest.

Risk aversion and uncertainty shocks have negative effects on stock markets, no matter what area they originate from, with only a few exceptions. Generally, the domestic effects are three to five times larger than the international spillover effects. Economically, the direct effects of both risk variable shocks on the stock market we observe are large. A one standard deviation positive shock to risk aversion or uncertainty generates a stock market drop of between 50 basis points and 1 percent. By contrast, the international spillover effects are not always statistically significant and sometimes have a surprising positive sign (e.g. risk aversion shocks from Japan to the US and the euro area). Note that these spillover effects are economically very small (less than 10 basis points for a one standard deviation move in risk aversion), and that they change sign when uncertainty shocks are dropped from the regression. US uncertainty and risk aversion shocks do have a statistically significant and negative effect on Japanese stock returns, and euro area risk aversion shocks lower Japanese stock returns.[15] Recall that some of the spillover effects are already absorbed by the direct risk spillover effects documented in Table 6.

In unreported results, the regressions also reveal that US monetary policy events (as measured by the dummy coefficients) are associated with positive US and Japanese equity returns. This is not so surprising as it may partially reflect the pre-FOMC meeting drift, documented in Lucca and Moench (2015). It is interesting that there is a strong effect on Japanese stock returns, but almost none on euro area stock returns, especially given that the correlation between Japanese stock returns and US and euro area returns is much lower than the correlation between US and European stock returns.

4.2.2 Effects on bond returns

Starting with the “pure” MP effects, Panel B of Table 8 shows that tightening US monetary policy decreases domestic bond returns, which is a natural consequence of its effects on interest rates documented before. Economically, a 10 basis point shock leads to a bond return of about minus 30 basis points. In fact, with a duration of around 7 years, and a 30% pass through for short term interest rates, the term premium does not need to increase much at all to generate such a result. The euro area pure MP effect on domestic bond returns is statistically significant as well but economically weaker than the US domestic effect. The Japanese MP effect is stronger, with a 10 basis points shock generating a more than 50 basis points drop

[15] Ehrmann, Fratzscher, and Rigobon (2011) also find stronger within-country than across-country shock transmission for various asset classes in the US and the Europe, but we do not confirm their finding that US-driven international spillover effects dominate.

in Japanese bond prices; this is to be expected as we use shocks to 10-year government bond futures as a measure of MP shocks for Japan.

In terms of spillover effects, tighter Japanese monetary policy also has a significant negative effect on US bond returns. There are no other significant spillover effects of pure MP shocks. However, we observe significant information shock effects emanating from the euro area, with positive shocks decreasing bond returns not only in the euro area but also in Japan and the US. Hence, the information shock spillover effects are consistent with the interest rate effects documented in Table 7, but the pure MP shock effects are not.

In terms of risk aversion shocks, the within-country effects are invariably positive. This may constitute a domestic flight-to-safety effect. The effects are statistically significant, but economically rather small with a 1 standard deviation shock entailing about a 2 (Japan) to 7 (US) basis point positive bond return. The international spillover effects of risk aversion shocks are occasionally statistically significant but the effects appear economically small. For uncertainty shocks, the domestic effects are slightly larger than the risk aversion effects, but no longer statistically significant for Japan. There are also three statistically significant spillover effects. US uncertainty shocks lower euro area bond returns, indicating that German government bonds do not act as an international safe asset. Euro area uncertainty shocks increase US bond returns, whereas Japanese uncertainty shocks lower US bond returns. From the perspective of a CAPM equilibrium in global bond markets, we would still expect spillovers from the US to be the largest, but the euro area and Japanese bond markets do each represent close to 20% of world bond market capitalization. However, as Xu (2019) shows, bond returns across countries are much less correlated than are equity returns, and the spillover effects also depend on which market serves as a safe haven.

4.3 The economics of the asset return results

We now summarize the results in economic terms by converting coefficients in standard deviation (SD) units and then averaging the transformed effects across the country regression results. That is, we ask how much do interest rates, stock and bond prices change contemporaneously (expressed in number of SDs) given a 1 SD pure monetary policy (MP), central bank (CB) information, risk aversion or uncertainty shock.

In the first row of Figure 2, we compare the relative *economic* importance of the four shocks across several dimensions. First, central bank information shocks have stronger domestic and spillover effects than do pure monetary policy shocks, with only one exception. Domestic information shocks have smaller effects (slightly less than 0.1 SD decrease) on domestic bond returns, than do domestic pure monetary policy shocks (a 0.34 SD decrease). For the latter, the effects of pure monetary policy shocks on domestic interest rate and bond returns are similar in magnitude, which is not surprising if we think of bond returns as roughly duration times interest rate changes. The smaller domestic bond return effect generated by one information shock, as compared to the corresponding interest rate effect, is certainly possible if information shocks are accompanied by opposite changes in uncertainty or risk aversion, which then change the term premium in the opposite direction. Such a channel is present in any habit model (see Pflueger and Rinaldi (2020)). Notably, information shocks seem to have stronger international spillover effects than do pure monetary policy shocks.

Second, domestic pure monetary policy shocks generate larger changes in interest rates and bond prices than in stock prices: a 1 SD increase in the pure MP shock leads to a 0.35 SD increase (decrease) in the interest rate (bond price), but only a 0.25 SD move in stock prices. In fact, information shocks are economically relatively more important for stock prices.

Third, risk shocks are relatively more important in explaining stock price changes, producing changes of a 0.5 (for RA) or 0.4 (for UC) SD magnitude, whereas they only produce 0.1 SD moves in interest rates or bond prices. Note that for risk shocks, we do not distinguish between domestic and foreign effects, as they represent one day changes in risk aversion or uncertainty and may partially reflect the realization of a global shock. Given that Section 6 establishes the existence of an important global component in risk aversion shocks, this may help explain why international bond return comovements are much smaller than international stock return comovements (Xu (2019)).

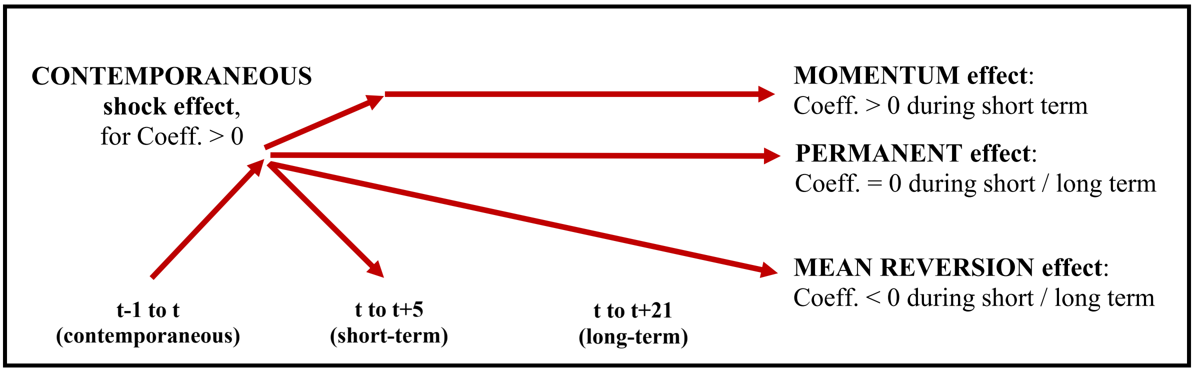
Finally, in terms of the relative importance of risk shocks versus monetary policy shocks, for interest rates and bond returns, the total effects of monetary policy shocks dominate the effects of risk shocks. However, for stock returns, pure monetary policy shocks (domestic+foreign) generate less than half the return impact than do risk shocks, whereas information shocks (domestic+foreign) generate return impact of the same order of magnitude. Together with our second point above, these results cast doubt on a risk channel playing an important role in the transmission of MP to equity markets. We revisit this issue in Section 4.4.

The lower set of plots provide a more detailed, country-by-country perspective on the monetary policy shocks. For each shock source (country; MP or CB) and asset class, we contrast the magnitude of domestic and spillover effects. A first important finding is that the immediate spillover effects for pure monetary shocks are quite small, at most half the magnitude of the domestic effects. Second, the US does not appear the hegemon country dominating spillover effects at all. From the "IR" plot, the US pure MP shock clearly is the only country shock with a meaningful interest rate spillover effect, but the effect is still only one quarter of its domestic effect. Not surprisingly, from the "BR" plot, the US pure MP spillover effect on bond returns is also dominant, but in relative terms Japanese monetary policy also generates a spillover effect of about half the magnitude of its domestic impact on bond returns. For stock returns (see the "SR" plot), the euro area pure MP spillover effect is much larger in economic terms than the effect emanating from the US. Finally, while the domestic interest rate effect of a US central bank information shock is very large, its spillover effect is smaller than that emanating from the euro area. In fact, the euro area's information shocks generate stronger spillover effects than those of the US for both stock and bond returns.

4.4 Dynamic effects

So far, we have solely discussed the well-identified high frequency effects. Of course, much of the literature has sought to assign a relatively large role to monetary policy shocks in explaining asset price comovements using relatively low frequency empirical settings, such as vector autoregressions with quarterly data. If we use our daily data regressions to provide a decomposition of the relative importance of the various shocks in determining the variation of the asset classes over the full sample (not only on monetary policy event days), monetary policy is only relatively important for interest rates. For bond and stock returns, the variation in monetary policy shocks is too small to generate meaningful variation over the full sample. Such a picture could change if monetary policy has persistent effects. Studying the persistence of the effects also helps interpret the economic channels behind the results.

To do so, we compare three price change responses: (1) the contemporaneous response or price changes from $t-1$ to t (as in our Tables 7 and 8); (2) short-term cumulative price changes from t to $t+5$; (3) long-term cumulative price changes from t to $t+21$. In practice, these changes represent the same day ((1)) or cumulative log returns ((2),(3)). Because there is a clear trade-off between identification and the horizon in the regressions, we do not go beyond the one month horizon. The diagram below demonstrates the corresponding channel interpretations, given various coefficient estimates of (1) versus (2) and (3):



Suppose a one unit shock has caused the price today to increase (Coeff.>0). The first possibility is that the effect on the first day does not represent a full response, and the effect continues in the same direction for a few days (momentum effect). A second possibility is that the first day effect is simply permanent, and further returns are simply noise. This would be the case, for example, for a pure cash flow effect; stock prices should increase and not change any further. Finally, discount rate effects naturally lead to mean reversion: higher prices today reflect lower future returns. This effect cannot be fully disentangled from a price pressure effect, apart from the fact that the latter should be reversed in the short run, whereas the former is likely to last longer, depending on the persistence of the interest rate or risk premium shock. We relegate a further discussion of the methodology and detailed dynamic effect results to the Online Appendix, but there are three key messages:

First, pure monetary policy shocks induce near permanent effects on interest rates, and their effects on stock and bond returns also show little mean reversion. This indirectly confirms that the effects of monetary policy on asset prices do not occur through a risk premium channel, but through a direct interest rate channel, which has become more potent given the unusually low interest rates in the last 10-15 years. This finding is consistent with Binsbergen (2020)'s recent assertion that equity returns in the US show little or no evidence of any risk premium over long term bonds. With interest rates highly persistent, such shocks may mean revert extremely slowly. Second, in direct contrast, the effects of risk aversion and uncertainty shocks on stock prices – to a lesser extent on bond prices – strongly mean revert, confirming the risk premium interpretation, where risk premiums may not be as persistent as previously thought (see e.g. Martin (2017), Bekaert, Engstrom, and Xu (2020)). Third, patterns observed for the central bank information shocks emanating from the Fed and the ECB are consistent with them representing permanent cash flow effects. However, the ECB information shock does exhibit week-long momentum.

5 Additional Results

In this section, we examine how monetary policy and risk shocks transmit to exchange rates and commodity and gold prices (Section 5.1, Table 9), and whether unconventional policies since 2008 change our results (Section 5.2, Online Appendix).

5.1 Effects of monetary policy and risk shocks on exchange rates and commodity and gold index returns

Table 9 considers the exchange rate effects of monetary policy and risk shocks. There is a vast literature regarding monetary policy effects on exchange rates which typically finds that monetary policy tightening leads to an appreciation of that country's nominal spot exchange rate. For example, Eichenbaum and Evans (1995) document evidence for the US, consistent with the prediction of the uncovered interest rate parity relationship (coupled with unchanged exchange rate expectations). Inoue and Rossi (2019) recently extend this evidence to the period of unconventional monetary policy following the Great Recession, confirming that tight monetary policy still appreciates the dollar. Dedola, Georgiadis, Gräb, and Mehl (2020) estimate the effect of ECB and Federal Reserve quantitative easing on the dollar-euro exchange rate and show that a typical ECB or Federal Reserve expansionary QE depreciates the exchange rate by 7%.

The first two columns in Table 9 (Panel A) confirm that, upon a 10 basis point monetary policy shock in the US, the dollar appreciates by 40 and 35 basis points relative to the euro and the yen (exchange rates changes are measured in logs), respectively. Monetary policy in the euro area and Japan fails to affect exchange rates. Information shocks also have no effect on exchange rates.

We also show the effects of cleansed risk aversion and uncertainty shocks on exchange rates. An increase in US risk aversion tends to appreciate the dollar relative to the euro *but not* the dollar relative to the yen; in fact, the yen appreciates. This seems somewhat at odds with recent work that suggests the dollar has become the premier safe haven currency post 2008, but this result even holds post 2008 (see Online Appendix). Avdjiev, Du, Koch, and Shin (2019) show a positive correlation between the post 2008 deviations from covered interest parity (CIP), a stronger dollar and contractions of cross-border bank lending in dollars. They argue that the dollar has become a key indicator of risk-taking capacity in global capital markets. Following euro area risk aversion shocks, both the dollar and the yen appreciate, and the euro depreciates, which is consistent with a safe haven status for the dollar and the yen relative to the euro. The dollar effect is insignificant however. Following Japanese risk aversion shocks, the Japanese yen appreciates both relative to the dollar and the euro, perhaps indicating a flight-to-safety effect. Interestingly, there are fewer significant exchange rate effects in response to uncertainty shocks. The yen appreciates relative to the dollar following US uncertainty shocks, and relative to the euro following euro area uncertainty shocks, confirming its international safe haven status.

The last asset class we consider is commodities, which were also included in the construction of the global financial cycle in Miranda-Agrippino and Rey (2020). While they exclude precious metals, our index represents an overall index as measured by the Goldman Sachs Commodity Index. Separately, we also consider gold prices. Both are measured in dollars at the end of the day, and we consider one day log returns. We expect gold to potentially behave as a safe haven asset (see e.g. Baur and Lucey (2010)). However, Huang and Kilic (2019) recently cast doubt on this conjecture, showing that, for example, gold prices did not increase in the Great Recession. Interestingly, they show that the log ratio of gold to platinum prices is a significant risk indicator, predicting future equity returns, and being high when equity valuation ratios are low. To better isolate the potential *risk* correlation of gold, we orthogonalize gold price changes with respect to general commodity price changes.

The results are in Table 9 (Panel B), with the gold effects in the left panel and the commodity price effects in the right panel. First, monetary policy shocks have no effect on commodity prices, and only the euro area information shock significantly increases commodity prices (perhaps because it houses positive information about the world economy). However, US monetary policy tightening, which leads to higher real interest rates, significantly decreases gold prices. This is consistent with the idea that gold competes with other safe assets (such as short-term money market deposits and perhaps bonds). This effect is also present but slightly weaker for raw gold prices. There are no such significant effects in the euro area or Japan, and, generally, no significant information shock effects.

Uncertainty and risk aversion shocks everywhere drive down commodity prices, with the results statistically significant in the US and euro area. This effect would counteract any safe haven type of effect for gold. Perhaps surprisingly, risk aversion shocks do not have an effect on the (cleansed) gold price changes. However, uncertainty shocks in the US do significantly increase gold prices.

5.2 Unconventional monetary policy, risk, and asset returns

While we do not detect significant structural breaks over our sample,[16] we examine the unconventional policies employed by central banks in the aftermath of the global financial crisis, and the robustness of our results to the zero lower bound. Tables containing the results are relegated to the Online Appendix; below, we summarize the main take-aways.

There is a large literature investigating the effects of unconventional monetary policies. A number of contributions analyze the impact on domestic asset prices in advanced economies.[17] As for the international effects of unconventional policies, many papers focus on the impact of US's large-scale asset purchases, finding spillovers to international bond and foreign exchange markets, as well as global portfolio flows. For example, Neely (2015) finds that the Federal Reserve's asset-purchase program had substantial international effects on bond and foreign exchange markets. Albagli, Ceballos, Claro, and Romero (2019) and Gilchrist, Yue, and Zakrajšek (2019) document significant US monetary policy spillovers to international bond markets. Fratzscher, Lo Duca, and Straub (2018) study the effects of the Federal Reserve's quantitative easing (QE) on global portfolio flows.

Our paper considers monetary policies spillovers across three large economies. To account for non-standard policies like forward guidance and asset purchases, we rely on measures developed by Rogers, Scotti, and Wright (2018) for the US.[18] They use intraday changes in interest rates or yields of different maturities, in the window from 15 minutes before to 1 hour and 45 minutes after the time of an FOMC or other monetary policy announcement. Specifically, target shocks are changes in yield on the current or next-month federal funds futures contracts, following Kuttner (2001). The forward guidance (FG) shocks are the residuals from a regression of the change in the yield for the fourth Eurodollar futures contract onto the target surprise, which is a bet on the level of 3-month interest rates about 1 year hence. These shocks thus approximately capture changes in the 1-year ahead expected 3-month Eurodollar interest rate. Asset purchase (AP) shocks are the residuals from a regression of the change in the 10-year Treasury futures yield onto the target and forward guidance surprises. They measure the

[16] In fact, we perform break tests on our main specifications linking risk aversion and uncertainty to monetary policy using the Bai, Lumsdaine, and Stock (1998) methodology. We perform the tests in several configurations (country by country, using UC and RA separately or jointly, and over all countries and both variables). We invariably find break dates, perhaps not surprisingly, in the October-November 2008 period, but the break tests mostly do not yield significant rejections of the no break null and the confidence intervals for the break dates are large. When we combine all variables and all countries in one test, the break date is November 21, 2008.

[17] See, e.g., Krishnamurthy and Vissing-Jorgensen (2011), Wright (2012), D'Amico and King (2013), Andrade, Breckenfelder, De Fiore, Karadi, and Tristani (2016), Ghysels, Idier, Manganelli, and Vergote (2017). Kuttner (2018) and Dell'Ariccia, Rabanal, and Sandri (2018) survey the literature.

[18] Note that for Japan, the non-conventional MP measures from Rogers, Scotti, and Wright (2014) are already used as our baseline measure of Japanese MP shocks. For the euro area, an expanded asset purchase programme (APP) only started in 2015, which is when our sample ends.

jumps in long-term interest rates that were associated with FOMC announcements related to large-scale asset purchases from 2008 onwards. As before, a positive (negative) shock indicates policy tightening (easing). For example, in case of the AP shocks, a positive shock could mean an earlier-than-expected termination of the QE program.

Starting with the effects of monetary policy on risk aversion in the post-2008 sample, we confirm our full sample findings: US monetary policy does not have a significant effect on risk aversion, domestically or internationally. For the US unconventional policies, only forward guidance shocks have a statistically significant effect on risk aversion (with positive shocks, which represent tightening of MP, increasing risk aversion).

Results for uncertainty are also largely confirmed in the post-2008 sample. For the effects of US unconventional policies on US uncertainty, the effect is only significant for asset purchase (AP) shocks, and is likely due to such shocks reflecting good news about the economy. This shock transmits relatively strongly to the euro area as well. However, forward guidance (FG) shocks in the US have a significant positive effect on euro area uncertainty.

For short-term interest rates, we do not observe pass-through from either standard MP shocks, or the unconventional FG and AP shocks in the US in the post-2008 sample. However, there are strong positive interest rate effects in Germany (a coefficient of 0.6 significant at the 1% level) and Japan (at 0.64 very close to the full sample estimate). For the standard MP shocks, we observe significant spillovers from the US to Japan; from the euro area to Japan and from Japan to the euro area.

For stock returns, US monetary policy tightening increases stock prices, and spills over in a statistically significant and economically important fashion to the euro area and Japanese stock markets. However, conversely, the US stock market is also strongly and positively affected by monetary policy tightening in the euro area and Japan. These results likely indicate that information shocks dominated monetary policy shocks post 2008. The US unconventional policy shocks produce mixed results. Forward guidance tightening reduces stock market prices in a significant fashion, but they have no international spillover effects. Asset purchase shocks on the other hand do not have a significant stock market effect. For bond returns, we largely confirm our full sample result that tightening US monetary policy decreases domestic bond returns. Moreover, US unconventional policy tightening also reduces bond prices, both domestically and internationally.

For exchange rates, both US forward guidance and asset purchase tightening shocks lead to appreciations in the dollar. For the euro area and Japan, we mostly do not find significant exchange rate effects in response to domestic monetary policy shocks (with the only exception being the euro-yen exchange rate response to 1-month euro area shocks). For the unconventional monetary policy period, the yen actually depreciates relative to the euro in response to Japanese policy tightening, whereas we observe the euro appreciating relative to the dollar following positive euro area monetary policy shocks post 2008. Risk aversion and uncertainty effects remain robust post-2008; we therefore confirm the safe-haven status of the Japanese yen.

6 A Simple Global Factor Model of Risk Aversion

Given the strong evidence that risk aversion shocks cleansed of monetary policy and macro news effects spill over to other countries, it is plausible that there is a "global"/common risk aversion component. Our previous results suggest that global risk may gradually realize as various markets open and close during the 24 hour day. We now formulate a "timezone" factor model, which, applied to our risk shocks data, allows us to filter a global risk aversion measure.

As in Sections 3 and 4, denote the non-MP, non-Macro risk aversion shock of country i as $\overline{ra_t^i}$; that is, the residual of a linear regression with the first-differenced risk aversion ΔRA_t^i on the left-hand-side and monetary policy shocks, macroeconomic shocks and event dummies on the right-hand-side. We denote $\overline{\boldsymbol{ra_t}} = \begin{bmatrix}\overline{ra_t^{JP}} & \overline{ra_t^{EA}} & \overline{ra_t^{US}}\end{bmatrix}'$. We assume that there exists a common, global component driving variation in these cleansed risk aversion shocks. Denote its realization during JP, EA, and US trading hours as follows:

$$\boldsymbol{gra_t} = \begin{bmatrix} gra_t^{JP} & gra_t^{EA} & gra_t^{US}\end{bmatrix}',$$

where gra_t^i denotes the realization of the global factor in time zone i. For simplicity, the global component is assumed to follow a multinormal distribution, $\boldsymbol{gra_t} \sim N(0, \boldsymbol{\Sigma})$ where $\boldsymbol{\Sigma}$ is a diagonal matrix with σ_i^2 on the diagonal.

Thinking of shocks realizing across time zones in different markets implies a natural factor structure:

$$\overline{\boldsymbol{ra_t}} = \boldsymbol{B}\ \boldsymbol{gra_t} + \boldsymbol{C}\ \boldsymbol{gra_{t-1}} + \boldsymbol{u_t}, \tag{9}$$
$$\boldsymbol{B} = \begin{bmatrix}\beta_{JP} & 0 & 0\\ \beta_{EA} & \beta_{EA} & 0\\ \beta_{US} & \beta_{US} & \beta_{US}\end{bmatrix},\ \boldsymbol{C} = \begin{bmatrix}0 & \beta_{JP} & \beta_{JP}\\ 0 & 0 & \beta_{EA}\\ 0 & 0 & 0\end{bmatrix},$$

where $\boldsymbol{u_t} \sim N(0, \boldsymbol{S})$, and $\boldsymbol{S}$ is a diagonal matrix with s_i^2 on the diagonal. Economically, this model essentially follows the global risk aversion shock, that is cleansed from international macro and monetary policy fluctuations around the globe. We start in Japan where the risk aversion shock must first digest global information from Europe and the US of the previous day, hence the dependence on the global shocks in the US and Europe from the day before. The euro area market responds to the Japanese global information, but also to the US global information of the day before. Finally, the US, being the last market to open, responds to earlier released global information in Japan and Europe. While the model is exceedingly simple, it captures the notion of a global risk aversion process. We assume that different countries have different exposures to global information (the βs are country specific) and that global information released within a country's time zone has different variances. Moreover, we assume country-specific idiosyncratic variances. Together these assumptions imply that the explanatory power of global sentiment shocks is country-specific. This is important because the correlations of risk shocks differ across countries. In the Appendix, we show that Japanese cleansed risk aversion shocks show very little correlation with US risk aversion shocks, but more correlation with euro area risk aversion shocks. The correlation between US and euro area risk aversion shocks is highest at 29%. This is somewhat reminiscent of stock market correlations between Japan and the other countries being much smaller than between the euro area and the US. There are a total of 9 unknown parameters:

$$\boldsymbol{\Theta} = [\beta_{JP}, \beta_{EA}, \beta_{US}, \sigma_{JP}, \sigma_{EA}, \sigma_{US}, s_{JP}, s_{EA}, s_{US}]'.$$

Because the $\boldsymbol{gra}$ processes are latent, we resort to moment matching to identify the parameters. We use the second moment matrix and first-order autocovariance matrix of $\overline{\boldsymbol{ra_t}}$ to obtain the parameters. The orthogonality conditions, $\boldsymbol{g_t}$ are given by:

$$\boldsymbol{g_t} = \left\{ \begin{array}{l} vec\left[E\left[\overline{\boldsymbol{ra_t}}\ \overline{\boldsymbol{ra_t}}'\right]\right] - vec(\boldsymbol{Z}) \\ vec\left[E\left[\overline{\boldsymbol{ra_t}}\ \overline{\boldsymbol{ra_{t-1}}}'\right]\right] - vec(\boldsymbol{W}) \end{array} \right. = \mathbf{0}, \tag{10}$$

$$\text{where} \left\{ \begin{array}{ll} \boldsymbol{Z} & = \boldsymbol{B\Sigma B}' + \boldsymbol{C\Sigma C} + \boldsymbol{S} \\ & = \begin{bmatrix} \beta_{JP}^2\left(\sigma_{JP}^2+\sigma_{EA}^2+\sigma_{US}^2\right)+s_{JP}^2 & \beta_{JP}\beta_{EA}\left(\sigma_{JP}^2+\sigma_{US}^2\right) & \beta_{JP}\beta_{US}\sigma_{JP}^2 \\ \beta_{JP}\beta_{EA}\left(\sigma_{JP}^2+\sigma_{US}^2\right) & \beta_{EA}^2\left(\sigma_{JP}^2+\sigma_{EA}^2+\sigma_{US}^2\right)+s_{EA}^2 & \beta_{EA}\beta_{US}\left(\sigma_{JP}^2+\sigma_{EA}^2\right) \\ \beta_{JP}\beta_{US}\sigma_{JP}^2 & \beta_{EA}\beta_{US}\left(\sigma_{JP}^2+\sigma_{EA}^2\right) & \beta_{US}^2\left(\sigma_{JP}^2+\sigma_{EA}^2+\sigma_{US}^2\right)+s_{US}^2 \end{bmatrix} \\ \boldsymbol{W} & = \boldsymbol{C\Sigma B}' \\ & = \begin{bmatrix} 0 & \beta_{JP}\beta_{EA}\sigma_{EA}^2 & \beta_{JP}\beta_{US}\sigma_{EA}^2+\beta_{JP}\beta_{US}\sigma_{US}^2 \\ 0 & 0 & \beta_{EA}\beta_{US}\sigma_{US}^2 \\ 0 & 0 & 0 \end{bmatrix} \end{array} \right. .$$

Given the zero-mean nature of country cleansed risk aversion shocks, the uncentered expectations in Equation (10), $E\left[\overline{\boldsymbol{ra_t}}\ \overline{\boldsymbol{ra_t}}'\right]$ and $E\left[\overline{\boldsymbol{ra_t}}\ \overline{\boldsymbol{ra_{t-1}}}'\right]$, simply represent the variance(s) and covariance(s) of risk aversion shocks. One obvious estimation challenge is that the system is non-linear; while the second moments can be matched, the respective magnitudes of the βs and σs are not uniquely identified given that they enter multiplicatively in model-implied moments. To obtain unique identification, we introduce one more moment, $\beta_{JP} + \beta_{EA} + \beta_{US} = 1$. There are now 10 moments for 9 unknown parameters, and we use a standard GMM algorithm to obtain the parameters (and associated standard errors), considering 250 sensible initial parameter values in sensible ranges (e.g., βs are between 0 and 1, and s^2 might be 0–60% of the total variance of risk aversion shocks in the data from previous regression results). The Appendix shows that most initial values successfully converge to the global optimum, and that the betas are within the 0.3 to 0.4 range. From the estimation, we can compute how much risk aversion variation the global component explains in each country according to Equation (9); these effective R^2's are 30.07% for Japan, 61.44% for the euro area and 47.69% for the US.

The global risk aversion shock by time zone (gra_t^i as defined above) can be filtered using the linear factor framework, assuming normality and no memory (with $\widehat{\ }$ denoting filtered estimates):

$$\widehat{\boldsymbol{gra_t}} = E[\boldsymbol{gra_t} | \overline{\boldsymbol{ra_t}}, \widehat{\boldsymbol{\Theta}}], \tag{11}$$

$$= \widehat{\boldsymbol{D}}\ \overline{\boldsymbol{ra_t}}, \tag{12}$$

where $\widehat{\boldsymbol{D}} = \widehat{\boldsymbol{\Sigma}}\widehat{\boldsymbol{B}}'(\widehat{\boldsymbol{B}}\widehat{\boldsymbol{\Sigma}}\widehat{\boldsymbol{B}}' + \widehat{\boldsymbol{C}}\widehat{\boldsymbol{\Sigma}}\widehat{\boldsymbol{C}}' + \widehat{\boldsymbol{S}})^{-1}$, according to the linear identity in Equation (9); that is, $\widehat{\boldsymbol{\Sigma}}\widehat{\boldsymbol{B}}'$ is the model-implied covariance between $\overline{\boldsymbol{ra_t}}$ and $\boldsymbol{gra_t}$, and $\widehat{\boldsymbol{B}}\widehat{\boldsymbol{\Sigma}}\widehat{\boldsymbol{B}}' + \widehat{\boldsymbol{C}}\widehat{\boldsymbol{\Sigma}}\widehat{\boldsymbol{C}}' + \widehat{\boldsymbol{S}}$ is the model-implied variance of $\overline{\boldsymbol{ra_t}}$.

Essentially, the filtering process constructs the global components from linear combinations of the realized risk aversion shocks. We report these $\widehat{\boldsymbol{D}}$ loadings in Table 10. The key message is that, according to the first column, Japanese risk aversion shocks contribute little to the global risk aversion shocks – even to the global risk aversion shock realized during Japanese trading hours – whereas, according to the last two columns, euro area and US risk aversion shocks contribute a lot to global risk aversion — especially during their trading hours. The US risk aversion shocks contribute positively to $\widehat{gra_t^{EA}}$, but the euro area shock negatively to $\widehat{gra_t^{US}}$, which reflects US risk aversion shocks containing information of the full trading day, including information released during European hours.

While these cleansed risk aversion *shocks* are of interest, the data used by Miranda-Agrippino and Rey (2020) to construct a global financial cycle are integrated time series (i.e., reflecting asset prices, rather than returns). We therefore simply aggregate the $\widehat{gra_t^i}$ over time. The sum of them each day is our global risk aversion measure $\widehat{GRA_t}$.[19]

Figure 3 shows that the relative contribution of the three global risk aversion components varies strongly over time. The global risk aversion realized during Japanese trading hours rarely exceeds 30%. The US component around 40% in the recession of the early 2000s, then its contribution drops to lower than 10% in the pre-crisis years before becoming increasingly important after the Great Recession.

Importantly, despite being derived using an entirely different methodology, our non-MP-driven, cleansed $\widehat{GRA_t}$ measure is also highly correlated with the measure in Miranda-Agrippino and Rey (2020). Figure 4 shows two measures of the global financial cycle downloadable at `http://www.helenerey.eu/`, which is a risk appetite measure, and our measure of risk aversion, sampled at the end of the month. The strong negative correlation is apparent and reaches -0.76 (-0.71) for their long (short) sample version, suggesting that the low frequency movements of our daily risk aversion measure capture a great deal of the variation of the global financial cycle. This is remarkable given that our measure is entirely cleansed of monetary policy influences and from the effects of macroeconomic announcements. It is therefore close to a pure sentiment shock. This is suggestive of an important role of common factors other than US monetary policy that drive global risk aversion variation.

7 Conclusion

This paper studies the effects of monetary policy and risk shocks on risk and asset prices in a global world. Unlike the extant literature (see e.g. Miranda-Agrippino and Rey (2020)), we focus on the immediate (one day) effects of monetary policy shocks, identified with high-frequency data, and also examine spillovers across the main economic areas (the US, euro area and Japan). To do so, we first construct measures of financial risk for each country by decomposing option-implied variances of the stock market into a conditional variance component (the amount-of-risk factor or "uncertainty") and a variance risk premium component (which we interpret as "risk aversion"). The conditional variance component is computed using a novel non-linear realized variance model.

Our first main result is that, in contrast to the extant literature focusing on a pre-2008 Global Financial Crisis sample (see e.g. Bekaert, Hoerova, and Lo Duca (2013)), the monetary

[19]Specifically, we initiate the global risk aversion series with $\widehat{GRA_{t=1}^i} = \widehat{gra_{t=1}^i} + 100$ and then calculate $\widehat{GRA_{t>1}^i} = \sum_{z=1}^t \widehat{gra_z^i}$ for subsequent days. Then, the global risk aversion (of interest) is calculated as $\widehat{GRA_t^{sum}} = \sum_i \widehat{GRA_t^i}$. We simply refer to $\widehat{GRA_t^{sum}}$ as $\widehat{GRA_t}$ in the remainder of this section for simplicity.

policy effects on risk and uncertainty have become weak or non-existent. Moreover, we do not find any evidence of spillovers through the risk channel. That is, we do not find a special role for the US "hegemon" affecting risk or uncertainty across large advanced economies or of monetary policy acting directly through risk or uncertainty to induce a global financial cycle. However, we do find evidence of significant spillovers through interest rates. In fact, as our second result, US monetary policy is not unique in its interest rate spillover effects; euro area monetary policy also has strong effects on US interest rates.

Our third main result regards the monetary policy effects on various domestic and international asset prices. The monetary policy effects for bond markets mostly mirror the results for interest rates. The results are weaker for information shocks than for pure monetary policy shocks. Monetary policy, especially in the US, has a strong domestic effect on stock market prices, but its international spillover effects are somewhat weak, and certainly not unexpected given the importance of the US stock market in global equity markets. Moreover, the spillover effects for equity prices from information shocks are stronger emanating from the euro area affecting the US than vice versa.

Fourth, we also separately examine alternative asset classes. Monetary policy effects are non-existent for commodity prices. For exchange rates, they are simply in line with the predictions of uncovered interest rate parity and thus work through the standard channel of interest rates. Monetary policy tightening decreases gold prices, with the results stronger when a "commodity effect" is removed from the gold price. Clearly, in assessing the own-country and spillover effects of monetary policy, it is important to differentiate across different asset classes.

Finally, we consider the effects of non-monetary policy-driven risk aversion and uncertainty shocks, which are highly correlated across countries. Not surprisingly, they have strong, mean reverting effects on stock prices, but weaker effects on interest rates and bond prices, where monetary policy effects are relatively more important.

In sum, our analysis mostly confirms Mr. Powell's conjecture that the role of US monetary policy in setting global financial conditions is exaggerated. However, we do not dispute the existence of a global financial cycle. In fact, when we extract a global factor from our risk aversion shocks, cleansed from the effects of monetary policy and macroeconomic announcements, it is highly correlated with Miranda-Agrippino and Rey (2020)'s global financial cycle variable. We defer further analysis of the real effects of such global risk conditions to future work.

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Table 1: Coefficient estimates of the best linear model.

This table reports the parameter estimates for the best linear model (L Model 5; henceforth L(5)); model selection is reported in the Appendix:

$$E_t\left[RV_{t+22}^{(22)}\right] = \widehat{\alpha} + \widehat{\beta^m} RV_t^{(22)} + \widehat{\beta^w} RV_t^{(5)} + \widehat{\beta^d} RV_t + \widehat{\gamma} IV_t \tag{13}$$

All regressions are based on daily observations. The standard errors reported in brackets are computed using 30 Newey-West lags. Bold values indicate that a coefficient is significant; ***, at the 1% significance level; **, 5%; *, 10%. Model specifications including AIC, BIC, MSE (mean squared errors) and adjusted R^2 are reported at the end of the table.

	US	EA	JP
$\hat{\alpha}$	**3.00E-04***	2.00E-04	**1.30E-03*****
	(2.00E-04)	(2.00E-04)	(2.00E-04)
$\hat{\beta^m}$	**0.215****	-0.0630	**0.344*****
	(0.100)	(0.074)	(0.117)
$\hat{\beta^w}$	**0.283****	**0.152*****	**0.142****
	(0.118)	(0.050)	(0.069)
$\hat{\beta^d}$	**0.093*****	0.0130	**0.021****
	(0.028)	(0.014)	(0.009)
$\hat{\gamma}$	**0.140*****	**0.514*****	0.0600
	(0.065)	(0.073)	(0.068)
AIC ($\times 10^4$)	-8.736	-5.667	-8.121
BIC ($\times 10^4$)	-8.733	-5.664	-8.117
MSE ($\times 10^{-4}$)	0.048	0.102	0.089
R2	0.586	0.507	0.338
UC-RA corr	0.559	0.928	0.429

Table 2: Coefficient estimates of the best non-linear model.

This table reports the parameter estimates for the best non-linear model for each country. The full local model with non-linear predictive coefficients (NL Model 1) is as follows:

$$E_t\left[RV_{t+22}^{(22)}\right] = \widehat{\alpha} + \frac{\exp(\widehat{\beta_0^m} - \widehat{\beta_1^m} Z_t^m)}{\exp(\widehat{\beta_0^m} - \widehat{\beta_1^m} Z_t^m) + 1} RV_t^{(22)} + \frac{\exp(\widehat{\beta_0^w} - \widehat{\beta_1^w} Z_t^w)}{\exp(\widehat{\beta_0^w} - \widehat{\beta_1^w} Z_t^w) + 1} RV_t^{(5)}$$
$$+ \frac{\exp(\widehat{\beta_0^d} - \widehat{\beta_1^d} Z_t^d)}{\exp(\widehat{\beta_0^d} - \widehat{\beta_1^d} Z_t^d) + 1} RV_t + \frac{\exp(\widehat{\gamma_0} - \widehat{\gamma_1} Z_t^{iv})}{\exp(\widehat{\gamma_0} - \widehat{\gamma_1} Z_t^{iv}) + 1} IV_t, \tag{14}$$

where instrument Z_t^m is $RV_t^{(22)}$; Z_t^w is $RV_t^{(5)}$; Z_t^d is RV_t; Z_t^{iv} is IV_t. The estimation is conducted by finding the minimum least squares of residuals. The model reported is the best out of 15 possible models according to the BIC criterion; the full model selection is reported in the Appendix. For a non-linear variance coefficient that is chosen, $\widehat{\beta^m}_0/\widehat{\beta^m}$ represents the $\widehat{\beta^m}_0$ parameter in the logistic function; if a linear variance coefficient is chosen, $\widehat{\beta^m}_0/\widehat{\beta^m}$ represents the actual variance coefficient $\widehat{\beta^m}$ and $\widehat{\beta^m}_1$ is not defined (analogously for the other coefficients). Bold values indicate that a coefficient is significant; ***, at the 1% significance level; **, 5%; *, 10%. Model specifications including AIC, BIC, MSE and adjusted R^2 are reported at the end of the table.

	US	EA	JP
$\widehat{\alpha}$	8.90E-05	-1.40E-04	**5.11E-04*** **
	(5.93E-05)	(9.30E-05)	(7.20E-05)
$\widehat{\beta^m}_0/\widehat{\beta^m}$	**0.223*** **	**-0.061*** **	-0.145
	(0.020)	(0.021)	(0.139)
$\widehat{\beta^m}_1$	–	–	**45.908*** **
			(6.486)
$\widehat{\beta^w}_0/\widehat{\beta^w}$	**-0.522*** **	**0.167*** **	**-0.442*** **
	(0.125)	(0.017)	(0.125)
$\widehat{\beta^w}_1$	**9.119*** **	–	**35.976*** **
	(2.015)		(4.217)
$\widehat{\beta^d}_0/\widehat{\beta^d}$	**-1.424*** **	**-1.468*** **	**-2.369*** **
	(0.128)	(0.226)	(0.244)
$\widehat{\beta^d}_1$	**7.758*** **	**26.674*** **	**18.652*** **
	(1.606)	(6.489)	(6.037)
$\widehat{\gamma}_0/\widehat{\gamma}$	**-1.880*** **	0.081	**0.047*** **
	(0.235)	(0.123)	(0.010)
$\widehat{\gamma}_1$	**36.514** **	**12.417*** **	–
	(16.541)	(3.239)	
AIC ($\times 10^4$)	-8.755	-5.676	-8.171
BIC ($\times 10^4$)	-8.749	-5.671	-8.166
MSE ($\times 10^{-4}$)	0.046	0.100	0.083
R2	0.597	0.516	0.384
UC-RA corr	0.536	0.873	0.316
UC corr with Best Linear	0.992	0.991	0.940
RA corr with Best Linear	0.980	0.979	0.978

Table 3: Economic magnitudes of non-linear coefficients.

This table shows the non-linear coefficients in the best non-linear model (Table 2) of several values of the (daily) Z-instruments. Given that all the variances used are positively skewed, we focus on the right-tail percentiles: mean, median, 75th, 90th, and 95th of the actual distributions of the Zs in the data.

	Mean	Median	75%	90%	95%
			US		
$RV^{(22)}$	0.223	0.223	0.223	0.223	0.223
$RV^{(5)}$	0.368	0.370	0.368	0.364	0.360
RV	0.192	0.193	0.192	0.189	0.186
IV	0.118	0.122	0.115	0.106	0.098
			EA		
$RV^{(22)}$	-0.061	-0.061	-0.061	-0.061	-0.061
$RV^{(5)}$	0.167	0.167	0.167	0.167	0.167
RV	0.173	0.180	0.173	0.159	0.144
IV	0.503	0.507	0.500	0.484	0.468
			JP		
$RV^{(22)}$	0.426	0.438	0.421	0.385	0.354
$RV^{(5)}$	0.363	0.375	0.361	0.334	0.302
RV	0.081	0.083	0.081	0.076	0.070
IV	0.047	0.047	0.047	0.047	0.047

Table 4: Properties of the risk variables.

This table reports statistical properties of country-level model-implied risk aversion (RA) and uncertainty (UC) using the longest country sample in Panel A, and cross-country correlation and Principal component analysis (PCA) in Panels B and C. To compute correlations of daily data, we consider all possible perspectives:

Time:	New York	Frankfurt	Tokyo
(1) US perspective:			
Trading Time	10:00 AM	4:00 PM	11:00 PM
Information set for US investor	(t)	(t)	(t)
(2) EA perspective:			
Trading Time	4:00 AM	10:00 AM	5:00 PM
Information set for US investor	(t-1)	(t)	(t)
(3) JP perspective:			
Trading Time	9:00 PM	3:00 AM	10:00 AM
Information set for US investor	(t-1)	(t-1)	(t)

The general rule of thumb is to consider the correctly-dated foreign information sets that enter the country representative investor's information by the time of trading. In Panel C, we report the 1st principal component (PC) result: the first three rows correspond to coefficient loadings on each country's risk variable, and the fourth row shows the % of total variance explained by the 1st PC.

	US	EA	JP		US	EA	JP
	1/2/1990	1/4/1999	11/1/1989		1/2/1990	1/4/1999	11/1/1989
N	7150	4954	7008		7150	4954	7008
Panel A. Country-level summary statistics							
	Risk Aversion				**Uncertainty**		
Mean	0.0017	0.0022	0.0022	Mean	0.0020	0.0035	0.0033
SD	0.0017	0.0022	0.0037	SD	0.0026	0.0033	0.0023
Skewness	3.7099	3.9078	5.4022	Skewness	5.8056	2.8503	1.9064
Panel B. Cross-country correlations (1/4/1999-5/31/2018)							
	Risk Aversion				**Uncertainty**		
(1) US perspective	US(t)	EA(t)	JP(t)	**(1) US perspective**	US(t)	EA(t)	JP(t)
US(t)	1.000	0.708	0.633	US(t)	1.000	0.774	0.724
EA(t)		1.000	0.730	EA(t)		1.000	0.581
JP(t)			1.000	JP(t)			1.000
(2) EA perspective	US(t-1)	EA(t)	JP(t)	**(2) EA perspective**	US(t-1)	EA(t)	JP(t)
US(t-1)	1.000	0.717	0.660	US(t-1)	1.000	0.760	0.730
EA(t)		1.000	0.730	EA(t)		1.000	0.582
JP(t)			1.000	JP(t)			1.000
(3) JP perspective	US(t-1)	EA(t-1)	JP(t)	**(3) JP perspective**	US(t-1)	EA(t-1)	JP(t)
US(t-1)	1.000	0.708	0.660	US(t-1)	1.000	0.774	0.730
EA(t-1)		1.000	0.724	EA(t-1)		1.000	0.582
JP(t)			1.000	JP(t)			1.000
Panel C. 1st principal component coefficients and explained variance (1/4/1999-5/31/2018)							
	Risk Aversion				**Uncertainty**		
Perspective:	(1) US	(2) EA	(3) JP	**Perspective:**	(1) US	(2) EA	(3) JP
Coeff on US	0.220	0.224	0.229	Coeff on US	0.780	0.781	0.780
Coeff on EA	0.800	0.799	0.721	Coeff on EA	0.499	0.495	0.498
Coeff on JP	0.558	0.558	0.654	Coeff on JP	0.378	0.382	0.380
1st PC %	85.2	85.4	90.9	1st PC %	83.9	83.6	84.0

Table 5: Summary statistics for monetary policy shocks.

This table reports summary statistics for our main MP shock measures from 2000 to 2015: for the US and the EA, the cleansed monetary policy shocks "MP cleansed" and central bank information shocks "CB info" (from Jarociński and Karadi (2020), JK for short); for Japan, a measure based on the 10-year Japanese government bond futures yields (from Rogers, Scotti, and Wright (2014), RSW for short).

	Mean	SD	5%	95%	N
US MP cleansed JK	-0.598	6.635	-11.209	7.186	137
US CB info JK	-0.930	6.594	-11.384	8.322	137
EA MP cleansed JK	0.254	6.095	-8.453	7.994	263
EA CB info JK	-0.451	5.910	-9.980	7.960	263
JP RSW shock	-0.116	1.459	-2.420	2.104	247

Table 6: Monetary Policy, Risk Aversion, and Uncertainty

This table reports the direct and spillover effects of monetary policy (MP), risk aversion (RA), and uncertainty (UC) in Section 3 (or Equation (4) in Section 2):

$$\Delta RA_{j,t} = \alpha_j + \sum_{i=US,EA,JP} \beta_j^{MP,i} MP_t^i + \sum_{i \neq j} \beta_j^{RA,i} \overline{ra_t^i} + \sum_{i=US,EA,JP} \gamma_j^i D_{t,i} + \sum_{k=Macro} \delta_j^k Macro_t^k + \sum_{k=Macro} r_j^k \theta_t^k + \varepsilon_{j,t}$$

$$\Delta UC_{j,t} = \alpha_j + \sum_{i=US,EA,JP} \beta_j^{MP,i} MP_t^i + \sum_{i \neq j} \beta_j^{UC,i} \overline{uc_t^i} + \sum_{i=US,EA,JP} \gamma_j^i D_{t,i} + \sum_{k=Macro} \delta_j^k Macro_t^k + \sum_{k=Macro} r_j^k \theta_t^k + \varepsilon_{j,t}$$

where $\Delta RA_{j,t}$ ($\Delta UC_{j,t}$) is the first-differenced risk aversion (uncertainty) measure of country j. $\beta_j^{MP,i}$ captures the effect of domestic or foreign MP shocks; $\beta_j^{RA,i}$ ($\beta_j^{UC,i}$) captures the effect of the foreign non-MP risk aversion (uncertainty) shocks denoted by $\overline{ra_t}$ ($\overline{uc_t}$). The $\beta_j^{MP,i}$ and $\beta_j^{RA,i}$ coefficients from the first equation above are reported in Panel A; the $\beta_j^{MP,i}$ and $\beta_j^{UC,i}$ coefficients from the second equation above are reported in Panel B. For each country, non-MP risk aversion or uncertainty shocks are obtained from the "direct effect" regression results. Risk aversion and uncertainty measures are in monthly percentages squared. MP shocks are in basis points. We consider MP cleansed shocks and information shocks from Jarociński and Karadi (2020). All regressions control for monetary policy and macro event dummies and macro news, as described in Section 2. Bold values indicate that a coefficient is significant; ***, at the 1% significance level; **, 5%; *, 10%.

	US, direct	US → EA	US →JP	EA, direct	EA→US	EA→JP	JP, direct	JP→US	JP→EA
Panel A. Risk Aversion									
MP cleansed shocks JK	0.201	0.021	0.096	-0.067	0.394	0.560**	0.077	-0.856	0.068
	(0.292)	(0.116)	(0.173)	(0.220)	(0.320)	(0.270)	(1.012)	(0.745)	(0.447)
CB info shocks JK	-0.146	0.065	-0.123	**-0.438****	**-0.705*****	-0.172			
	(0.394)	(0.225)	(0.249)	(0.172)	(0.250)	(0.301)			
non-MP RA shocks		**0.094***	**0.307*****		**0.468*****	0.114		-0.066	**0.157****
		(0.057)	(0.078)		(0.148)	(0.125)		(0.060)	(0.065)
Panel B. Uncertainty									
MP cleansed shocks JK	0.086	-0.020	-0.029	**0.358***	**-0.311*****	0.033	-0.151	0.341	0.495
	(0.115)	(0.078)	(0.029)	(0.190)	(0.117)	(0.035)	(0.315)	(0.357)	(0.558)
CB info shocks JK	-0.286	-0.086	-0.019	-0.152	**-0.142***	-0.056			
	(0.250)	(0.149)	(0.037)	(0.145)	(0.083)	(0.043)			
non-MP UC shocks		0.068	**0.079*****		**0.501*****	**0.053*****		0.090	**0.571*****
		(0.064)	(0.023)		(0.054)	(0.014)		(0.082)	(0.117)

Table 7: The effects of MP and risk shocks on interest rates

This table reports the regression in Section 4 for interest rates (or Equation (5) in Section 2):

$$Y_{j,t} = \alpha_j + \sum_{i=US,EA,JP} \beta_j^{MP,i} MP_t^i + \sum_{i=US,EA,JP} \beta_j^{RA,i} \overline{ra_t^i} + \sum_{i=US,EA,JP} \beta_j^{UC,i} \overline{uc_t^i} + \sum_{i=US,EA,JP} \gamma_j^i D_{t,i} + \sum_{k=Macro} \delta_j^k Macro_t^k + \sum_{k=Macro} r_j^k \theta_t^k + \varepsilon_{j,t} \tag{15}$$

where $Y_{j,t}$ denotes the first-differenced interest rates of country j; $\beta_j^{MP,i}$, $\beta_j^{RA,i}$ and $\beta_j^{UC,i}$ capture the effects of the MP, and non-MP risk aversion and uncertainty shocks (denoted as $\overline{ra_t^i}$ and $\overline{uc_t^i}$, respectively) of country i on the interest rate changes of country j. All regressions control for monetary policy and macro event dummies and macro news, as described in Section 2. Bold values indicate that a coefficient is significant; ***, at the 1% significance level; **, 5%; *, 10%.

	US, direct	US → EA	US→DE	US →JP	EA, direct	DE, direct	EA →US	EA →JP	JP, direct	JP→US	JP→EA	JP→DE
MP cleansed shocks JK	**0.323*****	**0.140*****	**0.152****	0.031	**0.216***	**0.276*****	-0.004	0.001	**0.628*****	-0.344	0.117	**0.433***
	(0.109)	(0.044)	(0.072)	(0.039)	(0.115)	(0.103)	(0.033)	(0.028)	(0.144)	(0.213)	(0.349)	(0.240)
CB info shocks JK	**0.634*****	0.086	0.120	0.017	0.044	**0.465*****	**0.091****	**0.071****				
	(0.168)	(0.053)	(0.082)	(0.038)	(0.158)	(0.103)	(0.046)	(0.032)				
non-MP RA shocks	-0.029	-0.026	**-0.066****	-0.000	-0.020	0.004	-0.043	**-0.013****	**-0.026*****	**0.027***	0.004	-0.003
	(0.022)	(0.023)	(0.027)	(0.006)	(0.015)	(0.034)	(0.038)	(0.006)	(0.006)	(0.016)	(0.014)	(0.014)
non-MP UC shocks	**-0.075****	**-0.077****	-0.060	-0.004	-0.003	-0.014	-0.011	0.001	-0.034	**-0.109*****	-0.070	-0.009
	(0.033)	(0.035)	(0.040)	(0.009)	(0.018)	(0.017)	(0.022)	(0.008)	(0.022)	(0.040)	(0.049)	(0.032)

Table 8: The effects of MP and risk shocks on stock and bond returns

This table reports the regression in Equation (5), where $Y_{j,t}$ is the stock returns of country j in Panel A, and bond returns in Panel B. Other details are in Table 6 and Table 7. Bold values indicate that a coefficient is significant; ***, at the 1% significance level; **, 5%; *, 10%.

	US, direct	US → EA	US→JP	EA, direct	EA→US	EA→JP	JP, direct	JP→US	JP→EA
Panel A. Stock Returns									
MP cleansed shocks JK	**-0.085*****	**-0.025****	-0.016	**-0.095*****	**-0.046*****	**-0.057*****	0.038	0.042	-0.030
	(0.017)	(0.011)	(0.018)	(0.013)	(0.011)	(0.010)	(0.096)	(0.045)	(0.082)
CB info shocks JK	**0.039****	**0.061*****	-0.029	**0.141*****	**0.116*****	**0.071*****			
	(0.019)	(0.016)	(0.049)	(0.012)	(0.011)	(0.012)			
non-MP RA shocks	**-0.081*****	0.005	**-0.013****	**-0.057*****	**0.009***	**-0.019*****	**-0.056*****	**0.005****	**0.009****
	(0.005)	(0.005)	(0.005)	(0.013)	(0.005)	(0.006)	(0.005)	(0.002)	(0.004)
non-MP UC shocks	**-0.076*****	0.008	**-0.012***	**-0.109*****	**-0.017*****	-0.007	**-0.076*****	0.002	0.008
	(0.007)	(0.005)	(0.007)	(0.007)	(0.004)	(0.005)	(0.012)	(0.008)	(0.010)
Panel B. Bond Returns									
MP cleansed shocks JK	**-0.029*****	-0.014	-0.004	**-0.013****	0.007	0.000	**-0.053*****	**-0.067****	-0.010
	(0.011)	(0.010)	(0.004)	(0.006)	(0.008)	(0.003)	(0.016)	(0.030)	(0.016)
CB info shocks JK	-0.000	-0.005	-0.001	**-0.009***	**-0.019*****	**-0.007****			
	(0.015)	(0.008)	(0.004)	(0.005)	(0.006)	(0.003)			
non-MP RA shocks	**0.007*****	-0.001	0.000	**0.005*****	0.002	**0.001****	**0.002*****	**-0.003****	-0.001
	(0.002)	(0.001)	(0.001)	(0.002)	(0.002)	(0.001)	(0.001)	(0.001)	(0.001)
non-MP UC shocks	**0.010*****	**-0.003*****	0.000	**0.008*****	**0.010*****	-0.000	0.003	**-0.005***	-0.004
	(0.003)	(0.001)	(0.001)	(0.001)	(0.002)	(0.001)	(0.002)	(0.003)	(0.002)

Table 9: The effects of MP and risk shocks on exchange rates, gold and commodity prices

This table reports the regression in Equation (5), where $Y_{j,t}$ is the log change in exchange rates of country j (defined in units of domestic currency per foreign currency) in Panel A, and log changes in gold and commodity prices in Panel B respectively. To be specific, in Panel A, only direct effects are considered (e.g., columns USD per EUR and USD per JPY display effects of US MP vs non-MP-driven RA and UC shocks on the dollar exchange rate vis-a-vis the euro and the yen, respectively). In Panel B, "GSCI" is the log change in the Goldman Sachs commodity price index (GSCI for short) while "Gold" is the gold price cleansed of the commodity price component (residuals from the regressions of log changes in the gold price on the GSCI; gold prices are quoted in US dollars and measured at the end of the trading day in the US). All regressions control for monetary policy and macro event dummies and macro news, as described in Section 2. Other details are in Table 6 and Table 7. Bold values indicate that a coefficient is significant; ***, at the 1% significance level; **, 5%; *, 10%.

Panel A. Exchange Rates	US → USD per EUR	US → USD per JPY	EA → EUR per USD	EA → EUR per JPY	JP → JPY per USD	JP → JPY per EUR
MP cleansed shocks JK	**-0.042****	**-0.033*****	-0.018	-0.012	-0.044	0.088
	(0.017)	(0.012)	(0.012)	(0.011)	(0.035)	(0.090)
CB info shocks JK	0.005	-0.011	-0.001	-0.009		
	(0.019)	(0.014)	(0.010)	(0.012)		
non-MP RA shocks	**-0.007*****	**0.016*****	0.003	**0.011*****	**-0.009*****	**-0.011*****
	(0.001)	(0.002)	(0.002)	(0.004)	(0.001)	(0.003)
non-MP UC shocks	-0.002	**0.016*****	0.002	**0.018*****	-0.006	0.000
	(0.003)	(0.003)	(0.002)	(0.003)	(0.005)	(0.006)

Panel B. Gold and Commodity Prices	US → Gold	EA → Gold	JP → Gold	US → GSCI	EA → GSCI	JP → GSCI
MP cleansed shocks JK	**-0.063*****	-0.010	-0.010	0.019	-0.019	-0.058
	(0.020)	(0.010)	(0.067)	(0.019)	(0.016)	(0.073)
CB info shocks JK	0.017	-0.018		-0.023	**0.054*****	
	(0.022)	(0.011)		(0.033)	(0.016)	
non-MP RA shocks	0.004	-0.005	-0.002	**-0.020*****	**-0.012****	-0.001
	(0.005)	(0.006)	(0.003)	(0.004)	(0.005)	(0.004)
non-MP UC shocks	**0.016****	0.003	-0.007	**-0.031*****	**-0.025*****	-0.004
	(0.006)	(0.004)	(0.007)	(0.007)	(0.005)	(0.008)

Table 10: The $\widehat{\boldsymbol{D}}$ matrix from the time-zone global risk aversion model

This table reports the construction of global risk aversion shocks (row) at the three time zones as linear functions of country risk aversion shocks (column) according to Equation (11), or matrix $\widehat{\boldsymbol{D}}$ in Equation (12).

Loadings on:	$\overline{ra^{JP}}$	$\overline{ra^{EA}}$	$\overline{ra^{US}}$
$\widehat{gra_t^{JP}}$	0.0930	0.1961	0.1154
$\widehat{gra_t^{EA}}$	-0.1144	0.8212	0.3998
$\widehat{gra_t^{US}}$	0.0047	-0.2242	0.4237

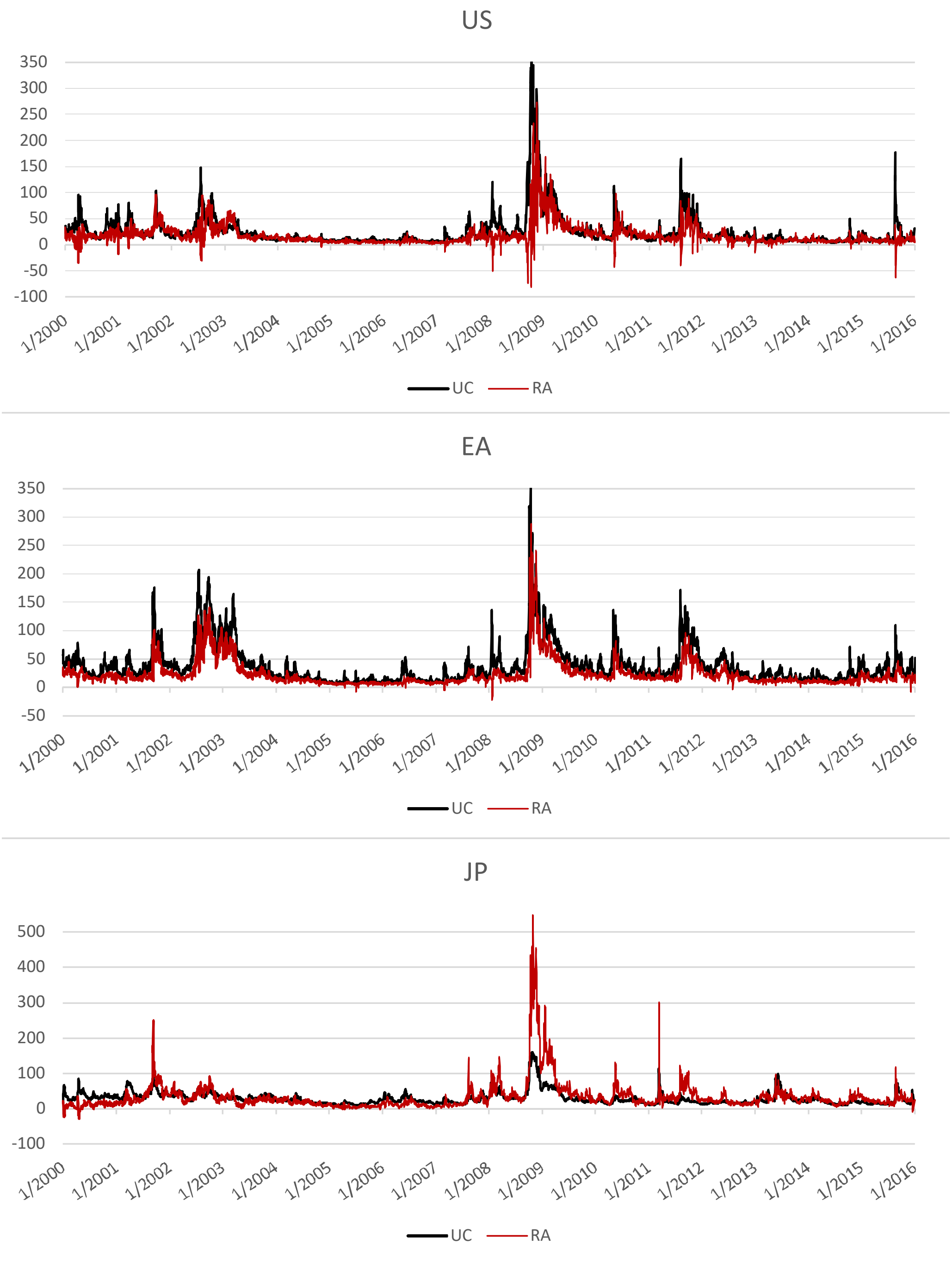

Figure 1: Time series of daily local UC (black) and RA (red) estimates based on best non-linear model in Table 2. For consistency, we show the sample that is used in the Monetary Policy part of the paper (2000-2015). Both RA and UC are in monthly percentages squared.

Figure 2: The effects of MP and risk shocks on interest rates (IR), stock returns (SR) and bond returns (BR).

The figure shows the change in IR, SR or BR expressed in terms of standard deviations given a 1 standard deviation (SD) increase in a pure MP, CB information, risk aversion or uncertainty shock. The top three plots compare the average domestic versus foreign effects of the monetary policy shocks, across the three country regressions; for the RA and UC shocks, bars represent the average total effect (domestic+foreign) averaged across the three country regressions. The bottom three plots show the domestic versus foreign effects of a pure MP or CB information shock emanating from a particular country; for instance, the first dark bar (0.4132) shows the change of US interest rate (in units of SD) in response a 1 SD US pure MP shock, while the first light bar (0.1285) shows the average change (in units of SD) of Japanese and euro interest in response to a 1 SD US pure MP shock.

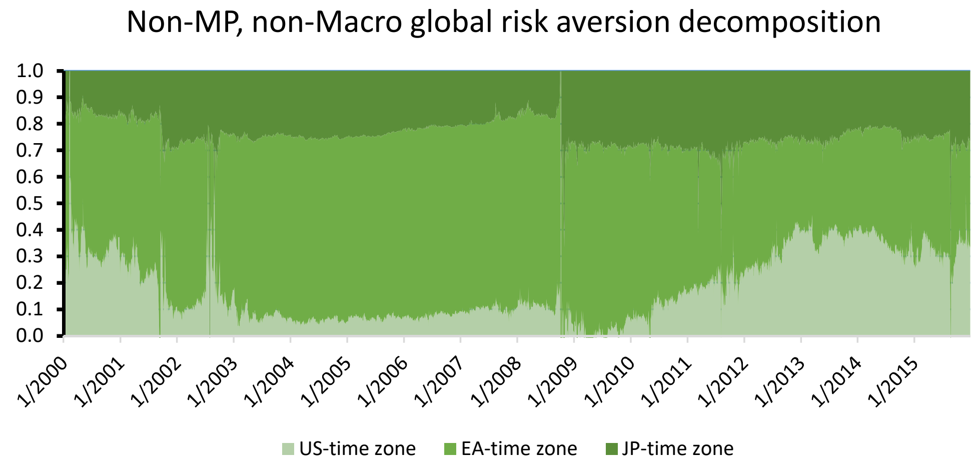

Figure 3: The time zone factor model: estimated daily global risk aversion decomposition, by time zones.

The fraction is calculated as $\widehat{GRA_t^i}/\widehat{GRA_t^{sum}}$. Note that this is not a decomposition of JP, EA, or US risk aversion levels, but a decomposition of how much global risk aversion from a particular time zone contributes to global risk aversion.

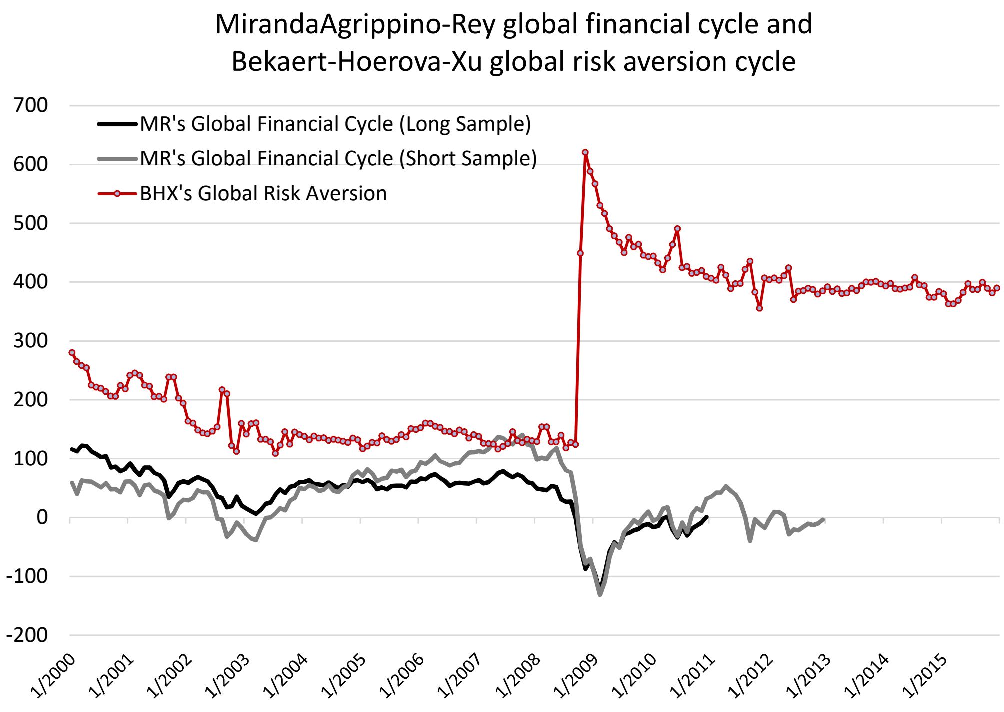

Figure 4: Comparisons to Miranda-Agrippino and Rey (2020)'s global financial cycle variables.

The correlation between our global risk aversion measure and Miranda-Agrippino and Rey (2020)'s global financial cycle using long sample (short sample) is -0.76 (-0.71).

A Appendix: A simple dynamic asset pricing model for Section 2

We set out a consumption-based asset pricing model, which is a variant of the model in Bekaert, Engstrom, and Xing (2009), BEX henceforth. The model features three key state variables, expected consumption growth (g_t), uncertainty (the conditional variance of consumption growth, UC_t), and stochastic risk aversion (RA_t). The modelling of consumption and dividend growth is simpler than in BEX, who assume they are cointegrated.

A.1 Fundamental and preferences

The dynamics of the state variables for consumption growth (Δc_{t+1}) and its conditional moments are given by:

$$\Delta c_{t+1} = \mu_c + g_t + \sqrt{UC_t}\varepsilon_{c,t+1}, \tag{A1}$$

$$UC_{t+1} = \mu_{UC} + \rho_{UC}UC_t + \sigma_{UC}\sqrt{UC_t}\varepsilon_{UC,t+1}, \tag{A2}$$

$$g_{t+1} = \rho_g g_t + \sigma_{gc}\underbrace{\sqrt{UC_t}\varepsilon_{c,t+1}}_{\Delta c_{t+1}-E_t[\Delta c_{t+1}]} + \sigma_{gg}\sqrt{UC_t}\varepsilon_{g,t+1}. \tag{A3}$$

The risk aversion process loads on the consumption growth shock, but also features an uncorrelated preference shock, which is heteroskedastic, that is, risk aversion becomes more variable as it increases in value:

$$RA_{t+1} = \mu_{RA} + \rho_{RA}RA_t + \sigma_{RAc}\sqrt{UC_t}\varepsilon_{c,t+1} + \sigma_{RA}\sqrt{RA_t}\varepsilon_{RA,t+1}. \tag{A4}$$

Dividend growth (Δd_{t+1}) similarly loads on consumption growth and an independent homoskedastic shock:

$$\Delta d_{t+1} = \mu_d + \rho_{dg}g_t + \sigma_{dc}\sqrt{UC_t}\varepsilon_{c,t+1} + \sigma_d\varepsilon_{d,t+1}. \tag{A5}$$

Shocks $\varepsilon_{c,t+1}, \varepsilon_{UC,t+1}, \varepsilon_{g,t+1}, \varepsilon_{RA,t+1}$ and $\varepsilon_{d,t+1}$ are independently and normally distributed $N(0,1)$.

The agent maximizes $E_t\left[\sum_{t=0}^{\infty}\beta^t\frac{(C_t-H_t)^{1-\gamma}}{1-\gamma}\right]$, with $C_t > H_t$ and H_t is the habit stock. Define $Q_t \equiv \frac{C_t}{C_t-H_t} > 1$. This is the inverse of Campbell and Cochrane (1999)'s surplus consumption ratio. The equilibrium pricing kernel is $M^*_{t+1} = \beta\frac{(C_{t+1}/C_t)^{-\gamma}}{(Q_{t+1}/Q_t)^{-\gamma}}$, and the equilibrium log real pricing kernel is,

$$\begin{aligned} m^*_{t+1} &= log\beta - \gamma\Delta c_{t+1} + \gamma(q_{t+1} - q_t) \\ &= log\beta - \gamma(\mu_c + g_t - \mu_{RA} + (1-\rho_{RA})RA_t) - \gamma(1-\sigma_{RA})\sqrt{UC_t}\varepsilon_{c,t+1} + \gamma\sigma_{RA}\sqrt{RA_t}\varepsilon_{RA,t+1}. \end{aligned} \tag{A6}$$

In this model q_t essentially represents stochastic risk aversion, so $q_t = RA_t$.

A.2 Asset price: Real interest rate

First, the real rate in equilibrium is, (using a superscript $*$ to denote equilibrium value)

$$\begin{aligned} rf^*_t &= -logE^*_t\left[\exp\left(m_{t+1}\right)\right], \\ &= k_0 + k_g g_t + k_{RA}RA_t + k_{UC}UC_t, \end{aligned} \tag{A7}$$

where

$$
\begin{aligned}
k_0 &= -log\beta + \gamma(\mu_c - \mu_{RA}) \\
k_g &= \gamma \\
k_{RA} &= \gamma(1-\rho_{RA}) - \frac{1}{2}\gamma^2\sigma_{RA}^2 \\
k_{UC} &= -\frac{1}{2}\gamma^2(1-\sigma_{RAc})^2.
\end{aligned}
$$

We do not model the monetary policy transmission function directly, instead assuming there exists a non-persistent monetary policy shock, $MP_t \sim N(0, \sigma_{MP})$, that can affect the various state variables directly and is uncorrelated with $\{\varepsilon_{c,t+1}, \varepsilon_{UC,t+1}, \varepsilon_{g,t+1}, \varepsilon_{RA,t+1}, \varepsilon_{d,t+1}\}$. This is tantamount to adding $\phi_x MP_{t+1}$, with $x = UC$, g, and RA, to Equations (A2), (A3), and (A4), respectively. We discuss the various channels through which such effects can occur in the main text in Section 2.

Because the shock is not persistent, it will not affect pricing equations. In addition, we must allow for monetary policy to affect interest rates directly. Assume that there is a wedge between the equilibrium real pricing kernel and the true pricing kernel, M_{t+1}, such that $M_{t+1} = M^*_{t+1}exp(-\phi_{MP}MP_t)$. This is equivalent to assuming that monetary policy affects liquidity in the market for short term securities; a contractionary shock decreases liquidity and drives up the liquidity premium and vice versa (see, e.g., Drechsler, Savov, and Schnabl (2018a)). Therefore, the actual real rate equals:

$$
rf_t = rf^*_t + \phi_{MP}MP_t. \tag{A8}
$$

With this structure, monetary policy potentially transmits to the real economy through an information shock/expected cash flow channel (through ϕ_g), though risk channels (through ϕ_{UC} and ϕ_{RA}) and directly through ϕ_{MP}. MP_t here acts as a pure term structure level factor.

For simplicity, we focus on the special case of $\phi_g = 0$, $\phi_{RA} = 0$, and $\phi_{UC} = 0$ to describe the model solutions, which are correct up to a constant term for the general case as well.

A.3 Asset prices: Long-term real bond prices

A.3.1 Two-period zero-coupon bond price

As derived above, the price for the one-period zero-coupon real bond is,

$$
P_{1,t} = E_t\left[\exp\left(m_{t+1}\right)\right] = \exp\left(A_1 + B_1 g_t + C_1 RA_t + D_1 UC_t - \phi_{MP}MP_t\right), \tag{A9}
$$

where

$$
\begin{aligned}
A_1 &= log\beta - \gamma(\mu_c - \mu_{RA}) \\
B_1 &= -\gamma - \rho_{\pi g} \\
C_1 &= -\gamma(1-\rho_{RA}) + \frac{1}{2}\gamma^2\sigma_{RA}^2 \\
D_1 &= \frac{1}{2}\gamma^2(1-\sigma_{RAc})^2
\end{aligned}
$$

The price for the two-period zero-coupon real bond is,

$$\begin{aligned} P_{2,t} &= E_t\left[M_{t+1}P_{1,t+1}\right] \\ &= E_t\left[\exp\left(m_{t+1} + \underbrace{A_1 + B_1 g_{t+1} + C_1 RA_{t+1} + D_1 UC_{t+1} - \phi_{MP} MP_{t+1}}_{\Delta_{t+1} \equiv -rf_{t+1}}\right)\right]. \end{aligned} \tag{A10}$$

We can rewrite m_{t+1} and Δ_{t+1} in matrix representations:

$$m_{t+1} = m_0 + \boldsymbol{m_1}\begin{bmatrix} g_t \\ RA_t \end{bmatrix} + \boldsymbol{m_2}\begin{bmatrix} \sqrt{UC_t}\varepsilon_{c,t+1} \\ \sqrt{RA_t}\varepsilon_{RA,t+1} \end{bmatrix} - \phi_{MP} MP_t,$$

$$\Delta_{t+1} \equiv -rf_{t+1} = \Delta_0 + \boldsymbol{\Delta_1}\begin{bmatrix} g_t \\ RA_t \\ UC_t \end{bmatrix} + \boldsymbol{\Delta_2}\begin{bmatrix} \sqrt{UC_t}\varepsilon_{c,t+1} \\ \sqrt{RA_t}\varepsilon_{RA,t+1} \\ \sqrt{UC_t}\varepsilon_{UC,t+1} \\ MP_{t+1} \end{bmatrix}.$$

Then, Equation (A10) can be solved as follows:

$$\begin{aligned} P_{2,t} &= \exp\left\{ \begin{array}{l} E_t(m_{t+1}) + \frac{1}{2}V_t(m_{t+1}) \\ +E_t(\Delta_{t+1}) + \frac{1}{2}V_t(\Delta_{t+1}) \\ +Cov_t(m_{t+1}, \Delta_{t+1}) \end{array} \right\} \\ &= \exp\left[A_2 + B_2 g_t + C_2 RA_t + D_2 UC_t - \phi_{MP} MP_t\right]. \end{aligned} \tag{A11}$$

A.3.2 Term premia

The yield rate for the two-period real bond, $y_{2,t} = -\frac{\log(P_{2,t})}{2}$, can be derived as:

$$\begin{aligned} y_{2,t} &= -\frac{1}{2}\left\{ \begin{array}{ll} E_t(m_{t+1}) + \frac{1}{2}V_t(m_{t+1}) & [= -rf_t^* - \phi_{MP} MP_t] \\ +E_t(\Delta_{t+1}) & [\text{1. Expectations Hypothesis terms}] \\ +\frac{1}{2}V_t(\Delta_{t+1}) & [\text{2. Jensen's inequality term}] \\ +\,Cov_t(m_{t+1}, \Delta_{t+1}) & [\text{3. Bond term premium channel}] \end{array} . \right\} \\ &= \frac{1}{2}(rf_t^* + \phi_{MP} MP_t) + \frac{1}{2}E_t(rf_{t+1}) - \frac{1}{4}V_t(rf_{t+1}) + \frac{1}{2}Cov_t(m_{t+1}, rf_{t+1}), \end{aligned} \tag{A12}$$

where the term premium component $tp_t = Cov_t(m_{t+1}, rf_{t+1})$ is given by:

$$tp_t = \underbrace{(-m_{2,c}\Delta_{2,c})}_{\eta_{UC}} UC_t + \underbrace{(-m_{2,RA}\Delta_{2,RA})}_{\eta_{RA}} RA_t, \tag{A13}$$

which was shown in Equation (2) in Section 2.

A.3.3 N-period zero-coupon real bond price

By induction, it can be easily shown that

$$P_{N,t} = \exp\left[A_N + B_N g_t + C_N RA_t + D_N UC_t - \phi_{MP} MP_t\right], \tag{A14}$$

where,

$$
\begin{aligned}
A_N &= log\beta - \gamma\mu_c + \gamma\mu_{RA} + A_{N-1} + C_{N-1}\mu_{RA} + D_{N-1}\mu_{UC} + \frac{1}{2}\phi_{MP}^2\sigma_{MP}^2 \\
B_N &= -\gamma - \rho_{\pi g} + B_{N-1}\rho_g \\
C_N &= -\gamma(1-\rho_{RA}) + C_{N-1}\rho_{RA} + \frac{1}{2}\left(\gamma\sigma_{RA} + C_{N-1}\sigma_{RA}\right)^2 \\
D_N &= D_{N-1}\rho_{UC} + \frac{1}{2}\left(-\gamma(1-\sigma_{RAc}) + B_{N-1}\sigma_{gc} + C_{N-1}\sigma_{RAc}\right)^2 + \frac{1}{2}(B_{N-1})^2\sigma_g^2 + \frac{1}{2}(D_{N-1})^2\sigma_{UC}^2
\end{aligned}
$$

Equation (A14) shows that the price of a N-period zero-coupon real bond is determined by expected growth, risk aversion, uncertainty, and the monetary policy shock. Intuitively, a positive MP shock leads to a lower long-term bond price today, with the pass-through depending on the persistence of the various shocks affecting short-term interest rate. Apart from this EH effect, the MP shock can also affect the state variables itself through an information (expected growth) or risk (risk aversion, uncertainty) channel.

A.3.4 Contemporaneous log long-term bond returns

Denote $\boldsymbol{Y_t} = \begin{bmatrix} g_t & RA_t & UC_t & MP_t \end{bmatrix}'$. The contemporaneous log bond return, $\tilde{r}_t^b = log\left(\frac{P_{N-1,t}}{P_{N,t-1}}\right)$, can be derived as follows:

$$
r_t^b = \xi_0^b + \boldsymbol{\xi_1^b Y_{t-1}} + \boldsymbol{\xi_2^b} \begin{bmatrix} g_t - E_{t-1}(g_t) \\ RA_t - E_{t-1}(RA_t) \\ UC_t - E_{t-1}(UC_t) \\ MP_t \end{bmatrix}, \tag{A15}
$$

where ξ_0^b, $\boldsymbol{\xi_1^b}$, and $\boldsymbol{\xi_2^b}$ are implicitly defined. This equation motivates the four shocks that the paper uses.

A.4 Asset prices: Stock price

A.4.1 Price-dividend ratio

The price-dividend ratio, $PD_t = E_t\left[M_{t+1}\left(\frac{P_{t+1}+D_{t+1}}{D_t}\right)\right]$, can be rewritten as,

$$
PD_t \;\; = \;\; \sum_{n=1}^{\infty} E_t\left[\exp\left(\sum_{j=1}^{n} m_{t+j} + \Delta d_{t+j}\right)\right]. \tag{A16}
$$

Let $F_{n,t}$ denote the n-th term in the summation:

$$
F_{n,t} = E_t\left[\exp\left(\sum_{j=1}^{n} m_{t+j} + \Delta d_{t+j}\right)\right], \tag{A17}
$$

and $F_{n,t}D_t$ can be interpreted as the price of zero-coupon equity that matures in n periods. We can rewrite $\Delta d_{t+1} = d_0 + d_1 g_t + \boldsymbol{d_2}\begin{bmatrix}\sqrt{UC_t}\varepsilon_{c,t+1} \\ \varepsilon_{d,t+1}\end{bmatrix}$. The first term, $F_{1,t}$, can be solved as follows:

$$
\begin{aligned}
F_{1,t} &= E_t\left[\exp\left(m_{t+1}+\Delta d_{t+1}\right)\right] \\
&= \exp\left\{\begin{array}{ll} E_t(m_{t+1})+\frac{1}{2}V_t(m_{t+1}) & [1.\text{ Interest rate channel, } = -rf_t^* - \phi_{MP}MP_t] \\ +E_t(\Delta d_{t+1})+\frac{1}{2}V_t(\Delta d_{t+1}) & [2.\text{ Cash flow channel}] \\ +Cov_t(m_{t+1},\Delta d_{t+1}) & [3.\text{ premium channel (from pure cash flow)}] \end{array}\right\} \\
&= \exp\left(e_{1,0}+\boldsymbol{e_{1,1}}\left[g_t \quad RA_t \quad UC_t\right]' - \phi_{MP}MP_t\right) \qquad \text{(A18)}
\end{aligned}
$$

Suppose $F_{N-1,t} = \exp\left(e_{N-1,0}+\boldsymbol{e_{N-1,1}}\left[g_t \quad RA_t \quad UC_t\right]' - \phi_{MP}MP_t\right) \equiv \exp(f_{N-1,t})$, and $f_{N-1,t+1}$ can be rewritten as $f_{N-1,0}+\boldsymbol{f_{N-1,1}}\left[g_t \quad RA_t \quad UC_t\right]' + \boldsymbol{f_{N-1,2}}\begin{bmatrix}\sqrt{UC_t}\varepsilon_{c,t+1} \\ \sqrt{RA_t}\varepsilon_{RA,t+1} \\ \sqrt{UC_t}\varepsilon_{UC,t+1} \\ MP_{t+1}\end{bmatrix}$.

By induction,

$$
\begin{aligned}
F_{N,t} &= E_t\left[\exp\left(m_{t+1}\right)\underbrace{E_{t+1}\left(\exp\left(\sum_{j=1}^{N-1} m_{t+j+1}-\pi_{t+j+1}+\Delta d_{t+j+1}\right)\right)}_{F_{N-1,t+1}}\right] \\
&= \exp\left\{\begin{array}{ll} E_t(m_{t+1})+\frac{1}{2}V_t(m_{t+1}) & [1.\text{ Interest rate channel, } = -rf_t^* - \phi_{MP}MP_t \\ +E_t(f_{N-1,t+1})+\frac{1}{2}V_t(f_{N-1,t+1}) & \\ +(m_{2,c}f_{N-1,2,c})UC_t+(m_{2,RA}f_{N-1,2,RA})RA_t & [2.\text{ risk premium channel}] \end{array}\right\} \\
&= \exp\left(e_{N,0}+\boldsymbol{e_{N,1}}\left[g_t \quad RA_t \quad UC_t\right]' - \phi_{MP}MP_t\right). \qquad \text{(A19)}
\end{aligned}
$$

Hence, the price-dividend ratio is approximately affine:

$$
\begin{aligned}
PD_t &= \sum_{n=1}^{\infty} E_t\left[\exp\left(\sum_{j=1}^{n} m_{t+j}+\Delta d_{t+j}\right)\right] \\
&= \sum_{n=1}^{\infty} F_{n,t} \\
&= \sum_{n=1}^{\infty} \exp\left(e_{n,0}+\boldsymbol{e_{n,1}}\left[g_t \quad RA_t \quad UC_t\right]' - \phi_{MP}MP_t\right), \qquad \text{(A20)}
\end{aligned}
$$

which implies that a positive MP shock could result in a lower stock price today (hence a lower contemporaneous stock return). Similarly, apart from this EH effect, the MP shock can also affect the state variables itself through an information or risk channel.

A.4.2 Contemporaneous log stock returns

As previously defined, $\boldsymbol{Y_t} = \left[g_t \quad RA_t \quad UC_t \quad MP_t\right]'$. We apply first-order Taylor approximations to the log stock return, from $t-1$ to t (as our paper focuses on contemporaneous changes), and obtain a linear system.

$$
\begin{aligned}
r_t^{eq} &= \Delta d_t + \ln\left[\frac{1+\sum_{n=1}^{\infty}\exp\left(e_{n,0}+\boldsymbol{e_{n,1}Y_t}\right)}{\sum_{n=1}^{\infty}\exp\left(e_{n,0}+\boldsymbol{e_{n,1}Y_{t-1}}\right)}\right] \\
&\approx \Delta d_t + \text{const.} + \frac{\sum_{n=1}^{\infty}\exp\left(e_{n,0}+\boldsymbol{e_{n,1}\bar{Y}}\right)\boldsymbol{e_{n,1}}}{\frac{1+\sum_{n=1}^{\infty}\exp\left(e_{n,0}+\boldsymbol{e_{n,1}\bar{Y}}\right)}{\sum_{n=1}^{\infty}\exp\left(e_{n,0}+\boldsymbol{e_{n,1}\bar{Y}}\right)}}\boldsymbol{Y_t} - \frac{\sum_{n=1}^{\infty}\exp\left(e_{n,0}+\boldsymbol{e_{n,1}\bar{Y}}\right)\boldsymbol{e_{n,1}}}{\sum_{n=1}^{\infty}\exp\left(e_{n,0}+\boldsymbol{e_{n,1}\bar{Y}}\right)}\boldsymbol{Y_{t-1}} \\
&= \xi_0^{eq} + \boldsymbol{\xi_1^{eq}Y_{t-1}} + \boldsymbol{\xi_2^{eq}}\begin{bmatrix} g_t - E_{t-1}(g_t) \\ RA_t - E_{t-1}(RA_t) \\ UC_t - E_{t-1}(UC_t) \\ MP_t \end{bmatrix},
\end{aligned}
\tag{A21}
$$

where ξ_0^{eq}, $\boldsymbol{\xi_1^{eq}}$, and $\boldsymbol{\xi_2^{eq}}$ are implicitly defined.

A.4.3 Equity risk premium

Given the no-arbitrage condition and that log stock return is quasi-linear and multinormal shock assumptions, the equity risk premium can be solved as follows:

$$
\begin{aligned}
E_t\left(r_{t+1}^{eq} - rf_t\right) + \frac{1}{2}V_t(r_{t+1}^{eq}) &\approx -Cov_t(m_{t+1}, r_{t+1}^{eq}) \\
&= \underbrace{\left(-m_{2,c}\xi_{2,c}^{eq}\right)}_{\kappa_{UC}} UC_t + \underbrace{\left(-m_{2,c}\xi_{2,RA}^{eq}\right)}_{\kappa_{RA}} RA_t,
\end{aligned}
\tag{A22}
$$

where $\xi_{2,c}^{eq}$ indicates the loading of r_{t+1}^{eq} on $\sqrt{UC_t}\varepsilon_{c,t+1}$ (which comes from dividend growth's exposure to consumption shock and the expected growth's exposure to consumption shock), and $\xi_{2,RA}^{eq}$ indicates the loading of r_{t+1}^{eq} on $\sqrt{RA_t}\varepsilon_{RA,t+1}$ (which comes from risk aversion). This equation proves Equation (3) in Section 2.

B Appendix: Tables and Figures

Table B1: Local model summary.

This table summarizes the 5 linear models and 15 non-linear models considered in estimating local physical variances (PVAR) at the daily frequency. "L" means that the coefficient of the predictor (see columns) is a constant; that is, local projected physical variance is linear to the predictors. "NL" means that the relationship between the projected physical variance and the predictor is non-linear. $RV_t^{(22)}$, $RV_t^{(5)}$, and RV_t denote the past cumulative 22-day, 5-day, and 1-day squared realized returns, respectively; IV_t denotes the 22-day implied variance. Other details of the models are listed in Table B2 and Table B3.

Linear model specifications					
	$RV_t^{(22)}$	$RV_t^{(5)}$	RV_t	IV_t	# of params.
L Model 1	1 (fix)				0
L Model 2	L				2
L Model 3	L	L	L		4
L Model 4	L			L	3
L Model 5	L	L	L	L	5
Non-linear model specifications					
	$RV_t^{(22)}$	$RV_t^{(5)}$	RV_t	IV_t	# of params.
NL Model 1	NL	NL	NL	NL	9
NL Model 2	L	NL	NL	NL	8
NL Model 3	NL	L	NL	NL	8
NL Model 4	NL	NL	L	NL	8
NL Model 5	NL	NL	NL	L	8
NL Model 6	NL	NL	L	L	7
NL Model 7	NL	L	NL	L	7
NL Model 8	NL	L	L	NL	7
NL Model 9	L	NL	NL	L	7
NL Model 10	L	NL	L	NL	7
NL Model 11	L	L	NL	NL	7
NL Model 12	NL	L	L	L	6
NL Model 13	L	NL	L	L	6
NL Model 14	L	L	NL	L	6
NL Model 15	L	L	L	NL	6

Table B2: Model selection: linear models.

This table reports model selection criteria across 5 forecasting models:

Model 1 $E_t\left[RV_{t+22}^{(22)}\right] = RV_t^{(22)}$

Model 2 $E_t\left[RV_{t+22}^{(22)}\right] = \widehat{\alpha} + \widehat{\beta^m} RV_t^{(22)}$

Model 3 $E_t\left[RV_{t+22}^{(22)}\right] = \widehat{\alpha} + \widehat{\beta^m} RV_t^{(22)} + \widehat{\beta^w} RV_t^{(5)} + \widehat{\beta^d} RV_t$

Model 4 $E_t\left[RV_{t+22}^{(22)}\right] = \widehat{\alpha} + \widehat{\beta^m} RV_t^{(22)} + \widehat{\gamma} IV_t$

Model 5 $E_t\left[RV_{t+22}^{(22)}\right] = \widehat{\alpha} + \widehat{\beta^m} RV_t^{(22)} + \widehat{\beta^w} RV_t^{(5)} + \widehat{\beta^d} RV_t + \widehat{\gamma} IV_t.$

Panels report AIC, BIC, MSE, R2, and in-sample unconditional correlation between model-implied risk aversion (RA) and uncertainty (UC). Bold numbers indicates the best model given the corresponding criterion.

	US	EA	JP
	AIC ($\times 10^4$):		
L Model 1	-8.514	-5.441	-7.914
L Model 2	-8.622	-5.549	-8.090
L Model 3	-8.729	-5.605	-8.117
L Model 4	-8.662	-5.656	-8.097
L Model 5	**-8.736**	**-5.667**	**-8.121**
	BIC ($\times 10^4$):		
L Model 1	-8.514	-5.441	-7.914
L Model 2	-8.621	-5.548	-8.089
L Model 3	-8.726	-5.603	-8.115
L Model 4	-8.660	-5.654	-8.095
L Model 5	**-8.733**	**-5.664**	**-8.117**
	MSE ($\times 10^{-4}$):		
L Model 1	0.065	0.162	0.120
L Model 2	0.056	0.130	0.093
L Model 3	0.048	0.116	0.090
L Model 4	0.053	0.104	0.093
L Model 5	**0.048**	**0.102**	**0.089**
	R2:		
L Model 1	0.433	0.219	0.109
L Model 2	0.514	0.373	0.307
L Model 3	0.581	0.441	0.334
L Model 4	0.540	0.496	0.314
L Model 5	0.586	0.507	0.338
	UC-RA Corr:		
L Model 1	-0.007	0.000	-0.098
L Model 2	0.480	0.563	0.333
L Model 3	0.444	0.577	0.318
L Model 4	0.784	1.000	0.494
L Model 5	0.559	0.928	0.429

Table B3: Model selection: non-linear models

The full local model with non-linear predictive coefficients (NL Model 1) is as follows:

$$E_t\left[RV_{t+22}^{(22)}\right] = \widehat{\alpha} + \frac{\exp(\widehat{\beta_0^m} - \widehat{\beta_1^m} Z_t^m)}{\exp(\widehat{\beta_0^m} - \widehat{\beta_1^m} Z_t^m) + 1} RV_t^{(22)} + \frac{\exp(\widehat{\beta_0^w} - \widehat{\beta_1^w} Z_t^w)}{\exp(\widehat{\beta_0^w} - \widehat{\beta_1^w} Z_t^w) + 1} RV_t^{(5)}$$
$$+ \frac{\exp(\widehat{\beta_0^d} - \widehat{\beta_1^d} Z_t^d)}{\exp(\widehat{\beta_0^d} - \widehat{\beta_1^d} Z_t^d) + 1} RV_t + \frac{\exp(\widehat{\gamma}_0 - \widehat{\gamma}_1 Z_t^{iv})}{\exp(\widehat{\gamma}_0 - \widehat{\gamma}_1 Z_t^{iv}) + 1} IV_t, \qquad \text{(A23)}$$

where instrument Z_t^m is standardized $RV_t^{(22)}$; Z_t^w, standardized $RV_t^{(5)}$; Z_t^d, standardized RV_t; Z_t^{iv}, standardized IV_t. The estimation is conducted by finding the minimum least squares:

$$\text{argmin}_{\alpha,\boldsymbol{\beta},\boldsymbol{\gamma}(\gamma)} ||RV_{t+22}^{(22)} - E_t\left[RV_{t+22}^{(22)}\right]||^2.$$

Other non-linear models allow 1∼3 predictors to have non-linear coefficients and the rest with linear coefficients as before. For each country, all 15 non-linear models are estimated using daily observations and standard model specifications (AIC, BIC, MSE) are obtained.

AIC, BM(5) ($\times 10^4$)	-8.736	-5.667	-8.121
BIC, BM(5) ($\times 10^4$)	-8.733	-5.664	-8.117
MSE, BM(5) ($\times 10^{-4}$)	0.048	0.102	0.089
	US	EA	JP
AIC ($\times 10^4$):			
NL Model 1	**-8.755**	-5.675	**-8.172**
NL Model 2	-8.755	-5.676	-7.775
NL Model 3	-8.753	-5.675	-8.148
NL Model 4	-8.752	-5.673	-8.169
NL Model 5	-8.753	-5.674	-8.171
NL Model 6	-8.749	-5.670	-8.169
NL Model 7	-8.750	-5.674	-8.158
NL Model 8	-8.746	-5.673	-7.806
NL Model 9	-8.753	-5.675	-8.166
NL Model 10	-8.752	-5.674	-8.163
NL Model 11	-8.753	**-5.676**	-8.119
NL Model 12	-8.741	-5.668	-8.149
NL Model 13	-8.750	-5.672	-8.163
NL Model 14	-8.749	-5.675	-8.137
NL Model 15	-8.746	-5.673	-8.126
BIC ($\times 10^4$):			
NL Model 1	-8.749	-5.669	-8.165
NL Model 2	**-8.749**	-5.671	-7.769
NL Model 3	-8.747	-5.670	-8.142
NL Model 4	-8.747	-5.667	-8.164
NL Model 5	-8.747	-5.669	**-8.166**
NL Model 6	-8.745	-5.666	-8.164
NL Model 7	-8.745	-5.669	-8.154
NL Model 8	-8.741	-5.669	-7.801
NL Model 9	-8.748	-5.670	-8.161
NL Model 10	-8.747	-5.669	-8.158
NL Model 11	-8.748	**-5.671**	-8.115
NL Model 12	-8.737	-5.665	-8.145
NL Model 13	-8.745	-5.668	-8.159
NL Model 14	-8.745	-5.671	-8.133
NL Model 15	-8.742	-5.670	-8.122
MSE ($\times 10^{-4}$):			
NL Model 1	**0.046**	0.100	**0.083**
NL Model 2	0.046	**0.100**	0.146
NL Model 3	0.046	0.100	0.086
NL Model 4	0.046	0.101	0.083

NL Model 5	0.046	0.101	0.083
NL Model 6	0.047	0.101	0.083
NL Model 7	0.047	0.101	0.085
NL Model 8	0.047	0.101	0.140
NL Model 9	0.046	0.100	0.084
NL Model 10	0.046	0.101	0.084
NL Model 11	0.046	0.100	0.089
NL Model 12	0.047	0.102	0.086
NL Model 13	0.047	0.101	0.084
NL Model 14	0.047	0.100	0.087
NL Model 15	0.047	0.101	0.089

Table B4: The time-zone factor model: parameter choices

The only constraint is $\beta_{JP}+\beta_{EA}+\beta_{US}=1$. Bold values indicate that a coefficient is significant; ***, at the 1% significance level; **, 5%; *, 10%.

(1)	(2)	(3)
β_{JP}	β_{EA}	β_{US}
0.3076***	0.2818***	0.4102***
(0.0528)	(0.0505)	(0.0889)
(4)	(5)	(6)
σ_{JP}	σ_{EA}	σ_{US}
7.2072*	13.5048***	11.0914***
(3.6839)	(1.6874)	(4.3840)
(7)	(8)	(9)
s_{JP}	s_{EA}	s_{US}
9.8965***	4.8095***	8.5575***
(0.9551)	(0.8368)	(1.2337)

Table B5: Moment matching results of the global factor model

	Moment	Empirical	(Boot.SE)	Model
(1)	$Var(\overline{ra_t^{JP}})$	140.90	(21.00)	**131.76**
(2)	$Cov(\overline{ra_t^{JP}}, \overline{ra_t^{EA}})$	24.23	(11.86)	**15.17**
(3)	$Cov(\overline{ra_t^{JP}}, \overline{ra_t^{US}})$	1.52	(8.68)	**6.56**
(4)	$Var(\overline{ra_t^{EA}})$	60.85	(13.60)	**51.51**
(5)	$Cov(\overline{ra_t^{EA}}, \overline{ra_t^{US}})$	26.47	(7.66)	**27.09**
(6)	$Var(\overline{ra_t^{US}})$	136.65	(19.37)	**133.36**
(7)	$Cov(\overline{ra_t^{JP}}, \overline{ra_{t-1}^{EA}})$	11.54	(8.05)	**15.81**
(8)	$Cov(\overline{ra_t^{JP}}, \overline{ra_{t-1}^{US}})$	43.96	(11.45)	**38.54**
(9)	$Cov(\overline{ra_t^{EA}}, \overline{ra_{t-1}^{JP}})$	-5.77	(6.01)	**0.00**
(10)	$Cov(\overline{ra_t^{EA}}, \overline{ra_{t-1}^{US}})$	18.72	(10.73)	**14.22**
(11)	$Cov(\overline{ra_t^{US}}, \overline{ra_{t-1}^{JP}})$	-13.20	(11.76)	**0.00**
(12)	$Cov(\overline{ra_t^{US}}, \overline{ra_{t-1}^{EA}})$	-3.01	(7.26)	**0.00**

Table B6: Correlations among the cleansed country risk aversion shocks, $\overline{ra_t^i}$; global risk aversion shocks realized at different time zones, $\widehat{gra_t^i}$; cumulative global risk aversion at different time zones, $\widehat{GRA_t^i}$; local risk aversion shock, $\widehat{u_t^i}$; cumulative local risk aversion, $\widehat{U_t^i}$. See detailed definitions of these variables in Section 6.

Panel A. country-level variable correlations

(1) US perspective:

	$\overline{ra_t^{US}}$	$\overline{ra_t^{EA}}$	$\overline{ra_t^{JP}}$		$\widehat{u_t^{US}}$	$\widehat{u_t^{EA}}$	$\widehat{u_t^{JP}}$		$\widehat{U_t^{US}}$	$\widehat{U_t^{EA}}$	$\widehat{U_t^{JP}}$
$\overline{ra_t^{US}}$	1.000	0.290	0.011	$\widehat{u_t^{US}}$	1.000	0.176	0.002	$\widehat{U_t^{US}}$	1.000	0.403	0.562
$\overline{ra_t^{EA}}$		1.000	0.262	$\widehat{u_t^{EA}}$		1.000	-0.270	$\widehat{U_t^{EA}}$		1.000	0.715
$\overline{ra_t^{JP}}$			1.000	$\widehat{u_t^{JP}}$			1.000	$\widehat{U_t^{JP}}$			1.000

(2) EA perspective:

	$\overline{ra_{t-1}^{US}}$	$\overline{ra_t^{EA}}$	$\overline{ra_t^{JP}}$		$\widehat{u_{t-1}^{US}}$	$\widehat{u_t^{EA}}$	$\widehat{u_t^{JP}}$		$\widehat{U_{t-1}^{US}}$	$\widehat{U_t^{EA}}$	$\widehat{U_t^{JP}}$
$\overline{ra_{t-1}^{US}}$	1.000	0.205	0.317	$\widehat{u_{t-1}^{US}}$	1.000	0.119	0.054	$\widehat{U_{t-1}^{US}}$	1.000	0.406	0.562
$\overline{ra_t^{EA}}$		1.000	0.262	$\widehat{u_t^{EA}}$		1.000	0.176	$\widehat{U_t^{EA}}$		1.000	0.715
$\overline{ra_t^{JP}}$			1.000	$\widehat{u_t^{JP}}$			1.000	$\widehat{U_t^{JP}}$			1.000

(3) JP perspective:

	$\overline{ra_{t-1}^{US}}$	$\overline{ra_{t-1}^{EA}}$	$\overline{ra_t^{JP}}$		$\widehat{u_{t-1}^{US}}$	$\widehat{u_{t-1}^{EA}}$	$\widehat{u_t^{JP}}$		$\widehat{U_{t-1}^{US}}$	$\widehat{U_{t-1}^{EA}}$	$\widehat{U_t^{JP}}$
$\overline{ra_{t-1}^{US}}$	1.000	0.290	0.317	$\widehat{u_{t-1}^{US}}$	1.000	-0.270	0.054	$\widehat{U_{t-1}^{US}}$	1.000	0.403	0.562
$\overline{ra_{t-1}^{EA}}$		1.000	0.125	$\widehat{u_{t-1}^{EA}}$		1.000	-0.073	$\widehat{U_{t-1}^{EA}}$		1.000	0.712
$\overline{ra_t^{JP}}$			1.000	$\widehat{u_t^{JP}}$			1.000	$\widehat{U_t^{JP}}$			1.000

Panel B. global variable correlations

	$\widehat{gra_t^{US}}$	$\widehat{gra_t^{EA}}$	$\widehat{gra_t^{JP}}$	SUM
$\widehat{gra_t^{US}}$	1.000	0.463	0.397	0.705
$\widehat{gra_t^{EA}}$		1.000	0.853	0.949
$\widehat{gra_t^{JP}}$			1.000	0.863
SUM				1.000

	$\widehat{GRA_t^{US}}$	$\widehat{GRA_t^{EA}}$	$\widehat{GRA_t^{JP}}$	SUM
$\widehat{GRA_t^{US}}$	1.000	0.710	0.732	0.789
$\widehat{GRA_t^{EA}}$		1.000	0.982	0.992
$\widehat{GRA_t^{JP}}$			1.000	0.988
SUM				1.000

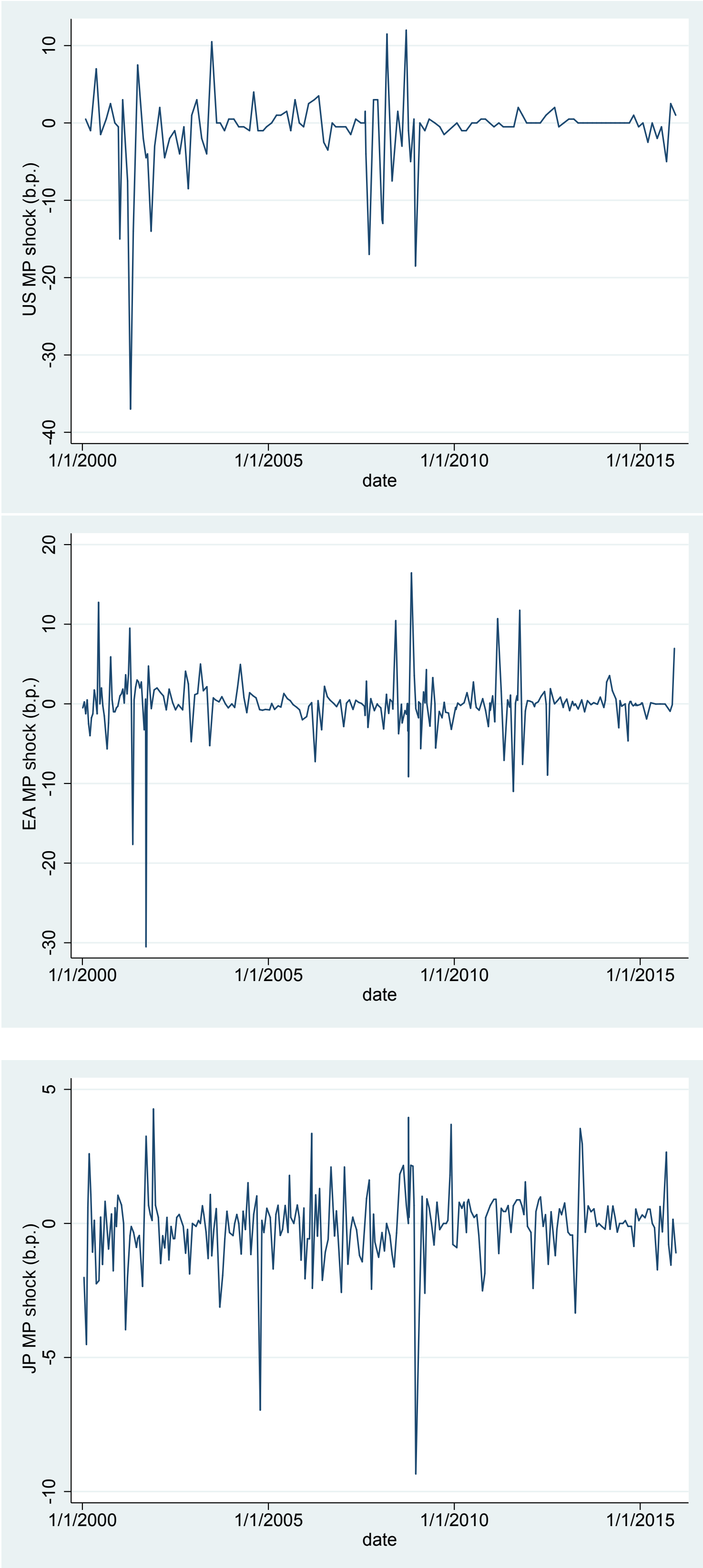

Figure B1: Top to bottom: Time series of US, EA, and JP monetary policy shocks in basis points.

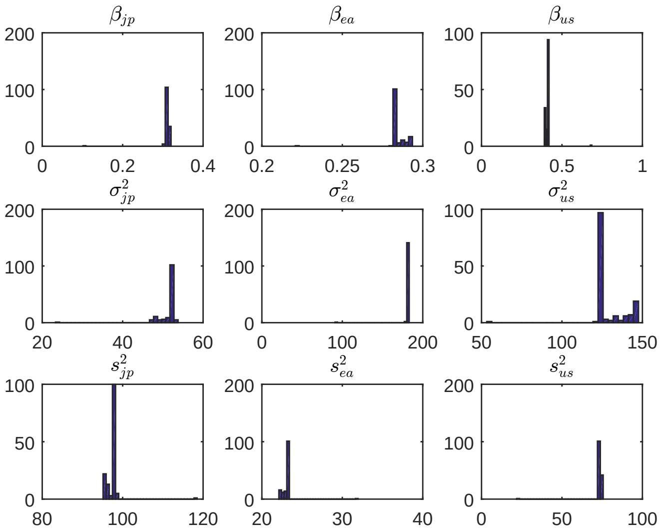

Figure B2: Histogram of final parameter estimates after applying the standard GMM procedure.

Notes: (1) We used 250 sensible initial values, e.g. β should, intuitively, be between 0 and 1, and s^2 might be 0–60% of the total variance of $\overline{ra_t}$ in the data. (2) As one can see, s^2 parameters are tightly estimated. (3) The chosen parameter set in Table B4 is identified from the mode-range β values; subsequently, σ^2 can be identified.

Online Appendix – NOT FOR PUBLICATION

Table OA1: Summary statistics for all monetary policy shocks.

This table reports summary statistics for our alternative MP shock measures used in this Online Appendix: for the US, 3M shock based on current-month (three-month) Federal Funds futures (from Gürkaynak, Sack, and Swanson (2005)); for the EA, 3M shock based on one-month (three-month) Eonia swap rates (from Jarociński and Karadi (2020)). The second panel contains statistics for the post-2008 subsample: for the US, RSW measures for the target rate, forward guidance (FG) and asset price (AP) shocks (from Rogers, Scotti, and Wright (2018)); for the EA and Japan, the measures are as in the baseline (first panel), only restricted to the post-2008 subsample. Shocks are measured in basis points.

	Mean	SD	5%	95%	N
US 3M shock	-0.975	5.344	-13.000	3.999	138
EA 3M shock	-0.054	3.575	-4.650	3.750	263
JP RSW shock	-0.116	1.459	-2.420	2.104	247
RSW US target rate shock	-0.393	1.971	-2.066	0.834	67
RSW FG shock	-1.576	6.493	-13.310	5.234	67
RSW AP shock	0.300	5.991	-7.935	8.030	67
EA 3M shock	-0.107	2.694	-5.550	3.550	118
JP RSW shock	0.018	1.459	-2.424	1.803	105

Table OA2: Summary statistics for dependent variables.

This table reports summary statistics for the dependent variables in the regressions with monetary policy shocks. Sample period is January 3, 2000 - December 31, 2015 (daily data). Risk aversion (RA) and uncertainty (UC) are expressed in monthly percentages squared, with statistics referring to the first-differences in those variables. Three-month (3M) and 10-year (10Y) interest rates are expressed in basis points, with statistics referring to the first-differences. All the other variables are expressed in percent (log first-differences multiplied by 100). GSCI is the Goldman Sachs commodity price index while Gold (residuals) is the gold price cleansed of the commodity price component (residuals from the regressions of log changes in the gold price on log changes in GSCI).

	Mean	SD	5%	95%	N
US RA (1st diff)	-0.003	11.739	-10.444	10.678	4014
EA RA (1st diff)	-0.001	8.330	-8.251	7.709	4075
JP RA (1st diff)	-0.002	12.019	-10.375	10.866	3910
US UC (1st diff)	-0.002	7.404	-7.679	7.709	4014
EA UC (1st diff)	-0.008	7.693	-9.473	10.281	4075
JP UC (1st diff)	0.001	3.295	-3.566	4.062	3910
US 3M rate (1st diff)	-0.128	5.155	-5.000	5.000	4162
DE 3M rate (1st diff)	-0.095	5.014	-6.300	6.300	4162
EA 3M rate (1st diff)	-0.087	4.746	-5.800	5.500	4162
JP 10Y rate (1st diff)	-0.033	2.647	-4.000	4.000	4162
stock ret US (log diff)	0.008	1.267	-1.973	1.812	4025
stock ret EA (log diff)	-0.010	1.521	-2.461	2.289	4097
stock ret JP (log diff)	0.000	1.545	-2.407	2.264	3929
bond ret US (log diff)	0.022	0.484	-0.768	0.756	4174
bond ret EA (log diff)	0.024	0.292	-0.457	0.472	4174
bond ret JP (log diff)	0.011	0.231	-0.353	0.360	4174
USD per EUR (log diff)	0.001	0.642	-1.075	1.050	4162
USD per JPY (log diff)	-0.004	0.637	-0.981	0.996	4162
EUR per USD (log diff)	-0.002	0.627	-1.017	1.033	4162
EUR per JPY (log diff)	-0.005	0.762	-1.168	1.223	4162
JPY per USD (log diff)	0.004	0.624	-1.013	1.001	4162
JPY per EUR (log diff)	0.006	0.781	-1.270	1.193	4162
Gold (residuals)	0.000	1.080	-1.735	1.612	4161
GSCI (log diff)	-0.006	1.478	-2.471	2.317	4173

Table OA3: Monetary Policy, Risk Aversion, and Uncertainty

We consider MP 3-month rate shocks and post-2008 subsample shocks (US: MP, FG and AP shocks are from Rogers, Scotti, and Wright (2018)). Other details are in Table 6. All regressions control for monetary policy and macro event dummies and macro news, as described in Section 2. Bold values indicate that a coefficient is significant; ***, at the 1% significance level; **, 5%; *, 10%.

	US, direct	US → EA	US →JP	EA, direct	EA→US	EA→JP	JP, direct	JP→US	JP→EA
Panel A. Risk Aversion									
MP 3-month rate shocks	0.024	0.005	0.085	-0.518	-0.232	0.558	0.016	-0.902	0.033
	(0.290)	(0.132)	(0.195)	(0.395)	(0.664)	(0.586)	(1.005)	(0.776)	(0.468)
non-MP RA shocks		**0.094***	**0.310*****		**0.478*****	0.113		-0.071	**0.157****
		(0.057)	(0.078)		(0.149)	(0.125)		(0.060)	(0.065)
MP shocks post-2008	-2.680	0.661	-0.848	**-0.802*****	-0.129	-0.764	1.326	**-1.033****	0.151
	(2.039)	(0.662)	(0.733)	(0.283)	(0.505)	(0.664)	(0.874)	(0.497)	(0.723)
US FG shocks post-2008	**1.701****	0.234	**0.455****						
	(0.779)	(0.283)	(0.219)						
US AP shocks post-2008	-0.139	-0.205	-0.090						
	(0.550)	(0.165)	(0.212)						
non-MP RA shocks		-0.003	**0.208*****		**0.767*****	**0.393*****		-0.031	**0.064****
		(0.037)	(0.066)		(0.110)	(0.107)		(0.045)	(0.029)
Panel B. Uncertainty									
MP 3-month rate shocks	-0.049	-0.029	-0.042	0.339	**-0.564*****	-0.013	-0.166	0.391	0.460
	(0.203)	(0.111)	(0.034)	(0.350)	(0.202)	(0.062)	(0.314)	(0.382)	(0.547)
non-MP UC shocks		0.071	**0.078*****		**0.497*****	**0.054*****		0.100	**0.571*****
		(0.065)	(0.023)		(0.053)	(0.014)		(0.083)	(0.117)
MP shocks post-2008	-0.260	-0.944	-0.088	-0.181	**-0.421***	-0.109	-0.230	0.468	0.716
	(0.531)	(0.595)	(0.232)	(0.477)	(0.232)	(0.130)	(0.609)	(0.562)	(0.552)
US FG shocks post-2008	-0.109	**0.431****	0.106						
	(0.220)	(0.202)	(0.086)						
US AP shocks post-2008	**-0.305***	**-0.320****	0.054						
	(0.175)	(0.158)	(0.126)						
non-MP UC shocks		0.017	**0.114*****		**0.527*****	**0.036****		0.051	**0.410*****
		(0.098)	(0.031)		(0.101)	(0.017)		(0.113)	(0.093)

Table OA4: Interest rates: MP vs non-MP-driven RA and UC

We consider MP 3-month rate shocks and post-2008 subsample shocks (US: MP, FG and AP shocks are from Rogers, Scotti, and Wright (2018)). Other details are in Table 6 and Table 7. All regressions control for monetary policy and macro event dummies and macro news, as described in Section 2. Bold values indicate that a coefficient is significant; ***, at the 1% significance level; **, 5%; *, 10%.

	US, direct	US → EA	US→DE	US →JP	EA, direct	DE, direct	EA →US	EA →JP	JP, direct	JP→US	JP→EA	JP→DE
MP 3-month rate shocks	**0.603*** **	**0.198*** **	**0.210** **	0.035	0.331	**0.826*** **	0.067	0.073	**0.642*** **	**-0.393* **	0.116	**0.454* **
	(0.175)	(0.060)	(0.086)	(0.044)	(0.257)	(0.199)	(0.069)	(0.057)	(0.144)	(0.215)	(0.351)	(0.233)
non-MP RA shocks	-0.032	-0.027	**-0.067** **	-0.001	-0.019	0.003	-0.040	**-0.013** **	**-0.026*** **	**0.028* **	0.003	-0.003
	(0.021)	(0.023)	(0.027)	(0.006)	(0.015)	(0.034)	(0.038)	(0.006)	(0.006)	(0.017)	(0.014)	(0.014)
non-MP UC shocks	**-0.081** **	**-0.076** **	-0.062	-0.004	-0.001	-0.016	-0.012	0.001	-0.034	**-0.111*** **	-0.071	-0.008
	(0.035)	(0.035)	(0.040)	(0.009)	(0.018)	(0.017)	(0.022)	(0.007)	(0.022)	(0.041)	(0.049)	(0.032)
MP shocks post-2008	0.048	0.390	0.023	**0.387*** **	-0.385	**1.583*** **	0.031	**0.121* **	**0.637*** **	-0.110	**1.052*** **	0.324
	(0.268)	(0.358)	(0.336)	(0.135)	(0.444)	(0.356)	(0.028)	(0.066)	(0.175)	(0.080)	(0.351)	(0.250)
US FG shocks post-2008	0.057	0.040	-0.076	0.043								
	(0.067)	(0.142)	(0.125)	(0.050)								
US AP shocks post-2008	0.045	0.077	**0.137* **	**0.094* **								
	(0.049)	(0.084)	(0.082)	(0.057)								
non-MP RA shocks	**0.012** **	-0.004	0.023	-0.006	-0.010	-0.026	-0.007	**-0.020* **	**-0.020** **	**0.006* **	-0.019	0.004
	(0.006)	(0.017)	(0.016)	(0.009)	(0.024)	(0.023)	(0.007)	(0.011)	(0.008)	(0.004)	(0.013)	(0.010)
non-MP UC shocks	0.008	-0.036	0.027	0.017	-0.007	-0.001	0.000	**-0.017** **	-0.026	-0.002	0.011	-0.039
	(0.010)	(0.026)	(0.024)	(0.011)	(0.023)	(0.021)	(0.006)	(0.009)	(0.026)	(0.014)	(0.032)	(0.027)

Table OA5: Stock and bond returns: MP vs non-MP-driven RA and UC

We consider MP 3-month rate shocks and post-2008 subsample shocks (US: MP, FG and AP shocks are from Rogers, Scotti, and Wright (2018)). Other details are in Table 8. All regressions control for monetary policy and macro event dummies and macro news, as described in Section 2. Bold values indicate that a coefficient is significant; ***, at the 1% significance level; **, 5%; *, 10%.

	US, direct	US → EA	US→JP	EA, direct	EA→US	EA→JP	JP, direct	JP→US	JP→EA
Panel A. Stock Returns									
MP 3-month rate shocks	**-0.057****	-0.003	-0.023	0.014	**0.063*****	-0.002	0.058	0.044	-0.013
	(0.023)	(0.016)	(0.025)	(0.023)	(0.017)	(0.019)	(0.097)	(0.045)	(0.076)
non-MP RA shocks	**-0.082*****	0.005	**-0.013****	**-0.058*****	**0.009***	**-0.019*****	**-0.057*****	**0.005****	**0.010****
	(0.005)	(0.005)	(0.005)	(0.013)	(0.005)	(0.007)	(0.005)	(0.002)	(0.005)
non-MP UC shocks	**-0.077*****	0.007	**-0.011***	**-0.111*****	**-0.018*****	-0.008	**-0.077*****	0.002	0.010
	(0.007)	(0.006)	(0.007)	(0.007)	(0.004)	(0.005)	(0.013)	(0.008)	(0.011)
MP shocks post-2008	**0.215****	**0.184***	**0.315*****	0.090	**0.104*****	0.039	0.099	0.093	**0.188***
	(0.105)	(0.097)	(0.102)	(0.062)	(0.034)	(0.030)	(0.129)	(0.071)	(0.102)
US FG shocks post-2008	**-0.158*****	-0.028	-0.042						
	(0.040)	(0.039)	(0.039)						
US AP shocks post-2008	0.041	0.023	0.040						
	(0.031)	(0.029)	(0.029)						
non-MP RA shocks	**-0.082*****	0.006	**-0.019*****	**-0.094*****	-0.006	-0.007	**-0.053*****	0.003	0.005
	(0.005)	(0.004)	(0.005)	(0.010)	(0.007)	(0.007)	(0.007)	(0.002)	(0.003)
non-MP UC shocks	**-0.074*****	0.006	0.001	**-0.123*****	**-0.020*****	**-0.025*****	**-0.066*****	-0.005	0.013
	(0.007)	(0.007)	(0.007)	(0.007)	(0.006)	(0.006)	(0.015)	(0.010)	(0.009)
Panel B. Bond Returns									
MP 3-month rate shocks	**-0.027*****	-0.015	-0.004	**-0.026****	-0.009	-0.006	**-0.054*****	**-0.071****	-0.009
	(0.010)	(0.011)	(0.004)	(0.010)	(0.013)	(0.005)	(0.016)	(0.034)	(0.016)
non-MP RA shocks	**0.007*****	-0.001	0.000	**0.005*****	0.003	**0.001****	**0.002*****	**-0.002****	-0.001
	(0.002)	(0.001)	(0.001)	(0.002)	(0.002)	(0.001)	(0.001)	(0.001)	(0.001)
non-MP UC shocks	**0.009*****	**-0.003*****	0.000	**0.008*****	**0.010*****	0.000	0.003	**-0.005***	-0.004
	(0.002)	(0.001)	(0.001)	(0.001)	(0.002)	(0.001)	(0.002)	(0.003)	(0.002)
MP shocks post-2008	**-0.075****	**-0.068****	**-0.032*****	**-0.075*****	**-0.050*****	**-0.012****	**-0.057*****	**-0.065****	-0.009
	(0.031)	(0.029)	(0.010)	(0.015)	(0.017)	(0.006)	(0.017)	(0.030)	(0.017)
US FG shocks post-2008	**-0.045*****	-0.008	**-0.008****						
	(0.017)	(0.011)	(0.003)						
US AP shocks post-2008	**-0.098*****	**-0.042*****	**-0.010*****						
	(0.011)	(0.009)	(0.003)						
non-MP RA shocks	**0.009*****	-0.001	0.001	**0.014*****	**0.010*****	**0.002****	**0.002****	-0.001	-0.001
	(0.002)	(0.001)	(0.001)	(0.002)	(0.003)	(0.001)	(0.001)	(0.001)	(0.001)
non-MP UC shocks	**0.009****	**-0.005*****	-0.001	**0.008*****	**0.014*****	**0.002****	0.002	-0.001	-0.001
	(0.003)	(0.002)	(0.001)	(0.002)	(0.003)	(0.001)	(0.002)	(0.003)	(0.003)

Table OA6: Exchange rates, gold, commodity: MP vs non-MP-driven RA and UC; e.g. USD per EUR

We consider MP 3-month rate shocks and post-2008 subsample shocks (US: MP, FG and AP shocks are from Rogers, Scotti, and Wright (2018)). Other details are in Table 9. All regressions control for monetary policy and macro event dummies and macro news, as described in Section 2. Other details are in Table 6 and Table 8. Bold values indicate that a coefficient is significant; ***, at the 1% significance level; **, 5%; *, 10%.

Panel A. Exchange Rates

	US → USD per EUR	US → USD per JPY	EA → EUR per USD	EA → EUR per JPY	JP → JPY per USD	JP → JPY per EUR
MP 3-month rate shocks	**-0.039****	**-0.035*****	-0.027	-0.028	-0.042	0.085
	(0.017)	(0.012)	(0.016)	(0.017)	(0.036)	(0.090)
non-MP RA shocks	**-0.007*****	**0.016*****	0.003	**0.011*****	**-0.009*****	**-0.011*****
	(0.001)	(0.002)	(0.002)	(0.004)	(0.002)	(0.003)
non-MP UC shocks	-0.002	**0.016*****	0.002	**0.018*****	-0.006	0.000
	(0.003)	(0.003)	(0.002)	(0.003)	(0.005)	(0.006)
MP shocks post-2008	0.074	**-0.119****	**-0.119*****	-0.068	-0.060	**0.206***
	(0.050)	(0.050)	(0.035)	(0.066)	(0.058)	(0.116)
US FG shocks post-2008	**-0.110*****	**-0.037***				
	(0.020)	(0.020)				
US AP shocks post-2008	**-0.041*****	**-0.066*****				
	(0.014)	(0.015)				
non-MP RA shocks	**-0.011*****	**0.011*****	**0.016*****	**0.024*****	**-0.010*****	**-0.017*****
	(0.003)	(0.002)	(0.005)	(0.005)	(0.002)	(0.003)
non-MP UC shocks	-0.006	**0.014*****	**0.011*****	**0.032*****	-0.000	0.012
	(0.006)	(0.005)	(0.004)	(0.004)	(0.008)	(0.010)

Panel B. Gold and Commodity Prices

	US → Gold	EA → Gold	JP → Gold	US → GSCI	EA → GSCI	JP → GSCI
MP 3-month rate shocks	**-0.053****	**-0.030***	-0.016	0.009	0.031	-0.047
	(0.021)	(0.016)	(0.068)	(0.021)	(0.027)	(0.072)
non-MP RA shocks	0.004	-0.005	-0.002	**-0.020*****	**-0.012****	-0.001
	(0.005)	(0.006)	(0.003)	(0.004)	(0.005)	(0.004)
non-MP UC shocks	**0.015****	0.003	-0.008	**-0.030*****	**-0.026*****	-0.004
	(0.006)	(0.004)	(0.007)	(0.007)	(0.005)	(0.008)
MP shocks post-2008	0.157	-0.012	0.051	0.172	**0.130****	0.047
	(0.134)	(0.037)	(0.065)	(0.128)	(0.057)	(0.074)
US FG shocks post-2008	**-0.149*****			-0.007		
	(0.050)			(0.070)		
US AP shocks post-2008	-0.032			0.028		
	(0.038)			(0.040)		
non-MP RA shocks	0.008	0.002	0.004	**-0.022*****	**-0.032*****	-0.000
	(0.006)	(0.007)	(0.003)	(0.007)	(0.010)	(0.005)
non-MP UC shocks	0.011	-0.001	0.004	**-0.033*****	**-0.037*****	0.007
	(0.008)	(0.006)	(0.010)	(0.008)	(0.008)	(0.012)

Table OA7: Dynamic effects (1): interest rate

This table presents country-level regression results of changes in interest rate on domestic and foreign shocks (MP=pure monetary policy shock; CB=CB information shock; RA=non-MP-driven risk aversion shock; UC=non-MP-driven uncertainty shock). Row “0d (t-1,t)” uses the contemporaneous changes in interest rate; row “5d (t,t+5)” (“21d (t,t+21)”) uses the cumulative future 5 (21)-day changes in interest rate. The shock coefficients are actual (not expressed in standard deviation term), and hence numbers in row “0d (t-1,t)” can also be found in the main paper tables. Highlighted regions reflect domestic effects.

US IR	US	US	EA	EA	JP	US	EA	JP	US	EA	JP
	MP	CB	MP	CB	MP	RA	RA	RA	UC	UC	UC
0d (t-1,t)	**0.323*****	**0.634*****	-0.004	**0.091****	-0.344	-0.029	-0.043	**0.027***	**-0.075****	-0.011	**-0.109*****
	(0.109)	(0.168)	(0.033)	(0.046)	(0.213)	(0.022)	(0.038)	(0.016)	(0.033)	(0.022)	(0.040)
5d (t,t+5)	-0.103	0.024	**0.191***	-0.037	-0.556	-0.041	**0.138****	-0.048	-0.028	-0.005	-0.014
	(0.200)	(0.216)	(0.109)	(0.132)	(0.473)	(0.026)	(0.056)	(0.030)	(0.048)	(0.034)	(0.087)
21d (t,t+21)	-0.15	**0.76****	-0.071	0.258	-1.254	**-0.05****	**0.064***	-0.033	**-0.078****	0.057	-0.171
	(0.375)	(0.364)	(0.163)	(0.225)	(1.150)	(0.021)	(0.039)	(0.028)	(0.039)	(0.057)	(0.151)
EA IR	US	US	EA	EA	JP	US	EA	JP	US	EA	JP
	MP	CB	MP	CB	MP	RA	RA	RA	UC	UC	UC
0d (t-1,t)	**0.14*****	0.086	**0.216***	0.044	0.117	-0.026	-0.02	0.004	**-0.077****	-0.003	-0.070
	(0.044)	(0.053)	(0.115)	(0.158)	(0.349)	(0.023)	(0.015)	(0.014)	(0.035)	(0.018)	(0.049)
5d (t,t+5)	-0.072	0.121	**0.257****	**0.574*****	-0.246	-0.005	0.022	0.003	-0.053	**-0.122***	-0.007
	(0.099)	(0.129)	(0.111)	(0.186)	(0.326)	(0.028)	(0.051)	(0.016)	(0.055)	(0.065)	(0.051)
21d (t,t+21)	-0.317	0.448	-0.187	**0.594*****	-0.567	-0.013	0.005	0.002	-0.104	-0.079	0.101
	(0.605)	(0.442)	(0.222)	(0.211)	(1.004)	(0.026)	(0.041)	(0.033)	(0.113)	(0.055)	(0.073)
JP IR	US	US	EA	EA	JP	US	EA	JP	US	EA	JP
	MP	CB	MP	CB	MP	RA	RA	RA	UC	UC	UC
0d (t-1,t)	0.031	0.017	0.001	**0.071****	**0.628*****	-3.05E-04	**-0.013****	**-0.026*****	-0.004	0.001	-0.034
	(0.039)	(0.038)	(0.028)	(0.032)	(0.144)	(0.006)	(0.006)	(0.006)	(0.009)	(0.008)	(0.022)
5d (t,t+5)	0.197	0.046	-0.002	-0.057	0.132	-0.009	0.003	**0.012****	0.010	0.020	0.001
	(0.155)	(0.157)	(0.041)	(0.048)	(0.271)	(0.006)	(0.009)	(0.006)	(0.013)	(0.013)	(0.031)
21d (t,t+21)	0.231	0.162	-0.002	-0.226	-0.567	-0.014	0.004	0.008	0.013	0.021	**0.118****
	(0.157)	(0.183)	(0.087)	(0.208)	(0.450)	(0.009)	(0.014)	(0.009)	(0.021)	(0.019)	(0.051)

Table OA8: Dynamic effects (2): stock returns

This table presents country-level regression results of log stock returns on domestic and foreign shocks (MP=pure monetary policy shock; CB=CB information shock; RA=non-MP-driven risk aversion shock; UC=non-MP-driven uncertainty shock). Row "0d (t-1,t)" uses the contemporaneous stock returns; row "5d (t,t+5)" ("21d (t,t+21)") uses the cumulative future 5 (21)-day log stock returns. The shock coefficients are actual (not expressed in standard deviation term), and hence numbers in row "0d (t-1,t)" can also be found in the main paper tables. Highlighted regions reflect domestic effects.

US SR	US	US	EA	EA	JP	US	EA	JP	US	EA	JP
	MP	CB	MP	CB	MP	RA	RA	RA	UC	UC	UC
0d (t-1,t)	**-0.085*****	**0.039****	**-0.046*****	**0.116*****	0.042	**-0.081*****	**0.009***	**0.005****	**-0.076*****	**-0.017*****	0.002
	(0.017)	(0.019)	(0.011)	(0.011)	(0.045)	(0.005)	(0.005)	(0.002)	(0.007)	(0.004)	(0.008)
5d (t,t+5)	-0.012	-0.047	0.028	0.058	-0.086	**0.018*****	0.006	**0.012***	**0.021***	**0.016***	-0.004
	(0.037)	(0.042)	(0.034)	(0.036)	(0.118)	(0.007)	(0.007)	(0.006)	(0.011)	(0.010)	(0.017)
21d (t,t+21)	-0.11	0.024	0.051	0.027	0.231	**0.013***	0	0.008	0.01	0.025	-0.039
	(0.105)	(0.067)	(0.047)	(0.070)	(0.274)	(0.007)	(0.008)	(0.007)	(0.018)	(0.018)	(0.031)
EA SR	US	US	EA	EA	JP	US	EA	JP	US	EA	JP
	MP	CB	MP	CB	MP	RA	RA	RA	UC	UC	UC
0d (t-1,t)	**-0.025****	**0.061*****	**-0.095*****	**0.141*****	-0.03	0.005	**-0.057*****	**0.009****	0.008	**-0.109*****	0.008
	(0.011)	(0.016)	(0.013)	(0.012)	(0.082)	(0.005)	(0.013)	(0.004)	(0.005)	(0.007)	(0.010)
5d (t,t+5)	-0.022	-0.061	-0.011	**0.079***	-0.12	**0.011***	0.009	0.012	**0.027*****	**0.027****	-0.019
	(0.032)	(0.040)	(0.040)	(0.043)	(0.133)	(0.007)	(0.009)	(0.008)	(0.009)	(0.011)	(0.017)
21d (t,t+21)	-0.059	0.033	-0.062	0.079	0.141	**0.015****	0.013	-0.001	**0.032****	**0.032*****	-0.035
	(0.096)	(0.075)	(0.066)	(0.081)	(0.380)	(0.007)	(0.013)	(0.007)	(0.016)	(0.012)	(0.038)
JP SR	US	US	EA	EA	JP	US	EA	JP	US	EA	JP
	MP	CB	MP	CB	MP	RA	RA	RA	UC	UC	UC
0d (t-1,t)	-0.016	-0.029	**-0.057*****	**0.071*****	0.038	**-0.013****	**-0.019*****	**-0.056*****	**-0.012***	-0.007	**-0.076*****
	(0.018)	(0.049)	(0.010)	(0.012)	(0.096)	(0.005)	(0.006)	(0.005)	(0.007)	(0.005)	(0.012)
5d (t,t+5)	-0.042	0.029	-0.034	**0.089***	**-0.334***	0.002	0.008	**0.019*****	0.006	**0.023***	0.019
	(0.043)	(0.056)	(0.031)	(0.048)	(0.174)	(0.006)	(0.010)	(0.006)	(0.010)	(0.012)	(0.021)
21d (t,t+21)	-0.061	0.058	-0.077	0.067	0.165	0.012*	0.016	0.012	-0.006	0.009	-0.005
	(0.102)	(0.085)	(0.060)	(0.077)	(0.440)	(0.006)	(0.022)	(0.008)	(0.018)	(0.017)	(0.035)

Table OA9: Dynamic effects (3): bond returns

This table presents country-level regression results of log bond returns on domestic and foreign shocks (MP=pure monetary policy shock; CB=CB information shock; RA=non-MP-driven risk aversion shock; UC=non-MP-driven uncertainty shock). Row “0d (t-1,t)” uses the contemporaneous bond returns; row “5d (t,t+5)” (“21d (t,t+21)”) uses the cumulative future 5 (21)-day log bond returns. The shock coefficients are actual (not expressed in standard deviation term), and hence numbers in row “0d (t-1,t)” can also be found in the main paper tables. Highlighted regions reflect domestic effects.

US BR	US	US	EA	EA	JP	US	EA	JP	US	EA	JP
	MP	CB	MP	CB	MP	RA	RA	RA	UC	UC	UC
0d (t-1,t)	**-0.029*****	-1.64E-04	0.007	**-0.019*****	**-0.067****	**0.007*****	0.002	**-0.003****	**0.01*****	**0.01*****	**-0.005***
	(0.011)	(0.015)	(0.008)	(0.006)	(0.030)	(0.002)	(0.002)	(0.001)	(0.003)	(0.002)	(0.003)
5d (t,t+5)	-0.012	-0.004	1.64E-04	0.001	0.028	-0.003	0.002	1.01E-04	**-0.008***	-0.003	-0.006
	(0.019)	(0.018)	(0.011)	(0.012)	(0.039)	(0.002)	(0.003)	(0.002)	(0.004)	(0.003)	(0.006)
21d (t,t+21)	0.029	-0.033	0.02	0.014	**0.214***	0.001	0.006	0.001	-0.004	-0.003	**-0.034*****
	(0.032)	(0.050)	(0.033)	(0.022)	(0.125)	(0.002)	(0.004)	(0.003)	(0.005)	(0.005)	(0.011)
EA BR	US	US	EA	EA	JP	US	EA	JP	US	EA	JP
	MP	CB	MP	CB	MP	RA	RA	RA	UC	UC	UC
0d (t-1,t)	-0.014	-0.005	**-0.013****	**-0.009***	-0.01	-0.001	**0.005*****	-0.001	**-0.003*****	**0.008*****	-0.004
	(0.010)	(0.008)	(0.006)	(0.005)	(0.016)	(0.001)	(0.002)	(0.001)	(0.001)	(0.001)	(0.002)
5d (t,t+5)	-0.004	-0.002	-0.002	0.001	-0.019	-0.001	0.002	-3.30E-04	**-0.005*****	**-0.007*****	-0.004
	(0.007)	(0.008)	(0.008)	(0.008)	(0.031)	(0.001)	(0.002)	(0.001)	(0.002)	(0.002)	(0.004)
21d (t,t+21)	0.01	**-0.039****	-0.003	0.007	0.051	0	**0.005***	-0.001	-0.003	-0.005	-0.014
	(0.018)	(0.019)	(0.017)	(0.014)	(0.072)	(0.001)	(0.003)	(0.002)	(0.004)	(0.003)	(0.008)
JP BR	US	US	EA	EA	JP	US	EA	JP	US	EA	JP
	MP	CB	MP	CB	MP	RA	RA	RA	UC	UC	UC
0d (t-1,t)	-0.004	-0.001	2.45E-04	**-0.007****	**-0.053*****	4.83E-06	**0.001****	**0.002*****	2.61E-04	-5.96E-05	6.93E-06
	(0.004)	(0.004)	(0.003)	(0.003)	(0.016)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.002)
5d (t,t+5)	-0.017	-0.005	3.42E-04	0.006	-0.014	0.001	-1.74E-04	**-0.001****	-0.001	-0.002	-6.43E-05
	(0.012)	(0.013)	(0.003)	(0.004)	(0.024)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.003)
21d (t,t+21)	-0.022	-0.017	-2.60E-04	**0.02***	0.055	0.001*	-4.81E-04	-0.001	-0.001	-0.002	**-0.009****
	(0.014)	(0.017)	(0.008)	(0.011)	(0.040)	(0.001)	(0.001)	(0.001)	(0.002)	(0.002)	(0.005)