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## ABSTRACT

We re-examine the Fama (1984) puzzle – the finding that ex post depreciation and interest differentials are negatively correlated, contrary to what theory suggests – for eight advanced country exchange rates against the US dollar, over the period up to February 2016. The rejection of the joint hypothesis of uncovered interest parity (UIP) and rational expectations – sometimes called the unbiasedness hypothesis – still occurs, but with much less frequency. Strikingly, in contrast to earlier findings, the Fama regression coefficient is positive and large in the period after the global financial crisis. However, using survey based measures of exchange rate expectations, we find much greater evidence in favor of UIP. Hence, the main story for the switch in Fama coefficients in the wake of the global financial crisis is mostly a change in how expectations errors and interest differentials co-move, though the risk premium also plays a critical role for safe haven currencies (Japanese yen and Swiss franc).

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## 1. Introduction

Uncovered interest parity – the proposition that anticipated exchange rate changes should offset interest rate differentials – is one of the most central concepts in international finance. At the same time, empirical validation of this concept has proven elusive. In fact, the failure of the joint hypothesis of uncovered interest rate parity (UIP) and rational expectations – sometimes termed the unbiasedness hypothesis – is one of the most robust empirical regularities in the literature. The most commonplace explanations – such as the existence of an exchange risk premium, which drives a wedge between forward rates and expected future spot rates – have little empirical verification.[1]

Several developments have prompted this revisit. First and foremost, the last decade includes a period in which short rates have effectively hit the zero interest rate bound. This point is clearly illustrated in Figure 1 where we plot one-year interest rates for a set of eight selected countries. This development affords us the opportunity to examine whether the Fama puzzle is a general phenomenon or one that is regime-dependent. Indeed, the jury is still out about the impact of the zero lower bound on the relationship between interest rate differentials and exchange rates, for example Fernald et al. (2017) find no clear evidence that the US dollar has become more sensitive since 2014. Second, we now have more indicators for risk aversion for extended periods of time. This potentially allows us to distinguish between competing explanations for the

[1] Engel (1996) surveys the failure of the portfolio balance models and consumption capital asset pricing models. See also Chinn (2006) and more recently Engel (2014) and Chinn and Frankel (2016).

failure of the unbiasedness hypothesis. Specifically, we can examine whether the inclusion of these risk proxies alters the Fama puzzle.[2]

To anticipate our results, we obtain the following findings. First, Fama's (1984) finding that interest rate differentials point in the wrong direction for subsequent ex-post changes in exchange rates is by and large replicated in regressions for the full sample, ranging from January 2000 to February 2016. However, the results change if the sample is truncated to apply to only the most recent decade, the period for which interest rates are essentially close to zero. For that period, interest differentials correctly signal the right direction of subsequent exchange rate changes, but with a magnitude that is altogether not reconcilable with the arbitrage interpretation of UIP. In other words, we obtain positive coefficients at exactly a time of high risk when it would seem less likely that UIP would hold.

We also find that the inclusion of a proxy variable for risk, namely the VIX, results in Fama regression coefficients that are overall similar to those obtained without accounting for risk aversion. This finding suggests that changes in the elevation of risk as measured by the VIX do not explain the Fama puzzle, at least not in a direct linear fashion.

The use of expectations data provides the following insights. First, interest differentials and anticipated exchange rate changes are overall positively correlated, consistent with the proposition that investors tend to equalize, at least partially, returns expressed in common currency terms. Second, in cases where the Fama coefficient

[2] The question of exchange rate developments in light of interest rate differentials is obviously important for policy makers in general (and central bankers in particular, see for instance Coeuré, 2017).

switches sign from negative to positive from pre- to post-crisis, the result arises because the correlation of expectations errors and interest differentials changes substantially. Hence, exchange risk does not appear to be the primary reason why the Fama coefficient has been so large in recent years (although the altered behaviour of exchange risk does play a role).[3]

In the next section we briefly lay out the theory underlying the UIP and Fama regressions, and review the existing literature. In Section 3, we examine the empirical results obtained from estimating the Fama regression. In Section 4 we explore the results dropping the rational expectations assumption. Section 5 presents a decomposition of the components driving the deviation of the Fama coefficient from the posited value of unity. Section 6 concludes.

## 2. Theory and Literature

One of the building blocks of international finance, the concept of uncovered interest parity (UIP) is incorporated into almost all theoretical models. UIP is a no arbitrage profits condition:

$$(1) \qquad E_t^M[s_{t+h} - s_t] = (i_{h,t} - i_{h,t}^*)$$

where $s_{t+h} - s_t$ is the depreciation of the reference currency with respect to the foreign currency from time $t$ to time $t+h$, $i_{h,t}$ and $i_{h,t}^*$ are the interest rates of horizon $h$ at time $t$ of the reference and the foreign country, respectively. $E_t^M$ denotes the market's expectation based on time $t$ information. To fix ideas and to anticipate on the empirical

[3] We performed robustness checks and obtained a broad range of additional results for different horizons, reference countries and time periods. These results can be accessed on https://heipertz.shinyapps.io/uipbcfh-app

results, let $i_{h,t}$ represent the US interest rate, $i^*_{h,t}$ the foreign interest rate (that of the UK, euro area, Japan, etc), and $s_t$ the number of US dollars per foreign currency unit, such that an increase in $s_t$ is a depreciation of the dollar. If the US interest rate, for any maturity *h*, is above Japan's interest rate, i.e. $i_t > i^*_t$, then we should expect the dollar to depreciate at horizon *h*.

In other words, the market's expectation of returns is equalized in common currency terms, so that excess returns are not anticipated *ex ante*. In practice, the most common way in which testing the validity of UIP has been implemented is by way of the Fama regression (Fama, 1984):[4]

(2) $$s_{t+h} - s_t = \alpha + \beta\left(i_{h,t} - i^*_{h,t}\right) + u_{t+h}$$

The OLS regression coefficient β is given by the following expression:

(3) $$\hat{\beta} = \frac{Cov(i_{h,t} - i^*_{h,t}, s_{t+h} - s_t)}{Var(i_{h,t} - i^*_{h,t})}$$

Under the joint null hypothesis of uncovered interest parity and rational expectations, $\beta = 1$, and the regression residual is a true random error term, orthogonal to the interest differential. Note that the intercept $\alpha$ may be non-zero while testing for UIP using equation (2). A non-zero $\alpha$ may reflect a constant risk premium (hence, tests for β = 1 are tests for a time-varying risk premium, rather than risk neutrality per se)

[4] For ease of exposition, log approximations are used. In the empirical implementation, exact formulas are used. We have examined data at three month and one year horizons (h ∈ [3,12]), using monthly data. This means the regression residuals are serially correlated under the null hypothesis of rational expectations and uncovered interest parity. We account for this issue by using robust standard errors. We report results for h=12, in order to conserve space, h=3 results are reported in the appendix.

and/or approximation errors stemming from Jensen's Inequality and from the fact that expectation of a ratio (the exchange rate) is not equal to the ratio of the expectation.

In order to understand the surprising nature of the results for empirical tests of uncovered interest parity, it is helpful to clarify what is to be expected from a Fama regression by isolating the key assumptions necessary to go from equation (1) to regression equation (2). There are three key assumptions for obtaining (2) from (1), as laid out in the following equations:

(4) $$f_{h,t} - s_t = \left(i_{h,t} - i^*_{h,t}\right) - \epsilon^{cip}_{h,t}$$

(5) $$f_{h,t} = E^M_t[s_{t+h}] + \epsilon^{rp}_{h,t}$$

(6) $$s_{t+h} = E^M_t[s_{t+h}] - \epsilon^f_{t+h}$$

When $\epsilon^{cip}_{h,t}$ is zero, then equation (4) indicates that there are no barriers to arbitrage using the forward rate $f_{h,t}$(of horizon h, at time t). In other words, covered interest parity holds, or equivalently, the covered interest differential is zero. This condition applies when capital controls are not relevant, and there are no regulatory or funding constraints.[5] For currency pairs of advanced economies, and for offshore yields, covered interest parity has held up, up until the global financial crisis. Equation (5) indicates that the forward rate is equal to the market's expectation of the future spot rate up to an exchange risk premium term, $\epsilon^{rp}_{h,t}$. This is tautology, unless greater structure is imposed.[6]

[5] See Dooley and Isard (1980) for discussion and Popper (1993) for a review of the pre-2008 experience, in which the covered interest differential is attributed to political risk.
[6] See Engel (1996) for a discussion of how the forward rate and the expected spot rate might deviate even under rational expectations and risk neutrality.

The combination of $\epsilon_{h,t}^{cip} = \epsilon_{h,t}^{rp} = 0$ in Equations (4) and (5) yields uncovered interest rate parity. Only when combined with the assumption of rational expectations, namely $E_t\left(\epsilon_{t+h}^{f}\right) = 0$ in equation (6)[7], does one obtain the regression equation (2), where the regression residual can be interpreted as the forecast error. In general, the $\beta = 1$ hypothesis can be seen to rely upon several moment conditions:

$$(7) \qquad plim(\hat{\beta}) = 1 - \frac{Cov\left(i_{h,t} - i_{h,t}^{*}, \epsilon_{h,t}^{cip}\right)}{Var\left(i_{h,t} - i_{h,t}^{*}\right)} - \frac{Cov\left(i_{h,t} - i_{h,t}^{*}, \epsilon_{h,t}^{rp}\right)}{Var\left(i_{h,t} - i_{h,t}^{*}\right)} - \frac{Cov\left(i_{h,t} - i_{h,t}^{*}, \epsilon_{t+h}^{f}\right)}{Var\left(i_{h,t} - i_{h,t}^{*}\right)}$$

When the covered interest differential is zero, the first covariance term is zero. This has been the conventional approach; however, recent work has documented the fact that covered interest differentials have increased in recent years, and so we do not impose this assumption in our analysis.[8] In the absence of covered interest differentials, as long as there is a time varying risk premium or biased expectations, then $plim(\hat{\beta})$ will deviate from unity.

The literature testing variants of the uncovered interest rate parity hypothesis is vast and varied. Most of the studies fall into the category employing the rational expectations hypothesis; in our lexicon, that means they are tests of the unbiasedness hypothesis. Estimates of equation (6) using horizons for up to one year typically reject the unbiasedness restriction on the slope parameter. For instance, the survey by Froot and

[7] Note that the definition of the expectation or forecast error is the negative of the convention, i.e., actual minus forecast.

[8] More recently, covered interest differentials have widened and remained wide (Borio et al., 2016; Du et al., 2017).

Thaler (1990), finds an average estimate for $\beta$ of -0.88.[9] Bansal and Dahlquist (2000) provide more mixed results, when examining a broader set of advanced and emerging market currencies. They also note that the failure of unbiasedness appears to depend upon whether the US interest rate is above or below the foreign interest rate.[10] Frankel and Poonawala (2010) document that for emerging markets more generally, the unbiasedness hypothesis coefficient is typically more positive.[11]

The poor performance of the interest differential shows up in other ways. At short horizons, the interest differential is outperformed by a random walk model of the exchange rate (Cheung et al., 2005; Cheung et al., 2017). However, at longer horizons, the interest differential does much better than a random walk, mirroring the fewer rejections of the unbiasedness hypothesis at longer horizons documented by Chinn and Meredith (2004).

There is an alternative approach that involves using survey-based data to measure exchange rate expectations. In this case, the error term in equation (6), $\epsilon_{t+h}^{f}$, need not be a true innovation. It could have a non-zero mean, be serially correlated, and perhaps correlated with the interest differential. Froot and Frankel (1989) were early expositors of this approach. In a related vein, Chinn and Frankel (1994) document that it was more difficult to reject UIP for a broad set of currencies when using survey based forecasts.

[9] Similar results are cited in surveys by MacDonald and Taylor (1992) and Isard (1995). Meese and Rogoff (1983) show that the forward rate is outpredicted by a random walk, which is consistent with the failure of the unbiasedness hypothesis.

[10] Flood and Rose (1996, 2002) note that including currency crises and devaluations, one finds more evidence for the unbiasedness hypothesis.

[11] Chinn and Meredith (2004) tested the UIP hypothesis at five year and ten year horizons for the Group of Seven (G7) countries, and found greater support for the UIP hypothesis holding at these long horizons than at shorter horizons of three to twelve months. The estimated coefficient on the interest rate differentials were positive and were closer to the value of unity than to zero in general.

Similar results were obtained by Chinn and Frankel (2016), when extending the data up to 2009, increasing the sample to about 24 years. This pattern of findings suggests that the assumption of rational expectations is not innocuous, and that the examination of the UIP condition both assuming and dispensing with the rational expectation assumption is warranted.

One approach we will not investigate is the bias arising from improper restrictions in the estimation methodology, such as coefficient restrictions when there is substantial persistence (Moore, 1994; Zivot, 2000), unbalanced regressions (Maynard and Phillips, 2001), nonlinearity due to thresholds (Baillie and Kilic, 2006), and issues of cointegration (Chinn and Meredith, 2005).

### 3. Fama Regressions

We collected monthly data for the interest rates and currencies of eight economies -- Canada, Switzerland, Japan, Denmark, Norway, Sweden, UK and the euro area – over the Jan. 2000 - Feb. 2016 period. We examined offshore interest rates of twelve month maturities; the use of offshore interest rates has historically obviated the need to account for the impact of capital controls.[12]

Figure 2 depicts twelve month maturity yield differentials, while Figure 3 shows twelve month depreciations, all over the 2000-2016 period. One of the contrasts clearly highlighted by the two figures is that while yield differentials have shrunk toward zero in

[12] To begin with, we adopt the standard assumption of no default risk. In general, this is believed to hold. During the height of the global financial crisis, counterparty risk was perceived as high (along with liquidity issues), so that covered interest parity did not hold (Coffey et al., 2009; Baba and Packer, 2009).

the wake of the global financial crisis, exchange rate depreciations have not exhibited a comparable compression.

Table 1 reports in Panel A the results from equation (2) at the twelve month horizon, for the full sample. The results are largely in accord with previous findings. In general, the slope coefficients on the interest differential (i.e., the “Fama coefficient”) are negative, although the coefficients are not statistically different from zero in most cases. Given that under the maintained hypothesis the coefficient should be unity, we also test if the coefficients are different from unity. It turns out that only the Swiss Franc differs significantly from one. Even when the coefficients are not significantly different from unity, it is important to recall that the proportion of variation explained is very small.

The Fama regression represents a non-structural relationship. There is little reason to believe the same results will hold over time, in the face of changes in the ways policies are implemented. For instance, as policy regimes change, the expectation formation process will change too. Changes in the general economic environment will also have an impact. The global financial crisis provides an obvious break-point to examine. We carried out various statistical tests to precisely identify the break date. All the eight currencies involved in our analysis exhibit a significant break over the sample, but there is no common date that immediately comes out of the analysis. However, all the currencies show a significant break around the years 2007-08, according to a Chow test. In this respect we decide to choose August 2007 as a common break date, having in mind that the summer 2007 can be considered as the beginning of the Global Financial Crisis, with some turmoil on the US housing market. Indeed, on August 9, 2007, BNP Paribas announced that it was closing three hedge funds that specialised in US mortgage debt.

This event is often considered as one of the first tangible signals of the financial crisis as it was followed by a freeze on the interbank lending market. According to the NBER Dating Committee, the US economic recession started three months later in December 2007. Thus separating the sample into pre- and post-crisis periods with August 2007 as break point, we obtain the results presented in Panel B and Panel C of Table 1. In the pre-crisis period (up to 2007M8), the coefficients are uniformly negative, significantly different from unity.

The most remarkable finding we obtain is that during the post-crisis period (after 2007M9), exchange rate depreciations are strongly, and positively, related to the interest differentials. The estimated coefficients range from 3.1 to 10.9. The null hypothesis of unity is uniformly rejected, except for the Danish krone. The proportion of variation explained is also substantially higher. To our knowledge, the only other study documenting something similar to our findings is Baillie and Cho (2014). However, their analysis only extends up to 2012, and -- unlike the results we obtain -- their estimates are not unambiguously positive at the end of their sample.

To highlight the change in how the relationship between interest differentials and *ex post* depreciations change over time, we focus on the British pound, in Figure 4. The stabilization of the interest differential, compared to pound depreciations, is now obvious. One way to illustrate the contrast pre- and post-crisis, not evident in Figure 4, is to show a scatterplot of depreciation against the yield differential. Figure 5 depicts the data for the two periods. In the pre-crisis period, the slope is negative (as in the conventional wisdom), while in the post-crisis period, it is clearly positive. Another way to illustrate this finding is to show the evolution of the beta coefficients from rolling Fama regressions. Figure 6 shows beta coefficients obtained from regressing the US dollar depreciation over twelve months on interest differentials for rolling windows of three years. Results confirm the switch of signs of coefficients from negative to positive in the post-crisis period. More importantly, most of beta coefficients stay positive in the aftermath of the global financial crisis (with the exception of the Japanese yen and the Norwegian krone.), therefore suggesting that a persistent change in correlations has occurred. Remarkably, this stylized fact holds for various base currencies (see table 1 in the appendix).

These results confront the researcher with at least two questions. The first is the longstanding puzzle of why the bias exists; the second is why the correlation changed so much after the crisis.

With respect to the first question, one approach is to allow for an exchange risk premium, i.e., drop the assumption of $\epsilon_t^{rp} = 0$ (but retain the assumption of $\epsilon_t^{cip} = 0$). Doing so means that the error $u_{t+h}$ in $s_{t+h} - s_t = \alpha + \beta\left(i_{h,t} - i_{h,t}^*\right) + u_{t+h}$ includes a term that is potentially correlated with the interest differential. A potential solution is to include as an additional regressor some variable that proxies for an exchange risk premium, $\epsilon_t^{rp}$. This suggests the following regression equation:[13]

(8) $$s_{t+h} - s_t = \alpha + \beta\left(i_{h,t} - i_{h,t}^*\right) + \gamma Z_t + u_{t+h},$$

[13] If the exchange risk premium is a mean zero random error term, there is no need to include a proxy variable. If however, there is a central bank reaction function that essentially makes the error term correlated with the interest differential (as in a Taylor rule), then the estimates obtained from a simple Fama regression will be biased. Variant of this approach include McCallum (1994), in which the central bank responds to exchange rate depreciation, and Chinn and Meredith (2004), in which exchange rate depreciation feeds into output and inflation gaps that determine central bank policy rates. See also Mark and Wu (1998) and Engel (2014).

where $Z$ is a proxy variable.

We evaluate the results using the VIX as a proxy measure [14]. The VIX is a commonly used measure of (inverse) risk appetite, and has been shown to have substantial explanatory power for exchange rates (Hossfeld and MacDonald, 2015, Ismailov and Rossi, forthcoming) and for excess returns (Brunnermeier et al., 2008, Habib and Stracca, 2012, or Husted *et al.*, forthcoming).[15]

The results of the VIX augmented Fama regressions are reported in Table 2 and are notable in the following sense. The inclusion of the VIX does not alter the basic pattern of results for the Fama coefficient estimates found in Panel A of Table 1. However, the estimate of the VIX coefficient is typically negative, though rarely significant, except for the Canadian dollar and the British pound. This means that when the VIX rises, the dollar appreciates relative to the foreign currency, even after controlling for the interest rate differential. Only in the case of the Japanese yen and the Swiss franc, well known safe haven currencies, does the reverse occur.[16]

## 4. Testing UIP with Survey Data

Another way of testing whether arbitragers equalize expected returns is by dropping the assumption of mean zero expectations error, namely $E_t(\epsilon_{t+1}^f) = 0$ in equation (6). It might be that agents are truly irrational, they use bounded rationality, or

[14] Note that we also evaluate inflation differentials (and industrial production growth differentials) as proxies for a premium, in this case a liquidity premium, in line with Engel et al.'s (2017) model of forward rate bias (and high interest-high value currencies). However, we do not obtain empirical evidence for the usefulness of those variables in explaining the Fama puzzle.

[15] See Berg and Mark (forthcoming) for discussion of uncertainty and the risk premium.

[16] The results are sensitive to the sample period selected. In other results, we have detected a sensitivity of the Fama coefficient to different levels of the VIX, using threshold regression. Hence, while augmenting the Fama regression with the VIX does not alter the estimates of the Fama coefficient, this result does not speak to whether the VIX enters in some nonlinear fashion.

have not completely learned the model governing the economy (or, as in Mark and Wu, 1998, some agents are noise traders).

This means we replace equation (6) with:

$$\hat{s}^M_{t+h} = E^M_t[s_{t+h}] - \epsilon^{Mf}_{t+h} \tag{9}$$

The *observed* survey based measure of the future spot rate, $\hat{s}^M_{t+1}$, equals the market's expectation, up to a mean zero random error.[17] There is no assumption, then, that the *ex-ante* measure will be an unbiased measure of the *ex post* measure.

This substitution leads to the following regression equation (where we have not suppressed the exchange risk premium):

$$\hat{s}^M_{t+h} - s_t = \alpha + \beta\left(i_{h,t} - i^*_{h,t}\right) + u_{t+h} \tag{10}$$

In this case, the regression error impounds the forecast error; there is no guarantee that this forecast error is mean zero, and uncorrelated with the interest differential -- or for that matter, the risk proxy.

We use as measures of expectations survey data sourced from Consensus Forecasts applying from 2003M1 to 2016M2. Notice that survey data availability necessitates a change in the sample period.[18]

The results of the regressions are reported in Table 3. One of the defining features of the results is (1) the point estimates are almost uniformly positive (except for the

[17] In other words, we are assuming Classical measurement error, in line with most other analyses. Constant bias would be impounded in the constant. Time varying bias would be much more problematic.

[18] An additional complication is that the interest rates and exchange rates do not align precisely in this data set. Interest rates are sampled at end-of-month, while exchange rates forecasts are sampled usually at the second Monday of the month by Consensus Forecasts.

Canadian dollar), and (2) coefficients for the Swiss franc and Japanese yen are significantly greater than one, confirming that those currencies are considered as safe havens by practitioners. These results are consistent with those obtained in previous studies using survey data, including Chinn and Frankel (1993) and Chinn and Frankel (2016)[19].

Why are the results so different going from the ex-post to ex-ante measures? The reason is that the two measures of exchange rate depreciation differ widely and that the variation in ex-ante measures is substantially smaller than that of ex-post measures. One way to highlight the difference in volatilities is to note that the scale typically ranges from -0.12 to +0.22 for ex ante depreciations, while for ex-post depreciations the range is -0.30 to +0.34.

Table 3 displays the beta coefficients for both horizons in the pre- and post-crisis periods. Interestingly, the point estimates for twelve month changes do not point to a switch in coefficients before and after the crisis. The Swiss franc and Japanese yen in particular retain their specific status on exchange rate markets.

## 5. Reconciling the Results

Thus far, we have documented the fact that Fama regressions tend to exhibit shifts in the estimated parameters, while the regressions using survey data are less subject to such

[19] Skeptics of survey based measures argue that reported forecasts are read off of interest differentials. Chinn and Frankel (1993) note the pattern of relationship between expected spot rates and forwards was consistent with the idea that survey respondents use other information in judging future exchange rate movements. In addition, Cheung and Chinn (2001) survey foreign exchange traders, and find that interest differentials are only one of the factors that go into forecasts.

shifts. This is suggestive of the idea that the characteristics of the expectations are critical in explaining the structural breaks in the Fama regressions.

To see this point explicitly, consider again the decomposition outlined in equation (7):

$$plim(\hat{\beta}) = 1 - \underbrace{\frac{Cov\left(i_{h,t} - i^*_{h,t}, \epsilon^{cip}_{h,t}\right)}{Var\left(i_{h,t} - i^*_{h,t}\right)}}_{A} - \underbrace{\frac{Cov\left(i_{h,t} - i^*_{h,t}, \epsilon^{rp}_{h,t}\right)}{Var\left(i_{h,t} - i^*_{h,t}\right)}}_{B} - \underbrace{\frac{Cov\left(i_{h,t} - i^*_{h,t}, \epsilon^{f}_{t+h}\right)}{Var\left(i_{h,t} - i^*_{h,t}\right)}}_{C}, \tag{7}$$

where the relevant interest differential correlations with the covered interest differential, exchange risk, and expectation errors are labelled A, B, and C, respectively. From this, it is clear that an increase in the estimated β coefficients could in principle be due to a decrease in A, B, or C. The fact that the use of survey expectations reduces the presence of structural breaks suggests that the C term, involving forecast errors, is of crucial importance.

In order to examine this conjecture more formally, we examine the regression coefficients conforming to A, B, and C, respectively, for the pre- and post-crisis period. Estimates at the twelve month horizon are presented in Figure 7 for five currencies. For the three currencies for which the Fama coefficient switches with the stronger amplitude from pre- to post-crisis – the euro, the sterling and the Canadian dollar, – the big change occurs in the expectations component. This is shown in Figure 7 (a), (d) and (e), respectively. To be concrete, in the pre-crisis period, forecast errors defined as $E_t^M[s_{t+h}] - s_{t+h}$ are positively correlated with $\left(i_{h,t} - i^*_{h,t}\right)$; that correlation is very negative over the last decade. Since these components are subtracted from the value of unity, that drives estimated Fama coefficients from negative to positive values.

Notice that the switch in the risk premium component – the B term -- is quite important in the case of the Swiss franc and Japanese yen. The negative correlation between risk premium and interest rate differentials contributes to about half to deviation to one for the yen (Figure 7(b)) and to two-thirds to deviation to one for the Swiss franc (Figure 7(c)).

The foregoing discussion suggests that the reason the Fama puzzle has evolved in the post-crisis period is mainly because of a change in how expectations errors co-move with interest differentials – the C component – during this specific period of time. However, for currencies identified as safe havens, namely Swiss franc and Japanese yen, we find that the way the risk premium behaves, insofar as it co-moves with the interest differential (the B component), is of primary importance.

What lies behind the change in the C component? For all of the currencies – save the Swiss franc and Japanese yen – the forecast errors as defined in equation (6) change from significantly negative in the pre-crisis period to insignificantly different from zero in the post-crisis period. In words, that means that in the 2003-2007M8 period, the dollar depreciated more than anticipated.

## 6. Conclusions

Our extensive cross-currency analysis of uncovered interest parity has yielded new empirical results that will establish a new set of stylized facts.

First, the bivariate relationship between ex-post depreciation and interest differentials, as summarized in the Fama regression, is subject to breaks. While such breaks have shown up in previous studies, the break associated with the global financial crisis and the subsequent period of low interest rates is quantitatively and qualitatively much more pronounced. The positive, albeit very large, Fama regression coefficient detected in the last decade is not consistent with uncovered interest parity. Moreover, even if the coefficient magnitude were consistent with UIP, the finding would run counter to the intuition that UIP should hold when risk is not important, either because the environment is not “risky”, or because agents are risk neutral.

Second, we find that the inclusion of a proxy variable for risk, in the form of the VIX, results in Fama regression coefficients that are largely unchanged. An elevated VIX typically appreciates the dollar, with few exceptions. Hence, the Fama puzzle is not explained by risk, at least when proxied by the VIX in a linear specification.

Third, uncovered interest parity regressions estimated using survey data are less indicative of breaks. That finding suggests that the breakdown in the Fama relationship is related to the nature of expectations errors. Surveys also confirm that practitioners consider the Swiss franc and the Japanese yen as safe haven currencies, including during the post-crisis period.

Fourth, a formal decomposition of deviations from the posited value of unity in the Fama regression indicates that the switch in signs from pre- to post-crisis can be attributed to a large extent to the switch in the nature of the co-movement between expectations errors and interest differentials. This finding implies that the change in the Fama coefficients is not necessarily a durable one. In contrast, the behaviour of safe haven currencies has also been sensitive to the way risk premium co-moves with the interest differential.

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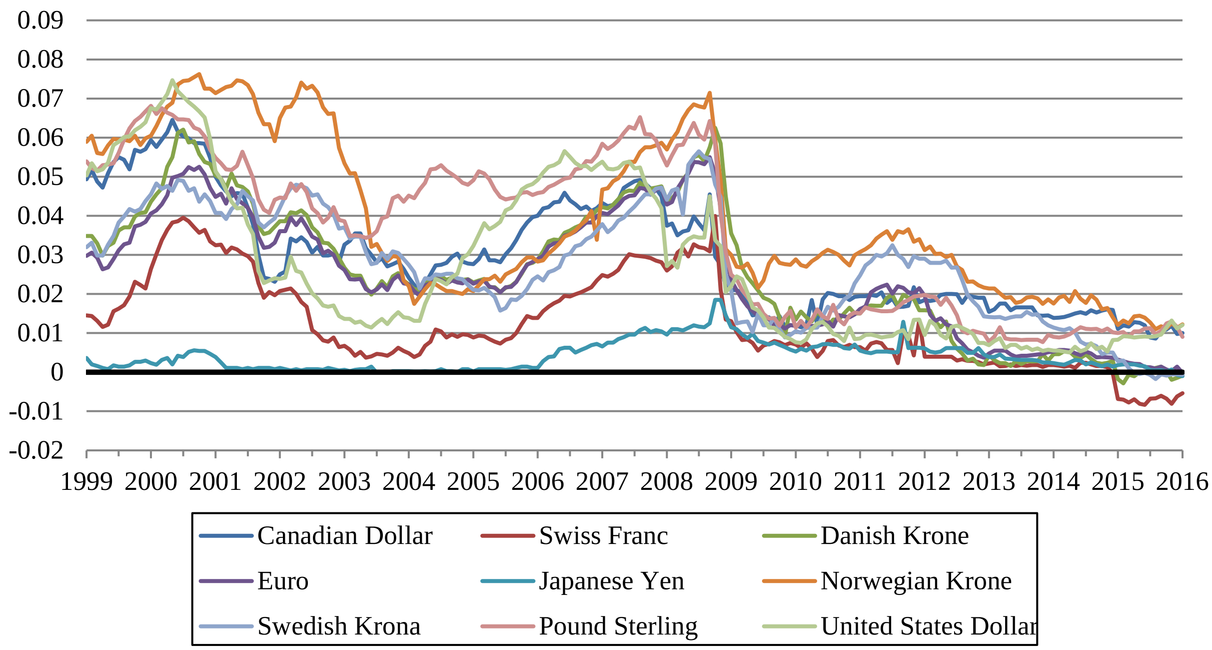



**Figure 1:** ***Interest Rates on 1Y-Eurocurrency Deposits***

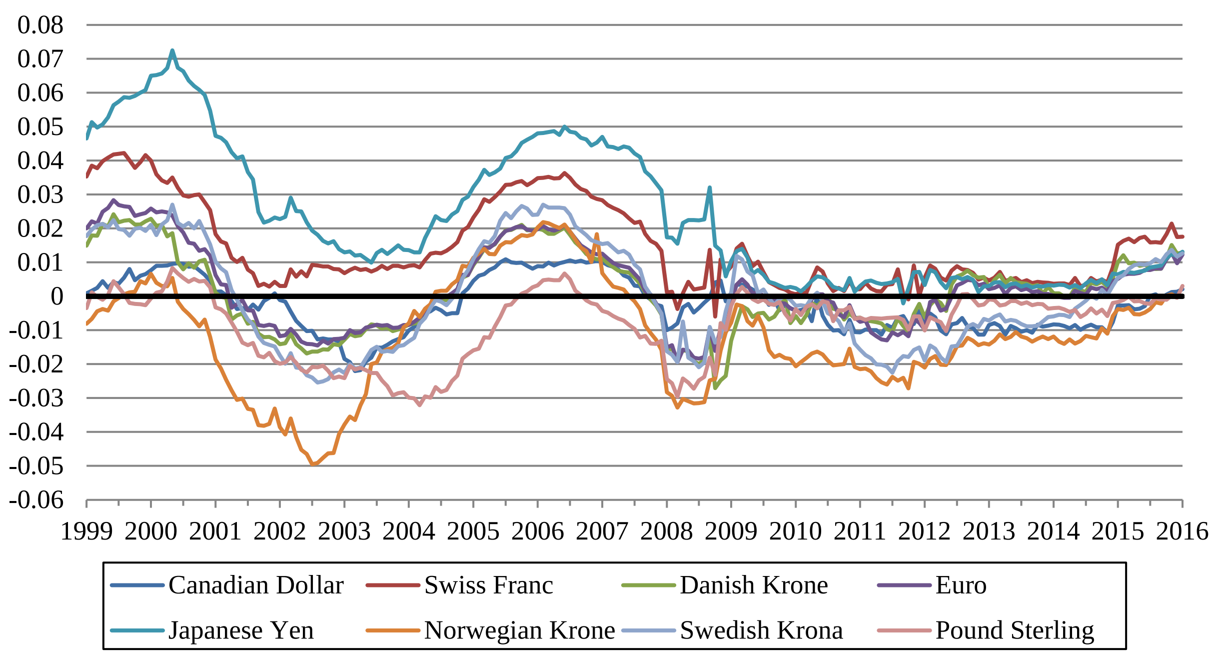



**Figure 2:** ***1Y-Eurocurrency Deposit Rates Differential (US Dollar minus Foreign Currency)***

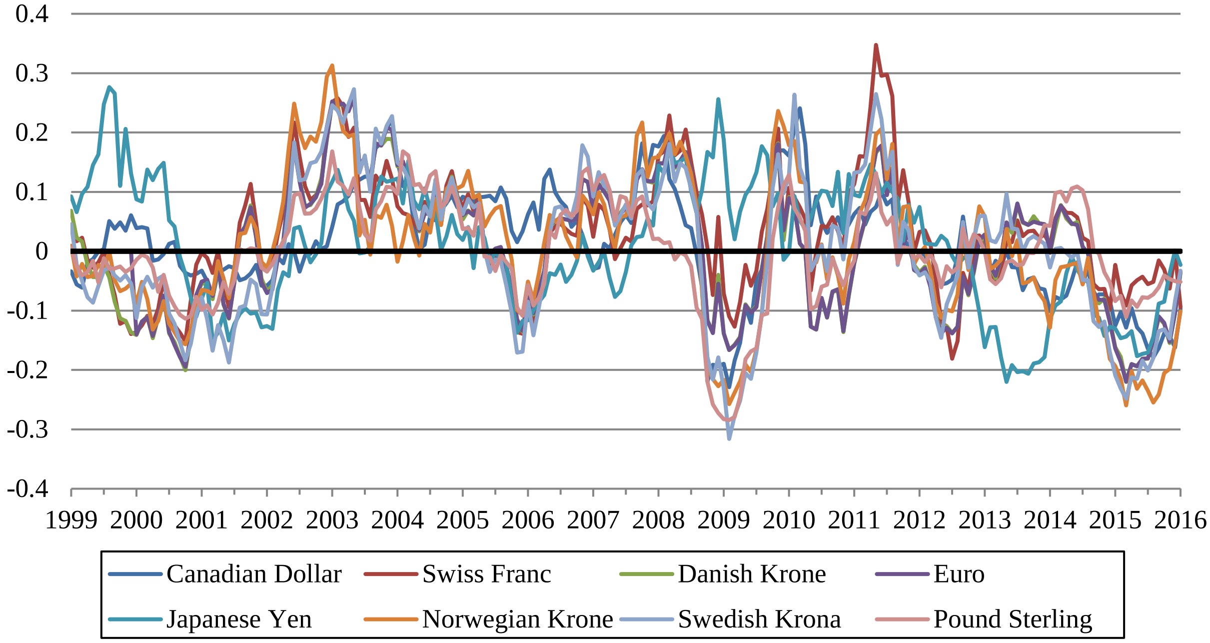



**Figure 3:** ***1Y-Ex-Post Depreciation Rate of the US Dollar w.r.t. Foreign Currency (Positive values indicate depreciations)***

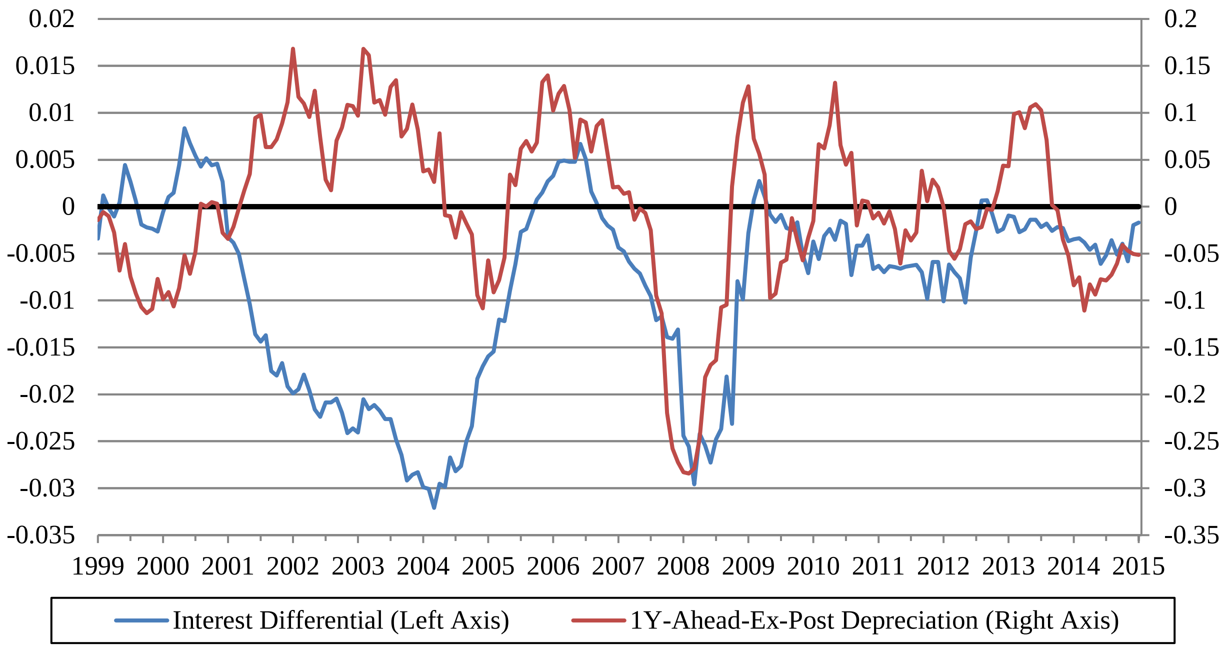
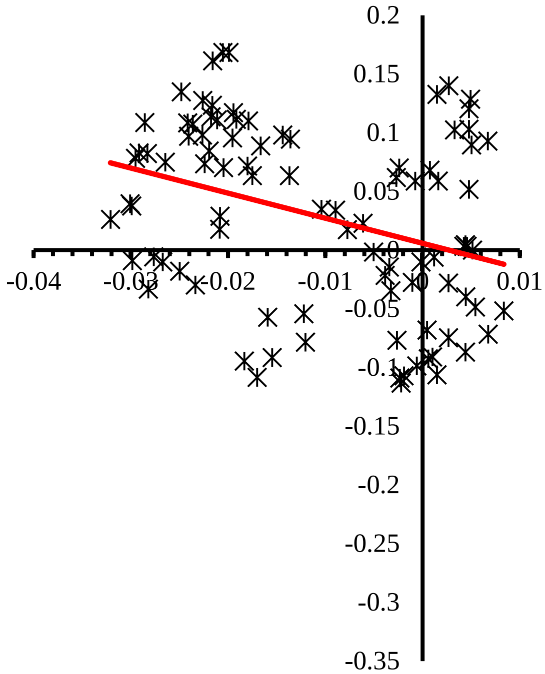
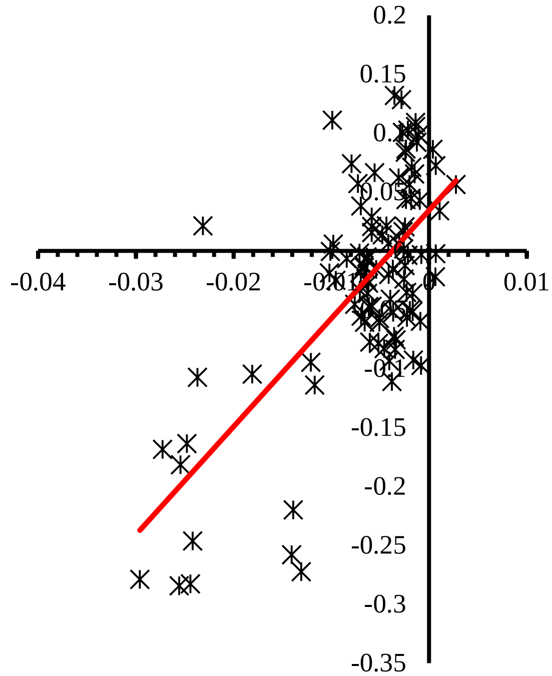



**Figure 4:** ***1Y-Eurocurrency Deposit Rates Differential and 1Y-Ex-Post Depreciation Rate of the US Dollar w.r.t. Pound Sterling***




*Panel (a): Pre-Crisis*




*Panel (b): Post-Crisis*

**Figure 5:** ***Linear Fit of the 1Y-Ex-Post Depreciation Rate (1Y-Ahead) on 1Y-Eurodeposit Rates Differential of US Dollar w.r.t. Pound Sterling***

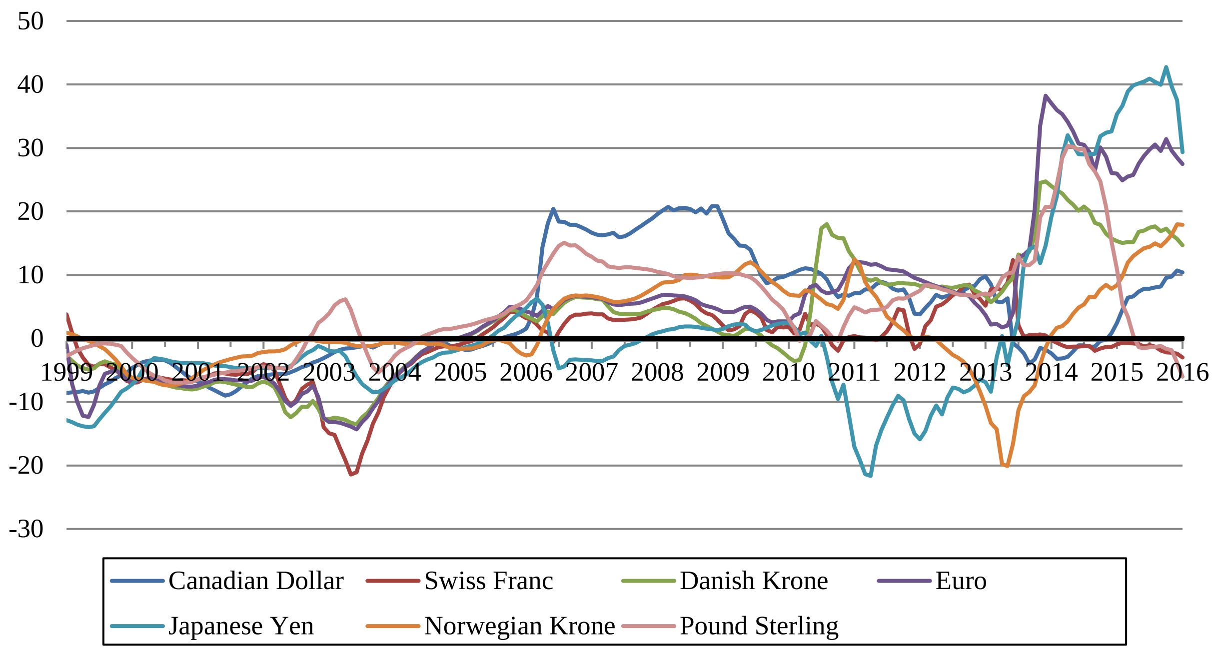
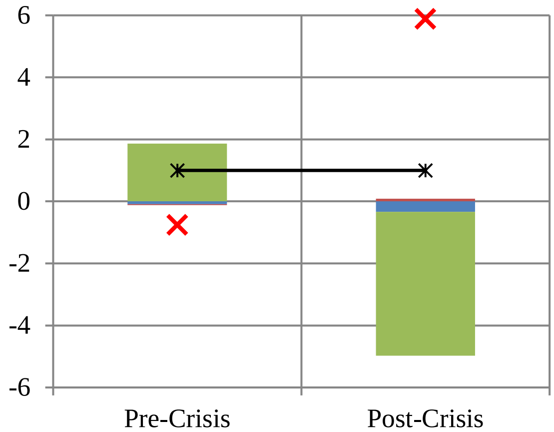
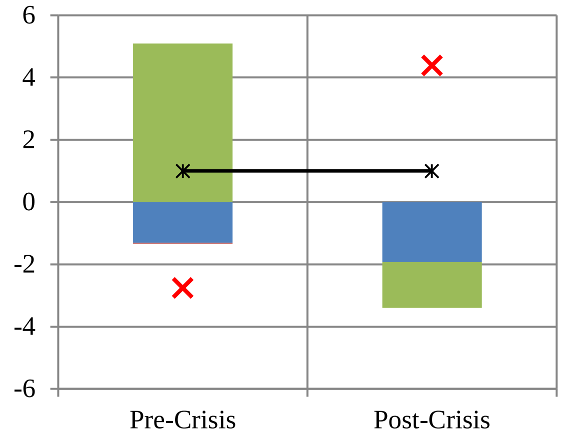


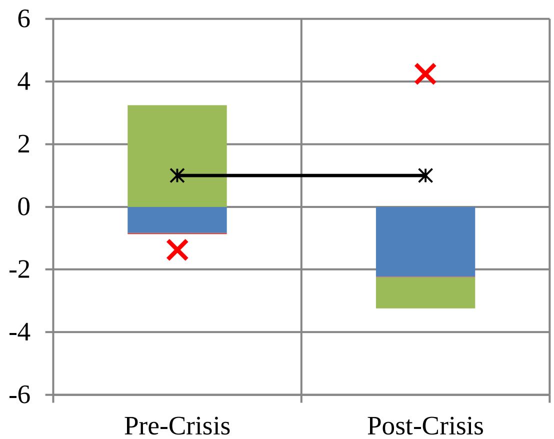

**Figure 6:** ***Estimates of Beta from a 1Y-horizon Fama Regression w.r.t. the US Dollar on Centred 3Y-Rolling Windows (timing refers to interest differentials)***



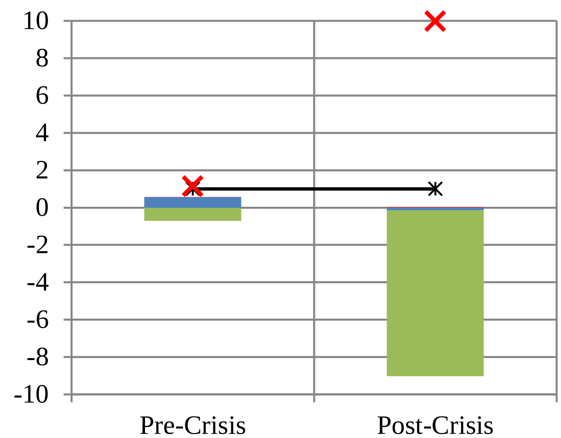

***(a) Euro***



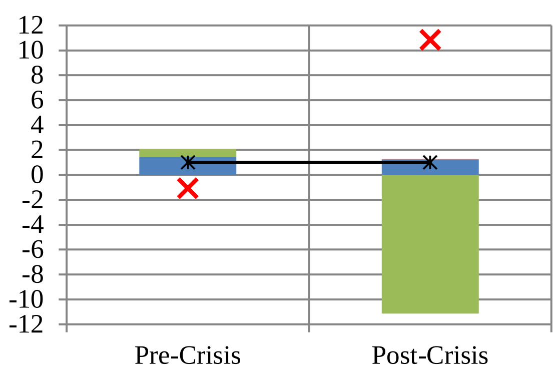

***(b) Japanese Yen***



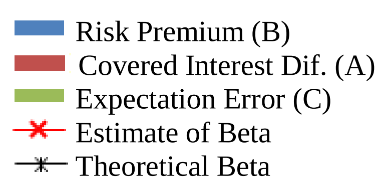

***(c) Swiss Franc***




***(d) Pound Sterling***




***(e) Canadian Dollar***




***Pre-Crisis: 2003M1 – 2007M8***
***Post-Crisis: 2007M9 – 2016M2***

**Figure 7:** ***Decomposition of the Deviation from Unity of Estimates of Beta from 1Y-horizon Fama Regressions w.r.t. the US Dollar***

**Table 1:** ***Fama Regression Results for the Full, Pre-Crisis and Post-Crisis Samples***

| | *Canadian Dollar* | *Swiss Franc* | *Danish Krone* | *Euro* | *Japanese Yen* | *Norwegian Krone* | *Swedish Krona* | *Pound Sterling* |
|---|---|---|---|---|---|---|---|---|
| ***PANEL (A): Full Sample 2000M1 – 2016M2*** | | | | | | | | |
| *Constant* | 0.013 | 0.049* | 0.003 | 0.005 | -0.007 | -0.006 | -0.002 | -0.001 |
| | (0.017) | (0.027) | (0.021) | (0.022) | (0.037) | (0.027) | (0.025) | (0.017) |
| *Beta* | 1.838 | -1.525* | -1.871 | -1.500 | 0.117 | -0.177 | -0.953 | 0.351 |
| | (2.274) | (1.332) | (1.876) | (1.887) | (1.134) | (1.448) | (1.709) | (2.290) |
| *Adj.R^2* | 0.020 | 0.035 | 0.035 | 0.021 | -0.005 | -0.005 | 0.007 | -0.004 |
| *P-Value of F-Statistic* | 0.026 | 0.005 | 0.005 | 0.024 | 0.758 | 0.748 | 0.121 | 0.599 |
| *Number of Observations* | 194 | 194 | 194 | 194 | 194 | 194 | 194 | 194 |
| ***PANEL (B): Pre-Crisis 2000M1 – 2007M8*** | | | | | | | | |
| *Constant* | 0.035*** | 0.129*** | 0.048*** | 0.061*** | 0.085*** | 0.013 | 0.039* | 0.003 |
| | (0.012) | (0.023) | (0.016) | (0.016) | (0.027) | (0.023) | (0.021) | (0.027) |
| *Beta* | -3.631*** | -4.85*** | -5.065*** | -5.107*** | -2.564*** | -2.000*** | -4.034*** | -2.089** |
| | (1.291) | (0.969) | (1.409) | (1.171) | (0.808) | (0.936) | (1.302) | (1.427) |
| *Adj.R^2* | 0.295 | 0.435 | 0.424 | 0.463 | 0.284 | 0.189 | 0.366 | 0.105 |
| *P-Value of F-Statistic* | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.001 |
| *Number of Observations* | 92 | 92 | 92 | 92 | 92 | 92 | 92 | 92 |
| ***PANEL (C): Post-Crisis 2007M9 – 2016M2*** | | | | | | | | |
| *Constant* | 0.031 | -0.007 | -0.016 | -0.013 | -0.054 | 0.047 | 0.003 | 0.034** |
| | (0.021) | (0.027) | (0.025) | (0.023) | (0.036) | (0.040) | (0.030) | (0.016) |
| *Beta* | 10.857*** | 4.239** | 3.147 | 5.888** | 4.387*** | 5.863*** | 5.145** | 9.983*** |
| | (1.975) | (1.505) | (2.875) | (2.166) | (1.244) | (1.782) | (1.777) | (2.325) |
| *Adj.R^2* | 0.379 | 0.115 | 0.055 | 0.168 | 0.253 | 0.166 | 0.150 | 0.473 |
| *P-Value of F-Statistic* | 0.000 | 0.000 | 0.010 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| *Number of Observations* | 102 | 102 | 102 | 102 | 102 | 102 | 102 | 102 |

***Note****: Significance tests relate to the null hypothesis that the intercept is null and slope equal to one. *(**)[***] denotes significance at the 10%(5%)[1%] marginal significance level. The F-statistic refers to the joint null hypothesis that the intercept is null and slope equal to one.*

**Table 2:** ***Augmented Fama Regression Results Using the VIX as Proxy for the Risk Premium for the Full Sample (2000M1 – 2016M2)***

| | ***Canadian Dollar*** | ***Swiss Franc*** | ***Danish Krone*** | ***Euro*** | ***Japanese Yen*** | ***Norwegian Krone*** | ***Swedish Krona*** | ***Pound Sterling*** |
|---|---|---|---|---|---|---|---|---|
| *Constant* | 0.021 | 0.049* | 0.004 | 0.006 | -0.004 | 0.001 | 0.003 | 0.006 |
| | (0.015) | (0.026) | (0.020) | (0.021) | (0.036) | (0.026) | (0.023) | (0.018) |
| *Beta* | 3.165 | -1.540* | -1.744 | -1.386 | -0.085 | 0.080 | -0.474 | 0.624 |
| | (2.103) | (1.318) | (1.888) | (1.880) | (1.090) | (1.555) | (1.754) | (1.990) |
| *Gamma (VIX)* | -0.099** | 0.003 | -0.022 | -0.026 | 0.04 | -0.061 | -0.086 | -0.087* |
| | (0.040) | (0.033) | (0.045) | (0.044) | (0.034) | (0.059) | (0.054) | (0.044) |
| *Adj.R^2* | 0.208 | 0.030 | 0.037 | 0.026 | 0.015 | 0.030 | 0.077 | 0.143 |
| *P-Value of F-Statistic* | 0.000 | 0.020 | 0.010 | 0.029 | 0.084 | 0.021 | 0.000 | 0.000 |
| *Number of Observations* | 194 | 194 | 194 | 194 | 194 | 194 | 194 | 194 |

***Note****: Significance tests relate to the null hypothesis that the intercept is null, the slope equal to one and the VIX coefficient is null. *(**)[***] denotes significance at the 10%(5%)[1%] marginal significance level. The F-statistic refers to the joint null hypothesis that the intercept is null and slope equal to one.*

**Table 3:** ***UIP Regressions Results Using Survey Data on Exchange Rate Expectations for the Full, Pre-Crisis and Post-Crisis Samples***

| | *Canadian Dollar* | *Swiss Franc* | *Danish Krone* | *Euro* | *Japanese Yen* | *Norwegian Krone* | *Swedish Krona* | *Pound Sterling* |
|---|---|---|---|---|---|---|---|---|
| ***PANEL (A): Full Sample 2003M1 – 2016M2*** | | | | | | | | |
| *Constant* | -0.007* | -0.061*** | -0.013** | -0.014** | -0.058*** | 0.027*** | 0.023*** | -0.006 |
| | (0.004) | (0.008) | (0.006) | (0.006) | (0.009) | (0.004) | (0.006) | (0.006) |
| *Beta* | -0.327*** | 3.481*** | 1.388 | 1.699 | 3.167*** | 1.332 | 1.395 | 0.424 |
| | (0.308) | (0.339) | (0.559) | (0.528) | (0.270) | (0.237) | (0.395) | (0.390) |
| *Adj.R^2* | 0.001 | 0.483 | 0.133 | 0.180 | 0.657 | 0.233 | 0.176 | 0.013 |
| *P-Value of F-Statistic* | 0.286 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.092 |
| *Number of Observations* | 146 | 146 | 146 | 146 | 146 | 146 | 146 | 146 |
| ***PANEL (B): Pre-Crisis 2003M1 – 2007M8*** | | | | | | | | |
| *Constant* | -0.003 | -0.008 | 0.012** | 0.012** | -0.018 | 0.026*** | 0.045*** | 0.005 |
| | (0.004) | (0.013) | (0.006) | (0.006) | (0.013) | (0.005) | (0.007) | (0.006) |
| *Beta* | -0.468*** | 1.839* | 1.127 | 1.096 | 2.337*** | 1.151 | 0.689 | 0.458* |
| | (0.268) | (0.497) | (0.375) | (0.374) | (0.341) | (0.200) | (0.316) | (0.284) |
| *Adj.R^2* | 0.028 | 0.304 | 0.198 | 0.182 | 0.612 | 0.272 | 0.085 | 0.044 |
| *P-Value of F-Statistic* | 0.144 | 0.000 | 0.001 | 0.002 | 0.000 | 0.000 | 0.031 | 0.093 |
| *Number of Observations* | 44 | 44 | 44 | 44 | 44 | 44 | 44 | 44 |
| ***PANEL (C): Post-Crisis 2007M9 – 2016M2*** | | | | | | | | |
| *Constant* | -0.009 | -0.067*** | -0.026*** | -0.025*** | -0.062*** | 0.033*** | 0.015*** | -0.005 |
| | (0.007) | (0.008) | (0.005) | (0.005) | (0.009) | (0.008) | (0.005) | (0.007) |
| *Beta* | -0.276* | 3.247*** | 0.666 | 1.245 | 2.922*** | 1.685 | 1.527 | 1.097 |
| | (0.714) | (0.454) | (0.639) | (0.620) | (0.363) | (0.451) | (0.467) | (0.536) |
| *Adj.R^2* | -0.007 | 0.329 | 0.019 | 0.068 | 0.546 | 0.174 | 0.135 | 0.046 |
| *P-Value of F-Statistic* | 0.598 | 0.000 | 0.091 | 0.005 | 0.000 | 0.000 | 0.000 | 0.017 |
| *Number of Observations* | 102 | 102 | 102 | 102 | 102 | 102 | 102 | 102 |

***Note****: Significance tests relate to the null hypothesis that the intercept is null and slope equal to one. *(**)[***] denotes significance at the 10%(5%)[1%] marginal significance level. The F-statistic refers to the joint null hypothesis that the intercept is null and slope equal to one.*

**Appendix Table 1:** ***Estimated Fama Coefficients for the Full, Pre-Crisis and Post-Crisis Samples for Various Base Currencies***

***(12 month horizon)***

| | *Canadian Dollar* | *Swiss Franc* | *Danish Krone* | *Euro* | *Japanese Yen* | *Norwegian Krone* | *Swedish Krona* | *Pound Sterling* | *US Dollar* |
|---|---|---|---|---|---|---|---|---|---|
| ***PANEL (A): Full Sample 2000M1 – 2016M2*** | | | | | | | | | |
| *US Dollar* | 1.838 | -1.525* | -1.871 | -1.500 | 0.117 | -0.177 | -0.953 | 0.351 | / |
| *Japanese Yen* | 0.121 | 0.323 | 0.304 | 0.231 | / | 0.238 | 1.238 | 1.548 | 0.117 |
| *Euro* | -1.524 | -4.868*** | 0.060*** | / | 0.231 | 1.132 | 0.061 | 1.008 | -1.500 |
| *Pound Sterling* | 4.464*** | -0.334 | 0.632 | 1.008 | 1.548 | 0.845 | 0.160 | / | 0.351 |
| ***PANEL (B): Pre-Crisis 2000M1 – 2007M8*** | | | | | | | | | |
| *US Dollar* | -3.631*** | -4.850*** | -5.065*** | -5.107*** | -2.564*** | -2.000*** | -4.034*** | -2.089** | / |
| *Japanese Yen* | 1.287 | -4.573*** | -2.086*** | -3.494*** | / | 0.044 | 0.234 | 2.010 | -2.564*** |
| *Euro* | -6.629*** | -6.172*** | -0.159*** | / | -3.494*** | 0.751 | -1.974* | -3.282*** | -5.107*** |
| *Pound Sterling* | 4.009 | -2.605*** | -2.655*** | -3.282*** | 2.010 | -0.262 | -3.162*** | / | -2.089** |
| ***PANEL (C): Post-Crisis 2007M9 – 2016M2*** | | | | | | | | | |
| *US Dollar* | 10.857*** | 4.239** | 3.147 | 5.888** | 4.387*** | 5.863*** | 5.145** | 9.983*** | / |
| *Japanese Yen* | 5.340 | 6.331** | 3.739 | 4.081 | / | 5.755* | 5.487 | 6.325*** | 4.387*** |
| *Euro* | 3.363 | -2.790* | 0.116*** | / | 4.081 | 5.897* | 1.648 | 10.328*** | 5.888** |
| *Pound Sterling* | 6.503*** | 5.552** | 6.205** | 10.328*** | 6.325*** | 5.295*** | 3.735** | / | 9.983*** |

***Note**: Significance tests relate to the null hypothesis that the slope equal to one. *(**)[***] denotes significance at the 10%(5%)[1%] marginal significance level.*

## **Appendix Table 2:** ***Fama Regression Results for the Full, Pre-Crisis and Post-Crisis Samples***

## ***(3 month horizon)***

| | *Canadian Dollar* | *Swiss Franc* | *Danish Krone* | *Euro* | *Japanese Yen* | *Norwegian Krone* | *Swedish Krona* | *Pound Sterling* |
|---|---|---|---|---|---|---|---|---|
| ***PANEL (A): Full Sample 2000M1 – 2016M2*** | | | | | | | | |
| *Constant* | 0.009 | 0.044 | -0.001 | 0.001 | -0.009 | -0.006 | -0.004 | 0.001 |
| | (0.021) | (0.033) | (0.025) | (0.026) | (0.034) | (0.030) | (0.030) | (0.023) |
| *Beta* | 1.755 | -1.663* | -1.200 | -1.441 | 0.431 | 0.098 | -0.9300 | 0.925 |
| | (2.758) | (1.503) | (2.189) | (2.217) | (1.189) | (1.848) | (1.867) | (2.960) |
| *Adj.R^2* | 0.001 | 0.007 | 0.000 | 0.002 | -0.003 | -0.005 | -0.001 | -0.002 |
| *P-Value of F-Statistic* | 0.266 | 0.124 | 0.311 | 0.236 | 0.544 | 0.92 | 0.367 | 0.451 |
| *Number of Observations* | 203 | 203 | 203 | 203 | 203 | 203 | 203 | 203 |
| ***PANEL (B): Pre-Crisis 2000M1 – 2007M8*** | | | | | | | | |
| *Constant* | 0.041** | 0.132*** | 0.041 | 0.052** | 0.065 | 0.017 | 0.040 | 0.002 |
| | (0.020) | (0.045) | (0.026) | (0.024) | (0.055) | (0.030) | (0.027) | (0.023) |
| *Beta* | -2.106 | -5.123*** | -4.672*** | -4.923*** | -1.910* | -1.685** | -3.808*** | -2.191** |
| | (2.230) | (1.714) | (1.749) | (1.421) | (1.506) | (1.236) | (1.293) | (1.494) |
| *Adj.R^2* | 0.011 | 0.110 | 0.107 | 0.125 | 0.024 | 0.029 | 0.121 | 0.030 |
| *P-Value of F-Statistic* | 0.151 | 0.000 | 0.000 | 0.000 | 0.065 | 0.049 | 0.000 | 0.046 |
| *Number of Observations* | 101 | 101 | 101 | 101 | 101 | 101 | 101 | 101 |
| ***PANEL (C): Post-Crisis 2007M9 – 2016M2*** | | | | | | | | |
| *Constant* | 0.018 | -0.023 | -0.013 | -0.004 | -0.051 | 0.129 | 0.014 | 0.039 |
| | (0.053) | (0.045) | (0.032) | (0.035) | (0.046) | (0.097) | (0.047) | (0.033) |
| *Beta* | 10.282 | 7.810 | 5.077 | 9.301* | 7.456*** | 11.011 | 6.386 | 13.27** |
| | (6.796) | (4.184) | (3.750) | (4.599) | (1.767) | (6.367) | (4.178) | (5.709) |
| *Adj.R^2* | 0.052 | 0.053 | 0.037 | 0.078 | 0.133 | 0.091 | 0.047 | 0.198 |
| *P-Value of F-Statistic* | 0.012 | 0.011 | 0.030 | 0.003 | 0.000 | 0.001 | 0.016 | 0.000 |
| *Number of Observations* | 102 | 102 | 102 | 102 | 102 | 102 | 102 | 102 |

***Note****: Significance tests relate to the null hypothesis that the intercept is null and slope equal to one. *(**)[***] denotes significance at the 10%(5%)[1%] marginal significance level. The F-statistic refers to the joint null hypothesis that the intercept is null and slope equal to one.*

**Appendix Table 3:** ***Augmented Fama Regression Results Using the VIX as Proxy for the Risk Premium for the Full Sample (2000M1 – 2016M2) (3 month horizon)***

| | ***Canadian Dollar*** | ***Swiss Franc*** | ***Danish Krone*** | ***Euro*** | ***Japanese Yen*** | ***Norwegian Krone*** | ***Swedish Krona*** | ***Pound Sterling*** |
|---|---|---|---|---|---|---|---|---|
| *Constant* | 0.020 | 0.047 | 0.005 | 0.008 | -0.012 | 0.000 | 0.006 | 0.005 |
| | (0.019) | (0.032) | (0.022) | (0.023) | (0.032) | (0.030) | (0.025) | (0.021) |
| *Beta* | 2.155 | -1.691* | -1.424 | -1.728 | 0.304 | -0.111 | -1.074 | 0.741 |
| | (2.487) | (1.533) | (1.918) | (1.948) | (1.108) | (1.713) | (1.511) | (2.164) |
| *Gamma (VIX)* | -0.250*** | -0.063 | -0.149* | -0.152* | 0.154*** | -0.240** | -0.263*** | -0.145* |
| | (0.048) | (0.068) | (0.089) | (0.090) | (0.039) | (0.110) | (0.095) | (0.081) |
| *Adj.R^2* | 0.221 | 0.013 | 0.052 | 0.055 | 0.058 | 0.100 | 0.133 | 0.063 |
| *P-Value of F-Statistic* | 0.000 | 0.100 | 0.002 | 0.001 | 0.001 | 0.000 | 0.000 | 0.001 |
| *Number of Observations* | 203 | 203 | 203 | 203 | 203 | 203 | 203 | 203 |

***Note****: Significance tests relate to the null hypothesis that the intercept is null and slope equal to one. *(**)[***] denotes significance at the 10%(5%)[1%] marginal significance level. The F-statistic refers to the joint null hypothesis that the intercept is null and slope equal to one.*

**Appendix Table 4:** ***UIP Regressions Results Using Survey Data on Exchange Rate Expectations for the Full, Pre-Crisis and Post-Crisis Samples (3 month horizon)***

| | *Canadian Dollar* | *Swiss Franc* | *Danish Krone* | *Euro* | *Japanese Yen* | *Norwegian Krone* | *Swedish Krona* | *Pound Sterling* |
|---|---|---|---|---|---|---|---|---|
| ***PANEL (A): Full Sample 2003M1 – 2016M2*** | | | | | | | | |
| *Constant* | -0.017 | -0.103*** | -0.033*** | -0.035*** | -0.056*** | 0.038*** | 0.022 | -0.046*** |
| | (0.012) | (0.020) | (0.012) | (0.012) | (0.018) | (0.014) | (0.016) | (0.013) |
| *Beta* | -0.676 | 5.410*** | 1.459 | 2.151 | 4.037*** | 1.323 | 1.692 | -1.325*** |
| | (1.151) | (0.962) | (1.071) | (1.006) | (0.765) | (0.845) | (0.926) | (0.712) |
| *Adj.R^2* | -0.004 | 0.172 | 0.011 | 0.024 | 0.242 | 0.014 | 0.020 | 0.012 |
| *P-Value of F-Statistic* | 0.536 | 0.000 | 0.105 | 0.031 | 0.000 | 0.074 | 0.043 | 0.095 |
| *Number of Observations* | 155 | 155 | 155 | 155 | 155 | 155 | 155 | 155 |
| ***PANEL (B): Pre-Crisis 2003M1 – 2007M8*** | | | | | | | | |
| *Constant* | -0.021 | -0.020 | 0.006 | 0.004 | -0.010 | 0.04*** | 0.072*** | -0.019 |
| | (0.015) | (0.042) | (0.019) | (0.020) | (0.041) | (0.010) | (0.022) | (0.014) |
| *Beta* | -0.170 | 2.869 | 1.449 | 1.380 | 3.426** | 2.077*** | 0.374 | -0.405* |
| | (1.303) | (1.473) | (1.243) | (1.254) | (1.155) | (0.345) | (0.941) | (0.761) |
| *Adj.R^2* | -0.019 | 0.062 | 0.014 | 0.010 | 0.189 | 0.069 | -0.017 | -0.015 |
| *P-Value of F-Statistic* | 0.880 | 0.040 | 0.195 | 0.223 | 0.001 | 0.032 | 0.705 | 0.649 |
| *Number of Observations* | 53 | 53 | 53 | 53 | 53 | 53 | 53 | 53 |
| ***PANEL (C): Post-Crisis 2007M9 – 2016M2*** | | | | | | | | |
| *Constant* | -0.019 | -0.109*** | -0.059*** | -0.057*** | -0.057*** | -0.009 | -0.004 | -0.056*** |
| | (0.025) | (0.023) | (0.009) | (0.014) | (0.021) | (0.034) | (0.022) | (0.019) |
| *Beta* | -1.699 | 3.983* | -0.341 | 0.566 | 1.986 | -1.448 | 1.196 | -1.832 |
| | (3.325) | (1.758) | (0.863) | (1.555) | (0.732) | (2.118) | (1.743) | (1.875) |
| *Adj.R^2* | -0.005 | 0.025 | -0.009 | -0.009 | 0.024 | -0.004 | -0.004 | 0.005 |
| *P-Value of F-Statistic* | 0.475 | 0.063 | 0.806 | 0.760 | 0.064 | 0.430 | 0.444 | 0.222 |
| *Number of Observations* | 102 | 102 | 102 | 102 | 102 | 102 | 102 | 102 |

***Note**: Significance tests relate to the null hypothesis that the intercept is null and slope equal to one. *(**)[***] denotes significance at the 10%(5%)[1%] marginal significance level. The F-statistic refers to the joint null hypothesis that the intercept is null and slope equal to one.*

**Appendix Table 5:** ***Data Sources***

| ***Variable*** | ***Source*** | ***Timing*** |
|---|---|---|
| Spot Exchange Rates, against U.S. Dollar | IMF, International Financial Statistics | Monthly, End-of-Period, Start: 1999M1 |
| Forward Exchange Rates (3M and 12M), against U.S. Dollar | Thomson Reuters Datastream | Daily, End-of-Period, Start: 29/01/1999 |
| Expected Exchange Rates (3M and 12M), against U.S. Dollar | Consensus Forecast Economics Inc. | Monthly, sampled at the second Monday of the month, Start: 2003M1 |
| Eurocurrency Deposit Rates (3M and 12M) | Thomson Reuters Datastream | Daily, End-of-Period, Start: 29/01/1999 |
| Volatility S&P 500 Index (VIX) | CBOE | Daily, End-of-Period, Start: 29/01/1999 |

Note: If applicable, series are obtained for the following currencies: Canadian Dollar, Danish Krone, Euro, Japanese Yen, Norwegian Krone, Pound Sterling, Swedish Krona, Swiss Franc, United States Dollar