Policy Announcements in FX Markets*

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Abstract

This paper studies empirically and theoretically the price of monetary policy uncertainty in the foreign exchange (FX) market. Using a large cross-section of foreign exchange data we document that carry returns are mainly earned on days of scheduled meetings of the Federal Open Market Committee (FOMC). More concretely, on announcement days, currencies with low interest rates produce an average daily return of 5.2 basis points (bp) compared to -1 bp on non-announcement days. This wedge becomes larger for high interest rate currencies and during bad economic times. We develop a model of an international long-run risk economy in which asset prices respond to revisions of the monetary policy. Monetary policy uncertainty commands a risk premium that is larger in weaker economic conditions when a revision of the monetary policy is more probable. A calibrated version is consistent with the cross-sectional pattern of carry trade risk premia observed in the data.

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1 Introduction

In this paper, we document that returns on foreign exchange portfolios are mainly earned on days when a meeting of the Federal Open Market Committee (FOMC) is scheduled. Specifically, we show that currency returns on interest rate sorted portfolios have very small Sharpe ratios outside FOMC meeting days and large and significant annualized Sharpe ratios on days when the FOMC makes monetary policy announcements. To illustrate the intuition for our findings, we can consider the FOMC meetings of June 2013, when Fed Chairman Bernanke hinted at a tapering of its quantitative easing (QE) program. QE has led the US dollar to be the funding currency in large scale carry 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 carry trades. These events illustrate the potential risk that materializes during days when the FOMC announces its future policy, a risk for which international investors 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 cross-section of currencies and 30 years of data, we find that returns on currency portfolios are significantly larger on days when the Federal Reserve is scheduled to make an announcement. Specifically, portfolios of currencies with low interest rates earn an average daily return of 5.2 basis points (bp) in these days, compared to -1 bp on non-announcement days. This difference becomes even larger for high interest rate currencies, with a daily return of 9.9 bp on announcement days compared to 1.5 bp, a 8.4 bp difference which is strongly statistically significant from zero. Interestingly, since the risk associated with these returns (measured by realized volatility) does not experience a corresponding increase, the resulting implied Sharpe ratio is an order of magnitude larger on announcement days.
We show that our findings are driven neither by the choice of currencies nor by a few outliers. When we winsorize the dataset, discarding the top and the bottom 1%-iles, results are virtually unchanged. Interestingly, the announcement premium appears to be exclusively FOMC-related, as we show that other macroeconomic announcements, such as consumer price index, total nonfarm payroll, and the producer price index, for instance, generate no significant currency return in excess of non announcement days returns. Moreover the direction of the policy revision, i.e., whether interest rates are raised or reduced, is immaterial for our results: Constructing a policy surprise component, we find that this variable has no significant effect on announcement returns. In contrast, conditioning on the state of the economy has a significant 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 explain our findings with an equilibrium model of an open economy. The model draws upon two strands of the literature: First, long-run risk 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 it in a Bayesian fashion until the date when a new policy is announced. 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 the unpredictable revision of the policy, and consequently of its impact on long-term growth, determine shocks to expected consumption growth which are more volatile than the updates following normal (i.e. non-announcement) news. To distinguish between the two types of news, we dub the former announcement-day volatility “announcement uncertainty” and the latter a “normal” volatility. Due to recursive preferences, both types of risk affecting long-run expectations are priced in equilibrium. Our main focus is on explaining the cross-sectional pattern of carry trade risk premia around announcements through countries’ heterogeneous exposure to shocks in long-run risk. Countries more (less) exposed to “normal” uncertainty have smaller
(larger) interest rate differential vis-à-vis the home country and smaller (larger) carry trade risk premia.

The intuition is similar to Lustig, Roussanov, and Verdelhan (2011): when “normal” long-run risk volatility is larger in the foreign country, carry trade returns hedge negative shocks to expected domestic consumption 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 carry trade 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 carry trade returns triggered by announcements do 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, equity premia and FX options. We obtain a good goodness-of-fit and, using calibrated parameters, we show that the cross-sectional pattern of model-implied carry trade 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 is the unpredictability of policy revisions related to informational frictions. The authors are mainly concerned with the (equity) asset pricing implications of this uncertainty: price risk, tail risk, and variance risk related to political events. We instead, focus on an international framework and the
timing of announcement events.\footnote{Pástor and Veronesi (2013) specify a single time at which an announcement is made, whereas we consider a regular schedule of announcements and introduce a calendar state variable similar to Savor and Wilson (2014). The reason for this is that FOMC announcement dates are well known in advance and occur at fairly regular intervals.} 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 carry-trade risk premia, rather than equity returns. Kelly, Pástor, and Veronesi (2014) inspect the effect 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 carry trade 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 carry 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. 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.

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)).\footnote{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 survey of the literature).} More closely related to our paper, Jones, La- mont, and Lumsdaine (1998) study unconditional fixed income returns around macroe-
economic 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 from 1958 through 2009. Lucca and Moench (2014) study 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. 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&P500. 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 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, the authors implement a daily re-balancing strategy. We invest over a monthly horizon and then track the portfolio day by day. Second, 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 neutral to the US dollar.

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 depreciation in the US dollar is higher than the probability of a large appreciation.

Carry returns display a large negative skewness which is induced by periods of rapid depreciation. This has lead researchers to ask whether these returns merely reflect the presence of a peso problem. For example, Burnside, Eichenbaum, Kleshchelski, and Rebelo (2011) show that standard risk factors cannot explain the returns to the carry trade and test the hypothesis of a potential peso problem by implementing a crash-hedged currency carry trade through FX options. More recently, Jurek (2014) implements a similar strategy but instead of relying on at-the-money options, he studies
a crash-hedged trading strategy using out-of-the-money put options. His findings reject the hypothesis of a peso problem and he concludes that crash risk premia tend to be low. These findings are confirmed in Jurek and Xu (2013) who calibrate a multi-country model of stochastic discount factor dynamics. Interestingly, we find that while carry returns are negatively skewed on non announcement days, the skewness on announcement days is actually positive. For example, on non FOMC announcement days, the realized skewness becomes increasingly negative with the interest rate differential. On announcement days, the skewness ranges from 0.3 to 1.6. While crashes are certainly an important feature of carry returns, they lie outside the scope of our model.

The importance of monetary policy using interest-rate rules in exchange rate determination and its role for the understanding of the uncovered interest rate parity puzzle has been first highlighted by McCallum (1994) and more recently by Backus, Gavazzoni, Telmer, and Zin (2013). The latter explore how different specifications of the Taylor rule affect exchange rates and find that a more procyclical Taylor rule implies higher risk premia on exchange rates. Benigno, Benigno, and Nistico (2011) study how monetary policy uncertainty shocks affect currency risk premia within an equilibrium model featuring Epstein-Zin preferences and stochastic volatility. They show that their model explains well the violation of the uncovered interest rate parity. Bansal and Shaliastovich (2013) develop a two-country economy with long-run risk with time-varying volatilities of expected growth and inflation and show that their model simultaneously accounts for the violation of the UIP and bond return predictability. 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 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.
Our paper proceeds as follows. In the next section, we introduce our data and we study carry returns on FOMC announcement and non-announcement days in Section 3. 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 start by describing the data and the construction of different dollar carry portfolios. We work with a daily frequency beginning in January 1980 and ending in December 2010. There are eight scheduled FOMC meetings in one year. 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. 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)). This hence leaves us with 7,481 days of non-FOMC announcements 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). Following the extant literature since Fama (1984), we will work in logarithms of spot and forward rates for ease of exposition and notation.

We denote spot and forward rates in logs as $s_t$ and $f_t$, respectively. The log excess return $rx_{t+1}$ on buying a foreign currency in the forward market and then selling it in the spot market after one month is simply $rx_{t+1} = f_t - s_{t+1}$. This excess return can also be stated as the log forward discount minus the change in the spot rate: $rx_{t+1} = f_t - s_t - \Delta s_{t+1}$. In normal conditions, forward rates satisfy the covered interest rate parity condition; the forward discount is equal to the interest rate differential: $f_t - s_t \approx i_t^* - i_t$, where $i^*$ and $i$ denote the foreign and domestic nominal risk-free rates over the maturity of the contract. Akram, Rime, and Sarno (2008) study high frequency deviations from covered interest rate parity (CIP). They conclude that CIP holds at daily and lower
frequencies. Hence, the log currency excess return equals the interest rate differential less the rate of depreciation: \( r_{x_{t+1}} = i_t^* - i_t - \Delta s_{t+1} \).

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 with a longer data history. 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.

**Option Data:** We use daily over-the-counter (OTC) currency options data from JP Morgan. The data starts in January 1996. Using OTC options data has several advantages over exchange-traded data. First, the trading volume in the OTC FX options market is several times larger than the corresponding volume on exchanges such as the Chicago Mercantile Exchange, and this leads to more competitive quotes in the OTC market. Second, the conventions for writing and quoting options in the OTC markets have several features that are appealing when performing empirical studies. In particular, every day sees the issuance of new options series with fixed times to maturity and fixed strike prices, defined by sticky deltas; in comparison, the time to maturity of an exchange-traded option series gradually declines with the approaching expiration date and so the moneyness continually changes as the underlying exchange rate moves. As a result the OTC options data allows for better comparability over time because the series’ main characteristics do not change from day to day. The options used in this study are plain-vanilla European calls and puts with five option series per exchange
rate. Specifically, we consider a one-month maturity and a total of five different strikes: at-the-money (ATM), 10-delta and 25-delta calls, as well as 10-delta and 25-delta puts.

**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). Negative numbers indicate a below average growth and positive numbers an above average growth. As in Bernanke and Kuttner (2005) we measure a surprise component in target rate changes from the change in the one-month Federal Funds Futures contract price relative to the day prior to the FOMC meeting.

### 2.2 Portfolio Construction

At the end of each month $t$, we allocate currencies to five portfolios based on their observed forward discounts $f_t - s_t$. Sorting on forward discounts is equivalent to sorting on interest rate differentials since covered interest parity holds closely in the data at the frequency analyzed in this paper. We re-balance portfolios at the end of each month. This is repeated month by month for the whole data period. Currencies are ranked from low to high interest rates. Portfolio 1 contains currencies with the lowest interest rate (or smallest forward discounts) and portfolio 5 contains currencies with the highest interest rates (or largest forward discounts). We assume that the interest rate differential is earned linearly over the month. To calculate daily excess log returns, we use the daily interest rate differential and daily log exchange rate changes. Portfolio returns are calculated as the average of the currency excess returns in each portfolio as in Lustig, Roussanov, and Verdelhan (2011). 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 returns on average than high interest rate currencies: In our sample,
the low interest rate portfolio earns a daily return of $-0.86$ bp (with a t-statistic of 1.74) while the high interest rate portfolio earns 1.82 bp (with a t-statistic of 2.56). Annualized Sharpe ratios are large and range from $-0.3$ for the low interest rate currencies to 0.46 for high interest rate currencies.

In Table 2, we present summary statistics of the option data sorted on their interest rate differential. To this end, we keep track of all the currencies at the end of each month and then calculate the average implied volatility, slope, and forward discount for each portfolio. We first note that the (annualized) implied volatility increases with the interest rate differential, namely, low interest rate currencies have an implied volatility of 8.8% whereas high interest rate currencies have an implied volatility of 14.6%. The slope, on the other hand, becomes more negative, the larger the interest rate differential. For funding currencies, we find that the slope is positive, which implies a right-skewed distribution, whereas for investment currencies, the slope is highly negative (see also Jurek (2014)).

[Insert Table 2 here.]

3 Empirical Analysis

In this section, we analyze the characteristics of returns on interest rate sorted currency portfolios. In particular, we compare announcement-day to non-announcement-day returns. A number of robustness checks confirms our main result: returns on a trading strategy that is short the US dollar and long any other currency are on average significantly larger on days when an FOMC policy announcement takes place.

3.1 Carry Portfolios on Announcement Days

Figure 1 and Table 3 present the main results. The average daily return on the low interest rate portfolio is 5.213 bp on announcement days compared to $-1.07$ bp on non-announcement days. The difference between the two mean returns (6.28 bp) is statistically significant, with a t-statistics of 2.25.\footnote{The difference in mean test allows for different variances.} For the high interest rate currency
portfolio, the average return is almost 10 bp on announcement days, compared to 1.5 bp on non-announcement days: an 8.34 bp difference, again with a t-statistic of 2.07. This dramatic surge of mean returns is not accompanied by a corresponding increase in realized risk (as measured by realized volatility), which is almost unaltered during announcement days. Consequently, implied Sharpe ratios portray a risk-return trade-off which is an order of magnitude larger: 2.68 vs 0.39 for high interest rate currencies, and 1.87 vs -0.39 for low interest rate currencies.

Two observations are noteworthy. First, average announcement day returns are much larger than unconditional averages taken over all days as reported in Table 1. This implies that on 3% of all trading days one could earn more than holding the portfolio for the whole sample period. 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.\(^4\)

Figure 2 plots the empirical density functions for low (left panel) and high (right panel) interest rate sorted currency portfolio returns on non FOMC and FOMC announcement days. We find that the latter distribution is characterized not only by a larger mean but also by positive skewness considerably smaller kurtosis. Note that carry returns on non-announcement days feature a large negative skewness, a feature which has been attributed to crash risk in carry returns (see e.g., Farhi, Fraiberger, Gabaix, Ranciere, and Verdelhan (2014)).

To address the concern that our results are driven by a peculiar choice of currencies, we next apply our empirical strategy only to the subset of developed countries. The lower panels of Table 3 summarize the results. The findings are largely unchanged: mean returns on the carry trade strategy are significantly larger on FOMC announcement days. The spread between announcement and non announcement days average returns is 6.4 bp for low interest rate currencies and 9.27 bp for high interest rate currencies, with t-statistics of 1.93 and 2.15, respectively.

\[\text{[Insert Figure 1 - 2 and Table 3 here.]}\]

\(^4\)In our model, we interpret this homogeneous positivity of carry excess returns as a risk premium for “announcement uncertainty” risk.
Importantly, we rule out the additional concern that our results depend on a few outliers, as the well known tendency to crash of carry returns might suggest. Table 4 reports summary statistics for interest rate sorted portfolios on FOMC and non FOMC days for a winsorized data sample. Specifically, we discard the top and the bottom 1%-iles the dataset. There is virtually no distinction between the mean and standard deviation of returns in Panel A (all observations) and those in Panel B (windsorized data). We note that while carry returns exhibit a well-known negative skewness for all observations, the skewness on announcement days is positive.

Overall, robustness checks on outliers and currency subsamples confirm our conclusion that carry trade returns are significantly larger when we condition on a FOMC announcement day, with little or no change in realized risk, hence much larger Sharpe ratios.

3.1.1 The Time-Series of Carry Portfolios

We now continue our investigation of carry trade returns related to FOMC announcements by taking a time-series perspective. As a first exercise, we regress carry returns, \( r_{it}, i = 1, \ldots, 5 \), onto an announcement dummy which takes the value of one on announcement days and zero otherwise:

\[
r_{it} = \alpha_0 + \alpha_1 \times \text{Announcement Dummy}_t + \epsilon_t.
\]

In this regression, the intercept \( \alpha_0 \) measures the unconditional mean excess return earned on all time periods other than FOMC announcement days and \( \alpha_1 \) measures the spread between announcement versus non-announcement day returns.

Our findings are reported in Panel A of Table 5. The results mirror the findings in Table 3. Estimated coefficients for the announcement dummy are positive for all portfolios. For low interest rates, the estimated coefficient for \( \alpha_1 \) is 6.3 with an associated \( t \)-statistic of 2.22. For high interest rate currencies, the difference becomes larger, i.e. the return differential is 8.34 with an associated \( t \)-statistic of 2.21. The estimates for
the intercept $\alpha_0$ are not significant except for portfolio 1 and 5, indicating that outside of FOMC meetings, there is little return to be earned.

[Insert Table 5 here.]

One natural question to ask is whether the time period we study is a particularly good sample to invest in carry strategies and that investors learned about this on days when the FOMC makes its announcement. To assess this good-news hypothesis in more detail, we add a monetary policy surprise component extracted from federal funds futures data to our regressions. The difference between the forecast of the fed funds rate and the current fed funds target is an estimate of how the market expects the Fed to change the target. Comparing actual changes in the fed funds rate to the change that the market expected then tells us how much was expected and how much was a surprise. We run the following regression:

$$r_{i,t} = \beta_0 + \beta_1 \times \text{Announcement Dummy}_t + \beta_2 \text{Policy Surprise}_t + \epsilon_t.$$  

The results are presented in Panel B. The effect of monetary policy surprises is fairly modest and not significant for the high and low interest rate portfolios. All estimated coefficients ($\beta_2$) are negative, implying that a surprise tightening of monetary policy should lead to higher returns to the carry portfolio.

As a last exercise, we check whether returns are generally higher during bad economic times compared to boom periods. To this end, we interact the announcement dummy with a dummy of economic activity. The economic dummy takes a value of one if the Chicago Fed National Activity Index is negative and zero otherwise:

$$r_{i,t} = \gamma_0 + \gamma_1 \times \text{Announcement Dummy}_t \times \text{CFNAI}_t + \epsilon_t.$$  

The results are reported in Panel C of Table 5. Interestingly, we find that the wedge between announcement and non-announcement returns is much larger during bad economic times: not only are coefficients $\gamma_1$ statistically significant, but they are also 50% larger than coefficients $\alpha_1$ of the first regression. For example, for low (high) interest rate
currencies, the wedge between announcement and non-announcement returns increases from 6.2 bp (8.3 bp) to 9.2 bp (12.57 bp) once we condition for bad states. We hence conclude that announcement return premia are larger in weaker economic states.

### 3.1.2 Other Macroeconomic Announcements

Are other macroeconomic announcements able to generate the same large returns earned around FOMC events? To answer this question, we consider three major US macroeconomic news releases: total non-farm payroll employment, the Producer Price Index, and the Consumer Price Index, all published by the Bureau of Labor Statistics (BLS). We build announcement dummy variables and we check for larger conditional returns by means of the following regression

\[
 r^i_t = \beta_0 + \beta_1 \times \text{Announcement Dummy}_{E, CPI, PPI}^t + \epsilon_t,
\]

where \( r^i_t \) are the carry returns and Announcement Dummy \(_{E, CPI, PPI}^t \) is a dummy variable that takes the value of one on days when employment (E), consumer price index (CPI), or the producer price index (PPI), are announced. Results, reported in Table 6, show that the only statistically significant estimate is for PPI announcements on the low interest rate portfolio. We find that announcement days returns are on average 6.3 bp smaller (with a t-statistic of 2.5). These results support the conclusion that the role of monetary policy is unique among macroeconomic announcements to generate the large size of announcement excess returns on carry trade portfolios.

[Insert Table Table 6 here.]

### 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)).
Following the latter, we assume that the policy action of some authority affects expected output growth. As the precise impact is not observed, agents learn 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 consumption growth which are more volatile than updates following “normal” (i.e. non-announcement) news. We dub this announcement-day volatility “announcement uncertainty”, to distinguish it from the “normal”, intra-announcements volatility. Due to a preference for early resolution of uncertainty, 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” ones.\(^5\) Each country is populated by a representative agent with Epstein and Zin (1989) preferences:

$$U_{i,t} = \left\{ (1 - \delta) (C_{i,t})^{1 - \frac{1}{\psi}} + \delta E_t \left[ (U_{i,t+1})^{1 - \gamma} \right]^{\frac{1}{1 - \psi}} \right\}^{\frac{1}{1 - \psi}} \text{ for } i = h, 1, \ldots, N,$$

where the subscript $h$ identifies the home country, and $\theta = (1 - \gamma)/(1 - 1/\psi)$. $\delta$ is the subjective discount rate, $\gamma$ the relative risk aversion (RRA) coefficient, and $\psi$ 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

We assume that the home country’s monetary policy is sufficiently influential to affect real quantities across all countries, in particular their output growth.\(^6\) Moreover, the

\(^5\)Consistently with the empirical analysis, we consider the United States as the home country.

\(^6\)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.
central bank revises its policy every $\bar{A}$ periods. The logarithmic output growth of a
generic country $i$ $\Delta y_{i,t+1} = \log Y_{i,t+1} - \log Y_{i,t}$, is assumed to evolve as follows:

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

$$x_{i,t+1} = \alpha_{i,x} + b_{i,x} x_{i,t} + \sigma_{i,x} \sqrt{x_{i,t}} \omega_{i,t+1}^x, \quad \text{corr}(\epsilon_i, \omega_i^x) = 0, \quad (2)$$

where coefficients ($\sigma_d, \alpha_{i,x}, b_{i,x}, \sigma_{i,x}$) are constant. The policy affects expected output
growth through the common component $\mu_t$, which is assumed constant between policy
announcement (i.e. revision) dates: $\mu_t = m(t_a)$, when $t \in [t_a, t_a + \bar{A})$, for some constant
$m(t_a)$ and announcement date $t_a$. Each country loads differently on the policy factor
with loading $\beta_i$. $\overline{\mu}_i$ is a constant country-specific component of expected growth. While
agents observe $\beta_i$ and the volatility of output growth $x_{i,t}$, which follows a square-root
random process, they do not observe the innovations $\epsilon_{i,t+1}$, nor the expected growth
components ($\overline{\mu}_i, \mu_t$). Between any two policy announcements dates, agents learn about
expected growth in a Bayesian fashion from the realizations of domestic output growth
and a common signal.\footnote{For simplicity we assume an autarchic learning process where agents ignore foreign output growth from their inference.}

Specifically, we assume that this common signal is informative
about the piece-wise constant monetary policy impact, and features stochastic volatility
of the square-root type:

$$s_{t+1} = \mu_t + \sigma_s \sqrt{x_{t}} \epsilon_{t+1},$$

$$x_{t+1} = \alpha_x + b_{x,t} x_{t} + \sigma_x \sqrt{x_{t}} \omega_{t+1}^x, \quad \text{corr}(\epsilon_t, \omega_t^x) = 0.$$

We think of $s$ as a monetary policy signal which captures the numerous speeches done
by officials at central banks. 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 the agents’ learning.

**Proposition 1.** Let $\overline{\mu}_{i,t} = \mathbb{E}[\mu_i^t | I_t] + \beta_i \mathbb{E}[\mu_{i,t} | I_t]$ denote the posterior estimate of country
$i$’s expected output growth.\footnote{$I_t$ is the agent actual information set, which includes present and past observations of $\{\Delta y_{i,t}, s_t\}$.} Under the assumption that agents apply a constant weighting
matrix (the Kalman gain) to the current estimate and new information in their updating
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.

Expected output growth given in (3) resembles the highly persistent growth process assumed by the long-run risk literature (see e.g., Bansal and Yaron (2004)). The dynamics follows a random-walk with two stochastic volatility factors, a global ($x_t$) and a country-specific one ($x_{i,t}$). The volatility of global shocks is affected by the parameter $\beta_i$, the loading of the expected country’s 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 produce consistent evidence for currency returns (see, e.g., Lustig, Roussanov, and Verdelhan (2011), Lustig, Roussanov, and Verdelhan (2014), and Hassan and Mano (2013)) and currency options (see, e.g., Bakshi, Carr, and Wu (2008) and Jurek and Xu (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 countries’

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9That is, $t \in [t_a, t_a + A]$, for a generic announcement date $t_a$.

10To allow for full asymptotic learning, i.e. letting the variance of posterior output growth go to zero, 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 prevents agents to learn fully.
expected output growth $\tilde{\mu}_{i,t}$. 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 has an unexpected component, modeled as a shock to $\tilde{\mu}_{i,t}$ with volatility distinct from and, in our intuition, typically larger than the “normal” news component $x$.\textsuperscript{11} We refer to this volatility, $z$, as “announcement uncertainty” and we model its evolution with a square root process. We assume it to be negatively correlated to the systematic non-announcement shock of $\tilde{\mu}_{i,t}$ ($\epsilon_t$), since negative shocks to expected consumption growth are symptoms of an ineffective current monetary policy, which will be more likely revised, thus raising the uncertainty about the effects of the upcoming announcement.

An important difference between the “normal” news shock $\epsilon$, and the announcement shock $\nu_i$, 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 $\beta_i$ on the monetary policy component of expected growth should react more pronouncedly to announcement news. In summary, the evolution of expected output growth and announcement uncertainty reads as follows:

$$
\tilde{\mu}_{i,t+1} = \tilde{\mu}_{i,t} + \sqrt{\sigma_1 + \sigma_2^2 x_t} \epsilon_{i,t+1} + \beta_i \left( \sqrt{\sigma_2 + \sigma_3^2 x_t} \epsilon_{i,t+1} \right.
+ A_{t+1}\sqrt{z_t \nu_{i,t+1}}) \\
\left. + A_{t+1}\sqrt{z_t \nu_{i,t+1}} \right) \\

z_{t+1} = \alpha_z + b_z z_t + \sigma_z \sqrt{\sigma_2 + \sigma_3^2 x_t} \omega_{i,t+1},
$$

In expression (4), $A_{t+1}$ is a dummy variable that takes a value of one if $t + 1$ is an announcement date, which means that a monetary policy revision has taken place between time $t$ and $t + 1$. The announcement uncertainty $z_t$ is a discrete-time square root process, the volatility of which is proportional to the volatility of the volatility of the “normal” uncertainty $x$, which preserves 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.

\textsuperscript{11}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 agents’ imperfect knowledge of the political cost associated with the implementation of a given policy.
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 carry trade risk premia, both in 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:\(^{12}\)

\[
\begin{align*}
\Delta y_{i,t+1} &= \bar{\mu} + \mu_{i,t} + \sigma_d \nu_{i,t+1}, \\
\mu_{i,t+1} &= b_m \mu_{i,t} + \sigma_m \sqrt{x_{i,t}} \epsilon_{i,t+1} + \beta_i (\sqrt{x_t} \epsilon_{t+1} + A_{\tau=1} \sqrt{z_t} \eta_{i,t+1}), \\
 x_{i,t+1} &= \alpha_x + b_x x_{i,t} + \sigma_x \sqrt{x_{i,t}} \omega_{i,t+1}, \\
 x_{t+1} &= \alpha_x + b_x x_t + \sigma_x \sqrt{x_t} \omega_{t+1}, \\
 z_{t+1} &= \alpha_z + b_z z_t + \sigma_z \sqrt{x_t} \omega_{t+1}, \quad \text{corr}(\omega_{i,t}^2, \epsilon_{t}) = \rho_z < 0
\end{align*}
\]

Our model is not time-homogeneous, because announcements take place on a regular basis, i.e. every \( A \) periods. This allows us to consider the “time to next announcement”, \( \tau \), as a state variable instead of 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 \( \eta_{h,t+1} \) is allowed to be country-specific and cross-sectionally correlated. In particular, for any two countries \( i \) and \( j \), we have \( \text{corr}(\eta_{h,t+1}, \eta_{j,t+1}) = \rho_{ij} \). All other correlations are assumed to be zero. Our model features three global state variables — uncertainty related to “normal” global news \((x)\), announcement uncertainty \((z)\) and time to next announcement \((\tau)\) — and two country-specific state variables — uncertainty related to “normal” local news \((x_i)\) and expected country’s output growth \((\mu_i)\).

\(^{12}\)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 \( \nu_i \) instead of \( \epsilon_i \), are independent from the innovations in long-run risks. The time-varying component of expected output growth is now denoted by \( \mu_i \) instead of \( \tilde{\mu}_{i,t} \). 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 \( \eta_i \) instead of \( \nu_i \). The stochastic volatility components of long-run risks have also been simplified: \( \sqrt{x_{i,t}} \) and \( \sqrt{x_t} \) replace \( \sqrt{\sigma_1 + \sigma_d^2 x_{i,t}} \) and \( \sqrt{\sigma_2^2 + \sigma_d^2 x_t} \), respectively.
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:

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

(10)

where $r^y_{i}$ 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:

$$pc_{i,t} = B_0(i) + B_1\mu_{i,t} + B_2\tau(i)z_t + B_3\tau(i)x_t + B_4x_{i,t},$$

(11)

where we have emphasized the dependence of the deterministic coefficients $B$ on time to next announcement and the country loading $\beta_i$.\textsuperscript{13} We impose the following set of mild assumptions.

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

\textsuperscript{13}Coefficients for $B$ are derived in the Appendix.

$\gamma > 1/\psi$ 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 ($B_3\tau(i) < 0$), 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 $z_t$, ($B_2\tau(i) < 0$). 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 $\beta_i$. As shown in the Appendix, we have $\frac{\partial B_{2,\tau}(i)}{\partial \beta_i} < 0$ and $\frac{\partial B_{3,\tau}(i)}{\partial \beta_i} < 0$, therefore countries with larger $\beta_i$ 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 is more affected by monetary policy are intuitively more sensitive to both “normal” news releases and announced policy revisions, thus 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 $\beta_i$.

4.4 Interest Rates and Carry Trade 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:

$$r_{i,t} = C_0 + C_1 \mu_{i,t} + A_{r-1} C_2(i) z_t + C_{3,\tau}(i)x_t + C_4 x_{i,t}, \quad (12)$$

where coefficients for $C$ are derived in the Appendix. $C_1$ is positive, thus, 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 risk-less asset. Conversely to the market risk premium, the risk-free rate is decreasing in “normal” uncertainty (global: $C_{3,\tau}(i) < 0$; 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, countries with more volatile long-run risks (larger $\beta_i$) have smaller interest rates, once we control for differences in expected growth.\textsuperscript{14}

\textsuperscript{14}Formally, we show in the Appendix that $\frac{C_{3,\tau}(i)}{\partial \beta_i} < 0$ and $\frac{C_2(i)}{\partial \beta_i} < 0$. \hfill 21
Note that index $h$ identifies the home country and $i$ denotes any of the $N$ foreign countries. We 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 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:

$$\Delta q_{i,t+1} = m_{h,t+1} - m_{i,t+1}, \quad (13)$$

where lower case letters denote logarithms. A carry-trade is a one-period zero-investment strategy that invests one unit of domestic currency, financed at the domestic risk-free rate, into the foreign risk-free asset. Thus the logarithmic return of the strategy, in terms of domestic currency, is:

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

Hence, realized carry-trade returns depend on the volatilities of home and foreign stochastic discount factors and on their correlation. 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 shocks to uncertainty in long-run risk. The next Proposition details the risk premium earned in equilibrium by a carry trade strategy.

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

$$\mathbb{E}_t[r_{c,i,t+1}^e] + \frac{1}{2} \text{Var}_t[r_{c,i,t+1}^e] = \gamma^2 \sigma_a^2 + g_x(i, \tau - 1)x_t + A_{\tau=1}g_z(i)z_t + g_h x_{h,t},$$

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

According to equation (2), the carry trade risk premium is linear in the global uncertainty factors and the domestic local 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.\(^{15}\) In particular, it takes the following form:

\[
g_z(i) = (\theta - 1)^2 \rho_i^2 B_h^2 \beta_h (\beta_h - \beta_i \rho_{h,i}).
\] (14)

\( g_z \) 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 \( \beta_h \) and \( \beta_i \), 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 \( \beta_h > \beta_i \), and bad (good) news for the home country conveyed by a policy revision is associated with an appreciation (depreciation) of the exchange rate, thus, a negative (positive) carry trade return. In other words, for low \( \beta_i \) foreign countries, carry trade returns do not hedge monetary (announcement) shocks to long-run risks, so that carry-trade 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 carry-trade announcement premium is negative, unless foreign and domestic announcement shocks are negatively correlated: if a policy revision conveys opposite news for foreign and domestic long-term growth perspective, the sign of the carry trade return is again in line with domestic news, and the risk premium may become positive.\(^{16}\)

The global component of the non-announcement premium is \( g_x(i, \tau - 1)x_t \) and we can decompose it into two parts. The first compensates for carry trade risk inherent to “normal” (i.e. non-announcement) long-run risk shocks:

\[
(\theta - 1)^2 \rho_i^2 B_h^2 \beta_h x_t (\beta_h - \beta_i).
\]

\(^{15}\)It is an abuse of terminology because the term is not solely responsible for the difference between the carry risk premium on announcement and non-announcement days, as the state variable \( \tau \) affects the other components. In the calibration exercise, we find that \( g_z(i)z_t \) is the quantitatively dominating term.

\(^{16}\)Indeed a calibration yields a negative \( \rho_{h,i} \) for high \( \beta_i \) countries, and an announcement carry-trade premium that is positive for all countries.
The second compensates the domestic agent for carry-trade risk due to shocks to uncertainty factors, both non-announcement and announcement, respectively:\(^\text{17}\)

\[
(\theta - 1)^2 \rho_i^2 B_{3,\tau-1}(h) \sigma_x^2 x_t [B_{3,\tau-1}(h) - B_{3,\tau-1}(i)],
\]

\[
(\theta - 1)^2 \rho_i^2 B_i^2 B_{2,\tau-1}(h)^2 \sigma_x^2 x_t [B_{2,\tau-1}(h) - B_{2,\tau-1}(i)].
\]

The intuition for the sign and cross-sectional pattern of these premium components is akin to the announcement premium, with the difference that shocks to foreign and domestic stochastic discount factors are now perfectly correlated, so that only the relative magnitude of \(\beta_i\) matters. In particular, positive carry-trade returns hedge negative “normal” shocks to domestic long-run risks, positive shocks to announcement and non-announcement uncertainty—all of which are considered bad news—when the foreign country has high \(\beta_i\). Conversely, the carry-trade investor in low \(\beta_i\) foreign countries obtains negative returns in conjunction to bad domestic news, thus requiring a positive risk premium. 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 \(g_h x_{h,t}\). It depends only on the home country characteristics, and it compensates the domestic agent for carry-trade risk inherent to local (domestic) long-run risk shocks, and shocks to long-run risk volatility.

Cross-sectional variation in carry-trade risk premia are hence related to a differential loading \(\beta_i\), the loading of expected output growth on the monetary policy effect. This intuition is in line with the data presented in Section 3, where it is shown that carry returns are increasing with the interest rate differential. The model we analyze is consistent with this evidence, because equilibrium risk-free rates are decreasing in \(\beta_i\), according to (12), while carry-trade premia are increasing in the parameter. The next Corollary summarizes this finding.

**Corollary 1.** A carry trade 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

\(^{17}\)That is, shocks to long-run risk volatility factors.
risk-premium component \( A_{t=1} z_t \) 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.

In the following, we consider 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 \((\beta_i > \beta_h)\). The return on this strategy is the \( hml \) factor of Lustig, Roussanov, and Verdelhan (2011). According to Corollary 1, \( hml \) is maximally exposed to non-announcement (global) carry trade risk among currency portfolios, in the sense that it has the highest covariance with the log stochastic discount factor, pertaining “normal” shocks to long-run risks, shocks to non-announcement uncertainty and shocks to announcement uncertainty. This means that \( hml \) has the largest possible non-announcement risk premium. Corollary 1 instead implies that the magnitude of \( hml \)'s announcement premium is ambiguous if announcement shocks of low interest rate countries are negatively correlated with the home country, as in this case carry trade premia of low interest rate countries may be positive and large. The empirical analysis of Section 3, and Figure 1 in particular, show that \( hml \) earns positive mean excess returns also in announcement periods.

5 Calibration

We now calibrate the model parameters by targeting moments of interest rates and exchange rates, carry trade returns, the U.S. equity premium, and foreign exchange option implied volatilities. The set of moments resembles that used in Lustig, Roussanov, and Verdelhan (2011), with the addition of options data and the fact 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 \( \beta \). 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 average carry-trade excess returns, both unconditionally and conditionally to an announcement event. There are eight FOMC announcement days per year, or approximately one every \( \bar{A} = 32 \) days. As
customary in the long-run risk literature, we consider a monthly decision problem, thus targeting monthly-equivalent moments.

We set the subjective discount rate $\delta$ to 0.998, as in Colacito and Croce (2011). Parameter $\overline{\mu}$ coincides with the unconditional U.S. monthly consumption growth equal to 0.0024. 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 U.S. consumption growth, and the country-wise average of the slope in the UIP regression:

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

where $\Delta q_{i,t+1}$ are changes in the exchange rate and $(r_{i,t} - r_{US,t})$ 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 carry-trade return.\(^\text{18}\) As of equity markets moments, we match the unconditional U.S. equity premium. For this purpose, we assume aggregate dividend dynamics of the form:

$$\Delta d_{i,t+1} = \overline{\mu} + \lambda \mu_{i,t} + \sigma_d \nu_{d,i,t+1},$$

with the leverage parameter $\lambda$ set to 3 as in Colacito and Croce (2011).\(^\text{19}\) Finally we match the average implied volatility of one-month European options on exchange rates, for five different strikes.

Table 7 reports the empirical moments, along with their model-implied counterparts and computational details.\(^\text{20}\)

\(^{18}\)This is akin to the unconditional mean of the Dollar factor in Lustig, Roussanov, and Verdelhan (2011).

\(^{19}\)Parameters $\overline{\mu}$ and $\sigma_d$ coincide with their consumption-growth counterparts, while consumption and dividend growth are assumed conditionally uncorrelated. The equilibrium price-dividend ratio is $(p - d)_t = \log \frac{P_t}{D_t} = A_0(i) + A_1 \mu_{i,t} + A_2(i) y_t + A_3(i) x_t + A_4 x_{i,t}$. Coefficients are obtained similarly to those of the price-consumption ratio and they are not reported.

\(^{20}\)We use Monte-Carlo simulations to price FX European put options, although Fourier inversion methods (Carr and Madan (1999)) are also possible, since the system (5)-(9) has affine dynamics (see e.g., Duffie, Pan, and Singleton (2000)).
In a second stage, we introduce heterogeneity in the policy loading $\beta$ and in announcement shock correlations $\rho_{usa,i}$ with the purpose of matching i) the mean return of a portfolio which is long currencies with high-interest rate differentials and short currencies with low interest rate differentials, i.e. the unconditional mean of the hml factor given in (A-42) and, most importantly, ii) the carry trade risk premium component due to announcement risk, both for high and low interest rate differential countries. To this end, we define an interval $[\beta, \overline{\beta}]$ centered around the first stage $\beta$, which is identified with the home-country (U.S.) loading. Countries’ loadings are assumed equally spaced on $[\beta, \overline{\beta}]$. We find corresponding values $\rho_{usa,i}$ and $\overline{\rho}_{usa,i}$, and assume that on $[\beta, \beta_{usa}]$ ([$\beta_{usa}, \overline{\beta}$]) the announcement shock correlation $\rho_{usa,i}$ increases linearly from $\rho_{usa,i}$ ($\overline{\rho}_{usa,i}$) to 1.\textsuperscript{21} Table 8 reports all calibrated parameter values, together with model-implied moments and targeted ones.\textsuperscript{22}

Calibrated values of both $\gamma$ and $\psi$, 6.7 and 10.1 respectively, are reasonably moderate and greater than one, so that the Epstein and Zin representative agent displays a preference for early resolution of uncertainty. The pronounced persistence of expected consumption growth ($b_m = 0.94$) is consistent with the long-run risk nature of the model. As expected, the announcement risk factor is negatively correlated with expected consumption growth ($\rho_y = -0.57$), as the revision of expectations due to an announcement is more uncertain in bad times, when a policy change is likely. The steady-state mean of the long-run risk volatility factor $x$ is an order of magnitude smaller than the corresponding figure for the announcement uncertainty factor $y$: 0.545 opposed to 7.05.\textsuperscript{23}

This feature is mainly due to matching the announcement carry-trade risk premium,

\textsuperscript{21}It is natural to associate the first stage announcement shock correlation to the home country.

\textsuperscript{22}Note that the restrictions $\beta, \overline{\beta} > 0$, and imposing symmetry around $\beta$ for these two parameters, prevents us from exactly matching the second-stage targeted moments.

\textsuperscript{23}We have $E[y_t] = \alpha_y/(1 - b_y)$ $E[x_t] = \alpha_x/(1 - b_x)$. 

27
both for high and low interest rate currencies, which is empirically proxied by the difference between average announcement-day and non announcement-day carry-trade returns. Positive announcement carry-trade premia imply that announcement jumps in the long-run risks of low interest rate (differential) currencies are negatively correlated with their US counterparts \( \rho_{\text{usa},i} = -0.18 \), while the opposite is true for high-interest rate (differential) countries \( \rho_{\text{usa},i} = 0.3 \). Therefore, our model imposes that the revision of long-run beliefs following a policy announcement differs across countries: good news for the home country and low \( \beta \) loading/ high forward discount countries have typically the opposite interpretation for high \( \beta \) loading/ low forward discount countries.

The model overshoots the volatility of the home country consumption growth (0.39% vs 0.22%). The model’s home country real interest rate only roughly reproduces the characteristics of the U.S. rate on announcement and non announcement days, since it overstates the mean and understates the standard deviation, while the persistence of the US real rate is well matched (0.83 vs 0.99). 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 overshoots the average volatility of exchange rates (8.88% vs 2.27%), whereas the average correlation between interest rates is closely matched (0.63 vs 0.67). The model slightly underpredicts both the average unconditional carry trade return (0.39% vs 0.17%), and the unconditional mean of the long-short carry trade portfolio (0.33 % vs 0.41%). Importantly, carry trade announcement risk premia are close to their empirical counterparts both for a portfolio of high forward discount currencies (1.68% vs 1.83%) and for a portfolio of low forward-discount currencies (1.37% vs 1.38%). As for the stock market, the model almost replicates the monthly U.S. equity premium with 0.22% as opposed to 0.26% in the data. The model is also consistent producing a negative slope of the average implied volatility curve of one-month FX-options, however, it overpredicts the implied volatility levels.
5.1 Model-Implied Carry Risk Premia

Using the calibrated parameters from Table 8, we can now study different quantities of interest in our economy. First, we study the carry risk premia in our economy which is the difference between the return of the carry trade on an announcement day and a non-announcement day. Figure 3 plots the risk premia for five interest rate sorted portfolios.

[Insert Figures 3 and 4 here.]

We note that carry return premia increase with the interest rate differential. Table 2 reports the average interest rate differentials for the five portfolios as observed in the data. Note that portfolio 2 has an interest rate differential very close to zero, hence the carry trade premium is almost non-existent. Portfolio 1 with a negative interest rate differential and portfolio 3 with a positive interest rate differential feature a carry premium which is much larger compared to portfolio 2: Each of them has a premium of almost 4 bp per day. As the interest rate differential increases, the carry premium gets larger and peaks at 8 bp per day for portfolio 5. Note that these magnitudes are very similar to those in the data.

The empirical results in Table 5 indicate that carry return premia tend to be higher during bad economic times compared to good times. In the following, we plot model-implied carry premia as a function of US output growth. The results are depicted in Figure 4. We note that carry return premia both for low and high forward discount currencies are larger in bad economic states (negative output growth). For example for low (high) forward discount countries, we find that carry return premia decrease from 6.51 bp to 6.09 bp (7.8 bp to 7.48 bp) or a 6.5% (4.1%) drop.

[Insert Figure 5 here.]

Using the calibrated parameters, we can also check whether the model is able to replicate stylized features found in FX option data. Table 2 reports the summary statistic
of implied volatility and the implied volatility slope for interest rate sorted currencies. In line with the findings in Farhi, Fraiberger, Gabaix, Ranciere, and Verdelhan (2014) and Jurek (2014), we observe that implied volatility increases with the interest rate differential and the slope becomes more negative. In line with the data, we replicate a similar feature in our model depicted in Figure 5. The upper panel plots the average implied volatility for different interest rate sorted portfolios, whereas the lower panel plots the implied volatility slope. We note that in line with the empirical counterparts, implied volatilities become larger and slopes become more negative the larger the interest rate differential. While the quantitative pattern is replicated quite nicely, the numbers for the implied volatilities are too large compared to the data, which is mainly due to the exchange rate volatility which we overshoot in the second stage of the calibration.

6 Conclusion

Using a large cross-section and long time-series of spot and option data on FX, we document the following findings: First, returns to the carry trade earned on FOMC announcement days are on average an order of magnitude larger than on every other day. Moreover, this difference is increasing in the forward discount of the currency. More specifically, we document a 6.28 bp difference (t-statistic of 2.25) for low interest rate currencies, which becomes 8.34 bp (t-statistic of 2.07) for high interest currencies. The wedge between announcement and non-announcement returns becomes even larger during bad economic times: the above differences jump to 9.23 bp (t-statistic of 2.11) and 12.57 bp (t-statistic of 2.03), respectively. The results are robust to the choice of the currency set and to the presence of outliers.

We then study the effect of monetary policy uncertainty on carry trade returns through the lens of an international general equilibrium model with long-run consumption risk. Agents in each country do not know future consumption growth and learn about it by observing past realizations. Each countries’ consumption growth is driven by a common component which is affected by the monetary policy decisions of the central bank. In the model, consumption growth jumps as a consequence of resolution of monetary policy uncertainty when the central banks makes an announcement. Ex-
pected returns to the carry trade earn a risk premium due to the uncertainty about the monetary policy.

Calibrating our model, we find that it fits salient moments of equity, exchange rates, interest rates, and options on exchange rates. Using the calibrated parameters, we find that our model replicates the cross-sectional pattern of carry trade expected returns conditional on announcement days. The model is also consistent with larger carry trade announcement risk premia conditional in bad economic times.

There are several potential avenues of future research. We only analyze the impact of monetary policy announcements but there are many other political uncertainty shocks that affect asset prices. For example, in the spirit of Pástor and Veronesi (2012), we could model both scheduled and unscheduled policy announcements and study in more detail the jump risk premia associated with these. 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 study the impact of announcements on investment flows in the spirit of Colacito, Croce, Ho, and Howard (2014).
Figure 1. Carry Trade Returns on FOMC and Non-FOMC Days

This figure plots average daily carry 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. 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. Data used is daily and running from January 1980 to January 2011.
This figure plots empirical densities for returns on interest rate sorted currency portfolios on FOMC and non FOMC announcement days. The left panel plots the density for low interest rate currencies and the right panel the density for high interest rate currencies. Data used is daily and running from January 1980 to January 2011. There are 250 announcement days and 7,481 non announcement days.
Figure 3. Model-Implied Carry Trade Premia

Using the parameters from Table 8, we plot daily carry-trade risk premia conditional on an announcement day and non-announcement days. The expression for the carry trade risk premium of country $i$ is given by equation (A-32), where the state-variables are evaluated at their unconditional mean.
Using the parameters from Table 8, we plot daily carry-trade risk premia as a function of the US output growth rate.

Figure 4. Model-Implied Carry Trade Premia and Output Growth
Using the parameters from Table 8, we plot the average implied volatilities for different interest rate sorted portfolios (upper panel). The lower panel plots the slope of the implied volatility curve for a European put option with one-month time to maturity. The slope is defined as the difference between the implied volatility for a 90∆ Put strike and the implied volatility of a 10∆ Put strike. All option prices are computed by Monte-Carlo simulation, initializing each state variable at its unconditional mean value. pf1 (pf5) is the portfolio with the lowest (highest) interest rate differential vis-à-vis the domestic country.
Table 1

Carry Trade 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. All returns are excess returns in USD. HML denotes a long-short portfolio that is long Portfolio 5 and short Portfolio 1. Returns are daily (in bp) and sampled over the period January 1980 to January 2011. The Sharpe ratio is annualized.

<table>
<thead>
<tr>
<th></th>
<th>pf1</th>
<th>pf2</th>
<th>pf3</th>
<th>pf4</th>
<th>pf5</th>
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</thead>
<tbody>
<tr>
<td>mean</td>
<td>-0.860</td>
<td>-0.223</td>
<td>0.917</td>
<td>0.156</td>
<td>1.822</td>
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<td>t-stat</td>
<td>-1.74</td>
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<td>1.74</td>
<td>0.26</td>
<td>2.56</td>
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<tr>
<td>stdev</td>
<td>43.440</td>
<td>46.403</td>
<td>46.435</td>
<td>53.287</td>
<td>62.597</td>
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<td>Sharpe ratio</td>
<td>-0.31</td>
<td>-0.08</td>
<td>0.31</td>
<td>0.05</td>
<td>0.46</td>
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<th>pf5</th>
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</thead>
<tbody>
<tr>
<td>mean</td>
<td>-0.837</td>
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<td>0.502</td>
<td>0.729</td>
<td>1.579</td>
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<td>t-stat</td>
<td>-1.43</td>
<td>-0.65</td>
<td>0.78</td>
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<td>stdev</td>
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<td>56.374</td>
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<td>67.005</td>
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<td>Sharpe ratio</td>
<td>-0.26</td>
<td>-0.12</td>
<td>0.14</td>
<td>0.21</td>
<td>0.37</td>
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Table 2
Option Data Summary Statistic

This table reports summary statistics of forward discounts, implied volatility, and the slope of the implied volatility curve for different portfolios sorted according to their forward discount. The slope of the implied volatility curve is defined as the difference between the implied volatility of a $10\Delta$-Call and a $10\Delta$-Put option. Pf1 is the portfolio with the lowest forward discount whereas pf5 is the portfolio of currencies with the highest forward discount. Data is running from January 1996 to January 2011.

<table>
<thead>
<tr>
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<th>pf5</th>
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<tr>
<td>Slope</td>
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<td>-40.03</td>
<td>-59.12</td>
<td>-123.52</td>
<td>-223.88</td>
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<tr>
<td>IV</td>
<td>8.81</td>
<td>9.55</td>
<td>10.09</td>
<td>11.54</td>
<td>14.66</td>
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<tr>
<td>Forward Discount</td>
<td>-2.06</td>
<td>-0.22</td>
<td>1.23</td>
<td>3.74</td>
<td>8.57</td>
</tr>
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</table>
Table 3
Carry Trade Summary Statistics

This table reports summary statistics of carry trade returns on announcement and non-announcement days. Announcement days are when the FOMC releases its interest rate decisions. The sample covers January 1980 to January 2011. All numbers are expressed in daily returns (in bp) except for Sharpe ratios which are annualized.

<table>
<thead>
<tr>
<th></th>
<th>pf1</th>
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<th>pf3</th>
<th>pf4</th>
<th>pf5</th>
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<tbody>
<tr>
<td><strong>ALL CURRENCIES</strong> (FOMC DAYS)</td>
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<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>mean</td>
<td>5.213</td>
<td>3.818</td>
<td>5.352</td>
<td>6.110</td>
<td>9.889</td>
</tr>
<tr>
<td>t-stat</td>
<td>(1.87)</td>
<td>(1.15)</td>
<td>(1.57)</td>
<td>(1.96)</td>
<td>(2.67)</td>
</tr>
<tr>
<td>stdev</td>
<td>44.168</td>
<td>52.416</td>
<td>54.057</td>
<td>49.197</td>
<td>58.652</td>
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<tr>
<td>Sharpe ratio</td>
<td>1.87</td>
<td>1.16</td>
<td>1.57</td>
<td>1.97</td>
<td>2.68</td>
</tr>
<tr>
<td><strong>ALL CURRENCIES</strong> (NON FOMC DAYS)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>mean</td>
<td>-1.068</td>
<td>-0.361</td>
<td>0.767</td>
<td>-0.044</td>
<td>1.546</td>
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<tr>
<td>t-stat</td>
<td>(-2.16)</td>
<td>(-0.69)</td>
<td>(1.46)</td>
<td>(-0.07)</td>
<td>(2.17)</td>
</tr>
<tr>
<td>stdev</td>
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<td>46.186</td>
<td>46.155</td>
<td>53.410</td>
<td>62.713</td>
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<td>Sharpe ratio</td>
<td>-0.39</td>
<td>-0.12</td>
<td>0.26</td>
<td>-0.01</td>
<td>0.39</td>
</tr>
<tr>
<td>diff means</td>
<td>6.28</td>
<td>4.18</td>
<td>4.59</td>
<td>6.15</td>
<td>8.34</td>
</tr>
<tr>
<td>t-stat</td>
<td>(2.25)</td>
<td>(1.40)</td>
<td>(1.54)</td>
<td>(1.80)</td>
<td>(2.07)</td>
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</tbody>
</table>

<table>
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<tr>
<th></th>
<th>pf1</th>
<th>pf2</th>
<th>pf3</th>
<th>pf4</th>
<th>pf5</th>
</tr>
</thead>
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<tr>
<td><strong>DEVELOPED CURRENCIES</strong> (FOMC DAYS)</td>
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<td></td>
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<td></td>
<td></td>
</tr>
<tr>
<td>mean</td>
<td>5.344</td>
<td>4.693</td>
<td>6.323</td>
<td>7.299</td>
<td>10.537</td>
</tr>
<tr>
<td>t-stat</td>
<td>(1.51)</td>
<td>(1.40)</td>
<td>(1.49)</td>
<td>(1.80)</td>
<td>(2.16)</td>
</tr>
<tr>
<td>stdev</td>
<td>56.131</td>
<td>52.871</td>
<td>66.925</td>
<td>64.044</td>
<td>77.249</td>
</tr>
<tr>
<td>Sharpe ratio</td>
<td>1.51</td>
<td>1.41</td>
<td>1.50</td>
<td>1.81</td>
<td>2.17</td>
</tr>
<tr>
<td><strong>DEVELOPED CURRENCIES</strong> (NON FOMC DAYS)</td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>mean</td>
<td>-1.056</td>
<td>-0.554</td>
<td>0.311</td>
<td>0.501</td>
<td>1.265</td>
</tr>
<tr>
<td>t-stat</td>
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<td>(-0.94)</td>
<td>(0.49)</td>
<td>(0.81)</td>
<td>(1.67)</td>
</tr>
<tr>
<td>stdev</td>
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<td>51.611</td>
<td>55.983</td>
<td>54.607</td>
<td>66.632</td>
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<td>Sharpe ratio</td>
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<td>-0.17</td>
<td>0.09</td>
<td>0.15</td>
<td>0.30</td>
</tr>
<tr>
<td>diff means</td>
<td>6.40</td>
<td>5.25</td>
<td>6.01</td>
<td>6.80</td>
<td>9.27</td>
</tr>
<tr>
<td>t-stat</td>
<td>(1.93)</td>
<td>(1.58)</td>
<td>(1.66)</td>
<td>(1.92)</td>
<td>(2.15)</td>
</tr>
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</table>
Table 4
Carry Trade Summary Statistics Winsorized Data

This table reports summary statistics of carry trade returns on announcement and non-announcement days for the whole sample (Panel A) and a winsorized sample (Panel B) 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 1980 to January 2011.

<table>
<thead>
<tr>
<th>pf1</th>
<th>pf2</th>
<th>pf3</th>
<th>pf4</th>
<th>pf5</th>
<th>pf1</th>
<th>pf2</th>
<th>pf3</th>
<th>pf4</th>
<th>pf5</th>
</tr>
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<tr>
<td></td>
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<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td><strong>Panel A: All Observations</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>mean</td>
<td>5.213</td>
<td>3.818</td>
<td>5.352</td>
<td>6.110</td>
<td>9.889</td>
<td>-1.068</td>
<td>-0.361</td>
<td>0.767</td>
<td>-0.044</td>
</tr>
<tr>
<td>t-stat</td>
<td>(1.87)</td>
<td>(1.15)</td>
<td>(1.57)</td>
<td>(1.96)</td>
<td>(2.67)</td>
<td>(-2.16)</td>
<td>(-0.69)</td>
<td>(1.46)</td>
<td>(-0.07)</td>
</tr>
<tr>
<td>stdev</td>
<td>43.168</td>
<td>52.416</td>
<td>54.057</td>
<td>49.197</td>
<td>58.652</td>
<td>43.406</td>
<td>46.186</td>
<td>46.155</td>
<td>53.410</td>
</tr>
<tr>
<td>skew</td>
<td>1.151</td>
<td>1.635</td>
<td>1.480</td>
<td>0.295</td>
<td>1.240</td>
<td>0.274</td>
<td>-0.051</td>
<td>-0.113</td>
<td>-0.502</td>
</tr>
<tr>
<td>N</td>
<td>250</td>
<td>250</td>
<td>250</td>
<td>250</td>
<td>250</td>
<td>7481</td>
<td>7481</td>
<td>7481</td>
<td>7481</td>
</tr>
</tbody>
</table>

| **Panel B: Excluding top and bottom 1%** |     |     |     |     |     |     |     |     |     |
| mean | 5.445 | 3.965 | 5.307 | 6.573 | 10.192 | -1.058 | -0.351 | 0.816 | -0.034 | 1.535 |
| t-stat | (1.91) | (1.17) | (1.52) | (2.07) | (2.69) | (-2.09) | (-0.65) | (1.51) | (-0.05) | (2.09) |
| stdev | 44.563 | 52.963 | 54.672 | 49.547 | 59.245 | 43.422 | 46.193 | 46.269 | 53.539 | 62.937 |
| skew | 1.138 | 1.616 | 1.468 | 0.279 | 1.240 | 0.274 | -0.073 | -0.121 | -0.513 | -2.354 |
| N | 244 | 244 | 244 | 244 | 244 | 7333 | 7333 | 7333 | 7333 | 7333 |
Table 5
Regression Carry Portfolios Announcement Dummy

This table reports estimated coefficients from regressing carry trade portfolios sorted on their interest rate differential on an announcement dummy which takes the value of one on an announcement day and zero otherwise (Panel A). Panel B reports the results when we regress carry portfolios on the announcement dummy and a policy surprise component as in Bernanke and Kuttner (2005). Panel C reports coefficients from regressing carry portfolio returns onto the announcement dummy which is interacted with the Chicago Fed National Activity Index. t-statistics are calculated using Newey and West standard errors and are given in parentheses. Data is daily and runs from January 1980 to January 2011.

<table>
<thead>
<tr>
<th></th>
<th>pf1</th>
<th>pf2</th>
<th>pf3</th>
<th>pf4</th>
<th>pf5</th>
</tr>
</thead>
<tbody>
<tr>
<td><strong>Panel A: Announcement Dummy</strong></td>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>constant</td>
<td>-1.068</td>
<td>-0.361</td>
<td>0.767</td>
<td>-0.044</td>
<td>1.546</td>
</tr>
<tr>
<td>t-stat</td>
<td>(-2.13)</td>
<td>(-0.68)</td>
<td>(1.44)</td>
<td>(-0.07)</td>
<td>(2.13)</td>
</tr>
<tr>
<td>t-stat</td>
<td>(2.22)</td>
<td>(1.25)</td>
<td>(1.33)</td>
<td>(1.94)</td>
<td>(2.21)</td>
</tr>
<tr>
<td>$R^2$ (in %)</td>
<td>0.053</td>
<td>0.012</td>
<td>0.018</td>
<td>0.029</td>
<td>0.043</td>
</tr>
</tbody>
</table>

|                  |       |       |       |       |       |
| **Panel B: Announcement Dummy and Policy Surprise** |       |       |       |       |       |
| constant         | -1.074| -0.384| 0.735 | -0.066| 1.540 |
| t-stat           | (-2.07)| (-0.69)| (1.33)| (-0.10)| (2.14)|
| t-stat           | (2.22) | (1.23) | (1.32) | (1.94) | (2.17) |
| Policy Surprise  | -0.164| -0.380| -0.491| -0.355| -0.191 |
| t-stat           | (-0.71) | (-1.93) | (-1.90) | (-1.82) | (-1.17) |
| $R^2$ (in %)     | 0.050 | 0.051 | 0.090 | 0.050 | 0.037 |

|                  |       |       |       |       |       |
| **Panel C: Announcement Dummy and Economic State** |       |       |       |       |       |
| constant         | -1.000| -0.371| 0.715 | 0.009 | 1.634 |
| t-stat           | (-1.94)| (-0.68)| (1.31)| (0.01)| (2.32)|
| t-stat           | (2.11) | (1.70) | (2.21) | (1.77) | (2.03) |
| $R^2$ (in %)     | 0.055 | 0.054 | 0.111 | 0.038 | 0.047 |
Table 6
Regression Carry Portfolios Other Announcement Dummy

This table reports estimated coefficients from the following regression:

\[ r_i^t = \beta_0 + \beta_1 \times \text{Other Announcement Dummy}_t + \epsilon_t, \]

where \( r_i^t \) are the carry returns on portfolios sorted on their interest rate differential and the announcement dummy is a variable that takes the value of one in an announcement day and zero otherwise. Announcements are total nonfarm payroll employment, the Producer Price Index (PPI), and Consumer Price Index (CPI) all published by the BLS. Data is daily and runs from January 1980 to January 2011.

<table>
<thead>
<tr>
<th></th>
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<th>pf3</th>
<th>pf4</th>
<th>pf5</th>
</tr>
</thead>
<tbody>
<tr>
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<td>-0.310</td>
<td>0.846</td>
<td>0.159</td>
<td>1.597</td>
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<td></td>
<td>(-1.92)</td>
<td>(-0.55)</td>
<td>(1.52)</td>
<td>(0.25)</td>
<td>(2.27)</td>
</tr>
<tr>
<td>Dummy Employment</td>
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<td>2.735</td>
<td>2.032</td>
<td>-0.397</td>
<td>5.194</td>
</tr>
<tr>
<td></td>
<td>(0.35)</td>
<td>(1.01)</td>
<td>(0.73)</td>
<td>(-0.11)</td>
<td>(1.44)</td>
</tr>
<tr>
<td>( R^2 ) (in %)</td>
<td>-0.01</td>
<td>0.00</td>
<td>0.00</td>
<td>-0.01</td>
<td>0.02</td>
</tr>
<tr>
<td>Constant</td>
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<td>-0.267</td>
<td>0.915</td>
<td>-0.130</td>
<td>1.862</td>
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<tr>
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<td>(-2.07)</td>
<td>(-0.48)</td>
<td>(1.65)</td>
<td>(-0.19)</td>
<td>(2.81)</td>
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<td>Dummy CPI</td>
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<tr>
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<td>(1.08)</td>
<td>(0.70)</td>
<td>(0.20)</td>
<td>(1.75)</td>
<td>(-0.08)</td>
</tr>
<tr>
<td>( R^2 ) (in %)</td>
<td>0.00</td>
<td>-0.01</td>
<td>-0.01</td>
<td>0.04</td>
<td>-0.01</td>
</tr>
<tr>
<td>Constant</td>
<td>-0.714</td>
<td>-0.054</td>
<td>1.105</td>
<td>0.378</td>
<td>1.979</td>
</tr>
<tr>
<td></td>
<td>(-1.29)</td>
<td>(-0.10)</td>
<td>(2.02)</td>
<td>(0.57)</td>
<td>(2.90)</td>
</tr>
<tr>
<td>Dummy PPI</td>
<td>-6.398</td>
<td>-2.698</td>
<td>-3.457</td>
<td>-5.025</td>
<td>-2.924</td>
</tr>
<tr>
<td></td>
<td>(-2.51)</td>
<td>(-1.00)</td>
<td>(-1.37)</td>
<td>(-1.60)</td>
<td>(-0.85)</td>
</tr>
<tr>
<td>( R^2 ) (in %)</td>
<td>0.08</td>
<td>0.00</td>
<td>0.01</td>
<td>0.03</td>
<td>0.00</td>
</tr>
</tbody>
</table>
This table reports target moments in the first stage of the calibration, where all countries share the same loading $\beta$. All model-implied moments (second column) are unconditional relative to all state variables, except to $\tau$ where specified. “Average” refers to an average across countries. The last row considers European put options on an ‘average’ exchange rate vis-à-vis the US Dollar. The Black-Scholes implied volatility $BS^{-1}_\sigma \left(x, r^t, r^{US}, q_t, 1m, K\right)$, $x$ refers to the option price, $r^{US}_t$ and $r^t$ to unconditional means of US and average (across-countries) foreign short rates, $q_t = 0$ is the initial average log exchange rate, 1m is the one-month time-to–maturity, and $K$ is the strike, expressed in percentage of put $\Delta$, with 50%$\Delta$ being at-the-money. In the Monte-Carlo simulation for the option price, the other state variables are initialized at their unconditional mean value. We use the following notation: $\Phi = \frac{1}{A}\sum_{\tau=1}^{A} \left( (\theta - 1)(\rho B_1 \beta) \right) \Phi^{1\tau=1} + \frac{1}{2}(\theta - 1)(\rho^2 f(\tau))$. Numbers are monthly.

<table>
<thead>
<tr>
<th>Moment</th>
<th>Model</th>
<th>Data</th>
</tr>
</thead>
<tbody>
<tr>
<td>Expected Consumption Growth US</td>
<td>$\varphi$</td>
<td>0.24%</td>
</tr>
<tr>
<td>Volatility Consumption Growth US</td>
<td>$(\sigma^2 + \frac{1}{4} \varphi \sigma_r [\mu_t</td>
<td>\tau = 1] + \frac{1}{A}\sum_{\tau=1}^{A} \varphi \sigma_r [\mu_t</td>
</tr>
<tr>
<td>Average UIP</td>
<td>$(\sigma^2 + \frac{1}{4} \varphi \sigma_r \sigma_y )^2$</td>
<td>-1.799</td>
</tr>
<tr>
<td>Regression Slope</td>
<td>$\frac{2 \sigma^2 \alpha_x}{\sqrt{\sigma^2 (1-\rho^2)(1-\beta^2) + \frac{1}{4} \sigma^4 \gamma^2}} + \frac{1}{2} \theta (\rhoB_1 \beta)^2 \sigma^2 \alpha_x \sigma_y$</td>
<td></td>
</tr>
<tr>
<td>Expected Real Short Rate US conditional</td>
<td>$(\theta - 1)^2 \rho^3 \left( f(\tau) + B^2 \sigma^2 \right)^2 \frac{\sigma_y}{\sigma_x}$</td>
<td>0.035%</td>
</tr>
<tr>
<td>Expected Real Short Rate US conditional</td>
<td>$(\theta - 1)^2 \rho^3 \left( f(\tau) + B^2 \sigma^2 \right)^2 \frac{\sigma_y}{\sigma_x}$</td>
<td>0.044%</td>
</tr>
<tr>
<td>Volatility Real Short Rate US</td>
<td>$(\sigma^2 + \frac{1}{4} \varphi \sigma_r \sigma_y )^2$</td>
<td>0.245%</td>
</tr>
<tr>
<td>Volatility Real Short Rate US</td>
<td>$(\sigma^2 + \frac{1}{4} \varphi \sigma_r \sigma_y )^2$</td>
<td>0.241%</td>
</tr>
</tbody>
</table>

| Table 7                                   | Calibration: First Stage                                              |       |


Table I (cont’d)

<table>
<thead>
<tr>
<th>Moment</th>
<th>Model</th>
<th>Data</th>
</tr>
</thead>
</table>
| Autocorrelation Real Short Rate US | \[\begin{align*}
&\frac{1}{\tau} \sum_{t=1}^{\tau} \{ \rho A \sigma A \mid \tau = 1 \} = 1 + \frac{\sigma A}{\sigma B} V a r \{ \mu A \mid \tau \neq 1 \} \\
+ &\frac{1}{(\theta - 1)^2} \sum_{t=1}^{\tau} f(r) f(r-\tau) + \frac{1}{(\theta - 1)^2} \sum_{t=1}^{\tau} (\theta - 1)(\rho B_1 B_2)^2 r_{t-1} = 1 \\
+ &\frac{1}{(\theta - 1)^2} f(r) + \frac{1}{(\theta - 1)^2} f(r-\tau) \\
- \tau^2 &/ \left( \frac{\sigma A}{\sigma B} V a r \{ r_{\tau} \mid \tau = 1 \} + \frac{\sigma A}{\sigma B} V a r \{ r_{\tau} \mid \tau \neq 1 \} + \frac{\sigma A}{\sigma B} (E[r_{\tau} \mid \tau = 1] - E[r_{\tau}])^2 \\
+ &\frac{\sigma A}{\sigma B} (E[r_{\tau} \mid \tau \neq 1] - E[r_{\tau}])^2 \right)
\end{align*}\] | 0.99 |
| Average Volatility FX Rates vs US | \[\begin{align*}
&\left( \frac{2\sigma A^2}{\tau} \rho A \right) + \frac{1}{(\theta - 1)^2} \sum_{t=1}^{\tau} f(r) f(r-\tau) + \frac{1}{(\theta - 1)^2} \sum_{t=1}^{\tau} (\theta - 1)(\rho B_1 B_2)^2 r_{t-1} = 1 \\
+ &\frac{1}{(\theta - 1)^2} f(r) + \frac{1}{(\theta - 1)^2} f(r-\tau) \\
- \tau^2 &/ \left( \frac{\sigma A}{\sigma B} V a r \{ r_{\tau} \mid \tau = 1 \} + \frac{\sigma A}{\sigma B} V a r \{ r_{\tau} \mid \tau \neq 1 \} + \frac{\sigma A}{\sigma B} (E[r_{\tau} \mid \tau = 1] - E[r_{\tau}])^2 \\
+ &\frac{\sigma A}{\sigma B} (E[r_{\tau} \mid \tau \neq 1] - E[r_{\tau}])^2 \right)
\end{align*}\] | 2.273% |
| Average Correlation Short Rates | \[\begin{align*}
&\sum_{t=1}^{\tau} \left( \theta - 1 \right)(\rho B_1 B_2)^2 r_{t-1} + \frac{1}{(\theta - 1)^2} f(r) f(r-\tau) + \frac{1}{(\theta - 1)^2} f(r-\tau) \\
- \tau^2 &/ \left( \frac{\sigma A}{\sigma B} V a r \{ r_{\tau} \mid \tau = 1 \} + \frac{\sigma A}{\sigma B} V a r \{ r_{\tau} \mid \tau \neq 1 \} + \frac{\sigma A}{\sigma B} (E[r_{\tau} \mid \tau = 1] - E[r_{\tau}])^2 \\
+ &\frac{\sigma A}{\sigma B} (E[r_{\tau} \mid \tau \neq 1] - E[r_{\tau}])^2 \right)
\end{align*}\] | 67.5% |
| Average Carry-Trade Risk Premium | \[\begin{align*}
&\sum_{t=1}^{\tau} \left( \theta - 1 \right)(\rho B_1 B_2)^2 r_{t-1} + \frac{1}{(\theta - 1)^2} f(r) f(r-\tau) + \frac{1}{(\theta - 1)^2} f(r-\tau) \\
- \tau^2 &/ \left( \frac{\sigma A}{\sigma B} V a r \{ r_{\tau} \mid \tau = 1 \} + \frac{\sigma A}{\sigma B} V a r \{ r_{\tau} \mid \tau \neq 1 \} + \frac{\sigma A}{\sigma B} (E[r_{\tau} \mid \tau = 1] - E[r_{\tau}])^2 \\
+ &\frac{\sigma A}{\sigma B} (E[r_{\tau} \mid \tau \neq 1] - E[r_{\tau}])^2 \right)
\end{align*}\] | 0.174% |
| US Equity risk premium (Ann.) | \[\begin{align*}
&\sum_{t=1}^{\tau} \left( \theta - 1 \right)(\rho B_1 B_2)^2 r_{t-1} + \frac{1}{(\theta - 1)^2} f(r) f(r-\tau) + \frac{1}{(\theta - 1)^2} f(r-\tau) \\
- \tau^2 &/ \left( \frac{\sigma A}{\sigma B} V a r \{ r_{\tau} \mid \tau = 1 \} + \frac{\sigma A}{\sigma B} V a r \{ r_{\tau} \mid \tau \neq 1 \} + \frac{\sigma A}{\sigma B} (E[r_{\tau} \mid \tau = 1] - E[r_{\tau}])^2 \\
+ &\frac{\sigma A}{\sigma B} (E[r_{\tau} \mid \tau \neq 1] - E[r_{\tau}])^2 \right)
\end{align*}\] | 2.27% |
| US Equity risk premium (No Ann.) | \[\begin{align*}
&\sum_{t=1}^{\tau} \left( \theta - 1 \right)(\rho B_1 B_2)^2 r_{t-1} + \frac{1}{(\theta - 1)^2} f(r) f(r-\tau) + \frac{1}{(\theta - 1)^2} f(r-\tau) \\
- \tau^2 &/ \left( \frac{\sigma A}{\sigma B} V a r \{ r_{\tau} \mid \tau = 1 \} + \frac{\sigma A}{\sigma B} V a r \{ r_{\tau} \mid \tau \neq 1 \} + \frac{\sigma A}{\sigma B} (E[r_{\tau} \mid \tau = 1] - E[r_{\tau}])^2 \\
+ &\frac{\sigma A}{\sigma B} (E[r_{\tau} \mid \tau \neq 1] - E[r_{\tau}])^2 \right)
\end{align*}\] | 0.19% |
| Avg. Implied Volatilities | \[BS_{\tau}^{-1} \left( \sum_{t=1}^{\tau} \left( \theta - 1 \right)(\rho B_1 B_2)^2 r_{t-1} + \frac{1}{(\theta - 1)^2} f(r) f(r-\tau) + \frac{1}{(\theta - 1)^2} f(r-\tau) \\
- \tau^2 &/ \left( \frac{\sigma A}{\sigma B} V a r \{ r_{\tau} \mid \tau = 1 \} + \frac{\sigma A}{\sigma B} V a r \{ r_{\tau} \mid \tau \neq 1 \} + \frac{\sigma A}{\sigma B} (E[r_{\tau} \mid \tau = 1] - E[r_{\tau}])^2 \\
+ &\frac{\sigma A}{\sigma B} (E[r_{\tau} \mid \tau \neq 1] - E[r_{\tau}])^2 \right) \right)^{-1} \left( K - e^{-\nu_{1+t+1}} \right)^2 \sum_{t=1}^{\tau} \left( r_{t+1}^{\alpha}, q_1, 1, m, K \right) \right] \]

\[10\Delta Put = 11.78\% \\
25\Delta Put = 11.08\% \\
75\Delta Put = 10.77\% \\
90\Delta Put = 11.24\% \]
Panel A summarizes calibrated parameter values after the first stage. The policy factor loadings $\beta_i$ of individual countries are linearly spaced in the interval $[\underline{\beta}, \overline{\beta}]$. Panel B reports theoretical moments evaluated at calibrated parameters and targeted (empirical) counterparts.

### Panel A: Moments 1st Stage

<table>
<thead>
<tr>
<th>Moment</th>
<th>Model</th>
<th>Data</th>
</tr>
</thead>
<tbody>
<tr>
<td>$\delta$</td>
<td>0.998</td>
<td>6.72</td>
</tr>
<tr>
<td>$\gamma$</td>
<td>10.10</td>
<td>0.0024</td>
</tr>
<tr>
<td>$\psi$</td>
<td>0.00087</td>
<td>0.0017</td>
</tr>
<tr>
<td>$\mu$</td>
<td>0.0017</td>
<td>0.0001</td>
</tr>
<tr>
<td>$\beta_{US}$</td>
<td>0.9391</td>
<td>0.0016</td>
</tr>
<tr>
<td>$\beta$</td>
<td>5.89</td>
<td></td>
</tr>
<tr>
<td>$b_m$</td>
<td>0.1648</td>
<td></td>
</tr>
<tr>
<td>$\sigma_m$</td>
<td>0.1708</td>
<td></td>
</tr>
<tr>
<td>$\alpha_y$</td>
<td>0.0355</td>
<td>0.9340</td>
</tr>
<tr>
<td>$\rho_{y}$</td>
<td>0.2612</td>
<td>-0.5768</td>
</tr>
<tr>
<td>$\sigma_d$</td>
<td>-0.176</td>
<td>0.3</td>
</tr>
<tr>
<td>$\rho_{usa,i}$</td>
<td>0.00001</td>
<td></td>
</tr>
</tbody>
</table>

### Panel B: Moments 2nd Stage

- **Expected Consumption Growth US**: 0.24% (0.24 %)
- **Volatility Consumption Growth US**: 0.41% (0.22 %)
- **Average UIP Regression Slope**: -1.805 (-1.799)
- **Expected Real Short Rate US (Ann.)**: 0.09% (0.035 %)
- **Expected Real Short Rate US (No Ann.)**: 0.19 % (0.044 %)
- **Volatility Real Short Rate US (Ann.)**: 0.08% (0.245 %)
- **Volatility Real Short Rate US (No Ann.)**: 0.05 % (0.241 %)
- **Autocorrelation Real Short Rate US**: 0.83 (0.99)
- **Average Volatility FX Rates vs US**: 8.88% (2.27 %)
- **Average Correlation btw US and Other Short Rates**: 0.63 (0.675)
- **Average Carry-Trade Risk Premium**: 0.39 % (0.17 %)
- **US Equity risk premium**: 0.24% (0.26 %)
- **Average Carry-trade premium (high int. rate)**: 1.68% (1.83 %)
- **Average Carry-trade premium (low int. rate)**: 1.37 % (1.38 %)
- **Average HML Carry-trade premium**: 0.33 % (0.41 %)

**Average Implied Volatilities**

- $10\Delta Put = 29.07\%$ (11.78%)
- $25\Delta Put = 28.77\%$ (11.08%)
- $ATM = 28.33\%$ (10.66%)
- $75\Delta Put = 27.74\%$ (10.77%)
- $90\Delta Put = 27.03\%$ (11.24%)

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Appendix A Proofs and derivations

Proof of Proposition 1

Let \( \mu_{i,t} = (\overline{\mu}^t, \mu_t)' \).

\[
\begin{align*}
\hat{\mu}_{i,t} &= E[\mu_{i,t}|I_t], \\
\hat{\sigma}^2_{i,t} &= E[(\mu_{i,t} - \hat{\mu}_{i,t})(\mu_{i,t} - \hat{\mu}_{i,t})'|I_t]
\end{align*}
\]

(A-1)  

(A-2)

denote the vector of posterior estimates and the matrix of estimation errors, respectively, and \( I_t \) denotes the information set that includes all past realizations of output growth \( \Delta y_i \) and the signal \( s \) until time \( t \). Also let

\[
B_i = \begin{pmatrix} 1 & \beta_i \\ 1 & 0 \end{pmatrix} \quad \text{and} \quad \Sigma_i = \begin{pmatrix} \sigma^2_{x,i,t} & 0 \\ 0 & \sigma^2_{t} \end{pmatrix}.
\]

The bivariate random process \((\Delta y_{i,t+1}, s_{t+1})'\) has a Gaussian one-step ahead conditional density function, so that we can apply the Kalman filtering procedure in discrete time. Writing the system dynamics in innovation form:

\[
\begin{align*}
(\Delta y_{i,t+1}, s_{t+1})' &= B_i \hat{\mu}_{i,t} + \epsilon_{i,t+1}, \\
\hat{\mu}_{i,t+1} &= \hat{\mu}_{i,t} + K_i \epsilon_{i,t+1}, \\
K_i &= \frac{\hat{\sigma}^2_{i,t} B_i^t (B_i \hat{\sigma}^2_{i,t} B_i + \Sigma_i)^{-1}}{\hat{\sigma}^2_{i,t} - \hat{\sigma}^2_{i,t} B_i^t (B_i \hat{\sigma}^2_{i,t} B_i + \Sigma_i)^{-1} B_i \hat{\sigma}^2_{i,t}}, \\
\hat{\sigma}^2_{i,t+1} &= \hat{\sigma}^2_{i,t} (I_2 - B_i' K_i') = \hat{\sigma}^2_{i,t} (I_2 - B_i' K_i'),
\end{align*}
\]

(A-3)  

(A-4)  

(A-5)  

(A-6)

where \( K_i \) is the weight given to current information in the update 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. This implies that the posterior variances decrease deterministically in time, because (A-6) can be rewritten as

\[
\hat{\sigma}^2_{t+1} = \hat{\sigma}^2_{i,t} (I_2 - B_i' K_i') = \hat{\sigma}^2_{i,t} (I_2 - B_i' K_i'),
\]

which implies

\[
\hat{\sigma}^2_{i,t} = \hat{\sigma}^2_{i,0} (I_2 - B_i' K_i')^t.
\]

(A-7)  

Since \( \text{Var}_t[\epsilon_{i,t+1}] = B_i \hat{\sigma}^2_{i,t} B_i' + \Sigma_i \), defining \( \hat{\mu}_{i,t} = E[\overline{\mu}^t|I_t] + \beta_i E[\mu_{i,t}|I_t] = (1,0)B_i \hat{\mu}_{i,t} \) we have

\[
\begin{align*}
\hat{\mu}_{i,t+1} &= \hat{\mu}_{i,t} + (1,0)B_i K_i \sqrt{B_i \hat{\sigma}^2_{i,t} B_i' + \Sigma_i} (\epsilon_{i,t+1}, \epsilon_{i,t+1})', \\
&= \hat{\mu}_{i,t} + \sqrt{\sigma^1 + \sigma^2_{x,i,t} \epsilon_{i,t+1} + \beta_i \sqrt{\sigma^2 + \sigma^2_{x} \epsilon_{i,t+1}},
\end{align*}
\]

(A-8)

24In this way agents assign constant weights to news and prior estimates in their updating rule:

\[
\hat{\mu}_{i,t+1} = (I_2 - K_i B_i) \hat{\mu}_{i,t} + K_i (\Delta y_{i,t+1}, s_{t+1})'
\]
where $\epsilon$ and $\epsilon_i$ 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 $B_i \sigma_{i,t}^2 B'_i$, and defined the main diagonal elements as the constants $\sigma_1$ and $\sigma_2$.

**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):

$$r_{i,t+1}^y = k + \Delta y_{i,t} + \rho_i pc_{i,t+1} - pc_{i,t}, \quad (A-9)$$

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

$$pc_{i,t} = B_0(i) + B_1 \mu_{i,t} + B_2(i) z_t + B_3(i) x_t + B_4 x_{1,t},$$

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

$$E_t \left[ \exp(m_{i,t+1} + r_{i,t+1}^y) \right] = 1, \quad (A-10)$$

where

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

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

$$B_1 = \frac{1 - \frac{1}{\psi} \rho_{i,b}}{1 - \rho_{i,b}},$$

$$B_{2,i}(i) = \frac{1}{2} \theta \rho_i^2 B_1^2 \beta_i^2 \left[ \sum_{j=1}^{\infty} (\rho_i b_z)^j A_{t+j} \right], \quad \rho_i b_z < 1$$

$$\rho b_x B_{3,t+1}(i) - B_{3,t}(i) + \frac{1}{2} \theta \rho_i^2 \left( B_1^2 \beta_i^2 + 2B_1 B_{2,t+1}(i) \beta_i \sigma_z \rho_z + B_{2,t+1}(i) \sigma_z^2 + B_{3,t+1}(i) \sigma_z^2 \right) = 0$$

$$B_4 = \left( \theta \rho_i^2 \sigma_x^2 \right)^{-1} \left[ \left( 1 - \rho_i b_x \right) \pm \sqrt{\left( 1 - \rho_i b_x \right)^2 - \theta^2 \rho_i^4 B_1^2 \sigma_m^2 \sigma_x^2} \right]$$

$$B_{0,i}(i) = \sum_{j=1}^{\infty} B_j \left[ \log \delta + \left( 1 - \frac{1}{\psi} \right) \mu_i + \theta \left( 1 - \frac{1}{\psi} \right)^2 \sigma_x^2 + k + \rho_i (B_{2,t+j}(i) \alpha_z + B_{3,t+j}(i) \alpha_x + B_{4}(\alpha_x)) \right]$$

(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 $\sigma_x$ around the deterministic case ($\sigma_x = 0$). The solution of the third equation then becomes:

$$B_{3,t}(i) = \frac{1}{2} \theta \rho_i^2 \sum_{j=1}^{\infty} (\rho_i b_z)^j f(t + j, i) \quad (A-13)$$

$$f(t + j, i) = \left[ B_1^2 \beta_i^2 + 2B_1 B_{2,t+j}(i) \beta_i \sigma_z \rho_z + B_{2,t+j}(i) \sigma_z^2 \right]$$

(A-14)
We now replace calendar time with the state variable $\tau$, time to next announcement, which takes values $1, 2, \ldots, A$, with $A$ the number of periods between announcements. In the sequel, we introduce time homogeneity in our framework by assuming that policy announcements take place regularly, once every $A$ periods. This allows us to replace calendar time with the state variable $\tau$: periods to next announcement. The time-varying coefficients of the price consumption ratio appearing in (A-12) become:

\[
B_{2,\tau}(i) = (\rho_ib_x)^\tau \frac{\theta \rho_i^2 B_1^2}{2[1 - (\rho_ib_x)^A]} \beta_i^2
\]  
(A-15)

\[
B_{3,\tau}(i) = \frac{\theta \rho_i^2}{2} \left( \frac{\sum_{j=1}^A (\rho_ib_x)^j f(\tau - j, i)}{1 - (\rho_ib_x)^A} \right)
\]  
(A-16)

\[
f(\tau, i) = \left[ B_1^2 \beta_i^2 + 2B_1 B_{2,\tau}(i) \beta_i \sigma_z \rho_z + B_{2,\tau}(i) \beta_i^2 \sigma_z^2 \right]
\]  
(A-17)

\[
B_{0,\tau}(i) = \frac{1}{1 - \rho} \left[ \log \delta \left( 1 - \frac{1}{\psi} \right) \pi + \theta \left( 1 - \frac{1}{\psi} \right)^2 \sigma_d^2 \frac{2}{2} + k + \rho B_4 \sigma_x \right] + \left( \frac{\sum_{j=1}^A \rho (B_{2,\tau-j}(i) \alpha_x + B_{3,\tau-j}(i) \alpha_x)}{1 - \rho^A} \right)
\]  
(A-18)

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

Cross-sectional variation of price-consumption ratio coefficients

We report the partial derivatives of coefficients $B_{2,\tau}(i)$ and $B_{3,\tau}(i)$ in (A-15)-(A-16) with respect to $\beta_i$. Their signs hold under Assumption 1.

\[
\frac{\partial B_{2,\tau}(i)}{\partial \beta_i} = (\rho_ib_x)^\tau \frac{\theta \rho_i^2 B_1^2}{[1 - (\rho_ib_x)^A]} \beta_i < 0
\]  
(A-20)

\[
\frac{\partial B_{3,\tau}(i)}{\partial \beta_i} = \frac{\theta \rho_i^2}{2} \left( \frac{\sum_{j=1}^A (\rho_ib_x)^j}{1 - (\rho_ib_x)^A} \right) \frac{\partial f(\tau - j, i)}{\partial \beta_i} < 0
\]  
(A-21)

\[
\frac{\partial f(\tau, i)}{\partial \beta_i} = \left[ 2B_1 \beta_i + 2B_1 \frac{\partial B_{2,\tau}(i)}{\partial \beta_i} \beta_i \sigma_z \rho_z + 2B_1 B_{2,\tau}(i) \beta_i \sigma_z \rho_z 
\right.
\]
\[
+ 2B_{2,\tau}(i) \frac{\sigma_z^2}{\partial \beta_i} > 0.
\]  
(A-22)

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

\[
r_t = -\mathbb{E}_t[m_{t+1}] - \frac{1}{2} \text{Var}_t[m_{t+1}].
\]  
(A-24)
From (A-11), and using Campbell-Shiller log-linearization of the return to aggregate wealth, we obtain:

\[
m_{i,t+1} = \log \delta - \frac{\bar{r}_t}{\psi} - (\theta)(\theta - 1) \left( 1 - \frac{1}{\psi} \right) \frac{\sigma^2}{2} - \frac{\mu_{i,t}}{\psi} - \frac{A_{r=1}=(\theta-1)\theta_{i,t}}{2} B_1^2 \beta^2 z_t
\]

\[-\frac{1}{2}(\theta - 1)\theta_{i,t}\beta_{f} \left( \tau_{i,1} \right) x_{i,t} - \frac{1}{2}(\theta - 1)\theta_{i,t} \left[ B_2^2 \sigma^2_m + B_3^2 \sigma^2_x \right] x_{i,t}
\]

\[-\gamma \sigma_d \nu_{i,t+1} + (\theta - 1) \rho_{B1} \left[ \sigma_m \sqrt{\bar{x}_{i,t}} \epsilon_{t+1} + \beta_i \sqrt{\bar{x}_{i,t}} \epsilon_{t+1} + A_{r=1} \beta_i \sqrt{\bar{x}_{i,t}} \eta_{i,t+1} \right]
\]

\[+ (\theta - 1) \rho_{B2} \left( \tau_{i,1} \right) \sigma_x \sqrt{\bar{x}_{i,t}} \omega_{i,t+1} + (\theta - 1) \rho_{B3} \left( \tau_{i,1} \right) \sigma_x \sqrt{\bar{x}_{i,t}} \omega_{i,t+1} + (\theta - 1) \rho_{i,t} \beta_{i,t} \sqrt{\bar{x}_{i,t}} \omega_{i,t+1}. \quad (A-25)\]

After computing moments in (A-24) and simplifying terms, we obtain expression (12), where coefficients are:

\[
C_0 = -\log \delta + \frac{\bar{r}_t}{\psi} + \frac{1 - \gamma(\psi+1)}{\psi} \frac{\sigma^2}{2}, \quad (A-26)
\]

\[
C_1 = \frac{1}{\psi} > 0, \quad (A-27)
\]

\[
C_2(i) = \frac{1}{2}(\theta - 1)\rho_{i,t}^2 B_1^2 \beta^2 < 0, \quad (A-28)
\]

\[
C_3(i) = \frac{1}{2}(\theta - 1)\rho_{i,t}^2 f(\tau - 1, i) < 0, \quad (A-29)
\]

\[
C_4 = \frac{1}{2}(\theta - 1)\rho_{i,t}^2 \left[ B_1^2 \sigma^2_m + B_2^2 \sigma^2_x \right] < 0. \quad (A-30)
\]

The signs hold under Assumption 1. It follows immediately that \( \frac{\partial C_2(i)}{\partial \beta} > 0 \) and \( \frac{\partial C_4(i)}{\partial \beta} > 0 \), because \( \frac{\partial f(\tau, i)}{\partial \beta} > 0 \).

**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:

\[
\Delta q_{i,t+1} = m_{h,t+1} - m_{i,t+1} = \frac{\mu_{i,t} - \mu_{h,t}}{\psi} - \frac{A_{r=1}=(\theta-1)\theta_{i,t}}{2} B_1^2 \left( \beta^2_h - \beta^2_i \right) z_t
\]

\[-\frac{1}{2}(\theta - 1)\theta_{i,t} \left[ f(\tau - 1, h) - f(\tau - 1, i) \right] x_{i,t} - \frac{1}{2}(\theta - 1)\theta_{i,t} \left[ \left( B_2^2 \sigma^2_m + B_3^2 \sigma^2_x \right) x_{i,t} \right.
\]

\[-\left( B_1^2 \sigma^2_m + B_3^2 \sigma^2_x \right) x_{i,t}] - \gamma \sigma_d (\nu_{i,t+1} - \nu_{i,t+1}) + (\theta - 1) \rho_{B1} [(\beta_h - \beta_i) \sqrt{\bar{x}_{i,t}} \epsilon_{t+1}]
\]

\[+ A_{r=1} \sqrt{\bar{x}_{i,t}} (\beta_h \eta_{h,t+1} - \beta_i \eta_{i,t+1}) + \sigma_m \sqrt{\bar{x}_{i,t}} \epsilon_{h,t+1} - \sigma_m \sqrt{\bar{x}_{i,t}} \epsilon_{i,t+1}]
\]

\[+ (\theta - 1) \rho_{i} (B_{2,\tau-1}(h) - B_{2,\tau-1}(i)) \sigma_x \sqrt{\bar{x}_{i,t}} \omega_{i,t+1} + (\theta - 1) \rho_{i} (B_{3,\tau-1}(h) - B_{3,\tau-1}(i)) \sigma_x \sqrt{\bar{x}_{i,t}} \omega_{i,t+1}
\]

\[+ (\theta - 1) \rho_{i} B_1 (\sigma_x \sqrt{\bar{x}_{i,t}} \omega_{i,t+1} - \sigma_x \sqrt{\bar{x}_{i,t}} \omega_{i,t+1}). \quad (A-31)\]

The Euler equation for the equilibrium pricing of carry-trade returns in the home country, and the joint conditional Normality of log-stochastic discount factor and carry-trade log re-
Proof of Corollary

Using (A-25) and (A-31) to compute conditional covariances leads to expression (2), where the coefficients read:

\[ g_x(i, \tau - 1) = (\theta - 1)^2 \rho_1^2 B_1^2 \beta_h \beta_i + B_1 \beta_h (B_{2,\tau-1}(h) - B_{2,\tau-1}(i)) \sigma_z \rho_z + B_{2,\tau-1}(h) (B_{2,\tau-1}(h) - B_{2,\tau-1}(i)) \sigma_z^2 + B_{3,\tau-1}(h) (B_{3,\tau-1}(h) - B_{3,\tau-1}(i)) \sigma_z^3 \]

\[ g_z(i) = (\theta - 1)^2 \rho_1^2 B_1^2 \beta_h (\beta_i - \beta_i) \]

\[ g_h = (\theta - 1)^2 \rho_1^2 (B_1^2 \sigma_m^2 + B_1^2 \sigma_x^2) \]

Linearity of the conditional carry trade risk premium in the state variables \( z, x, \) and \( i \) implies that the unconditional premium reads:25

\[ g_x(i, \tau - 1)E[x_t] + A_{\tau-1}g_z(i)E[z_t] + g_hE[x_{i,t}] \] (A-36)

where \( E[x_t] = E[x_{i,t}] = \alpha_x/(1 - b_x) \), and \( E[z_t] = \alpha_z/(1 - b_z) \). We have also made use of the expression of the carry trade 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:

\[ E[z_t|\mu_{h,t}] = E[z_t] + \frac{Cov[\mu_{h,t}, z_t]}{Var[\mu_{h,t}]} (\mu_{h,t} - E[\mu_{h,t}]) \] (A-37)

where

\[ E[\mu_{h,t}] = 0 \] (A-38)

\[ Cov[\mu_{h,t}, z_t] = \frac{\sigma_z \rho_z \beta_h \alpha_x}{(1 - b_x)(1 - b_y b_m)} \] (A-39)

\[ Var[\mu_{h,t}] = \frac{\alpha_x \sigma_m^2 + \beta_i^2}{(1 - b_x)(1 - b_y b_m)} + \frac{1}{A} \frac{\alpha_z \beta_i^2}{(1 - b_z)(1 - b_m)} \] (A-40)

The carry trade risk premium conditional on the home country’s current economic conditions (output growth) is thus

\[ g_x(i, \tau - 1)E[x_t] + A_{\tau-1}g_z(i)E[z_t|\mu_{h,t}] + g_hE[x_{i,t}] \] (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 (12) for the equilibrium interest rate, where, in particular, \( \frac{\partial g_x(i, \tau - 1)}{\partial \beta_i} < 0 \) and \( \frac{\partial g_z(i)}{\partial \theta_i} < 0 \).

\[ ^{25}\text{The only conditioning variable in the expression below is } \tau. \]
The hml and Dollar 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

$$hml_{t+1} = \frac{1}{N_H} \sum_{j \in H} r_{j,t+1}^c - \frac{1}{N_L} \sum_{j \in L} r_{j,t+1}^c.$$  

(A-42)

The innovation component of hml reads

$$hml_{t+1} - E_t[hml_{t+1}] = (\theta - 1)\rho B_1 \left[ (\beta_H - \beta_L) \sqrt{x_t} \epsilon_{t+1} + \sqrt{z_t} A_{\tau = 1}(\eta_{t+1} - \eta_{L,t+1}) \right] +$$

$$+ (\theta - 1)\rho (B_{2,\tau - 1}(H) - B_{2,\tau - 1}(L)) \sigma_z \sqrt{x_t} w_{t+1}^x$$

(A-43)

where $\eta_{H,t+1} = \frac{1}{N_H} \sum_{j \in H} \beta_i \eta_{h,t+1}$, $\overline{\beta} = \frac{1}{N_H} \sum_{j \in H} \beta_j$, $\overline{B_{2,\tau - 1}(H)} = \frac{1}{N_H} \sum_{j \in H} B_{2,\tau - 1}(j)$, $\overline{B_{3,\tau - 1}(H)} = \frac{1}{N_H} \sum_{j \in H} B_{3,\tau - 1}(j)$, and similarly for $L$ terms. We have assumed that all country-specific shocks, which are IID, average out. Exposure of the carry-trade strategy to home-specific long-run risk shocks and shocks to their volatility $x_i$ also demand a time-varying risk premium component in (2), driven by the term $g_h$. As in Lustig, Roussanov, and Verdelhan (2014), a Dollar factor, defined as the return on an equally weighted portfolio long all countries’ carry-trade strategies, proxies for these shocks. If we assume that $\beta_h \approx \frac{1}{N} \sum_{i=1}^N \beta_i$ and $\frac{1}{N} \sum_{i=1}^N \beta_i \eta_{h,t+1} \to \beta_h \eta_{h,t+1}$, the innovation component of the Dollar factor reads:

$$\frac{1}{N} \sum_{i=1}^N r_{i,t+1}^c - E_t \left[ \frac{1}{N} \sum_{i=1}^N r_{i,t+1}^c \right] = \gamma \sigma_d \nu_{h,t+1} - (\theta - 1)\rho B_1 \sqrt{x_{h,t}} \epsilon_{h,t+1}$$

$$+ (\theta - 1)\rho B_4 \sigma_x \sqrt{x_{h,t}} w_{h,t+1}^x.$$  

(A-44)
References


