Policy Announcements in FX Markets*

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Abstract

A strategy that is short the US dollar and long the rest of the world earns large excess returns on days of scheduled meetings of the Federal Open Market Committee (FOMC). Moreover, the difference between announcement and non announcement returns becomes larger during periods of high uncertainty and bad economic conditions. To reconcile our findings, we develop a model of an international long-run risk economy in which asset prices respond to revisions of monetary policy. Monetary policy uncertainty commands a risk premium that is larger in weaker economic conditions. A calibrated version is consistent with the cross-sectional pattern of currency risk premia observed in the data.

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1 Introduction

In this paper, we document that returns on foreign exchange portfolios which are short the US dollar and long the rest of the world are mainly earned on days when a meeting of the Federal Open Market Committee (FOMC) is scheduled. The intuition for this finding can be illustrated by the May 22, 2013 testimony of Chairman Bernanke to the Congress, when he raised the possibility of tapering the Fed's quantitative easing (QE) program. QE has led the US dollar to be the funding currency in large scale currency trades with emerging markets as target currencies. The perspective of an imminent US monetary tightening triggered not only a 15 basis point rise of the yield on 10-year bonds but also an unwinding of these currency trades.

This example highlights the potential risk that materializes during days when the FOMC announces its future policy, a risk for which investors in foreign exchange markets seek ex-ante compensation. Thus, in equilibrium, the uncertainty about future monetary policy should command a risk premium or, in other words, expected foreign exchange returns should be higher upon announcements. In this paper, we study both theoretically and empirically how monetary policy announcements by the Federal Reserve affect currency returns.

Using a large panel data set, we find that portfolios of currencies with low interest rates earn an average daily return of 7.43 basis points (bps) during days when the Federal Reserve makes an announcement, compared to -1.01bps on non-announcement days. This difference becomes larger for high interest rate currencies, with a daily return of 14.44bps on announcement days compared to 2.14bps on non-announcement days, a 12.30bps difference which is strongly statistically significant (t-statistic of 2.63).

We also study whether announcement returns behave differently for the dollar and HML (carry) factors. The former is the average return of all interest rate sorted portfolios, whereas the latter is the return of a portfolio long high interest rate currencies and short low interest rate currencies. We find that the dollar strategy has average returns of 9.63bps on announcement days and statistically insignificant average returns on non-announcement days. The difference between the two, which is 9.12bps, has a t-statistic of 2.81. On the other hand, the HML factor, a US dollar neutral strategy, displays the opposite behavior, with highly significant returns on non-announcement days only.

One might suspect that these findings are driven by a few outliers, as some FOMC meetings are more anticipated than others. However, we show that our results neither depend on the choice of currencies nor on outliers. When we winsorize the data set, discarding the top and the bottom percentiles, results remain virtually unchanged. Moreover, the direction of the policy revision, i.e. whether interest rates are raised or reduced relative to the expected value, is immaterial for our findings, as we show that a policy surprise variable has no significant effect on announcement returns. In contrast, conditioning on the state of the economy has a sizable impact: when we control for the economic condition, we find that the wedge between announcement and non-announcement returns is larger in bad times. We also inspect the relationship between these returns and uncertainty, of which we consider four proxies: the VIX, implied volatility extracted from options on Treasury futures, an uncertainty measure from survey forecasts of the target Fed Funds rate, and the economic policy uncertainty index of Baker, Bloom, and Davis (2013). A regression analysis shows strong evidence of a positive relationship between announcement currency returns and all four uncertainty indices.

We explain our findings with an equilibrium model of an open economy. The model draws upon two strands of the literature: First, long-run risks in an international framework (Colacito and Croce (2011)) and second, the political uncertainty literature pioneered by Pástor and Veronesi (2012). Following the latter, we assume that the policy action of some authority—the central bank—affects expected output growth. As the precise impact of the action is not observed, agents learn about it in a Bayesian fashion from observing realized output growth and signals such as speeches by officials at the central bank. Policy revisions take place at regular intervals, consistent with the fact that FOMC announcement dates are made public in advance. An unspecified informational friction prevents agents from solving the monetary authority's optimization, so that there remains uncertainty about the future stance of monetary policy. Thus, unpredictable revisions of the policy and their impact on long-term growth determine shocks to expected output growth. These shocks are more volatile than the Bayesian updates following normal (i.e. non-announcement) news. We therefore distinguish between volatility driven by "announcement uncertainty" and "normal" volatility. Due to recursive preferences, both normal and announcement risks affecting long-run expectations are priced in equilibrium. Our main focus is on explaining the cross-sectional pattern of currency risk premia around announcements through countries' heterogeneous exposure to shocks in long-run risk. Countries which are more (less) exposed to "normal" uncertainty have smaller (larger) interest rate differentials vis-à-vis the home country and smaller (larger) currency risk premia.

The intuition is similar to Lustig, Roussanov, and Verdelhan (2011): when "normal" long-run risk volatility is larger in the foreign country, currency returns hedge negative shocks to expected domestic output growth and positive shocks to uncertainty (which are both considered bad news), thus commanding a smaller premium. When we condition on policy announcement days, the data suggest that currency risk premia are larger in magnitude and less heterogeneous across foreign/domestic interest rates differentials. We reconcile our model with this evidence by postulating that in countries with larger (smaller) exposure to announcement uncertainty, announcement shocks to long-run risk are negatively (positively) correlated with their home country's counterpart. This different interpretation of the monetary policy revisions across countries implies that currency returns triggered by announcements never hedge the corresponding shock to domestic long-run risk.

We then calibrate the model to match a number of empirical moments related to interest rates, exchange rates, and equity premia. We obtain modest pricing errors and, using calibrated parameters, we show that the cross-sectional pattern of model-implied currency risk premia resembles its empirical counterpart qualitatively and quantitatively, both on announcement and on non-announcement days.

Literature Review: Our paper is related to the literature that studies the effect of political uncertainty on asset prices. The paper most closely linked is Pástor and Veronesi (2013). In their model, a firm's expected growth rate is affected by the current government policy in an unobserved way. Both the government and investors learn about the impact in a Bayesian fashion by observing realized profitability. At pre-specified dates, the government decides whether to revise the policy, thus incurring political costs which are unknown to investors. Political uncertainty stems from the unpredictability of policy revisions due to informational frictions. The authors are mainly concerned with the asset pricing implications of this uncertainty for equity markets: price, tail, and variance risks related to political events. We, instead, focus on an international framework and the timing of announcement events.[1] We use a similar learning mechanism to endogenously obtain a long-run risk economy but we model the shock to expected growth following the policy announcement in reduced form. Moreover, we focus on the cross-section of currency risk premia, rather than equity returns. Kelly, Pástor, and Veronesi (2014) inspect the effects of political uncertainty on an international set of equity index options. Their findings confirm their theoretical predictions, namely that equity options spanning political events are significantly more expensive as they provide a protection against the risk of political events. Croce, Kung, Nguyen, and Schmid (2012) study the impact of tax uncertainty on asset prices when the representative agent features recursive preferences. In their model, fiscal policies resemble Taylor rules and they show that tax uncertainty is a first order concern to explain sizable risk premia.

The paper is further related to a large literature in international finance on currency and dollar risk premia. Using a reduced form model Lustig, Roussanov, and Verdelhan (2011) show that asymmetric exposure to a common or global factor is key to understanding the global currency trade premium. Lustig, Roussanov, and Verdelhan (2014) extend this model to explain excess returns on the "dollar carry trade", a strategy that compensates US investors for taking on aggregate risk during bad times, i.e. when the US price of risk is high. Our model is a long-run risk analogue to their reduced form approach, meant to price the political risk of FOMC announcements. Maggiori (2013) shows that the US dollar earns a safety premium against a basket of foreign currencies that is particularly high in times of global financial distress. Verdelhan (2013) argues that similarly to the carry factor, the dollar factor has a risk based explanation. He also

[1] Pástor and Veronesi (2013) specify a single time at which an announcement is made, whereas we consider a regular schedule of announcements. The reason is that FOMC announcement dates are well known in advance and occur at fairly regular intervals.

shows that both factors explain a large variation of individual exchange rate movements. Our paper is also related to Lustig and Verdelhan (2007) who show that low interest rate currencies provide US investors with a hedge against US aggregate consumption risk. In our model, because of long-run risk, low interest rate currencies provide a hedge not only to bad states of US consumption growth, but also to expected consumption growth, the volatility of consumption growth, and announcement uncertainty.

Our paper also draws upon the literature that studies the implications of long-run risks for international finance. Colacito and Croce (2011) show that global long-run risk shocks drive most of the variation in pricing kernels and that their model successfully addresses the Backus and Smith (1993) puzzle. Colacito and Croce (2013) propose a general equilibrium model with recursive preferences and highly correlated long-run components in output and show that their model is able to address both the failure of the UIP and the lack of correlation between consumption growth differentials and exchange rate movements while matching salient moments of asset prices in the US and the UK. Colacito (2009) extends the framework of Colacito and Croce (2011) by adding a second stochastic component into the consumption growth, which affects consumption growth in two countries asymmetrically. Among other things, he then shows that the model successfully explains the forward premium puzzle.

A related literature has documented sizable conditional responses of various asset classes to macroeconomic news announcements (Fleming and Remolona (1999), Andersen, Bollerslev, Diebold, and Vega (2003)).[2] More closely related to our paper, Jones, Lamont, and Lumsdaine (1998) study unconditional fixed income returns around macroeconomic releases (inflation and labor market), and Savor and Wilson (2013) find positive unconditional excess equity returns on days of inflation, labor market and FOMC releases. Lucca and Moench (2014) study S&P 500 index returns ahead of scheduled announcements and their results indicate that the unconditional announcement day returns are due to a pre-FOMC drift rather than returns earned at the announcement.

[2]A large empirical literature studies the impact of monetary policy announcements on second moments in foreign exchange markets. The main finding is that surprising policy actions, such as changes in interest rates or currency parities increase volatility and that more precise policy announcements usually lead to less volatility (see Neely (2011) for a survey of the literature).

Amengual and Xiu (2013) posit a non-affine term structure model which includes jumps to study the effect of FOMC announcements on variance swaps on the S&P 500 index. They find that downward jumps are mostly associated with the resolution of political uncertainty after an announcement day. Savor and Wilson (2014) find that systematic market risk prices risky assets well, including foreign exchange portfolios, on announcement days. Moreover, the authors find that a portfolio which is long high interest rate currencies and short low interest rate currencies has positive returns on announcement days and negative returns on non-announcement days. Our paper is different along several dimensions. First, our focus is on strategies which are short the US dollar and long any other currencies, whereas the authors focus on a portfolio which is US dollar neutral (the carry portfolio). Second, we provide a theoretical motivation for these returns and interpret them as a compensation for monetary policy uncertainty.

Our results are also related to Chernov, Graveline, and Zviadadze (2014), who study jumps in exchange rates in reduced form. The authors find that jumps in exchange rates mostly coincide with important macroeconomic and political announcements, among others FOMC announcements. Moreover, when the interest rate differential is positive, the probability of a large appreciation of the US dollar is higher.

Our paper proceeds as follows. We introduce our data in the next Section, and in Section 3 we study currency returns on FOMC announcement and non-announcement days. Section 4 sets up a model which reconciles our empirical findings and Section 5 presents a calibration. Section 6 concludes. Proofs are deferred to the Appendix.

2 Data

We work with daily data which begins in January 1994 and ends in December 2010. There are eight scheduled FOMC meetings in one year. This leaves us with 4,107 days without a pre-scheduled FOMC announcement and 136 FOMC announcement days. Prior to 1994, the FOMC did not disclose policy actions and market participants could infer those from the size and type of the open market operations (OMOs). As a robustness exercise, we run our analysis using data starting in January 1980. For data before January 1994, we assume that the FOMC decision became public one day after the meeting (see Kuttner (2001) and Gürkaynak, Sack, and Swanson (2005)). For the expanded sample we have 7,481 non-announcements days and 250 announcement days.

2.1 Data

Spot and Forward Data: The data for spot exchange rates and one-month forward exchange rates versus the US dollar (USD) are obtained from BBI and Reuters (via Datastream).

We denote spot and forward rates in logs as s_t and f_t, respectively. The log excess return rx_{t+1} of buying a foreign currency in the forward market and then selling it in the spot market after one month is $rx_{t+1} = f_t - s_{t+1}$. This excess return can also be stated as the log forward discount minus the change in the spot rate: $rx_{t+1} = f_t - s_t - \Delta s_{t+1}$. Since covered interest rate parity (CIP) holds at daily and lower frequency, the forward discount is equal to the interest rate differential: $f_t - s_t \approx i_t^\star - i_t$, where $i^\star$ and i denote the foreign and domestic nominal risk-free rates over the maturity of the contract (see, e.g., Akram, Rime, and Sarno (2008)). Hence, the log currency excess return equals the interest rate differential less the rate of depreciation: $rx_{t+1} = i_t^\star - i_t - \Delta s_{t+1}$.

Our total sample consists of the following 35 countries: Australia, Austria, Belgium, Canada, Czech Republic, Denmark, Euro, Finland, France, Germany, Greece, Hungary, India, Indonesia, Ireland, Italy, Japan, Kuwait, Malaysia, Mexico, Netherlands, New Zealand, Norway, Philippines, Poland, Portugal, Singapore, South Africa, South Korea, Spain, Sweden, Switzerland, Taiwan, Thailand, and the UK. Based on large failures of covered interest rate parity (see Lustig, Roussanov, and Verdelhan (2011)), we delete the following observations from our sample: Malaysia from the end of August 1998 to the end of June 2005; Indonesia from the end of December 2000 to the end of May 2007. We also study a smaller sub-sample consisting only of 15 developed countries. This sample includes: Australia, Belgium, Canada, Denmark, Euro area, France, Germany, Italy, Japan, Netherlands, New Zealand, Norway, Sweden, Switzerland, and the United Kingdom. Since the introduction of the Euro in January 1999, the sample of developed countries covers ten currencies only.

Other Data: US Consumption growth is calculated from personal consumption expenditures on non-durables and services available from FRED. Consumption growth volatility is calculated as the rolling standard deviation of consumption growth using a 36-month window (see Lustig, Roussanov, and Verdelhan (2014)). As an indicator of the economic state, we use the Chicago Fed National Activity Index (CFNAI), which is a monthly index designed to gauge overall economic activity and related inflationary pressure in the US. Negative numbers indicate a below average growth and positive numbers an above average growth. As in Kuttner (2001) and Bernanke and Kuttner (2005) we measure a surprise component in target rate changes using the change in the one-month Federal Funds Futures contract price on the FOMC meeting day. We use four proxies of policy uncertainty: the VIX, the TIV, the economic policy uncertainty index of Baker, Bloom, and Davis (2013), and an uncertainty proxy from the Bloomberg survey of the target Fed Funds rate. The economic policy uncertainty index is based on the frequency of newspaper references to economic policy uncertainty and other indicators. TIV is an implied volatility index extracted from options on 30-year Treasury futures akin to the VIX (see Choi, Mueller, and Vedolin (2014)). Finally, the uncertainty proxy is calculated as the cross-sectional standard deviation of all forecasts divided by the consensus forecast.[3] We also use GDP growth forecasts from Blue Chip for the G10 countries.

2.2 Portfolio Construction

At the end of each month t, we allocate currencies to five portfolios based on their observed forward discounts $f_t - s_t$, or equivalently their interest rate differentials. Portfolios are ranked in increasing interest rate order, so that "pf1" denotes the portfolio with the lowest interest rate currencies. We calculate daily excess log returns on individual currencies using the daily interest rate differential and daily log exchange rate changes. We assume that the interest rate differential is earned linearly over the month. Portfolio returns are calculated as the average of the currency excess returns in each portfolio as in Lustig, Roussanov, and Verdelhan (2011). The average excess return on all currency

[3] Bloomberg conducts surveys of international financial market institutions and professional forecasters regarding their expectations for the target Fed Funds rate. One important feature of this survey is that they are conducted only a couple of days before each FOMC meeting.

portfolios is denoted by DOL in line with Lustig, Roussanov, and Verdelhan (2011). HML denotes the portfolio which is long portfolio 5 and short portfolio 1. Summary statistics are presented in Table 1.

[Insert Table 1 here.]

The summary statistics confirm the well-known empirical pattern that low interest currencies earn lower average returns than high interest rate currencies: In our sample, the first portfolio earns a daily return of -0.73 bps (with a t-statistic of -1.18) while the last earns 2.54bps (with a t-statistic of 3.09). Corresponding annualized Sharpe ratios are large in absolute value: -0.28 and 0.75, respectively. We also note that the average forward discount ranges from -2.31% (pf1) to 9.56% (pf5). Table 1 also presents summary statistics for the dollar (DOL) and the HML factor, showing that while the average return of the former is not statistically significant (t-statistic 1.43), the latter's is highly significant (t-statistic 4.03).

3 Empirical Analysis

In this Section, we analyze the characteristics of returns on interest rate sorted currency portfolios on FOMC announcement and non announcement days. A number of robustness checks confirm the main finding: returns on a trading strategy that is short the US dollar and long any other currency are on average significantly larger on days when a policy announcement takes place.

3.1 Currency Portfolios on Announcement Days

Figure 1 and Table 2 present the main results. The average daily return on the low interest rate portfolio is 7.43bps on announcement days compared to -1.01bps on non-announcement days. This 8.44 bps difference is statistically significant, with a t-statistic of 2.38.[4] For the high interest rate currency portfolio, the average return is 14.43bps on announcement days compared to 2.13bps on non-announcement days; a 12.30bps

[4]The difference-in-mean test allows for different variances.

difference which is again significantly different from zero (t-statistic of 2.63). This large increase in mean returns around announcement days is not accompanied by a corresponding increase in realized risk, as measured by realized volatility, because for three out of the five portfolios annualized Sharpe ratios are significantly larger on announcement days compared to their non-announcement counterpart.[5]

Two observations are noteworthy. First, a sizable portion of the portfolios' average yearly returns is earned on FOMC announcement days. For instance, the average yearly return on portfolio 5 is 641bps (641.34bps = 2.545bps × 252, see Table 1) of which almost 20% (or 115.51bps = 14.439bps × 8, see Table 2) are earned on the eight FOMC announcement days. This proportion is even higher for the other portfolios (for example 47% for portfolio 4). Second, announcement day returns are always positive whereas returns calculated over the whole sample are negative for low interest rate currencies and positive for high interest rate currencies.[6]

We now want to explore in more detail the properties of the DOL and HML factors. Table 3 depicts summary statistics for the two portfolios conditional on announcement and non-announcement days. We note that in contrast to the unconditional average taken over the whole sample (see Table 1), the DOL factor features a large and statistically significant return on announcement days—9.63bps, with a t-statistic of 2.31—whereas the return of the HML is not significant. On the other hand, the DOL factor mean return becomes insignificant on non-announcement days (t-statistic of 1.24) whereas the HML return is highly significant (t-statistic of 5.23). Figure 2 plots average returns together with the associated t-statistic of a test for difference-in-means between announcement and non-announcement returns. The DOL mean return is 9.12bps larger on announcement days (t-statistic of 2.81). On the other hand, the 3.85bps earned in excess by the HML portfolio on announcement days are not statistically different from zero. We also note that the Sharpe ratio of the DOL portfolio more than doubles con-

[5] Annualized announcement and non-announcement Sharpe ratios are obtained by adjusting daily values for the yearly frequency of FOMC announcements (eight out of 252 trading days). Thus, the adjustment factor is $\sqrt{8}$ and $\sqrt{244}$ for the announcement and non-announcement Sharpe ratio, respectively.

[6] In our model, we interpret this homogeneous positivity of carry excess returns as a risk premium for "announcement uncertainty" risk.

ditional on an announcement, whereas the Sharpe ratio of the HML portfolio drops by almost two thirds.

Figure 3 plots the empirical density functions of returns for the different portfolios. We find that conditioning on FOMC days, the distribution displays not only a larger mean but also a positive skewness, especially for the high interest rate portfolio. Note that currency dollar returns on non-announcement days are negatively skewed, a feature that has been attributed to crash risk in currency returns (see, e.g., Farhi, Fraiberger, Gabaix, Ranciere, and Verdelhan (2014)).

To verify that our results are not driven by emerging market currencies alone, we repeat our empirical analysis for the sample of developed currencies. The lower panel of Table 2 summarizes the results. The findings remain largely unchanged, in that mean returns on the currency dollar strategy are significantly larger on FOMC announcement days. The spread between announcement and non-announcement days average returns is 9.15bps for low interest rate currencies and 14.25bps for high interest rate currencies, with t-statistics of 2.11 and 2.33, respectively.

[Insert Table 2 and Figures 1–3 here.]

One might suspect that our results are determined by a few outliers, as some announcements are more anticipated than others, and it is well-known that currency returns occasionally experience large crashes. Table 4 reports summary statistics for a winsorized data sample. Specifically, we discard the top and the bottom percentile of the data. We find that there is virtually no distinction between the mean and standard deviation of winsorized and non-winsorized returns across interest rate sorted portfolios, both on announcement and non-announcement days.

Before 1994, market participants had to infer policy decisions through the size and type of open market operations in the days after the FOMC meetings. As an additional robustness check, we study an extended sample starting in 1980 and report results in Table 5.[7] The results are consistent with those obtained for the post-1994 sample: The DOL factor has a significant announcement premium component of 5.91bps (t-statistic

[7]Additional results for the sample starting in 1980 are presented in the Online Appendix.

2.14), while there is no statistically significant difference for the HML mean return between non-announcement and announcement days. Moreover, the pattern of currency returns across interest rates is qualitatively unchanged.

Overall, these robustness checks confirm our conclusion that currency returns are significantly larger when we condition on a FOMC announcement day, with little or no change in realized risk, resulting in larger Sharpe ratios. Moreover, this announcement premium component is prominent in the average (DOL) currency portfolio, but it is not significant in the HML.

[Insert Tables 4 and 5 here.]

3.2 The Time-Series of Currency Portfolios

We now continue our investigation of currency dollar returns related to FOMC announcements by taking a time-series perspective. As a first exercise, we consider a regression of the five interest rate portfolios as well as the dollar portfolio onto a dummy which takes the value of one on announcement days and zero otherwise:

$$r_t^i = \alpha_0 + \alpha_1 \times \text{Announcement Dummy}_t + \epsilon_t, \qquad i = 1, \dots, 5, \text{DOL}. \tag{1}$$

In this regression, the intercept α_0 measures the mean return on non-announcement days, while α_1 measures the spread between announcement and non-announcement mean returns.[8]

Our findings, reported in Table 6, mirror those in Table 2 with positive coefficients for the announcement dummy for all portfolios indicating a positive spread between announcement and non-announcement days. In particular, α_1 is 8.44 (t-statistic of 2.06) for the low interest rate portfolio and 12.30 (t-statistic of 2.19) for the high interest rate portfolio, respectively. The estimates for the intercept α_0 are not significant except for portfolios 3 and 5, implying that outside FOMC meetings there is little return to

[8] To address the issue of statistical inference in small samples, we also perform a bootstrap exercise. The results are reported in Appendix B.

be earned. Similarly, for the DOL portfolio, the announcement dummy ($\alpha_1 = 9.12$) is statistically significant whereas the intercept is not.

[Insert Table 6 here.]

To investigate the relation between currency returns and the business cycle, we interact the announcement dummy with a dummy for weak economic activity. We measure economic activity using the CFNAI and set the dummy to one for negative values of the index, which indicates growth below trend. We run the following regression

$$r_t^i = \gamma_0 + \gamma_1 \times \text{Announcement Dummy}_t \times \text{CFNAI}_t + \epsilon_t, \qquad i = 1, \ldots, 5, \text{DOL}.$$

The estimates of γ_1 reported in Table 6 are statistically significant for all portfolios and much larger than the coefficients in the previous regression, which does not condition on the state of the economy. For instance, the slope coefficient increases from 8.44 and 12.30 to 11.29 and 20.84 for the low and high interest rate portfolio, respectively. For the dollar portfolio, the coefficient jumps from 9.12 to 15.62 with an associated t-statistic of 2.06. Thus, the announcement premium appears to be more pronounced during bad times.

One might argue that the time period we study is a particularly good sample to invest in currency strategies and that investors learned about this on days when the FOMC makes its announcements. To inspect this "good-news" hypothesis in more detail, we add to our regression a monetary policy surprise component extracted from Federal funds futures data as in Kuttner (2001). A negative surprise component is consistent with an unexpected tightening of monetary policy. We run the following regression:

$$r_t^i = \beta_0 + \beta_1 \times \text{Announcement Dummy}_t \times \text{Policy Surprise}_t + \epsilon_t.$$

We find that the effect of monetary policy surprises is fairly modest and not significant for most portfolios, including the DOL portfolio. Estimated coefficients (β_1) are negative for the DOL portfolio and portfolios 2 through 5, implying that an unexpected tightening of monetary policy should lead to higher currency returns.[9]

Finally, we want to inspect the relationship between currency announcement returns and uncertainty, for which we consider four proxies: the VIX, implied volatility from options on Treasury futures (TIV), an uncertainty proxy from forecasts of the target Fed Funds rate, and the economic policy uncertainty index by Baker, Bloom, and Davis (2013).[10] We interact the announcement dummy with the uncertainty proxy, and consider the following regression:

$$r_t^i = \delta_0 + \delta_1 \times \text{Announcement Dummy}_t \times \text{Uncertainty}_t + \epsilon_t.$$

The results are in the last four panels of Table 6. We find that the coefficient on the DOL portfolio is positive and significant for all four indicators of uncertainty (the t-statistics range between 1.95 (TIV) and 2.97 (VIX)). Furthermore, all of the coefficient estimates for the individual portfolios are positive and most of them are highly statistically significant. Overall, these results support the existence of a positive relationship between announcement excess returns and uncertainty.

Our results are specifically related to FOMC announcements and they do not carry over to other major US macroeconomic news releases. In the Online Appendix we repeat our empirical investigation to include a wide variety of additional announcements. We find no evidence that any of these trigger positive and significant mean currency returns. We thus conclude that the existence of announcement premia in FX markets is a specific property of FOMC meetings.

To conclude this section, we use a limited sample of high frequency data to inspect how a DOL factor built from four foreign currencies – the Euro, the Japanese Yen, the Swiss Franc, and the British Pound – behaves intra-daily on announcement

[9] In unreported results, we also test for the hypothesis that surprises mainly matter during bad economic times (see, e.g., Gilbert (2011)). However, we find that an interaction term between the announcement dummy, the policy surprise, and the CFNAI produces coefficients which are not statistically significant.

[10] For example Bekaert, Hoerova, and Duca (2013) study the link between the VIX and monetary policy, whereas the economic policy uncertainty index has been extensively used to study the effect of uncertainty on risk premia, see, e.g., Pástor and Veronesi (2013) and Amengual and Xiu (2013).

days. In the Online Appendix, we plot average (across all announcements events) DOL cumulative returns at five-minutes intervals, within a 72-hour window spanning the announcement. Interestingly, we find evidence of a positive drift in the 12 hours preceding the announcement, which accounts for approximately 80% of the daily return. Although quantitatively different, this evidence resembles the findings of Lucca and Moench (2014) concerning S&P500 announcement returns. The fact that most of the return is earned in the run-up to the announcement, before any new information is released, is consistent with the interpretation of a risk premium for the ex-ante compensation of monetary policy uncertainty. We formalize this interpretation in the next section.

4 Theory

In this Section, we describe a general equilibrium model of an open economy consistent with the empirical evidence previously discussed. The model draws upon two strands of the literature: First, long-run risks in an international framework (Colacito and Croce (2011)) and second, the political uncertainty literature (Pástor and Veronesi (2012)). In the spirit of the latter, we assume that the policy action of the central bank affects expected output growth. As the precise impact is not observed, agents learn about it in a Bayesian fashion until the date when a new policy is announced. The unpredictable revision of the policy, and consequently of its impact, determines shocks to the expected output growth, which are more volatile than updates following "normal" (i.e. non-announcement) news. We thus distinguish between "announcement uncertainty" which captures the announcement-day volatility and the "normal" intra-announcement volatility. Due to recursive preferences, the announcement risk affecting long-run expectations is priced in equilibrium.

4.1 Preferences

Time is discrete and the horizon is infinite. There are $N+1$ consumption goods and $N+1$ countries, the "home" country and N "foreign" countries.[11] Each country is populated by a representative agent with Epstein and Zin (1989) preferences:

$$U_{i,t} = \left\{(1-\delta)(C_{i,t})^{1-\frac{1}{\psi}} + \delta \mathbb{E}_t \left[(U_{i,t+1})^{1-\gamma}\right]^{\frac{1}{\theta}}\right\}^{\frac{1}{1-\frac{1}{\psi}}} \quad i = h, 1, \ldots, N,$$

where the subscript h identifies the home country, and $\theta = (1-\gamma)/(1-1/\psi)$. δ is the subjective discount rate, γ the relative risk aversion (RRA) coefficient, and ψ is the intertemporal elasticity of substitution (IES). Preference parameters are assumed to be common to all countries. As in Colacito and Croce (2011), we assume that in equilibrium agents consume the entire domestic output $Y_{i,t}$, and representative agents exclusively hold the claims to domestic endowments. In addition, markets are complete.

4.2 Dynamics of Fundamentals

A key assumption in our setup is that the home country's monetary policy affects real quantities both at home and abroad. The first part of the assumption, i.e. that monetary policy can have real effects, can be justified for example through nominal rigidities.[12] The second part of the assumption is related to the special role given to the US, which is due to two main reasons. First, dollar holdings feature prominently in official foreign exchange reserves. In 2013, 61% of allocated foreign reserves were held in USD compared to 24% held in EUR (see International Monetary Fund (2014)). Second, in international trade the dollar is widely used for invoicing and settling import and export transactions around the world. For instance, according to Bank of International Settlements (2013), 87% of all currency transactions take place in USD, more than twice as much as in EUR.[13]

[11] Consistently with the empirical analysis, we consider the United States as the home country.

[12] See, e.g., Cochrane (2014) who discusses a sticky-price model where monetary policy affects output and real interest rates.

[13] A large macroeconomic literature studies the spillover effects of US monetary policy onto output and asset prices internationally. See, e.g., Ilzetzki and Jin (2013) for a recent comprehensive empirical study on how US monetary policy affects output, exchange and interest rates for a large cross-section of countries.

We model the real effects of US monetary policy in reduced-form and we neither consider the decision making process of the monetary authority, nor its objective function. Furthermore, we are silent about the specific transmission mechanism from domestic monetary policy to foreign output growth. In line with a regular FOMC meeting calendar, we assume that the central bank can revise its policy every $\overline{A}$ periods.

The logarithmic output growth of country i, $\Delta y_{i,t+1} = \log Y_{i,t+1} - \log Y_{i,t}$, is assumed to evolve as follows:

$$\Delta y_{i,t+1} = \overline{\mu}_i + \beta_i \mu_t + \sigma_d \sqrt{x_{i,t}} \epsilon_{i,t+1}, \quad (2)$$

$$x_{i,t+1} = \alpha_{i,x} + b_{i,x} x_{i,t} + \sigma_{i,x} \sqrt{x_{i,t}} \omega^x_{i,t+1}, \quad \text{corr}(\epsilon_i, \omega^x_i) = 0, \quad (3)$$

where coefficients $(\sigma_d, \alpha_{i,x}, b_{i,x}, \sigma_{i,x})$ are constant. Monetary policy affects expected output growth through the common component μ_t, which is assumed constant between policy announcement (i.e. revision) dates: $\mu_t = m(t_a)$, when $t \in [t_a, t_a + \overline{A})$, for some constant $m(t_a)$ and announcement date t_a. A large macroeconomic literature (see, e.g., Christiano, Eichenbaum, and Evans (2005)) has documented a relatively fast decay in the response of output growth to a monetary policy shock. Our assumption about μ_t (i.e. that it takes a piece-wise constant form) is not inconsistent with this evidence because the time period between policy announcement dates is fairly short (FOMC meetings are scheduled every six or seven weeks). Each country has a different loading β_i on this policy factor, and without loss of generality μ_t is defined such that the domestic loading (β_h) is not necessarily one. $\overline{\mu}_i$ is a constant country-specific component of expected growth. While agents observe β_i and the volatility of output growth $x_{i,t}$, which follows a square-root random process, they observe neither the innovations $\epsilon_{i,t+1}$, nor the expected growth components $(\overline{\mu}_i, \mu_t)$. Between any two policy announcement dates, agents learn about expected growth in a Bayesian fashion observing both the realizations of domestic output growth and a common signal.[14] Specifically, we assume that this common

[14] For simplicity we assume an autarchic learning process where agents ignore foreign output growth for their inference.

signal is informative about the piece-wise constant monetary policy impact, and that it features stochastic volatility of the square-root type:

$$\begin{aligned} s_{t+1} &= \mu_t + \sigma_s \sqrt{x_t} \epsilon_{t+1}, \\ x_{t+1} &= \alpha_x + b_x x_t + \sigma_x \sqrt{x_t} \omega^x_{t+1}, \quad \operatorname{corr}(\epsilon, \omega^x) = 0. \end{aligned}$$

We think of s as a monetary policy signal which captures the numerous speeches done by officials at the central bank. We refer to it as a "normal" signal because it occurs during non-announcement times and it is informative about the impact of the monetary policy in place. The following proposition outlines agents' learning.

Proposition 1. *Let $\widetilde{\mu}_{i,t} = \mathbb{E}[\overline{\mu}^i | \mathcal{I}_t] + \beta_i \mathbb{E}[\mu_{i,t} | \mathcal{I}_t]$ denote the posterior estimate of country i's expected output growth.*[15] *Under the assumption that in the updating rule agents apply a constant weighting matrix (the Kalman gain) to the current estimate and new information, we can describe the evolution of $\widetilde{\mu}_{i,t}$ between any two monetary policy revision dates as follows:*[16]

$$\widetilde{\mu}_{i,t+1} = \widetilde{\mu}_{i,t} + \sqrt{\overline{\sigma}_1 + \sigma_d^2 x_{i,t}}\, \epsilon_{i,t+1} + \beta_i \sqrt{\overline{\sigma}_2 + \sigma_s^2 x_t}\, \epsilon_{t+1}, \tag{4}$$

where ϵ and ϵ_i are standard Gaussian innovations, while $\overline{\sigma}_1$ and $\overline{\sigma}_2$ are constants.[17]

The hypothesis of a constant weighting matrix in the posterior updating rule is called "recency-biased learning" in Bansal and Shalistovich (2010) because the agent tends to overweight the present news in the posterior update relative to the optimal Kalman gain, especially at times of large uncertainty about the signal or output growth. Collin-Dufresne, Johannes, and Lochstoer (2014) study an equilibrium model where young generations are affected by this learning bias and they report supporting empirical evidence.[18]

[15] $\mathcal{I}_t$ is the agents' actual information set, which includes present and past observations of $(\Delta y_{i,t}, s_t)$.

[16] That is, $t \in [t_a, t_a + \overline{A})$, for a generic announcement date t_a.

[17] To allow for full asymptotic learning, i.e., letting the variance of posterior output growth vanish asymptotically, the Kalman gain could be approximated as a deterministic function of time. In our setup, however, policy revisions happen in finite time intervals. Hence, the policy impact is piece-wise, which always prevents agents to learn fully.

[18] Rather than taking a stance on the suitability of this behavioral hypothesis, we introduce it to simplify the model and obtain closed-form expressions, at the cost of a reasonable approximation error. We could also assume that the volatility of output growth is large relative to the volatility of the signal s_t, so that the Kalman gain of the latter would be large relative to the former, and both would display little time variation, consistent with our hypothesis.

The expected output growth given in equation (4) resembles the highly persistent process assumed by the long-run risk literature (see, e.g., Bansal and Yaron (2004)). We note that the persistence is not critical for our purpose, and it is a consequence of the constant intra-announcement policy impact μ_t.[19] The dynamics (4) follow a random-walk with two stochastic volatility factors, one that is global (x_t) and one that is country-specific ($x_{i,t}$). The volatility of global shocks is affected by the parameter β_i, the loading of country i's expected growth on the monetary policy component. Intuitively, the larger the sensitivity to the policy, the larger the volatility of long-run risks related to policy news releases. In the international finance literature, asymmetric global factor loadings are known to produce a violation of uncovered interest rate parity (see Backus, Foresi, and Telmer (2001)) and to produce consistent evidence for currency returns (see, e.g., Lustig, Roussanov, and Verdelhan (2011), Lustig, Roussanov, and Verdelhan (2014), and Hassan and Mano (2013)).

We introduce announcement risk into the economy by considering a sequence of announcement dates, at which the monetary authority can revise its policy as in Savor and Wilson (2013). Furthermore, we model the effects of an announcement on country i's expected output growth $\widetilde{\mu}_{i,t}$. As in the political uncertainty literature (Pástor and Veronesi (2013)), we implicitly assume that an informational friction prevents the agents from exactly solving the policy maker's optimization. Thus, the announcement contains an unexpected component, modeled as a shock to $\widetilde{\mu}_{i,t}$ with volatility distinct from and, in our intuition, typically larger than the "normal" news component x.[20] We refer to this volatility, z, as "announcement uncertainty" and we model it with a square root process.

[19] In fact, we could assume a time-varying form for the unobserved component μ_t and intuitively obtain less persistent dynamics for the posterior $\widetilde{\mu}_{i,t+1}$. For simplicity, we opt for the constant form.

[20] In a continuous-time framework, we would think of non-announcement shocks as Brownian motions, and of announcement shocks as jumps. Pástor and Veronesi (2013) obtain this feature by assuming that agents have imperfect knowledge of the political cost associated with the implementation of a given policy.

The overall evolution of expected output growth and announcement uncertainty reads as follows:

$$
\begin{aligned}
\widetilde{\mu}_{i,t+1} &= \widetilde{\mu}_{i,t} + \sqrt{\overline{\sigma}_1 + \sigma_d^2 x_{i,t}}\,\epsilon_{i,t+1} + \beta_i\Big(\sqrt{\overline{\sigma}_2 + \sigma_s^2 x_t}\,\epsilon_{t+1} \\
&\quad + A_{t+1}\sqrt{z_t}\,\eta_{i,t+1}\Big) \qquad (5)\\
z_{t+1} &= \alpha_z + b_z z_t + \sigma_z\sqrt{\overline{\sigma}_2 + \sigma_d^2 x_t}\,\omega_{t+1}^z, \quad \mathrm{corr}(\omega_t^z, \epsilon_t) = \rho_z < 0.
\end{aligned}
$$

In expression (5), A_{t+1} is a dummy variable that takes a value of one if $t+1$ is an announcement date, meaning that a monetary policy revision – possibly also no action – has taken place between time t and $t+1$. The announcement uncertainty z_t is typically negatively correlated with the systematic non-announcement shock of $\widetilde{\mu}_{i,t}$ (ϵ_t). The reason is that negative shocks to expected output growth are generally symptoms of an ineffective current monetary policy, which is more likely to be revised, thus raising the uncertainty about the ensuing effects of an upcoming announcement. An important difference between the "normal" news shock ϵ, and the announcement shock η_i, is that the latter is country-specific, with an imperfect cross-sectional correlation. This means that an announcement can be interpreted as good news for long-run risk by some countries and as bad news by others. Moreover, countries with a larger loading β_i on the monetary policy component of expected growth should react more strongly to announcement news. Finally, the volatility of z_t is proportional to "normal" uncertainty $\sqrt{x}$ in order to preserve the affine nature of the model.

Using the learning framework of the political uncertainty literature, we have thus motivated a long-run risk model with announcement shocks to agents' expectations arising from monetary policy revisions. The uncertainty arising from these revisions, z, and the "normal" uncertainty factors, both global (x) and local (x_i), play an important role in reconciling the model with the empirical evidence about currency risk premia, both during announcement and non-announcement periods. In the interest of clarity and tractability, we consider the following slightly re-parameterized state-space model, which inherits the main elements of the specification discussed so far and shares its economic motivation:[21]

$$\Delta y_{i,t+1} = \overline{\mu}^i + \mu_{i,t} + \sigma_d \nu_{i,t+1}, \tag{6}$$

$$\mu_{i,t+1} = b_m \mu_{i,t} + \sigma_m \sqrt{x_{i,t}}\, \epsilon_{i,t+1} + \beta_i \left(\sqrt{x_t}\, \epsilon_{t+1} + A_{\tau=1} \sqrt{z_t}\, \eta_{i,t+1} \right), \tag{7}$$

$$x_{i,t+1} = \alpha_x + b_x x_{i,t} + \sigma_x \sqrt{x_{i,t}} \omega^x_{i,t+1}, \tag{8}$$

$$x_{t+1} = \alpha_x + b_x x_t + \sigma_x \sqrt{x_t} \omega^x_{t+1}, \tag{9}$$

$$z_{t+1} = \alpha_z + b_z z_t + \sigma_z \sqrt{x_t}\, \omega^z_{t+1}, \quad \mathrm{corr}(\omega^z_t, \epsilon_t) = \rho_z < 0. \tag{10}$$

Our model is not time-homogeneous. Since announcements take place on a regular basis, we can consider the "time to next announcement", τ, as a state variable that replaces calendar time. In particular, $A_{\tau=1}$ is the indicator function of the event $\tau = 1$, such that a monetary policy announcement takes place between this date and the next. As mentioned before, the announcement jump innovation $\eta_{i,t+1}$ is allowed to be country-specific and cross-sectionally correlated. In particular, for any two countries i and j, we have $\mathrm{corr}(\eta_{i,t+1}, \eta_{j,t+1}) = \rho_{ij}$. All other correlations are assumed to be zero. Our model features three global and two country-specific state variables. The global state variables are (i) uncertainty related to "normal" global news x, (ii) announcement uncertainty z and (iii) time to next announcement τ; the country-specific variables are (i) uncertainty related to "normal" local news x_i and (ii) country i's expected output growth μ_i.

4.3 Asset Prices

The log stochastic discount factor (intertemporal marginal rate of substitution) of country i, which can be inferred from the first-order conditions of the representative agent's optimization, is:

$$m_{i,t+1} = \theta \log \delta - \frac{\theta}{\psi} \Delta y_{i,t+1} + (\theta - 1) r^y_{i,t+1}, \tag{11}$$

[21] In particular, we have made the following parametric and notational modifications: log output growth $\Delta y_{i,t}$ has constant volatility and its innovations, now denoted by ν_i instead of ϵ_i, are independent from the innovations in long-run risks. The time-varying component of expected output growth is now denoted by $\mu_{i,t}$ instead of $\widetilde{\mu}_{i,t}$. The latter has an autoregressive coefficient b_m, which, in line with the long-run risk literature, is intended to be almost equal to one. The announcement shock to long-run risks is now denoted by η_i, instead of ν_i. The stochastic volatility components of long-run risks have also been simplified: $\sqrt{x_{i,t}}$ and $\sqrt{x_t}$ replace $\sqrt{\overline{\sigma}_1 + \sigma_d^2 x_{i,t}}$ and $\sqrt{\overline{\sigma}_2 + \sigma_s^2 x_t}$, respectively.

where r_i^y is the log return on the claim to aggregate output. Following the standard approach to solve for asset prices in long-run risk models (see, e.g., Bansal and Yaron (2004)), we show in the Appendix that the equilibrium price-dividend ratio of the claim to the aggregate output of country i reads as follows:

$$pc_{i,t} = B_0(i,\tau) + B_1\,\mu_{i,t} + B_2(i,\tau)\,z_t + B_3(i,\tau)\,x_t + B_4\,x_{i,t}, \tag{12}$$

where we have emphasized the dependence of the deterministic coefficients B on time to next announcement and on country i.[22] We impose the following set of mild assumptions.

Assumption 1. *We assume the parameter set satisfies the following restrictions:* $\gamma > 1/\psi$, $\psi > 1$, $\rho_z < 0$, $\beta_i > 0$.

$\gamma > 1/\psi$ is a standard assumption and implies that the Epstein-Zin agent has a preference for early resolution of uncertainty, while $\psi > 1$ implies that the substitution effect dominates the income effect, well in line with earlier literature (see Bansal and Yaron (2004)). $\rho_z < 0$ allows announcement uncertainty to be larger in bad times, on average, whereas $\beta_i > 0$ is a normalization. Under these assumptions, the log price-consumption ratio of a given country is increasing in its expected output growth ($B_1 > 0$). Conversely, larger long-run risk uncertainty arising from "normal" (non-announcement) news, both local ($B_4 < 0$) and global ($B_3(i,\tau) < 0$), decreases the price-consumption ratio: when $\psi > 1$, the desire to down-weight risky assets prevails over the additional precautionary savings demand for all assets. A similar effect occurs with larger announcement uncertainty z_t, ($B_2(i,\tau) < 0$). Intuitively, the more imminent the announcement, the larger the price drop caused by a positive announcement uncertainty shock. This is due to the persistence of z as there is less scope for uncertainty to resolve in time before the announcement. It is useful to analyze the response of the price-consumption ratio to the main driver of country-wise heterogeneity, the monetary-policy loading β_i. As shown in the Appendix, we have $\frac{\partial B_2(i,\tau)}{\partial \beta_i} < 0$ and $\frac{\partial B_3(i,\tau)}{\partial \beta_i} < 0$. Therefore, countries with larger β_i are more sensitive to both announcement and (global) non-announcement uncertainty shocks, as their price-consumption ratio drops (increases) more in response to positive (negative) shocks to z and x. The reason is that countries whose expected output growth

[22] The coefficients B are derived in Appendix A.

is more affected by monetary policy are intuitively more sensitive to both "normal" news releases and announced policy revisions. Thus, they have more volatile long-run risks, which translates into a more pronounced price response to uncertainty shocks. For the same reason, the positive relation between long-run risk volatility and the equity risk premium is enhanced by a larger β_i.

4.4 Interest Rates and Currency Risk Premia

It is easy to show that the equilibrium risk-free (continuously compounded) interest rate of a given country i is affine in the country's expected output growth and uncertainty state variables:

$$r_{i,t} = C_0 + C_1 \mu_{i,t} + A_{\tau=1} C_2(i) z_t + C_3(i,\tau) x_t + C_4 x_{i,t}, \tag{13}$$

where coefficients C are derived in Appendix A. As C_1 is positive, a negative shock to long-run risk reduces the interest rate, because both the desire to save more to transfer consumption to future periods (wealth effect) and the desire to dump risky for safer assets (substitution effect) increases the demand for the riskless asset. Conversely to the market risk premium, the risk-free rate is decreasing in "normal" uncertainty (global: $C_3(i,\tau) < 0$; and local: $C_4 < 0$). Moreover, the risk-free rate experiences a negative jump during announcement periods, proportional to the announcement uncertainty ($C_2(i) < 0$). More important for the analysis to follow, there is an inverse relation between the interest rate and β_i, the loading on the output growth component driven by monetary policy.[23] In light of the dynamics for μ_i given in equations (6) to (10), this implies that the cross-sectional correlation between output growth expectations decreases in the interest rate differential. We provide empirical support for this claim by considering the following regression:

$$\widehat{\mu}_{i,t+1} = \alpha_0 + \alpha_1 \widehat{\mu}_{h,t+1} + \alpha_2 (r_{i,t} - r_{h,t}) \widehat{\mu}_{h,t+1} + \epsilon_{i,t+1},$$

where h is the home country (the US) and i denotes any of the N foreign countries. The expectation of country i's output growth is approximated by the corresponding

[23] Formally, we show in the Appendix that $\frac{C_3(i,\tau)}{\partial \beta_i} < 0$ and $\frac{C_2(i)}{\partial \beta_i} < 0$.

consensus analysts' forecast $\widehat{\mu}_{i,t+1}$. Table 7 reports regression results for nine developed countries.

[Insert Table 7 here.]

We are mainly interested in the slope coefficient of the interaction term: α_2 is negative for all countries and statistically significant in all but two cases. This evidence supports the prediction of our model, namely that the correlation between the foreign and domestic GDP growth expectations decreases with the interest rate differential.

The intuition for this result is similar to Lustig and Verdelhan (2007), who show that currencies with low (high) US consumption betas have low (high) interest rates. The authors also present empirical support for this mechanism using realized consumption growth data. Note that in our model β_i measures the sensitivity of foreign expected output growth to the domestic counterpart, rather than the sensitivity of currency returns to realized consumption growth.

We next define the exchange rate $Q_{i,t}$ as the number of units of country i's currency exchanged for a unit of domestic currency. Thus, an increase of $Q_{i,t}$ corresponds to an appreciation of the US dollar and a depreciation of the foreign currency. Assuming a complete market setting, no-arbitrage implies that the change of the logarithmic exchange rate is equal to the difference between domestic and foreign log-stochastic discount factors: $\Delta q_{i,t+1} = m_{h,t+1} - m_{i,t+1}$, where lower case letters denote logarithms. A currency dollar trade is a one-period zero-investment strategy that invests one unit of domestic currency (US dollar), financed at the domestic risk-free rate, into the foreign risk-free asset. Thus, the log return of the strategy expressed in the domestic currency, is:

$$r^c_{i,t+1} = r_{i,t} - r_{h,t} - \Delta q_{i,t+1}.$$

Hence, realized USD returns depend on the volatilities of the home and foreign stochastic discount factors and on their correlations. Expression (A-31) in the Appendix, which reports the exchange rate dynamics, shows that because of Epstein-Zin preferences these volatilities are driven by shocks to long-run risks, both due to announcement and non-announcement news, and by shocks to uncertainty in long-run risk. The next Proposition details the risk premium earned in equilibrium by a currency trading strategy.

Proposition 2. *The equilibrium risk-premium (inclusive of the Jensen inequality adjustment) of a currency strategy long country i's currency and short the home currency (USD) is*

$$\mathbb{E}_t[r^c_{i,t+1}] + \frac{1}{2}Var_t[r^c_{i,t+1}] = \gamma^2\sigma_d^2 + g_x(i,\tau-1)x_t + A_{\tau=1}g_z(i)z_t + g_h x_{h,t}, \tag{14}$$

where the deterministic functions $g_x(i,\tau-1)$, $g_z(i)$, g_h are given by equations (A-33) to (A-35) in the Appendix.

According to equation (14), the currency risk premium is linear in the global uncertainty factors and the domestic one. With a slight abuse of terminology, we call the component $g_z(i)z_t$ the "announcement premium" because it is a jump component occurring on monetary announcement dates.[24] In particular, it takes the following form:

$$g_z(i) = (\theta-1)^2\rho_i^2 B_1^2 \beta_h(\beta_h - \beta_i\rho_{h,i}). \tag{15}$$

Note from (15) that g_z and its cross-sectional pattern depend on two elements: First, the announcement volatility differential between foreign and domestic long-run risks, driven by the relative magnitudes of β_h and β_i, and second, the correlation between foreign and domestic announcement news, $\rho_{h,i}$. If domestic long-run risk is more sensitive to an announcement, that is $\beta_h > \beta_i$, then bad (good) news for the home country conveyed by a policy revision leads to an appreciation (depreciation) of the exchange rate, thus, to a negative (positive) currency return. In other words, if foreign countries with low β_i are involved, currency returns do not hedge monetary announcement shocks to long-run risks, so that currency announcement premia are positive. Conversely, when domestic long-run risks are less sensitive to announcements ($\beta_h < \beta_i$), the hedge takes place and the currency announcement premium is negative, unless foreign and domestic announcement shocks are negatively correlated. If a policy revision conveys opposite

[24] It is an abuse of terminology because the term is not solely responsible for the difference between the currency risk premium on announcement and non-announcement days, as the state variable τ affects the other components. In the calibration exercise, however, we find that $g_z(i)z_t$ is the quantitatively dominating term.

news to foreign and domestic expected growth, the sign of the currency dollar return is again in line with domestic news, and the risk premium may become positive.[25]

The global component of the non-announcement premium is $g_x(i,\tau-1)x_t$ and it can be decomposed into three parts:

$$\begin{aligned} g_x(i,\tau-1)x_t &= \underbrace{(\theta-1)^2\rho_i^2 B_1^2 \beta_h x_t \left(\beta_h - \beta_i\right)}_{\text{long-run risk shocks}} \\ &\quad + \underbrace{(\theta-1)^2\rho_i^2 B_3(h)\sigma_x^2 x_t \left[B_3(h,\tau-1) - B_3(i,\tau-1)\right]}_{\text{volatility of long-run risk shocks}} \\ &\quad + \underbrace{(\theta-1)^2\rho_i^2 B_1^2 B_2(h,\tau-1)^2\sigma_z^2 x_t \left[B_2(h,\tau-1) - B_2(i,\tau-1)\right]}_{\text{announcement uncertainty shocks}}. \end{aligned}$$

The intuition for the sign and cross-sectional pattern of each of these components is similar to the intuition given for the announcement premium, with the difference that shocks to foreign and domestic stochastic discount factors are now perfectly correlated, so that only the magnitude of the foreign policy loading β_i relative to β_h matters. In particular, assume that the foreign country has larger β_i: then negative "normal" shocks to domestic long-run risks (the first component), or positive shocks to non-announcement (second component) and announcement (third component) uncertainty, are on average associated to positive currency returns. Since these shocks are all considered bad news by the domestic agent, any investment in currency i serves as a hedge which motivates a negative sign for each of the components above. Conversely, an investment in low β_i currencies behaves cyclical, in the sense that bad domestic news from long-run risk and volatility of long-run risks are on average coupled to negative currency returns. In this case, the cyclical behavior motivates a positive sign for each of the components above. This mechanism is the long-run risk counterpart of the reduced-form setup presented in Lustig, Roussanov, and Verdelhan (2011).

The local component of the non-announcement premium is $g_h x_{h,t}$, which compensates the agent for US (i.e. domestic) specific shocks to long-run risk and to its volatility.

Note that the model-implied cross-sectional variation of the currency premium (14) – controlled by β_i, the loading of expected output growth on the systematic monetary

[25] Indeed a calibration yields a negative $\rho_{h,i}$ for high β_i countries, and an announcement currency premium that is positive for all countries.

policy effect – is consistent with the empirical evidence. Section 3 shows that currency mean excess returns are increasing in the foreign/domestic interest rate differential: consistently, in our model both the equilibrium risk-free rate – see (13) – and the currency risk premium are decreasing in β_i. The next Corollary summarizes this finding.

Corollary 1. *A currency strategy long countries with negative (positive) interest rate differential compared to the home country—i.e., $\beta_i > \beta_h$ ($\beta_i < \beta_h$)—earns a negative (positive) risk premium component $g_x(i, \tau - 1)x_t$. This global non-announcement component is increasing in the interest rate differential. Moreover, the announcement risk-premium component $A_{\tau=1}g_z(i)z_t$ is positive for countries with positive interest rate differential, while its sign depends on the foreign/domestic announcement shock correlation ($\rho_{h,i}$) if the differential is negative.*

4.5 Carry and Dollar factors on announcement days

Lustig, Roussanov, and Verdelhan (2011) conclude that two factors, HML and DOL, explain most of the cross-section of currency returns. We now explore in more detail the model-implied equivalents of these factors, focusing in particular on their properties on announcement days.

The HML factor is the return on a strategy that is long an equally-weighted portfolio of high interest rate currencies ($\beta_i < \beta_h$) and short an equally-weighted portfolio of low interest rate currencies ($\beta_i > \beta_h$). On non-announcement days, Corollary 1 implies that HML has maximal exposure to global currency risk. The market price of HML risk coincides with the HML risk premium, and it is positive as given in equation (A-48) in Appendix A.[26] Not surprisingly, the factor betas increase in the interest rate differentials, so that betas line up with risk premia.[27]

Focusing on announcement days, the HML premium is given by

$$\begin{aligned}\mathbb{E}_t[\text{HML}_{t+1}] + \frac{1}{2}Var_t[\text{HML}_{t+1}]|_{\tau=1} &= \underbrace{-Cov_t\left[\text{HML}_{t+1}, m_{h,t+1}\right]|_{\tau\neq 1}}_{\text{premium on non-announcement days}} \\ &\quad +(\theta-1)^2\rho^2 B_1^2\beta_h(\overline{\beta_L\rho_L} - \overline{\beta_H\rho_H})z_t. \quad (16)\end{aligned}$$

[26]The linear factor model $\mathbb{E}_t\left[r^c_{i,t+1}\right] + \frac{1}{2}Var_t\left[r^c_{i,t+1}\right] = \beta_i^{HML}\lambda_{HML}$ applied to the factor mimicking portfolio (for which $\beta^i_{HML} = 1$) leads to $\lambda_{HML} = \mathbb{E}_t\left[r^{HML}_{i,t+1}\right] + \frac{1}{2}Var_t\left[r^{HML}_{i,t+1}\right]$.

[27]See the Appendix for a proof.

This expression crucially depends on the average correlation between announcement shocks in the US and high ($\overline{\rho_L}$) or low ($\overline{\rho_H}$) interest rate countries, respectively. These correlations also influence the HML factor betas of currency returns, as (A-51) in the Appendix shows. It turns out that on announcement days HML betas line up inversely to risk premia if the HML announcement premium is negative.

The DOL factor is the return on an equally weighted portfolio long all foreign currencies. On non-announcement days, it is exposed to US-specific shocks. Verdelhan (2013) shows that the DOL factor should proxy for a global risk component because factor betas display significant cross-sectional variation, which cannot be obtained when DOL loads only on a local shock. Our model is consistent with this feature on announcement days, as DOL proxies for US announcement shocks to long-run risks, which are systematic. The DOL premium on announcement days is given by

$$\mathbb{E}_t[\mathrm{DOL}_{t+1}] + \frac{1}{2}Var_t[\mathrm{DOL}_{t+1}]|_{\tau=1} = \underbrace{-Cov_t\left[\mathrm{DOL}_{t+1}, m_{h,t+1}\right]|_{\tau=1}}_{\text{premium on non-announcement days}} + (\theta-1)^2\rho^2 B_1^2 \beta_h(\beta_h - \overline{\beta\rho})z_t. \quad (17)$$

The announcement component of the DOL premium—the announcement market price of DOL risk—depends on the average correlation between announcement shocks in the US and the rest of the world. If we impose the normalizing assumption that the US has the average β_i, then the DOL premium is positive, which is not guaranteed for the HML premium.[28] On the other hand, DOL factor betas depend on currencies' exposure to (or correlation with) US announcement risk, as expression (A-57) shows, consistently with the cross-sectional dispersion observed by Verdelhan (2013).

In the data, we find that the DOL premium earned on announcement days is significantly different from non-announcement days, whereas this feature does not hold for the HML premium. Interestingly, this finding can be explained by our model, since the DOL announcement premium must be positive, as noted above, while the HML announcement premium need not.

[28] The assumption $\beta_h = \frac{1}{N}\sum_{j=1}^{N}\beta_j$ implies that $\beta_h > \overline{\beta\rho} = \frac{1}{N}\sum_{j=1}^{N}\beta_j\rho_{h,j}$. Note that the systematic impact of US monetary policy, μ_t, is defined such that the US does not have β_i equal to one.

We conclude this section with two remarks. First, in our model currency returns are driven by real quantities and inflation is not accounted for. In fact, the cross-sectional variation of currency premia is related to real, rather than nominal interest rate differentials. While the model could easily be extended to include an inflation process, empirically inflation does not seem to drive currency returns. For example, Hollifield and Yaron (2001) document that nearly all of the variation in currency returns is due to real variables, with little nominal impacts. More recently, Lustig, Roussanov, and Verdelhan (2011) find that nominal interest rates and real interest rates spreads produce equally large dispersion in currency returns. Second, note that our model implies that the announcement risk premium should be equal to zero if there is no monetary policy uncertainty. After the 2008 crisis, the US Federal Reserve along with other central banks have started to provide more information about future monetary policy through increased forward guidance.[29] This has arguably reduced monetary policy uncertainty and, thus, the level of the state variable z_t in our model. We test the model predictions during periods of increased forward guidance in Appendix C and indeed find evidence that announcement risk premia are much smaller during the more recent period.

5 Calibration

We now calibrate the model parameters by targeting moments of interest rates, exchange rates, currency returns, and the US equity premium. The set of moments resembles that used in Lustig, Roussanov, and Verdelhan (2011) except that we condition on both announcement and non-announcement dates. As in their paper, we adopt a two-step procedure. First, a symmetric version of the model is calibrated where all countries share the same loading β. Moreover, we also assume a perfect correlation among countries' announcement shocks ($\rho_{ij} = 1$). Then, heterogeneity in loadings and announcement shock correlations is introduced to match some feature of the cross-section of currency returns, both unconditionally and conditionally to an announcement event. There are eight FOMC announcement days per year, or approximately one every $\overline{A} = 32$ days.

[29] Forward guidance can be implemented for example by announcing numerical guidelines for the forward path of the policy interest rate, or through more qualitative verbal statements.

The frequency of our model is daily, consistently with our aim to capture policy announcement effects, even though, we target monthly-equivalent moments, as customary in the long-run risk literature.

We set the subjective discount rate δ to 0.999, or the daily equivalent of the value used in Colacito and Croce (2011). Parameter $\overline{\mu}$ coincides with an unconditional US consumption growth of 0.00015, in daily equivalents. In the first stage of the calibration, the remaining 13 parameters minimize the root mean squared error (RMSE) obtained from matching 16 moments: the volatility of US consumption growth, and the country-wise average of the slope in the UIP regression:

$$\Delta q_{i,t+1} = a + b_i \left(r_{i,t} - r_{US,t} \right) + \epsilon_{i,t+1},$$

where $\Delta q_{i,t+1}$ are changes in the exchange rate and $(r_{i,t} - r_{US,t})$ is the interest rate differential between country i and the US. We also match mean, standard deviation, and autocorrelation of the US real short rate, both conditionally and unconditionally to an announcement date, the country-wise average correlation of real interest rates, the standard deviation of changes in exchange rates, and the country-wise average of the mean currency return.[30] Note that we target empirical moment of real interest rates, consistently with the focus of our model on real quantities. As of equity markets moments, we match the unconditional US equity premium. For this purpose, we assume aggregate dividend dynamics of the form:

$$\Delta d_{i,t+1} = \overline{\mu}_i + \lambda \mu_{i,t} + \sigma_d \nu^d_{i,t+1},$$

with the leverage parameter λ set to 3 as in Colacito and Croce (2011).[31]

Table 8 reports the empirical moments, along with their model-implied counterparts and computational details.

[Insert Table 8 here.]

[30] This is akin to the unconditional mean of the dollar factor in Lustig, Roussanov, and Verdelhan (2011).

[31] Parameters $\overline{\mu}_i$ and σ_d coincide with their consumption-growth counterparts, while consumption and dividend growth are assumed conditionally uncorrelated. The equilibrium price-dividend ratio is $(p-d)_t = \log \frac{P_t}{D_t} = A_0(i,t) + A_1 \mu_{i,t} + A_2(i,t) z_t + A_3(i,t) x_t + A_4 x_{i,t}$. Coefficients are obtained similarly to those of the price-consumption ratio and they are not reported.

In a second stage, we introduce heterogeneity in the policy loading β and in announcement shock correlations $\rho_{h,i}$ with the purpose of matching (i) the unconditional mean of the HML factor, and, importantly, (ii) the currency risk premium component due to announcement risk, both for high and low interest rate differential countries. To this end, we define an interval $[\underline{\beta}, \overline{\beta}]$ centered around the first stage β, which is identified with the home-country (US) loading. Countries' loadings are assumed equally spaced on $[\underline{\beta}, \overline{\beta}]$. We find corresponding values $\underline{\rho}_{h,i}$ and $\overline{\rho}_{h,i}$, and assume that on $[\underline{\beta}, \beta_h]$ ($[\beta_h, \overline{\beta}]$) the announcement shock correlation $\rho_{h,i}$ increases (decreases) linearly from $\underline{\rho}_{h,i}$ (1) to 1 ($\overline{\rho}_{h,i}$).[32] Finally, we target (iii) the announcement risk premium component of the dollar factor conditional on good economic times, and the spread between bad and good times. We proxy for good (bad) economic times by setting US long-run risks $\mu_{h,t}$ equal to the mean analyst forecast of the US GDP growth at times when the CFNAI activity index is above (below) trend.[33] The empirical counterparts of these moments are the mean excess returns of the DOL factor on announcement days, conditional on positive and, respectively, negative CFNAI values. Table 9 reports the calibrated parameter values, together with model-implied moments and targeted ones.

[Insert Table 9 here.]

Calibrated values of both γ and ψ, 4.08 and 6.58 respectively, are reasonably small and consistent with the rest of the literature. They are greater than one, so that the Epstein and Zin representative agent displays a preference for early resolution of uncertainty. The persistent expected output growth ($b_m = 0.98$) process is consistent with the long-run risk nature of the model. As expected, the announcement risk factor is highly negatively correlated with expected output growth ($\rho_z = -0.73$), as a policy announcement is expected to create more volatility in bad times, when a policy change is likely. The steady-state mean of the "normal" volatility factor x is almost two orders of magnitude smaller than the corresponding figure for the announcement uncertainty factor z: 0.088 opposed to 6.1.[34] This feature is mainly due to matching the pure announcement

[32] It is natural to associate the first stage announcement shock correlation to the home country.

[33] The expression of the currency risk premium conditional on expected output growth is reported in the Appendix.

[34] We have $\mathbb{E}[z_t] = \alpha_z/(1 - b_z)$, $\mathbb{E}[x_t] = \alpha_x/(1 - b_x)$.

currency risk premium components – for the high and the low interest rate portfolios – which is empirically proxied by the difference between average announcement-day and non announcement-day currency returns. Positive announcement currency risk premia imply that announcement jumps in the long-run risks of low interest rate (differential) currencies are negatively correlated with the domestic (US) counterpart ($\overline{\rho}_{h,i} = -0.18$), while the opposite is true for high-interest rate (differential) countries ($\underline{\rho}_{h,i} = 0.3$). Therefore, our model mandates that the revision of long-run beliefs following a policy announcement differs across countries: good news for the home country and low β loading (i.e. high forward discount) countries have typically the opposite interpretation for high β loading (i.e. low forward discount) countries.

The model overshoots the volatility of the home country consumption growth (0.70% vs 0.23%). The model's home country real interest rate accurately matches the volatility of the US rate – both on announcement and non announcement days – and its autocorrelation, but less accurately its mean. Turning the attention to the moments of the cross-section of countries, the model is consistent with the dynamic relation between exchange rates and interest rate differentials, because the theoretical slope of the UIP regression coincides with its empirical counterpart. The model overstates the average volatility of exchange rates (6.3% vs 2.15%), whereas the average correlation between interest rates is exactly reproduced. The risk premia of the dollar factor (0.20% vs 0.18%), and of the carry (HML) factor (0.77% vs 0.71%) are both closely matched. Importantly, announcement currency premia components are almost exactly replicated, both for the high (2.88% vs 2.71%) and the low forward discount portfolios (1.91% vs 1.86%). As for the stock market, the model's prediction for the US equity premium is accurate: 0.24% monthly compared to 0.30% in the data. Finally, as seen in the empirical section, conditioning on different levels of an economic activity index, such as the CFNAI, gives rise to significant variability in currency premia. Our model is consistent with heterogeneity of risk premia across the business cycle, though not to the extent as observed in the data: while the dollar announcement premium in good times is almost matched (0.61% vs 0.83%), the spread between good and bad times is underpredicted (0.841% vs 2.72%).

5.1 Model-Implied Currency Risk Premia

Using the calibrated parameters reported in Table 9, we generate model-implied currency risk premia and compare them to their empirical counterparts. We focus solely on the announcement component of expected excess returns, unconditional relative to all other state variables. Figure 4 plots the model-implied announcement returns for the five interest rate sorted portfolios, as well as the DOL and HML factors.

[Insert Figures 4 and 5 here.]

We note that the announcement premium increases with the interest rate differential, whenever this quantity is positive. When it is close to zero, as for portfolio 2 (see also Table 1), the announcement premium is also small. Since the premia of the first (8.7bps) and the last (13.1bps) interest rate portfolios have been exactly matched in the calibration, it is not surprising that the announcement component of the HML factor premium is as small as in the data (4.4bps). The announcement premium for the DOL portfolio (8.05bps) is also close to the empirical value (see Table 2), even though it is not targeted in the calibration.

Finally, the empirical results in Table 6 indicate that average announcement currency returns tend to be larger in bad economic times than in good times. The model is consistent with this feature as depicted in Figure 5, which shows that model-implied currency announcement premia are decreasing in US expected output growth and display a sizable range of variation from good to bad times.

6 Conclusion

We document the following findings: First, returns to a strategy that is short the US dollar and long the rest of the world are on average an order of magnitude larger on FOMC announcement days compared to non-announcement days. This difference is increasing in the forward discount of the currency and the wedge between announcement and non-announcement mean returns becomes significantly larger during bad economic times. Moreover, using different proxies of uncertainty, we find that announcement currency returns increase when uncertainty is high.

We then study the effect of monetary policy uncertainty on currency returns through the lens of an international general equilibrium model with long-run consumption risk. In our model, expected excess returns compensate investors for the unknown impact of monetary policy uncertainty on long run risk. A calibration exercise shows that the model fits salient moments of equity, exchange, and interest rates. Moreover, it is consistent with the cross-sectional pattern of announcement premia observed in the data and the fact that these premia increase in bad times.

There are several potential avenues of future research. For example, in the spirit of Croce, Kung, Nguyen, and Schmid (2012), we could study the impact of fiscal and monetary policy uncertainty. Exchange rate changes also impact a nation's international investment flows, as well as export and import prices. Within our international framework, we could also explore the impact of announcements on investment flows in the spirit of Colacito, Croce, Ho, and Howard (2014).

7 Figures

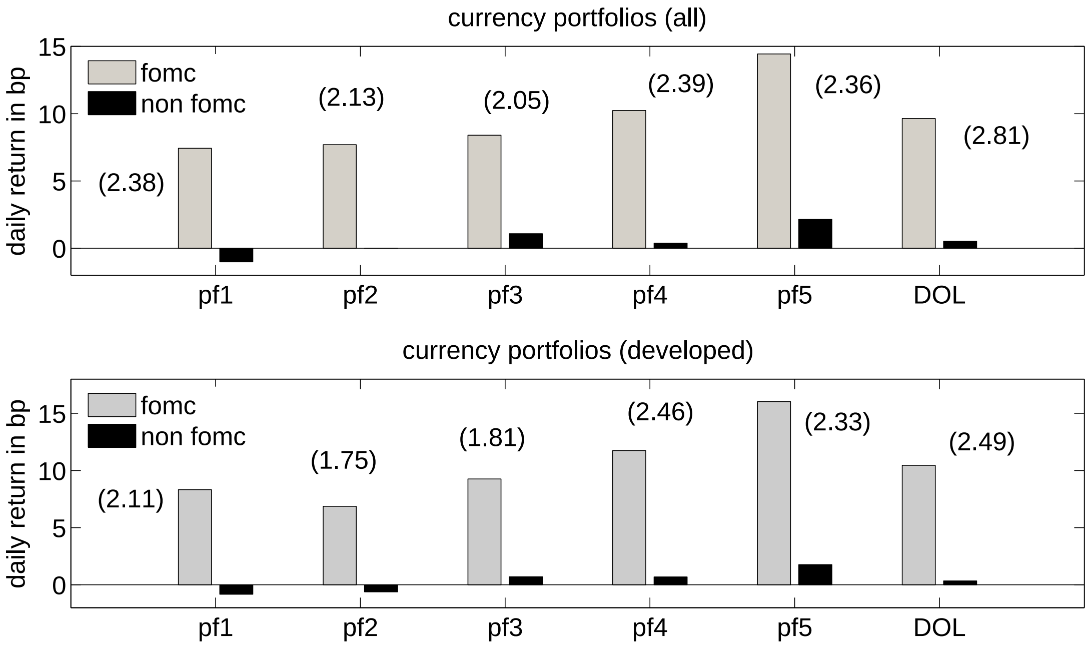

Figure 1. Currency Returns on FOMC and Non-FOMC Days

This figure plots average daily currency returns (in basis points) for portfolios sorted on their interest rate differential. Pf1 is the portfolio with the lowest interest rate differential, while pf5 the portfolio with the highest differential. DOL denotes the average return of the five currency portfolios. The upper panel includes all 35 currencies and the lower panel includes developed currencies only. The numbers in parentheses are t-values of a test of equal means between FOMC and non-FOMC announcement returns. Data used is daily and running from January 1994 to January 2011.

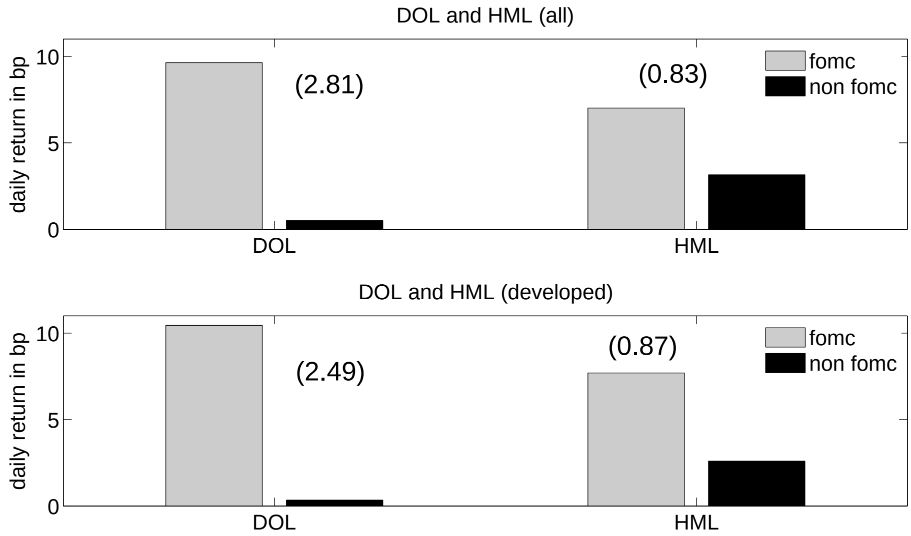

Figure 2. DOL and HML Factor

This figure plots average daily currency returns (in basis points) for the dollar and HML factor. DOL denotes the average return of the five currency portfolios and HML is the portfolio which is long pf5 and short pf1. The upper panel includes all currencies and the lower panel includes developed currencies only. The numbers in parentheses are t-values of a test of equal means between FOMC and non-FOMC announcement returns.. Data used is daily and running from January 1994 to January 2011.

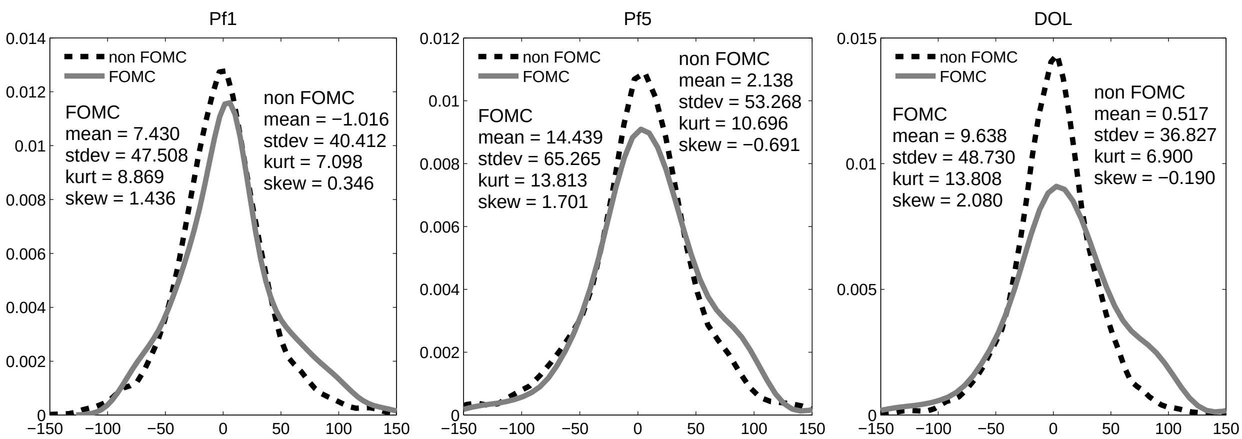

Figure 3. Empirical FX Return Densities on FOMC and Non-FOMC Days

This figure plots empirical densities for returns on interest rate sorted currency portfolios on FOMC and non FOMC announcement days. The left (middle) panel plots the density for low (high) interest rate currencies and the right panel the density for the DOL factor. Data used is daily and running from January 1994 to January 2011.

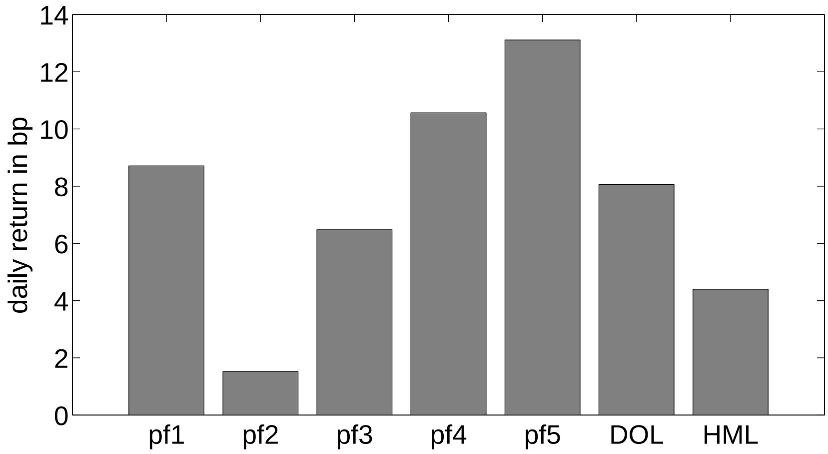

Figure 4. Model-Implied Announcement Currency Risk Premia

Using the parameters from Table 9, we plot daily currency return risk premia conditional on an announcement day. The expression for the currency return risk premium of country i is given by equation (A-32), where the state-variables are evaluated at their unconditional mean. Pf1 (pf5) denotes the portfolio with the lowest (highest) interest rate differential. DOL is the average return of the five portfolios. HML is a portfolio which is long pf5 and short pf1.

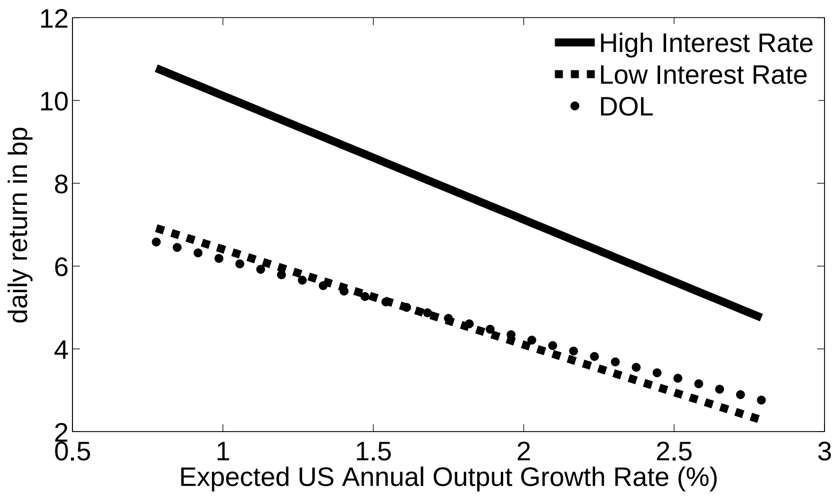

Figure 5. Model-Implied Announcement Currency Risk Premia conditional on US Expected Output Growth

Using the parameters from Table 9, we plot daily currency return risk premia, conditional on a FOMC announcement day, as a function of the US expected output growth rate for low (dashed line), high interest rates portfolios (bold line) and the DOL factor (dotted line).

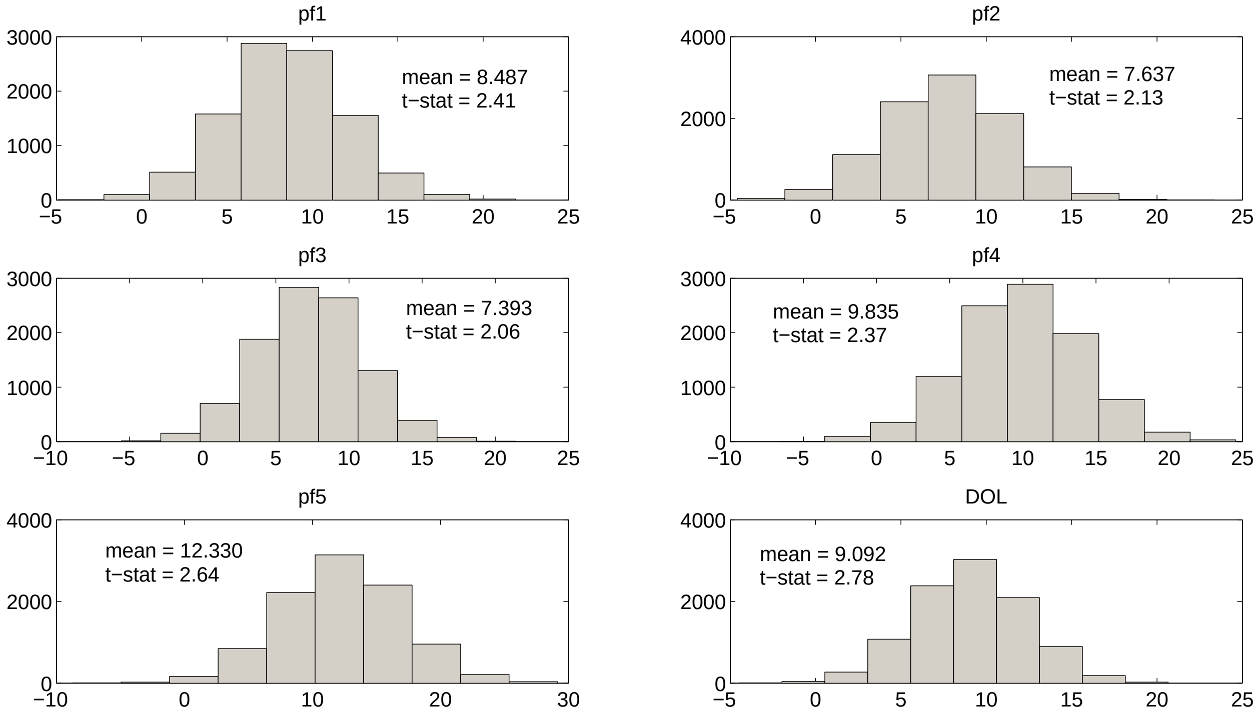

Figure 6. Empirical distribution bootstrapped regression coefficients

This figure plots the empirical distribution for regression coefficients $\hat{\alpha}_1$ in regression (1) for the interest rate sorted portfolios 1 to 5 and the DOL portfolio. The sample period is running from January 1994 to January 2011.

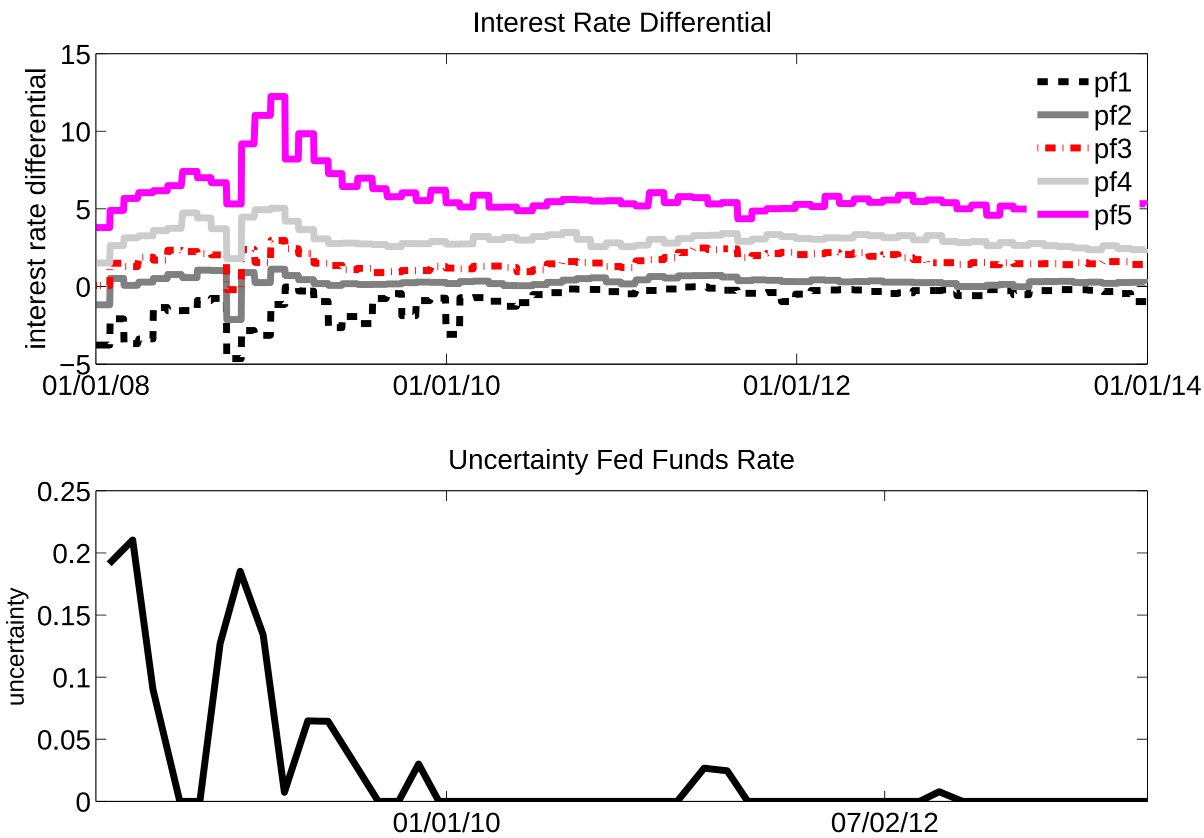

Figure 7. Interest Rate Differential and Uncertainty about Fed Funds Rate

The upper panel plots the average interest rate differential for the 35 different currencies sorted into five different bins according to their interest rate differential. The lower panel plots the cross-sectional standard deviation of the Bloomberg survey forecast on the target Fed Funds rate. Data runs from January 2008 to January 2014.

8 Tables

Table 1
Currency Returns Summary Statistics

This table reports summary statistics of currency portfolios sorted monthly on time $t-1$ forward discounts. Portfolio 1 contains the 20% of all currencies with the lowest forward discounts whereas Portfolio 5 contains currencies with the highest forward discount. DOL denotes the average return of the five portfolios. HML is a portfolio which is long pf5 and short pf1. All returns are excess returns in USD. Returns are daily (in bp) and sampled over the period January 1994 to January 2011. The forward discount is annualized and expressed in percent. The Sharpe ratio is annualized.

	pf1	pf2	pf3	pf4	pf5	DOL	HML
			All Currencies				
mean	-0.737	0.253	1.321	0.697	2.545	0.816	3.282
t-stat	(-1.18)	(0.40)	(2.10)	(0.96)	(3.09)	(1.43)	(4.03)
stdev	40.680	41.357	40.910	47.310	53.725	37.297	53.053
Sharpe ratio	-0.287	0.097	0.513	0.234	0.752	0.347	0.982
forward discount	-2.31	-0.34	1.09	3.70	9.56		
			Developed Currencies				
mean	-0.514	-0.369	0.971	1.056	2.250	0.679	2.764
t-stat	(-0.67)	(-0.49)	(1.16)	(1.33)	(2.09)	(0.95)	(2.68)
stdev	49.780	49.163	54.388	51.573	70.153	46.662	67.071
Sharpe ratio	-0.164	-0.119	0.283	0.325	0.509	0.231	0.654
forward discount	-2.61	-0.74	-0.02	1.06	3.56		

Table 2
Currency Return Summary Statistics

This table reports summary statistics of currency returns on announcement and non-announcement days. Portfolios are sorted according to their interest rate differential. Pf1 (pf5) has the lowest (highest) interest rate differential. DOL denotes the average return of the five portfolios. Announcement days are when the FOMC releases its interest rate decisions. Diff mean indicates the difference in average returns between FOMC and non FOMC returns, and the corresponding t-statistic is presented in parentheses below. The sample covers January 1994 to January 2011. All numbers are expressed in daily returns (in bp) except for Sharpe ratios, which are annualized taking into account the annual frequency of FOMC announcements (8/252). Thus FOMC SR = [daily mean return FOMC/daily std returns FOMC]$\sqrt{8}$, and NON FOMC SR = [daily mean return NON FOMC/daily std returns NON FOMC]$\sqrt{244}$.

	pf1	pf2	pf3	pf4	pf5	DOL
All currencies (FOMC)						
mean	7.430	7.693	8.396	10.232	14.439	9.638
t-stat	(1.82)	(1.57)	(1.71)	(2.60)	(2.58)	(2.31)
stdev	47.508	57.091	57.243	45.894	65.165	48.730
Sharpe ratio	0.442	0.381	0.415	0.631	0.627	0.559
All Currencies (non FOMC)						
mean	-1.016	0.003	1.083	0.379	2.138	0.517
t-stat	(-2.21)	(0.01)	(2.37)	(0.70)	(3.53)	(1.24)
stdev	40.412	40.714	40.240	47.323	53.268	36.827
Sharpe ratio	-0.393	0.001	0.421	0.125	0.627	0.219
diff mean	8.447	7.690	7.313	9.853	12.301	9.121
t-stat	(2.38)	(2.13)	(2.05)	(2.39)	(2.63)	(2.81)
Developed Currencies (FOMC)						
mean	8.330	6.872	9.256	11.753	16.027	10.448
t-stat	(1.52)	(1.38)	(1.44)	(1.97)	(2.07)	(1.89)
stdev	63.807	58.142	74.771	69.476	90.414	64.364
Sharpe ratio	0.369	0.334	0.350	0.478	0.501	0.459
Developed Currencies (non FOMC)						
mean	-0.829	-0.614	0.702	0.687	1.768	0.343
t-stat	(-1.48)	(-1.11)	(1.15)	(1.19)	(2.24)	(0.66)
stdev	49.244	48.823	53.566	50.848	69.360	45.936
Sharpe ratio	-0.263	-0.197	0.205	0.211	0.398	0.117
diff mean	9.159	7.486	8.554	11.066	14.259	10.105
t-stat	(2.11)	(1.75)	(1.81)	(2.46)	(2.33)	(2.49)

Table 3
DOL and HML Summary Statistics

This table reports summary statistics of the dollar (DOL) and HML factor conditional on FOMC announcement and non-announcement days. DOL denotes the average return of the five currency portfolios and HML denotes a long-short portfolio that is long in portfolio 5 and short in portfolio 1. Numbers are in daily basis points except for the Sharpe ratio which is annualized. The sample covers January 1994 to January 2011.

All Currencies				
	FOMC		non FOMC	
	DOL	HML	DOL	HML
mean	9.638	7.009	0.517	3.154
t-stat	(2.31)	(1.51)	(1.24)	(5.23)
stdev	48.730	54.017	36.827	53.017
Sharpe ratio	0.559	0.367	0.219	0.929
Developed Currencies				
	FOMC		non FOMC	
	DOL	HML	DOL	HML
mean	10.448	7.697	0.343	2.596
t-stat	(1.89)	(1.20)	(0.66)	(3.42)
stdev	64.364	74.656	45.936	66.801
Sharpe ratio	0.459	0.292	0.117	0.607

Table 4
Currency Return Summary Statistics Winsorized Data

This table reports summary statistics of currency returns on announcement and non-announcement days for a winsorized sample where we delete outliers at the bottom and top 1%. Announcement days are when the FOMC releases its interest rate decisions. The sample covers January 1994 to January 2011.

	pf1	pf2	pf3	pf4	pf5	DOL
FOMC						
mean	7.494	8.227	8.904	10.762	14.848	10.047
t-stat	(1.81)	(1.67)	(1.80)	(2.72)	(2.62)	(2.38)
stdev	47.861	57.027	57.319	45.739	65.482	48.826
skew	1.422	2.393	2.353	1.138	1.688	2.087
kurt	8.734	15.847	17.876	9.025	13.707	13.821
N	134	134	134	134	134	134
non FOMC						
mean	-1.092	-0.110	0.961	0.307	2.064	0.426
t-stat	(-1.70)	(-0.17)	(1.51)	(0.41)	(2.44)	(0.73)
stdev	40.745	40.810	40.379	47.607	53.705	37.039
skew	0.350	0.020	-0.242	-0.450	-0.684	-0.184
kurt	7.009	6.772	9.253	12.971	10.558	6.868
N	4026	4026	4026	4026	4026	4026

Table 5

Currency Return Summary Statistics since 1980

This table reports summary statistics of currency returns on announcement and non-announcement days. Portfolios are sorted according to their interest rate differential. Pf1 (pf5) has the lowest (highest) interest rate differential. DOL denotes the average return of the five portfolios. Announcement days are when the FOMC releases its interest rate decisions. Diff mean indicates the difference in average returns between FOMC and non FOMC returns, and the corresponding t-statistic is presented in parentheses below. The sample covers January 1980 to January 2011. All numbers are expressed in daily returns (in bp) except for Sharpe ratios which are annualized.

	pf1	pf2	pf3	pf4	pf5	DOL
All currencies (FOMC)						
mean	5.213	3.818	5.352	6.110	9.889	6.077
t-stat	(1.87)	(1.15)	(1.57)	(1.96)	(2.67)	(2.07)
stdev	44.168	52.416	54.057	49.197	58.652	46.483
Sharpe ratio	0.334	0.206	0.280	0.351	0.477	0.370
All Currencies (non FOMC)						
mean	-1.068	-0.361	0.767	-0.044	1.546	0.168
t-stat	(-2.16)	(-0.69)	(1.46)	(-0.07)	(2.17)	(0.34)
stdev	43.406	46.186	46.155	53.410	62.713	42.786
Sharpe ratio	-0.384	-0.122	0.259	-0.013	0.385	0.061
diff mean	6.282	4.179	4.585	6.154	8.344	5.909
t-stat	(2.25)	(1.40)	(1.54)	(1.80)	(2.07)	(2.14)
Developed Currencies (FOMC)						
mean	5.344	4.693	6.323	7.299	10.537	6.839
t-stat	(1.51)	(1.40)	(1.49)	(1.80)	(2.16)	(1.88)
stdev	56.131	52.871	66.925	64.044	77.249	57.407
Sharpe ratio	0.269	0.251	0.267	0.322	0.386	0.337
Developed Currencies (non FOMC)						
mean	-1.056	-0.554	0.311	0.501	1.265	0.094
t-stat	(-1.80)	(-0.94)	(0.49)	(0.81)	(1.67)	(0.17)
stdev	51.438	51.611	55.983	54.607	66.632	49.354
Sharpe ratio	-0.321	-0.168	0.087	0.143	0.297	0.030
diff mean	6.400	5.247	6.012	6.797	9.271	6.745
t-stat	(1.93)	(1.58)	(1.66)	(1.92)	(2.15)	(2.11)

Table 6
Regression Currency Portfolios Announcement Dummy

This table reports estimated coefficients from regressing currency portfolios sorted on their interest rate differential on an announcement dummy which takes the value of one on an announcement day and zero otherwise. In the other regressions we interact the announcement dummy with the Chicago Fed National Activity Index (CFNAI), a policy surprise component as in Bernanke and Kuttner (2005), VIX, TIV, a disagreement proxy from forecasts on the target Fed Funds rate (DiB), and the economic policy index (EPU) by Baker, Bloom, and Davis (2013). t-statistics are calculated using Newey and West standard errors and are given in parentheses. Data is daily and runs from January 1994 to January 2011.

	pf1	pf2	pf3	pf4	pf5	DOL
constant	-1.016	0.003	1.083	0.379	2.138	0.517
t-stat	(-1.61)	(0.00)	(1.73)	(0.51)	(2.57)	(0.90)
Announcement	8.447	7.690	7.313	9.853	12.301	9.121
t-stat	(2.06)	(1.56)	(1.48)	(2.47)	(2.19)	(2.17)
R^2 (in %)	0.11	0.08	0.08	0.11	0.14	0.16
constant	-0.916	0.027	1.049	0.492	2.218	0.574
t-stat	(-1.46)	(0.04)	(1.69)	(0.68)	(2.69)	(1.01)
Announcement × CFNAI	11.293	14.729	17.805	13.449	20.845	15.624
t-stat	(1.68)	(1.67)	(1.97)	(1.81)	(2.23)	(2.06)
R^2 (in %)	0.09	0.16	0.26	0.10	0.20	0.24
constant	-0.745	0.248	1.316	0.693	2.530	0.808
t-stat	(-1.19)	(0.39)	(2.10)	(0.95)	(3.07)	(1.41)
Announcement × Surprise	22.591	-32.820	-60.614	-62.585	-58.387	-38.363
t-stat	(0.79)	(-1.13)	(-1.99)	(-1.90)	(-1.12)	(-1.28)
R^2 (in %)	-0.01	0.00	0.04	0.03	0.01	0.01
constant	-1.106	-0.273	0.904	0.334	1.869	0.346
t-stat	(-1.77)	(-0.42)	(1.45)	(0.46)	(2.22)	(0.60)
Announcement × VIX	1.549	2.243	1.777	1.550	2.853	1.995
t-stat	(3.75)	(2.35)	(2.77)	(2.45)	(2.73)	(2.97)
R^2 (in %)	0.52	1.07	0.68	0.38	1.03	1.04
constant	-1.163	-0.183	0.872	0.344	1.995	0.373
t-stat	(-1.83)	(-0.28)	(1.35)	(0.46)	(2.38)	(0.64)
Announcement × TIV	7.653	7.923	8.183	6.434	9.868	8.012
t-stat	(2.31)	(1.60)	(1.62)	(1.65)	(2.09)	(1.95)
R^2 (in %)	0.42	0.43	0.47	0.21	0.40	0.55
constant	-0.971	-0.066	1.038	0.397	2.165	0.512
t-stat	(-1.55)	(-0.10)	(1.66)	(0.54)	(2.62)	(0.90)
Announcement × EPU	0.056	0.078	0.069	0.073	0.091	0.073
t-stat	(2.15)	(2.32)	(2.27)	(2.69)	(2.34)	(2.58)
R^2 (in %)	0.12	0.24	0.19	0.16	0.19	0.27
constant	-0.978	0.017	1.050	0.522	2.293	0.581
t-stat	(-1.57)	(0.03)	(1.68)	(0.72)	(2.80)	(1.02)
Announcement × DiB	3.493	3.489	4.030	2.597	3.593	3.440
t-stat	(2.76)	(1.69)	(1.77)	(1.76)	(2.04)	(2.04)
R^2 (in %)	0.78	0.75	1.03	0.30	0.46	0.90

Table 7
Regression Country-Specific Analysts' GDP Forecasts

This table reports estimated coefficients for the following regression:

$$\widehat{\mu}_{i,t+1} = \alpha_0 + \alpha_1 \widehat{\mu}_{h,t+1} + \alpha_2 (r_{i,t} - r_{h,t}) \widehat{\mu}_{h,t+1} + \epsilon_{i,t+1}$$

where $\widehat{\mu}_{i,t+1}$ is analysts' consensus forecast of country i's GDP growth (h being the US) and $(r_{i,t} - r_{h,t})$ is the foreign-domestic interest rate differential. Countries are: Australia (AUS), Canada (CAN), Switzerland (CH), Germany (GER), United Kingdom (UK), Japan (JP), Norway (NOR), New Zealand (NZ), and Sweden (SWE). Data is monthly and runs from January 1994 to January 2011.

	AUS	CAN	CH	GER	UK	JP	NOR	NZ	SWE
α_1	0.521	0.744	0.388	0.621	0.808	0.752	0.414	0.654	0.715
t-stat	(-17.09)	(26.91)	(9.41)	(13.28)	(17.59)	(10.73)	(7.69)	(17.10)	(18.41)
α_2	-12.020	-35.860	-30.790	-87.690	-12.520	-8.180	-60.890	-29.820	-61.420
t-stat	(-1.74)	(-3.32)	(-2.58)	(-6.41)	(-0.80)	(-0.62)	(-4.57)	(-3.50)	(-6.68)
R^2 (in %)	63.51	78.32	53.53	54.39	62.61	42.84	34.59	62.96	69.30

Table 8
Calibration: First Stage

This table reports target moments in the first stage of the calibration, where all countries share the same loading β. All model-implied moments (second column) are unconditional relative to all state variables, except to τ where specified. "Average" refers to an average across countries. We use the following notation: $Var[\mu_t|\tau=1] = \frac{(\sigma_m^2+\beta^2)\alpha_x}{(1-b_x)(1-b_m^2)} + \frac{\beta^2\alpha_z}{(1-b_z)(1-b_m^2)}$, $Var[\mu_t|\tau \neq 1] = \frac{(\sigma_m^2+\beta^2)\alpha_x}{(1-b_x)(1-b_m^2)}$, $Var[z_t] = \frac{\alpha_x\sigma_z^2}{(1-b_x)(1-b_z^2)}$, $Cov[\mu_t, y_t] = \frac{\alpha_x\sigma_z\beta\rho_z}{(1-b_x)(1-b_z b_m)}$, $Var[x_t] = \frac{\alpha_x\sigma_x^2}{(1-b_x)(1-b_x^2)}$. $\mathbb{E}[r_t|\tau=1]$ ($\mathbb{E}[r_t|\tau\neq 1]$) is the steady-state expected US short rate conditional on a (non) announcement period (reported below). The variances $Var[r_t|\tau=1]$ and $Var[r_t|\tau\neq 1]$ have a similar interpretation (also reported below), while the fully unconditional expected short rate is $\mathbb{E}[r_t] = \frac{\mathbb{E}[r_t|\tau=1]}{\overline{A}} + \frac{\overline{A}-1}{\overline{A}}\mathbb{E}[r_t|\tau\neq 1]$. Moreover, $\overline{\Phi} = \frac{1}{\overline{A}}\sum_{\tau=1}^{\overline{A}}\left[(\theta-1)(\rho B_1\beta)^2\mathbf{1}_{\tau=1} + \frac{1}{2}(\theta-1)\rho^2 f(\tau)\right]$. Theoretical moments are daily, while empirical moments are monthly equivalents.

Moment	Model	Data
Expected Consumption Growth US	$\overline{\mu}\times 22$	0.254 %
Volatility Consumption Growth US	$\left(\sigma_d^2 + \frac{1}{\overline{A}}Var[\mu_t|\tau=1] + \frac{\overline{A}-1}{\overline{A}}Var[\mu_t|\tau\neq 1]\right)^{1/2}\times\sqrt{22}$	0.232%
Average UIP Regression Slope	$\dfrac{\frac{2\sigma_m^2\alpha_x}{\psi^2(1-b_x)(1-b_m^2)} + \frac{1}{4}(\theta-1)^2\theta\rho^4(B_1^2\sigma_m^2+B_4^2\sigma_x^2)^2\frac{2\sigma_x^2\alpha_x}{\psi^2(1-b_x)(1-b_x^2)}}{\frac{2\sigma_m^2\alpha_x}{\psi^2(1-b_x)(1-b_m^2)} + \frac{1}{4}(\theta-1)^2\rho^4(B_1^2\sigma_m^2+B_4^2\sigma_x^2)^2\frac{2\sigma_x^2\alpha_x}{\psi^2(1-b_x)(1-b_x^2)}}$	-2.46
Expected Real Short Rate US conditional (Ann.)	$\Big(-\log\delta + \frac{\overline{\mu}}{\psi} + \frac{\sigma_d^2(1-\gamma(\gamma+1))}{2\psi} + \frac{1}{2}(\theta-1)\rho^2B_1^2\beta^2\frac{\alpha_z}{1-b_z} + \frac{1}{2}(\theta-1)\rho^2\left(f(1)+B_1^2\sigma_m^2+B_4^2\sigma_x^2\right)\frac{\alpha_x}{1-b_x}\Big)\times 22$	0.076%
Expected Real Short Rate US conditional (No Ann.)	$\Big(-\log\delta + \frac{\overline{\mu}}{\psi} + \frac{\sigma_d^2(1-\gamma(\gamma+1))}{2\psi} + \frac{1}{2}(\theta-1)\rho^2\left(\frac{1}{\overline{A}-1}\sum_{\tau=2}^{\overline{A}}f(\tau)+B_1^2\sigma_m^2+B_4^2\sigma_x^2\right)\frac{\alpha_x}{1-b_x}\Big)\times 22$	0.075%
Volatility Real Short Rate US (Ann.)	$\Big(\frac{Var[\mu_t|\tau=1]}{\psi^2} + \frac{1}{4}(\theta-1)^2(\rho B_1\beta)^4Var[y_t] + \frac{\theta-1}{\psi}(\rho B_1\beta)^2Cov[\mu_t,y_t] + \frac{1}{4}(\theta-1)^2\rho^4\left(f(1)+B_1^2\sigma_m^2+B_4^2\sigma_x^2\right)^2Var[x_t]\Big)^{1/2}\times\sqrt{22}$	0.241%
Volatility Real Short Rate US (No Ann.)	$\Big(\frac{Var[\mu_t|\tau\neq 1]}{\psi^2} + \frac{1}{4}(\theta-1)^2\rho^4\left[\frac{1}{\overline{A}-1}\sum_{\tau=2}^{\overline{A}}\left(f(\tau)+B_1^2\sigma_m^2+B_4^2\sigma_x^2\right)^2\right]Var[x_t] + \frac{1}{4}(\theta-1)^2\rho^4\frac{\alpha_x^2}{(1-b_x)^2}\frac{1}{\overline{A}-1}\sum_{\tau=2}^{\overline{A}}\left(f(\tau)-\frac{1}{\overline{A}-1}\sum_{j=2}^{\overline{A}}f(j)\right)^2\Big)^{1/2}\times\sqrt{22}$	0.235%

Table 8 (cont'd)

Moment	Model	Data
Autocorrelation Real Short Rate US	$\Big\{ \frac{b_m}{\psi^2} \left(\frac{1}{\overline{A}} Var[\mu_t \mid \tau = 1] + \frac{\overline{A}-1}{\overline{A}} Var[\mu_t \mid \tau \neq 1] \right) + \frac{1}{4}(\theta - 1)^2 \rho^2 \left(B_1^2 \sigma_m^2 + B_4^2 \sigma_x^2 \right)^2 Var[x_t] + \frac{(\theta-1)(\rho B_1 \beta)^2}{2\psi} Cov(\mu_t, y_t) \times \times \frac{1}{\overline{A}} \sum_{\tau=1}^{\overline{A}} \left(b_z \mathbf{1}_{\tau=1} + b_m \mathbf{1}_{\tau-1=1} \right) + \frac{(\theta-1)^2 \rho^4 b_x}{4} \frac{1}{\overline{A}} \sum_{\tau=1}^{\overline{A}} f(\tau) f(\tau - 1) + \frac{1}{\overline{A}} \sum_{\tau=1}^{\overline{A}} \left[(\theta - 1)(\rho B_1 \beta)^2 \mathbf{1}_{\tau=1} + \frac{1}{2}(\theta - 1)\rho^2 f(\tau) \right] \left[(\theta - 1)(\rho B_1 \beta)^2 \mathbf{1}_{\tau-1=1} + \frac{1}{2}(\theta - 1)\rho^2 f(\tau - 1) \right] - \overline{\Phi}^2 \Big\} / \Big[\frac{1}{\overline{A}} Var[r_t \mid \tau = 1] + \frac{\overline{A}-1}{\overline{A}} Var[r_t \mid \tau \neq 1] + \frac{1}{\overline{A}} \left(\mathbb{E}[r_t \mid \tau = 1] - \mathbb{E}[r_t] \right)^2 + \frac{\overline{A}-1}{\overline{A}} \left(\mathbb{E}[r_t \mid \tau \neq 1] - \mathbb{E}[r_t] \right)^2 \Big]$	0.98
Average Volatility FX Rates vs US	$\left(\frac{2\sigma_m^2 \alpha_x}{\psi^2 (1-b_x)(1-b_m^2)} + \frac{1}{4}(\theta - 1)^2 \theta \rho^4 (B_1^2 \sigma_m^2 + B_4^2 \sigma_x^2)^2 \frac{2\sigma_x^2 \alpha_x}{\psi^2 (1-b_x)(1-b_x^2)} + 2\gamma^2 \sigma_d^2 + (\theta - 1)^2 \rho^2 2 \left(B_1^2 \sigma_m^2 + B_4^2 \sigma_x^2 \right) \frac{\alpha_x}{1-b_x} \right)^{1/2} \sqrt{22}$	2.158%
Average Correlation Short Rates	$\Big(\frac{1}{\psi^2 (1-b_m^2)} \left(\beta^2 \frac{\alpha_x}{1-b_x} + \frac{1}{\overline{A}} \beta^2 \frac{\alpha_z}{1-b_z} \right) + \frac{(\theta-1)\rho^2 B_1^2 \beta^2}{2\overline{A}\psi} Cov[\mu_t, y_t] + \frac{(\theta-1)^2 \rho^4 B_1^4 \beta^4}{4\overline{A}} Var[y_t] + \frac{(\theta-1)^2 \rho^4}{4} \frac{\sum_{\tau=1}^{\overline{A}} f(\tau)^2}{\overline{A}} Var[x_t] + \frac{1}{\overline{A}} \sum_{\tau=1}^{\overline{A}} \left[(\theta - 1)(\rho B_1 \beta)^2 \mathbf{1}_{\tau=1} + \frac{1}{2}(\theta - 1)\rho^2 f(\tau) \right]^2 - \overline{\Phi}^2 \Big) / / \Big[\frac{1}{\overline{A}} Var[r_t \mid \tau = 1] + \frac{\overline{A}-1}{\overline{A}} Var[r_t \mid \tau \neq 1] + \frac{1}{\overline{A}} \left(\mathbb{E}[r_t \mid \tau = 1] - \mathbb{E}[r_t] \right)^2 + \frac{\overline{A}-1}{\overline{A}} \left(\mathbb{E}[r_t \mid \tau \neq 1] - \mathbb{E}[r_t] \right)^2 \Big]$	62.2%
Average Currency Return (Dollar) Risk Premium	$\left(\gamma \sigma_d^2 + (\theta - 1)^2 \rho^2 (B_1^2 \sigma_m^2 + B_4^2 \sigma_x^2) \frac{\alpha_x}{1-b_x} \right) \times 22$	0.18%
US Equity risk premium (Ann.)	$\Big(- (\theta - 1) \frac{A_1 B_1 \rho^2 \beta^2 \alpha_z}{1-b_z} - (\theta - 1) \frac{A_1 B_1 \rho^2 \sigma_m^2 \alpha_x}{1-b_x} - (\theta - 1) \frac{A_1 B_1 \rho^2 \beta^2 \alpha_x}{1-b_x} - (\theta - 1) \frac{A_1 B_2(1) \rho^2 \beta \rho_z \sigma_z \alpha_x}{1-b_x} - A_2(1)(\theta - 1)\rho^2 B_1 \beta \sigma_z \rho_z \frac{\alpha_x}{1-b_x} - \rho A_2(1)(\theta - 1)\rho B_2(1) \sigma_z^2 \frac{\alpha_x}{1-b_x} - \rho A_3(1)(\theta - 1)\rho B_3(1) \sigma_x^2 \frac{\alpha_x}{1-b_x} - (\theta - 1) A_4 B_4 \rho^2 \sigma_x^2 \frac{\alpha_x}{1-b_x} \Big) \times 22$	3.73%
US Equity risk premium (No Ann.)	$\Big(- (\theta - 1) \frac{A_1 B_1 \rho^2 \sigma_m^2 \alpha_x}{1-b_x} - (\theta - 1) \frac{A_1 B_1 \rho^2 \beta^2 \alpha_x}{1-b_x} - (\theta - 1) A_4 B_4 \rho^2 \sigma_x^2 \frac{\alpha_x}{1-b_x} - \frac{1}{\overline{A}-1} \sum_{\tau=2}^{\overline{A}} \Big[(\theta - 1) \frac{A_1 B_2(\tau) \rho^2 \beta \rho_z \sigma_z \alpha_x}{1-b_x} - A_2(\tau)(\theta - 1)\rho^2 B_1 \beta \sigma_z \rho_z \frac{\alpha_x}{1-b_x} - \rho A_2(\tau)(\theta - 1)\rho B_2(\tau) \sigma_z^2 \frac{\alpha_x}{1-b_x} - \rho A_3(\tau)(\theta - 1)\rho B_3(\tau) \sigma_x^2 \frac{\alpha_x}{1-b_x} \Big] \Big) \times 22$	0.19%

Table 9
Calibration: Second Stage

Panel A summarizes the final set of calibrated parameters. The country-specific loadings (β_i) of the policy effect on growth are linearly spaced in the interval $[\underline{\beta}, \overline{\beta}]$. For countries with $\beta_i < \beta_h$ ($\beta_i > \beta_h$), the announcement shock correlation increases (decreases) linearly from $\underline{\rho}_{h,i}$ (1) to 1 ($\overline{\rho}_{h,i}$) when β_i increases from $\underline{\beta}$ (β_h) to β_h ($\overline{\beta}$). Panel B reports theoretical moments evaluated at calibrated parameters and targeted (empirical) counterparts.

Panel A: Calibrated Parameters

δ	γ	ψ	$\overline{\mu}$	β_h	$\overline{\beta}$	$\underline{\beta}$	b_m	σ_m	α_z
0.999	4.08	6.58	0.00015	0.00028	0.00057	0.00001	0.987	0.00061	4.19

b_z	σ_z	α_x	β_x	σ_x	ρ_z	σ_d	$\overline{\rho}_{h,i}$	$\underline{\rho}_{h,i}$
0.313	0.122	0.00003	0.9996	0.239	-0.73	0.00012	-0.176	0.3

Panel B: 1st and 2nd Stage Moments

Moment	Model	Data
Expected Consumption Growth US	0.254%	0.254 %
Volatility Consumption Growth US	0.704%	0.231%
Average UIP Regression Slope	-2.46	-2.46
Expected Real Short Rate US (Ann.)	-0.092%	0.076%
Expected Real Short Rate US (No Ann.)	0.22%	0.075%
Volatility Real Short Rate US (Ann.)	0.35%	0.241%
Volatility Real Short Rate US (No Ann.)	0.171%	0.235%
Autocorrelation Real Short Rate US	0.99	0.98
Average Volatility FX Rates vs US	6.38%	2.15%
Average Correlation btw US and Other Short Rates	0.620	0.622
Average Dollar Risk Premium	0.20%	0.18%
US Equity risk premium	0.24%	0.30%
Announcement Currency Premium (high int. rate)	2.88%	2.71%
Announcement Currency Premium (low int. rate)	1.91%	1.86%
Average HML premium	0.77%	0.71%
Announcement Dollar Premium (good econ. times)	0.607%	0.837%
Announcement Dollar Premium (Δ bad-good econ. times)	0.841%	2.72%

Appendix A Proofs and derivations

Proof of Proposition 1

For $t \in [t_a, t_a + \overline{A}]$—where t_a is any FOMC announcement time and $\overline{A}$ the time length between announcements—the growth component controlled by monetary policy, μ_t, is an unobservable constant. Let $\boldsymbol{\mu_{i,t}} = (\overline{\mu}^i, \mu_t)'$, and

$$\widehat{\boldsymbol{\mu_{i,t}}} = \mathbb{E}\left[\boldsymbol{\mu_{i,t}}|\mathcal{I}_t\right], \quad \text{(A-1)}$$
$$\widehat{\boldsymbol{\sigma^2_{i,t}}} = \mathbb{E}\left[(\boldsymbol{\mu_{i,t}} - \widehat{\boldsymbol{\mu_{i,t}}})(\boldsymbol{\mu_{i,t}} - \widehat{\boldsymbol{\mu_{i,t}}})'|\mathcal{I}_t\right] \quad \text{(A-2)}$$

denote the vector of posterior means of output growth and the matrix of estimation errors, respectively, where $\mathcal{I}_t$ is the information set that includes all past realizations of output growth Δy_i and the signal s, until time t . Also let

$$B_i = \begin{pmatrix} 1 & \beta_i \\ 1 & 0 \end{pmatrix} \quad \text{and} \quad \Sigma_t^i = \begin{pmatrix} \sigma_d^2 x_{i,t} & 0 \\ 0 & \sigma_s^2 x_t \end{pmatrix}.$$

The bivariate random process $(\Delta y_{i,t+1}, s_{t+i})$ has a Gaussian one-step ahead conditional density function, so that we can apply the Kalman filtering procedure in discrete time. In particular, we can write the system dynamics in innovation form:

$$(\Delta y_{i,t+1}, s_{t+i})' = B\widehat{\boldsymbol{\mu_{i,t}}} + \widehat{\boldsymbol{\epsilon_{i,t+1}}}, \quad \text{(A-3)}$$
$$\widehat{\boldsymbol{\mu_{i,t+1}}} = \widehat{\boldsymbol{\mu_{i,t}}} + K_{i,t}\widehat{\boldsymbol{\epsilon_{i,t+1}}}, \quad \text{(A-4)}$$
$$K_{i,t+1} = \widehat{\boldsymbol{\sigma^2_{i,t}}} B_i' \left(B_i \widehat{\boldsymbol{\sigma^2_{i,t}}} B_i' + \Sigma_t^i\right)^{-1}, \quad \text{(A-5)}$$
$$\widehat{\boldsymbol{\sigma^2_{i,t+1}}} = \widehat{\boldsymbol{\sigma^2_{i,t}}} - \widehat{\boldsymbol{\sigma^2_{i,t}}} B_i' \left(B_i \widehat{\boldsymbol{\sigma^2_{i,t}}} B_i' + \Sigma_t^i\right)^{-1} B_i \widehat{\boldsymbol{\sigma^2_{i,t}}}, \quad \text{(A-6)}$$

where $K_{i,t+1}$ is the weight given to current information in the updating rule, the so-called Kalman gain. We follow Bansal and Shalistovich (2010) and simplify the model by assuming a constant Kalman gain matrix $K_{i,t+1} = K_i$ which is set equal to its steady-state value.[35] After this assumption the posterior variances decrease deterministically in time, because (A-6) can be rewritten as

$$\widehat{\boldsymbol{\sigma^2_{i,t+1}}} = \widehat{\boldsymbol{\sigma^2_{i,t}}}(I_2 - B_i' K_{i,t+1}') = \widehat{\boldsymbol{\sigma^2_{i,t}}}(I_2 - B_i' K_i'),$$

which implies

$$\widehat{\boldsymbol{\sigma^2_{i,t}}} = \widehat{\boldsymbol{\sigma^2_{i,0}}}(I_2 - B_i' K_i')^t.$$

Since $Var_t\left[\widehat{\boldsymbol{\epsilon_{i,t+1}}}\right] = B_i \widehat{\boldsymbol{\sigma^2_{i,t}}} B_i' + \Sigma_t^i$, defining $\widetilde{\mu}_{i,t} = \mathbb{E}\left[\overline{\mu}^i|\mathcal{I}_t\right] + \beta_i \mathbb{E}\left[\mu_{i,t}|\mathcal{I}_t\right] = (1,0)B_i\widehat{\boldsymbol{\mu_{i,t}}}$ we have, by virtue of (A-4)

$$\widetilde{\mu}_{i,t+1} = \widetilde{\mu}_{i,t} + (1,0)B_i K_i \sqrt{B_i \hat{\sigma}^2_{i,t} B_i' + \Sigma_t^i}\,(\epsilon_{i,t+1}, \epsilon_{t+1})', \quad \text{(A-7)}$$
$$= \widetilde{\mu}_{i,t} + \sqrt{\overline{\sigma}_1 + \sigma_d^2 x_{i,t}}\,\epsilon_{i,t+1} + \beta_i \sqrt{\overline{\sigma}_2 + \sigma_s^2 x_t}\,\epsilon_{t+1}, \quad \text{(A-8)}$$

[35] In this way agents assign constant weights to news and prior estimates in their updating rule:

$$\widehat{\boldsymbol{\mu}_{i,t+1}} = (I_2 - K_i B_i)\widehat{\boldsymbol{\mu}_{i,t}} + K_i(\Delta y_{i,t+1}, s_{t+1})'$$

where ϵ and ϵ_i are standard Gaussian innovations. In expression (A-8) we have assumed that $K_i = I_2$ and we have ignored the off-the-main-diagonal elements of the deterministic matrix $B_i \widehat{\boldsymbol{\sigma^2_{i,t}}} B_i'$, and defined the main diagonal elements as the constants $\overline{\sigma}_1$ and $\overline{\sigma}_2$.

Equilibrium price-consumption ratio

Let $pc_{i,t}$ denote the log price-dividend ratio of the claim to the aggregate output of country i, y_i. The return on this claim can be log-linearized as in Campbell and Shiller (1988):

$$r^y_{i,t+1} = k + \Delta y_{i,t+1} + \rho_i \, pc_{i,t+1} - pc_{i,t}, \tag{A-9}$$

where $\rho_i = [1+\exp(-\overline{pc_i})]^{-1} < 1$ is determined endogenously, as it depends on $\overline{pc_i}$, the long-run mean of pc_i. We conjecture the following linear expression for $pc_{i,t}$:

$$pc_{i,t} = B_0(i,t) + B_1\mu_{i,t} + B_2(i,t)z_t + B_3(i,t)x_t + B_4 x_{i,t},$$

where we have emphasized the dependence of deterministic functions B on time (t) and on the country-specific loading β_i (i). Plugging this expression into the Euler equation for $r^y_{i,t+1}$, or

$$\mathbb{E}_t\left[\exp(m_{i,t+1} + r^y_{i,t+1})\right] = 1, \tag{A-10}$$

where

$$m_{i,t+1} = \theta \log\delta - \frac{\theta}{\psi}\Delta y_{i,t+1} + (\theta-1) r^y_{i,t+1}, \tag{A-11}$$

and computing conditional expectations in (A-10) , yields the following restrictions that coefficients B must satisfy:

$$\begin{aligned}
&B_1 = \tfrac{1-\frac{1}{\psi}}{1-\rho_i b_m}, \\
&B_2(i,t) = \tfrac{1}{2}\theta\rho_i^2 B_1^2\beta_i^2\left[\textstyle\sum_{j=1}^{\infty}(\rho_i b_z)^j A_{t+j}\right], \qquad \rho_i b_z < 1 \\
&\rho b_x B_3(i,t+1) - B_3(i,t) + \tfrac{1}{2}\theta\rho_i^2\left(B_1^2\beta_i^2 + 2B_1B_2(i,t+1)\beta_i\sigma_z\rho_z\right. \\
&\qquad\qquad\qquad\qquad \left. + B_2(i,t+1)^2\sigma_z^2 + B_3(i,t+1)^2\sigma_x^2\right) = 0 \\
&B_4 = \left(\theta\rho_i^2\sigma_x^2\right)^{-1}\left[(1-\rho_i b_x) \pm \sqrt{(1-\rho_i b_x)^2 - \theta^2\rho_i^4 B_1^2\sigma_m^2\sigma_x^2}\right] \\
&B_0(i,t) = \textstyle\sum_{j=1}^{\infty}\rho_i^j\left[\log\delta + \left(1-\frac{1}{\psi}\right)\overline{\mu}^i + \theta\left(1-\frac{1}{\psi}\right)^2\frac{\sigma_d^2}{2} + k + \rho_i\left(B_2(i,t+j)\alpha_z\right.\right. \\
&\qquad\qquad\qquad\qquad \left.\left. + B_3(i,t+j)\alpha_x + B_4\alpha_x\right)\right]
\end{aligned} \tag{A-12}$$

Since quadratic difference equations have no tractable solution in general, in the third equation we apply a first order Taylor series expansion with respect to σ_x around the deterministic case ($\sigma_x = 0$). The solution of the third equation then becomes:

$$B_3(i,t) = \frac{1}{2}\theta\rho_i^2\sum_{j=1}^{\infty}(\rho_i b_x)^j f(t+j,i) \tag{A-13}$$

$$f(i,t+j) = \left[B_1^2\beta_i^2 + 2B_1B_2(i,t+j)\beta_i\sigma_z\rho_z + B_2(i,t+j)^2\sigma_z^2\right] \tag{A-14}$$

We now replace calendar time with the state variable τ, time to next announcement, which takes values $1, 2, \ldots, \overline{A}$, with $\overline{A}$ the number of periods between announcements. The time-varying coefficients of the price consumption ratio appearing in (A-12) become:

$$
\begin{aligned}
B_2(i,\tau) &= (\rho_i b_z)^\tau \frac{\theta \rho_i^2 B_1^2}{2[1-(\rho_i b_z)^{\overline{A}}]} \beta_i^2 && \text{(A-15)} \\
B_3(i,\tau) &= \frac{\theta \rho_i^2}{2} \left(\sum_{j=1}^{\overline{A}} (\rho_i b_x)^j \frac{f(i,\tau-j)}{1-(\rho_i b_x)^{\overline{A}}} \right) && \text{(A-16)} \\
f(i,\tau) &= [B_1^2 \beta_i^2 + 2 B_1 B_2(i,\tau) \beta_i \sigma_z \rho_z + B_2(i,\tau)^2 \sigma_z^2] && \text{(A-17)} \\
B_0(i,\tau) &= \frac{1}{1-\rho} \left[\log \delta + \left(1 - \frac{1}{\psi}\right) \overline{\mu}^i + \theta \left(1 - \frac{1}{\psi}\right)^2 \frac{\sigma_d^2}{2} + k + \rho B_4 \alpha_x \right] + && \text{(A-18)} \\
& \sum_{j=1}^{\overline{A}} \rho_i^j \frac{\rho \left(B_2(i,\tau-j)\alpha_z + B_3(i,\tau-j)\alpha_x\right)}{1-\rho^{\overline{A}}} && \text{(A-19)}
\end{aligned}
$$

with the convention $\tau - j = \overline{A} - (j - \tau)$ if $j \geq \tau$.

Cross-sectional variation of price-consumption ratio coefficients

We report the partial derivatives with respect to β_i of the coefficients $B_2(i,\tau)$ and $B_3(i,\tau)$ appearing in (A-15)-(A-16). The signs of the partial derivatives hold under Assumption 1.

$$
\begin{aligned}
\frac{\partial B_2(i,\tau)}{\partial \beta_i} &= (\rho_i b_z)^\tau \frac{\theta \rho_i^2 B_1^2}{[1-(\rho_i b_z)^{\overline{A}}]} \beta_i < 0 && \text{(A-20)} \\
\frac{\partial B_3(i,\tau)}{\partial \beta_i} &= \frac{\theta \rho_i^2}{2} \left(\sum_{j=1}^{\overline{A}} (\rho_i b_x)^j \frac{1}{1-(\rho_i b_x)^{\overline{A}}} \right) \frac{\partial f(i,\tau-j)}{\partial \beta_i} < 0 && \text{(A-21)} \\
\frac{\partial f(i,\tau)}{\partial \beta_i} &= \left[2 B_1^2 \beta_i + 2 B_1 \frac{\partial B_2(i,\tau)}{\partial \beta_i} \beta_i \sigma_z \rho_z + 2 B_1 B_2(i,\tau) \sigma_z \rho_z \right. && \text{(A-22)} \\
& \left. + 2 B_2(i,\tau) \frac{\partial B_2(i,\tau)}{\partial \beta_i} \sigma_z^2 \right] > 0. && \text{(A-23)}
\end{aligned}
$$

Equilibrium Interest Rate

The conditional Normality of state variables, hence the conditional log-normality of the SDF, and the Euler equation for the one-period bond yield imply that the (continuously compounded) one-period interest rate is given by

$$
r_t = -\mathbb{E}_t[m_{t+1}] - \frac{1}{2} Var_t[m_{t+1}]. \qquad \text{(A-24)}
$$

From (A-11), and using Campbell-Shiller log-linearization of the return on aggregate wealth, we obtain:

$$
\begin{aligned}
m_{i,t+1} \quad = \quad & \log\delta - \frac{\overline{\mu}^i}{\psi} - (\theta)(\theta-1)\left(1-\frac{1}{\psi}\right)^2\frac{\sigma_d^2}{2} - \frac{\mu_{i,t}}{\psi} - \frac{A_{\tau=1}}{2}(\theta-1)\theta\rho_i^2B_1^2\beta_i^2 z_t \\
& -\frac{1}{2}(\theta-1)\theta\rho_i^2 f(i,\tau-1)x_t - \frac{1}{2}(\theta-1)\theta\rho_i^2\left[B_1^2\sigma_m^2 + B_4^2\sigma_x^2\right]x_{i,t} \\
& -\gamma\sigma_d\nu_{i,t+1} + (\theta-1)\rho_i B_1\left[\sigma_m\sqrt{x_{i,t}}\,\epsilon_{i,t+1} + \beta_i\sqrt{x_t}\,\epsilon_{t+1} + A_{\tau=1}\beta_i\sqrt{z_t}\,\eta_{i,t+1}\right] \\
& +(\theta-1)\rho_i B_2(i,\tau-1)\sigma_z\sqrt{x_t}\omega_{t+1}^z + (\theta-1)\rho_i B_3(i,\tau-1)\sigma_x\sqrt{x_t}\omega_{t+1}^x + \\
& (\theta-1)\rho_i B_4\sigma_x\sqrt{x_{i,t}}\omega_{i,t+1}^x. \qquad \text{(A-25)}
\end{aligned}
$$

After computing moments in (A-24) and simplifying terms, we obtain expression (13), where coefficients are:

$$C_0 \quad = \quad -\log\delta + \frac{\overline{\mu}_i}{\psi} + \frac{1-\gamma(\psi+1)}{\psi}\frac{\sigma_d^2}{2}, \qquad \text{(A-26)}$$

$$C_1 \quad = \quad \frac{1}{\psi} > 0, \qquad \text{(A-27)}$$

$$C_2(i) \quad = \quad \frac{1}{2}(\theta-1)\rho_i^2B_1^2\beta_i^2 < 0, \qquad \text{(A-28)}$$

$$C_3(i,\tau) \quad = \quad \frac{1}{2}(\theta-1)\rho_i^2 f(i,\tau-1) < 0, \qquad \text{(A-29)}$$

$$C_4 \quad = \quad \frac{1}{2}(\theta-1)\rho_i^2\left[B_1^2\sigma_m^2 + B_4^2\sigma_x^2\right] < 0. \qquad \text{(A-30)}$$

The signs hold under Assumption 1. It follows immediately that $\frac{\partial C_2(i)}{\partial\beta_i} > 0$ and $\frac{\partial C_3(i,\tau)}{\partial\beta_i} > 0$, because $\frac{\partial f(i,\tau)}{\partial\beta_i} > 0$.

Proof of Proposition 2

Taking into account expression (A-11) for the log-stochastic discount factor, the logarithmic change of the exchange rate of country i's vs home currency is reads explicitly:

$$
\begin{aligned}
\Delta q_{i,t+1} = m_{h,t+1} - m_{i,t+1} = & \frac{\mu_{i,t}-\mu_{h,t}}{\psi} - \frac{A_{\tau=1}}{2}(\theta-1)\theta\rho_i^2B_1^2\left(\beta_h^2-\beta_i^2\right)z_t \\
& -\frac{1}{2}(\theta-1)\theta\rho_i^2\left[f(h,\tau-1)-f(i,\tau-1)\right]x_t - \frac{1}{2}(\theta-1)\theta\rho_i^2\left[(B_1^2\sigma_m^2+B_4^2\sigma_x^2)x_{h,t}\right. \\
& \left. -(B_1^2\sigma_m^2+B_4^2\sigma_x^2)x_{i,t}\right] - \gamma\sigma_d(\nu_{h,t+1}-\nu_{i,t+1}) + (\theta-1)\rho B_1\left[(\beta_h-\beta_i)\sqrt{x_t}\,\epsilon_{t+1}\right.
\end{aligned}
$$

$$
\begin{aligned}
& \left. +A_{\tau=1}\sqrt{z_t}(\beta_h\,\eta_{h,t+1} - \beta_i\,\eta_{i,t+1}) + \sigma_m\sqrt{x_{h,t}}\,\epsilon_{h,t+1} - \sigma_m\sqrt{x_{i,t}}\,\epsilon_{i,t+1}\right] \\
& + (\theta-1)\rho_i\left(B_2(h,\tau-1)-B_2(i,\tau-1)\right)\sigma_z\sqrt{x_t}\omega_{t+1}^z + (\theta-1)\rho_i\left(B_3(h,\tau-1)\right. \\
& \left. -B_3(i,\tau-1)\right)\sigma_x\sqrt{x_t}\omega_{t+1}^x + (\theta-1)\rho_iB_4(\sigma_x\sqrt{x_{h,t}}\omega_{h,t+1}^x - \sigma_x\sqrt{x_{i,t}}\omega_{i,t+1}^x) \qquad \text{(A-31)}
\end{aligned}
$$

The Euler equation for the equilibrium pricing of currency returns in the home country, and the joint conditional Normality of log-stochastic discount factor and currency log returns leads to the well know formula of the risk premium (inclusive of the Jensen inequality adjustment):

$$\mathbb{E}_t[r^c_{i,t+1}] + \frac{1}{2}Var_t[r^c_{i,t+1}] = -Cov_t[r^c_{i,t+1}, m_{h,t+1}] = Cov_t[\Delta q_{i,t+1}, m_{h,t+1}]. \tag{A-32}$$

Using (A-25) and (A-31) to compute conditional covariances leads to expression (14), where the coefficients read:

$$\begin{aligned} g_x(i,\tau-1) &= (\theta-1)^2\rho_i^2\left[B_1^2\beta_h(\beta_h-\beta_i) + B_1\beta_h\left(B_2(h,\tau-1) - B_2(i,\tau-1)\right)\sigma_z\rho_z\right. \\ &\quad + B_1B_2(h,\tau-1)(\beta_h-\beta_i)\sigma_z\rho_z + B_2(h,\tau-1)\left(B_2(h,\tau-1)-B_2(i,\tau-1)\right)\sigma_z^2 \\ &\quad \left.+ B_3(h,\tau-1)\left(B_3(h,\tau-1)-B_3(i,\tau-1)\right)\sigma_x^2\right], & \text{(A-33)} \\ g_z(i) &= (\theta-1)^2\rho_i^2B_1^2\beta_h(\beta_h-\beta_i\rho_{h,i}), & \text{(A-34)} \\ g_h &= (\theta-1)^2\rho_i^2(B_1^2\sigma_m^2+B_4^2\sigma_x^2). & \text{(A-35)} \end{aligned}$$

Linearity of the conditional currency risk premium in the state variables z, x, and x_i implies that the unconditional premium reads:[36]

$$g_x(i,\tau-1)\mathbb{E}[x_t] + A_{\tau=1}g_z(i)\mathbb{E}[z_t] + g_h\mathbb{E}[x_{i,t}] \tag{A-36}$$

where $\mathbb{E}[x_t] = \mathbb{E}[x_{i,t}] = \alpha_x/(1-b_x)$, and $\mathbb{E}[z_t] = \alpha_z/(1-b_z)$. We have also made use of the expression of the currency risk premium conditional only on the current economic conditions of the home country, $\mu_{h,t}$. z_t is the only state-variable correlated with $\mu_{h,t}$. Using the joint Normality of the steady-state distribution of $(z_t, \mu_{h,t})$, we can write:

$$\mathbb{E}\left[z_t|\,\mu_{h,t}\right] = \mathbb{E}\left[z_t\right] + \frac{\mathrm{Cov}[\mu_{h,t}, z_t]}{\mathrm{Var}[\mu_{h,t}]}\left(\mu_{h,t} - \mathbb{E}\left[\mu_{h,t}\right]\right) \tag{A-37}$$

where

$$\begin{aligned} \mathbb{E}\left[\mu_{h,t}\right] &= 0 & \text{(A-38)} \\ \mathrm{Cov}[\mu_{h,t}, z_t] &= \frac{\sigma_z\rho_z\beta_h\alpha_x}{(1-b_x)(1-b_zb_m)} & \text{(A-39)} \\ \mathrm{Var}[\mu_{h,t}] &= \frac{\alpha_x(\sigma_m^2+\beta_h^2)}{(1-b_x)(1-b_m^2)} + \frac{1}{A}\frac{\alpha_z\beta_h^2}{(1-b_z)(1-b_m^2)} & \text{(A-40)} \end{aligned}$$

The currency risk premium conditional on the home country's current economic conditions (output growth) is thus

$$g_x(i,\tau-1)\mathbb{E}[x_t] + A_{\tau=1}g_z(i)\mathbb{E}\left[z_t|\,\mu_{h,t}\right] + g_h\mathbb{E}[x_{i,t}]. \tag{A-41}$$

Proof of Corollary 1

The Corollary readily follows from Assumption 1, the fact that $g_x(h,\tau-1)=0$, $\frac{\partial g_x(i,\tau-1)}{\partial\beta_i} < 0$, the expression for $g_z(i)$, and expression (13) for the equilibrium interest rate, where, in particular, $\frac{C_3(i,\tau)}{\partial\beta_i} < 0$ and $\frac{C_2(i)}{\partial\beta_i} < 0$.

[36]The only conditioning variable in the expression below is τ.

HML and DOL factors

Let H (L) be the set of all the N_H (N_L) countries with positive interest rate differential over the home country , i.e. $\beta_i < \beta_h$ ($\beta_i > \beta_h$). The HML factor is defined as

$$\mathrm{HML}_{t+1} = \frac{1}{N_H}\sum_{j\in H} r^c_{j,t+1} - \frac{1}{N_L}\sum_{j\in L} r^c_{j,t+1}. \tag{A-42}$$

The innovation component of *hml* reads

$$\begin{aligned}\mathrm{HML}_{t+1} - \mathbb{E}_t[\mathrm{HML}_{t+1}] \quad = \quad & (\theta-1)\rho B_1\left[(\overline{\beta_H}-\overline{\beta_L})\sqrt{x_t}\epsilon_{t+1} + \sqrt{z_t}A_{\tau=1}(\overline{\beta_H\rho_H}-\overline{\beta_L\rho_L})\eta_{h,t+1}\right] + \\ & (\theta-1)\rho(\overline{B_2(H,\tau-1)}-\overline{B_2(L,\tau-1)})\sigma_z\sqrt{x_t}w^z_{t+1} \\ & +(\theta-1)\rho(\overline{B_3(H,\tau-1)}-\overline{B_3(L,\tau-1)})\sigma_x\sqrt{x_t}w^x_{t+1}\end{aligned} \tag{A-43}$$

where

$$\overline{\beta_H\rho_H} = \frac{1}{N_H}\sum_{j\in H}\beta_i\rho_{h,i} \tag{A-44}$$

$$\overline{\beta_H} = \frac{1}{N_H}\sum_{j\in H}\beta_j \tag{A-45}$$

$$\overline{B_2}(H,\tau-1) = \frac{1}{N_H}\sum_{j\in H}B_2(j,\tau-1) \tag{A-46}$$

$$\overline{B_3}(H,\tau-1) = \frac{1}{N_H}\sum_{j\in H}B_3(j,\tau-1) \tag{A-47}$$

and similarly for L terms. We have used the fact that $\eta_{i,t+1} = \rho_{h,i}\eta_{h,t+1} + \tilde{\eta}_{i,t+1}\sqrt{1-\rho^2_{h,i}}$, where the $\tilde{\eta}_{i,t+1}$ are cross-sectionally independent Gaussian white noise.

We can apply the law of large numbers to conclude that all country-specific innovations average out. The risk premium of the HML portfolio conditional on a non-announcement day is hence:

$$\begin{aligned}&\mathbb{E}_t[\mathrm{HML}_{t+1}] + \frac{1}{2}Var_t[\mathrm{HML}_{t+1}]\Big|_{\tau\neq 1} = -Cov_t\left[\mathrm{HML}_{t+1}, m_{h,t+1}\right]\big|_{\tau\neq1} = \\ &(\theta-1)^2\rho^2B_1^2\beta_h(\overline{\beta_L}-\overline{\beta_H})x_t + (\theta-1)^2\rho^2(\overline{B_2(L,\tau-1)}-\overline{B_2(H,\tau-1)})B_{2,\tau-1}(h)\sigma^2_z x_t \\ &+(\theta-1)^2\rho^2(\overline{B_3(L,\tau-1)}-\overline{B_3(H,\tau-1)})B_{3,\tau-1}(h)\sigma^2_x x_t + (\theta-1)\rho B_1(\overline{\beta_L}-\overline{\beta_H})B_2(h,\tau-1)\sigma_z\rho_z x_t \\ &+(\theta-1)^2\rho^2(\overline{B_2(L,\tau-1)}-\overline{B_2(H,\tau-1)})\beta_h\sigma_z\rho_z x_t.\end{aligned} \tag{A-48}$$

It is straightforward to see that this premium is always positive. The same quantity conditional on an announcement day then is:

$$\begin{aligned}&\mathbb{E}_t[\mathrm{HML}_{t+1}] + \frac{1}{2}Var_t[\mathrm{HML}_{t+1}]\Big|_{\tau=1} = -Cov_t\left[\mathrm{HML}_{t+1}, m_{h,t+1}\right]\big|_{\tau=1} = \\ &-Cov_t\left[\mathrm{HML}_{t+1}, m_{h,t+1}\right]\big|_{\tau\neq1} + (\theta-1)^2\rho^2B_1^2\beta_h(\overline{\beta_L\rho_L}-\overline{\beta_H\rho_H})z_t\end{aligned} \tag{A-49}$$

Thus the second term on the RHS of the last expression in the announcement risk component of the HML premium. The HML beta in a linear factor model for currency returns is proportional to the covariance between HML and exchange rate variations changed in sign:

$$\begin{aligned} Cov_t[HML_{t+1}, -\Delta q_{i,t+1}]|_{\tau \neq 1} &= (\theta-1)^2\rho^2 B_1^2(\beta_h - \beta_i)(\overline{\beta_L} - \overline{\beta_H})x_t \\ &\quad + \text{other terms decreasing in } \beta_i \end{aligned} \tag{A-50}$$

$$\begin{aligned} Cov_t[HML_{t+1}, -\Delta q_{i,t+1}]|_{\tau = 1} &= Cov_t[HML_{t+1}, -\Delta q_{i,t+1}]|_{\tau \neq 1} \\ &\quad + (\theta-1)^2\rho^2 B_1^2(\beta_h - \beta_i\rho_{h,i})(\overline{\beta_L\rho_L} - \overline{\beta_H\rho_H})z_t \end{aligned} \tag{A-51}$$

Since $\beta_L > \beta_H$ and the interest rate differential (foreign minus home) is decreasing in β_i, (A-50) is increasing in the interest rate differential.

The DOL factor is defined as:

$$\text{DOL}_{t+1} = \frac{1}{N}\sum_{i=1}^{N} r^c_{i,t+1} \tag{A-52}$$

Since we have assumed that the home (US) loading coincides with the average loading, i.e. $\beta_h = \left(\sum_{i=1}^N \beta_i\right)/N$, the innovation component of the DOL factor reads:

$$\begin{aligned} \text{DOL}_{t+1} - \mathbb{E}_t\left[\text{DOL}_{t+1}\right] &= \gamma\sigma_d\nu_{h,t+1} - (\theta-1)\rho B_1\sigma_m\sqrt{x_{h,t}}\epsilon_{h,t+1} - \\ &\quad (\theta-1)\rho B_4\sigma_x\sqrt{x_{h,t}}w^x_{h,t+1} \\ &\quad + A_{\tau=1}(\theta-1)\rho B_1(\overline{\beta\rho} - \beta_h)\sqrt{z_t}\eta_{h,t+1} \end{aligned} \tag{A-53}$$

where $\overline{\beta\rho} = \frac{1}{N}\sum_{j=1}^N \beta_i\rho_{h,i}$. The risk premium of the DOL portfolio conditional on a non-announcement day is:

$$\begin{aligned} \mathbb{E}_t[\text{DOL}_{t+1}] + \frac{1}{2}Var_t[\text{DOL}_{t+1}]\Big|_{\tau\neq 1} &= -Cov_t\left[\text{DOL}_{t+1}, m_{h,t+1}\right]|_{\tau\neq 1} = \gamma^2\sigma_d^2 + (\theta-1)^2\rho^2B_1^2\sigma_m^2 x_{h,t} \\ &\quad + (\theta-1)^2\rho^2 B_4^2\sigma_x^2 x_{h,t}, \end{aligned} \tag{A-54}$$

while conditional on an announcement day it reads:

$$\begin{aligned} \mathbb{E}_t[\text{DOL}_{t+1}] + \frac{1}{2}Var_t[\text{DOL}_{t+1}]\Big|_{\tau= 1} &= -Cov_t\left[\text{DOL}_{t+1}, m_{h,t+1}\right]|_{\tau\neq 1} = -Cov_t\left[\text{DOL}_{t+1}, m_{h,t+1}\right]|_{\tau=1} \\ &\quad + (\theta-1)^2\rho^2 B_1^2\beta_h(\beta_h - \overline{\beta\rho})z_t \end{aligned} \tag{A-55}$$

Notice that the DOL announcement premium component (second term on the RHS of the last expression) is positive, since $\beta_h = \left(\sum_{i=1}^N \beta_i\right)/N$ implies that $\beta_h > \overline{\beta\rho}$. The DOL beta in a linear factor model for currency returns is proportional to the covariance between DOL and exchange rate variations changed in sign.

$$\begin{aligned} Cov_t[DOL_{t+1}, -\Delta q_{i,t+1}]|_{\tau\neq 1} &= \gamma^2\sigma_d^2 + (\theta-1)^2\rho^2B_1^2\sigma_m^2x_{h,t} \\ &\quad + (\theta-1)^2\rho^2B_4^2\sigma_x^2x_{h,t} \end{aligned} \tag{A-56}$$

$$\begin{aligned} Cov_t[DOL_{t+1}, -\Delta q_{i,t+1}]|_{\tau= 1} &= Cov_t[DOL_{t+1}, -\Delta q_{i,t+1}]|_{\tau\neq 1} \\ &\quad + (\theta-1)^2\rho^2B_1^2(\beta_h - \beta_i\rho_{h,i})(\beta_h - \overline{\beta\rho})z_t \end{aligned} \tag{A-57}$$

Note that while expression (A-56) is cross-sectionally invariant (it does not depend on β_i), expression (A-57) is not.

Appendix B Robustness Checks

A natural concern for the results reported in Table 6 is that asymptotic theory may not provide a good approximation for the distribution of the estimates, due to the small number of sample observations. We address this concern with a bootstrap exercise. Namely, we compute small sample standard errors for the point estimates of the dummy variable regression. In particular, we draw with replacement from the observed distribution. We then estimate regression (1) and store the estimated coefficients for each portfolio. The empirical distributions are plotted in Figure 6. We note that the mean is very similar to the estimated coefficients for $\hat{\alpha}_1$ reported in Table 6. Moreover, standard errors are also similar to those reported for the regression.

[Insert Figure 6 here.]

Appendix C Forward Guidance

Forward guidance about the Fed's interest rate policy has led to a large decline in monetary policy uncertainty. Since March 2011, the FOMC introduced regular post-meeting press conferences by the chairman. The press conferences coincide with the committee's publication of the Summary of Economic Projections (SEP) and are intended to "further enhance the clarity and timeliness of the Federal Reserve's monetary policy communication". Moreover, since January 2012, the FOMC included Federal Funds Rate projections by FOMC participants in the SEP (so called "dot" graphs).

The decrease in monetary policy uncertainty is manifested for example in surveys conducted by Bloomberg about the target Fed Funds rate. Before each FOMC meeting, Bloomberg conducts a survey among international professional forecasters and asks them about their expected target rate. To get a measure of uncertainty, we construct a cross-sectional standard deviation from each survey.

The lower panel of Figure 7 plots the uncertainty proxy since 2008. We note that since the early 2010, the uncertainty is essentially zero.

At the same time, we find that interest rate differentials have decreased since the crisis as central banks around the world have lowered their target rates. Figure 7 (upper panel) plots the average interest rate differential for portfolio 1 to 5 focusing on the years 2008 to 2014. We note that since the late 2008, interest rate differentials for portfolios 1 and 2 are essentially zero, rendering a currency strategy for these particular currencies unattractive.

[Insert Figure 7 here.]

Indeed, if we look at the currency returns for this particular period, we find that the difference in currency returns between FOMC and non FOMC days is statistically not distinguishable from zero.

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