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RESEARCH ARTICLE

The New Fama Puzzle

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Abstract

We re-examine the historically common finding that *ex post* depreciation and the forward premium are negatively correlated, usually termed the forward premium puzzle. When covered interest differentials are zero, this finding is equivalent to the rejection of the joint hypothesis of uncovered interest parity (UIP) and full information rational expectations. We term this result the Fama puzzle (1984), given the difficulty in identifying a time-varying risk premium that could rationalize this result. In our analysis, the rejection occurs for eight exchange rates against the US dollar, but does not survive into the period during and in the decade after the financial crisis. Strikingly, in contrast to earlier findings, the Fama coefficient—the coefficient on the interest differential—then becomes large and positive; this is what we term the New Fama Puzzle. Using survey based measures of exchange rate expectations, we find much more consistant evidence in favor of UIP. Hence, the explanation for the switch in the Fama coefficient in the wake of the global financial crisis is mostly a change in how expectations errors and interest differentials co-move.

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1 Introduction

The commonplace finding that *ex post* changes in exchange rates do not offset interest differentials so as to equalize expected returns constitutes one of the durable puzzles in the international finance literature. Empirically, this condition manifests itself in a negative coefficient in a regression of exchange rate depreciation on interest differentials, which is often termed the forward premium puzzle,[1] documented in Fama (1984) and Tryon (1979), and what we term the old Fama puzzle.

This finding seemingly contradicts one of the most central concepts in international finance, namely uncovered interest rate parity. However, uncovered interest parity (UIP) relates *expected* exchange rate changes to interest rate differentials. It's only the joint hypothesis of UIP and full information rational expectations—sometimes termed the unbiasedness hypothesis—that leads to the implied value of unity for the regression coefficient in the Fama (1984) regression. The most commonplace explanation for the rejection of the unit coefficient—such as the existence of a time-varying exchange risk premium, which drives a wedge between forward rates and expected future spot rates—has found little empirical verification, despite numerous studies.[2]

We revisit this puzzle for several reasons, the most important of which is the finding that the Fama coefficient has switched sign during the period starting with the global financial crisis, and subsequently flipped sign again. It is this switching back and forth in a persistent fashion that prompts our investigation of this "new" Fama puzzle.

Even without this back-and-forth result, one would have wanted to re-examine the Fama result. First and foremost, interest rates in many advanced economies experienced a prolonged period in which short rates effectively hit the effective lower bound, with a corresponding compression of interest differentials, while *ex post* depreciations have not exhibited a comparable reduction. Moreover, some measures of risk and uncertainty have risen to record levels, raising the possibility that the effects of risk might be more easily detected than in previous periods. The first point is clearly illustrated in Fig. 1: where we plot one-year interest rates for a set of eight selected countries and the United States. The commensurate decline in interest differentials is shown in Fig. 2. Figure 3 depicts the corresponding one year exchange rate depreciations. These developments motivate us to re-examine whether the Fama result is a general phenomenon or one that is regime-dependent.

[1] If there are no covered interest differentials (as should be the case in the absence of capital controls and capital requirements), then the forward premium equals the interest differential. A regression of depreciation on the forward premium is equivalent to a regression of depreciation on interest differentials. We re-examine this point in the theoretical section.

[2] In fact, Fama did not interpret the negative coefficient as a puzzle, as he attributed the result to the presence of a time varying risk premium. Engel (1996) surveys the failure of the portfolio balance models and consumption capital asset pricing models to provide a risk premium basis for the Fama result. See also Chinn (2006) and more recently Engel (2014). Most recently, Corsetti and Marin (2020) argue no puzzle exists given the role of "Peso events'.

The second point is illustrated by the plot of the VIX and the Economic Policy Uncertainty Index, shown in Fig. 4. This development potentially allows us to distinguish between competing explanations for the failure of the unbiasedness hypothesis. Specifically, we can examine whether the inclusion of these risk proxies alters the Fama result.[3]

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 1999 to September 2021, but are really only replicated for the period 1999-2006. That is, the results change if the sample is broken into three periods—one before the global financial crisis, one during and after, encompassing the effective lower bound era, and another one largely corresponding to the period after the lift-off of US rates. For the middle period, interest differentials correctly signal the right direction of subsequent exchange rate changes, but with a magnitude that is not reconcilable with the conventional interpretation of UIP. In fact, we obtain positive coefficients at exactly a time of high risk when it would seem less likely that UIP would hold, presuming risk aversion explains deviations from UIP. Some months after US rates rise above zero, the old Fama finding re-appears, and persists into the second episode of zero lower bound rates.

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.

It is the use of expectations data that provides the following key insights. First, interest differentials and anticipated exchange rate changes are overall positively correlated throughout sample periods, consistent with the proposition that investors tend to equalize, at least partially, returns expressed in common currency terms. The relationship between expected depreciation and interest differentials also exhibits more stability than that involving *ex post* depreciation. Second, in cases where the Fama coefficient switches sign from negative to positive, and subsequently positive to negative, 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 that factor does play a role for certain currencies).

In the next section we briefly lay out the theory underlying the UIP and Fama regressions, and review the existing literature. In Sect. 3, we examine the empirical results obtained from estimating the Fama regression over different samples, and augmented with a risk proxy. In Sect. 4 we explore the results dropping the full information rational expectations assumption, and rely instead upon survey data on expectations. Section 5 presents a decomposition of the components driving the

[3] 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).

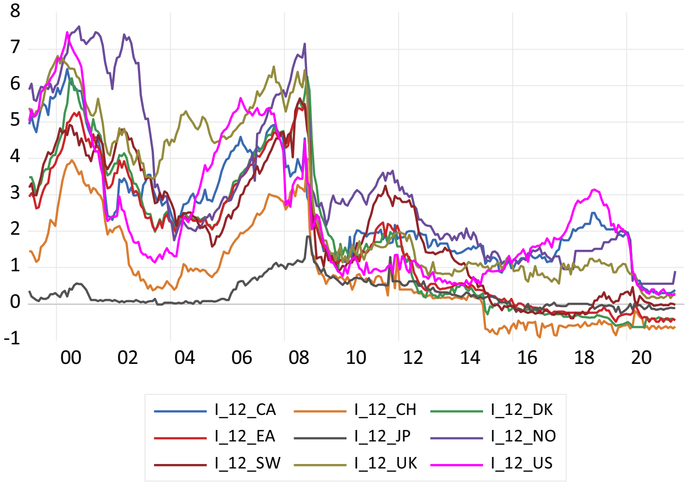


Fig. 1 Interest Rates on 1Year Eurocurrency Deposits, end of period, % *Source* Thomson Reuters Datastream

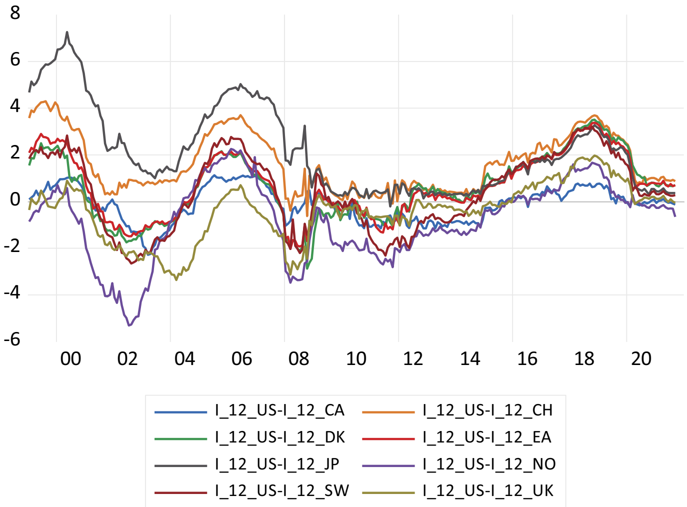


Fig. 2 1Year Eurocurrency Deposit Rates Differential (US Dollar minus Foreign Currency), end of period, % *Source* Thomson Reuters Datastream, and authors' calculations

deviation of the Fama coefficient from the posited value of unity, and an economic interpretation for the changes we observe. Section 6 concludes.

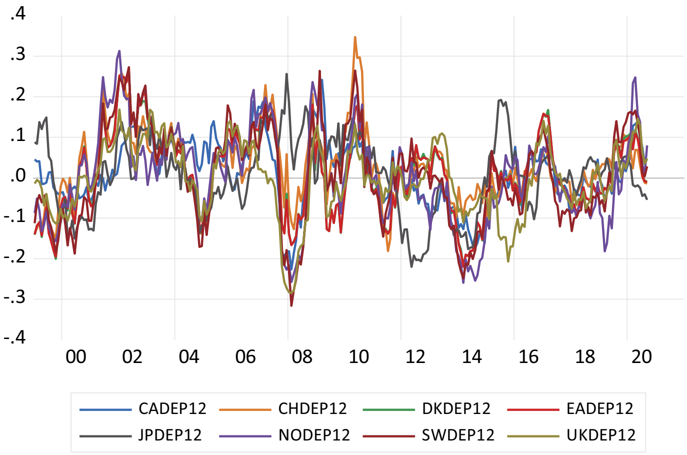


Fig. 3 1 Year *Ex-Post* Depreciation Rate of the US Dollar with respect to Foreign Currency (Positive values indicate dollar depreciations), decimal format, end of period, % *Source* International Financial Statistics and authors' calculations

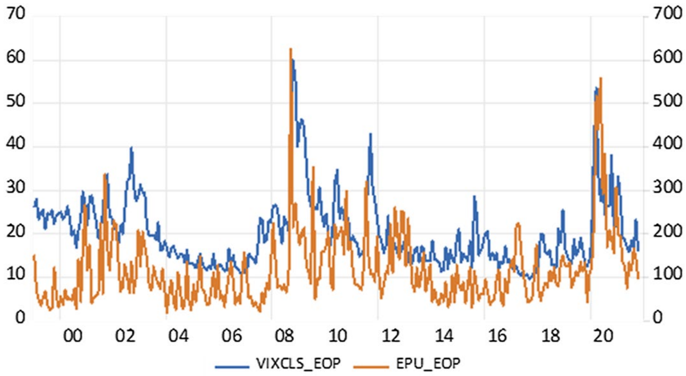


Fig. 4 VIX (left scale) and US Economic Policy Uncertainty index (right scale), both end-of-period *Source* policyuncertainty.com and CBOE

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:

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

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 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, for example, Japan's interest rate, i.e. $i_t > i^*_t$, then we should expect the dollar to depreciate with respect to the Japanese yen 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), where the forward premium is treated as being equivalent to the interest differential[4]:

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

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

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

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 α may be non-zero while testing for UIP using Eq. (2). A non-zero α may reflect a constant risk premium (hence, tests for $\beta = 1$ are tests for a time-varying risk premium, rather than risk neutrality per se) 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 Eq. (1) to regression Eq. (2). There are three key assumptions, as laid out in the following equations:

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

$$f_{h,t} = E_t^M\left[s_{t+h}\right] + \epsilon^{rp}_{h,t}, \tag{5}$$

[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 ($h \in [3, 12]$), 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 Tables 2-4.

$$s_{t+h} = E_t^M \left[s_{t+h} \right] - \epsilon^f_{t+h}. \tag{6}$$

When $\epsilon^{cip}_{h,t}$ is zero, then Eq. (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 (which we use),[6] covered interest parity has held up, up until the global financial crisis.Eq. (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.[7]

The combination of $\epsilon^{cip}_{h,t} = \epsilon^{rp}_{h,t} = 0$ in Eqs. (4) and (5) yields uncovered interest rate parity. Only when combined with the assumption of full information rational expectations, namely $E_t\left(\epsilon^f_{t+h}\right) = 0$ in Eq. (6)[8], does one obtain the regression Eq. (2), where the regression residual can be interpreted as the forecast error. In general, the $\beta = 1$ hypothesis relies upon three moment conditions:

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

When the covered interest differential is zero, the first covariance term is zero. This has been the approach adopted historically; however, recent work has documented the fact that covered interest differentials have increased in recent years even when using offshore rates (Borio et al. 2016; Du et al. 2018), and so we do not impose this assumption in our analysis. 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 full information rational expectations hypothesis; in our lexicon, that means they are tests of the unbiasedness hypothesis. Estimates of Eq. (6) using horizons for up to one year typically reject the unbiasedness restriction on the slope parameter. For instance, the survey by Froot and Thaler (1990), finds an average estimate for β of -0.88.[9] Bansal

[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] Note that we use offshore yields rather than sovereign bond yields, thereby mitigating the convenience yield channel emphasized by Engel (2016).

[7] 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.

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

[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.

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] [11] Frankel and Poonawala (2010) document that for emerging markets more generally, the unbiasedness hypothesis coefficient is typically more positive.[12]

The poor performance of the interest differential as a predictor 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, 2019). 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 relaxes the rational expectations approach involving the use of survey-based data to measure exchange rate expectations. In this case, the error term in Eq. (6), ϵ^{f}_{t+h}, 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. Similar results were obtained by Chinn and Frankel (2020), when extending the data up to 2018, increasing the sample to about 32 years. This pattern of findings suggests that the assumption of full information rational expectations is not innocuous, and that the examination of the UIP condition both 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. 1999—September 2021 period. (All data sources are detailed in Table 8). We examine twelve month exchange rate depreciation and the

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

[11] See Hassan and Mano (2017) for a different perspective on how the Fama puzzle relates to the carry trade.

[12] 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.

corresponding 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.[13]

Figure 2 depicts twelve month maturity yield differentials, while Fig. 3 shows twelve month depreciations, all over the 1999–2021 period. One of the contrasts clearly highlighted by the two figures is that while yield differentials have shrunk toward zero in the wake of the global financial crisis—at least until about 2015 --, exchange rate depreciations have not exhibited a comparable compression.

Table 1 reports in Panel A the results from Eq. (2) at the twelve month horizon, for the full sample.[14] The results are largely in accord with previous findings. The slope coefficients on the interest differential (i.e., the "Fama regression slope coefficient") are negative, with the exception of Canada. Under the maintained hypothesis the coefficient should be unity, which we test. In four cases, including the euro, one can reject the unit coefficient null at the 1% level. The Canadian dollar, the Japanese yen, the Norwegian krone and British pound fail to reject.[15] 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. For instance, as policy regimes change, the expectation formation process will change as well. Changes in the general economic environment will also have an impact, possibly through regulations or global risk.

In order to identify break points in the Fama regression, we used the Bai (1997) and Bai-Perron (1998) Sequential L+1 breaks vs. L test for structural breaks. While different break points are identified for different exchange rates, the euro exchange rate (against the dollar) is illustrative. Restricting the number of breaks to two, and using a 5% significance level, we identify three periods: 1999M01-2005M04, 2005M05-2017M04, and 2017M05-2020M09. For other exchange rates, we also identify two breakpoints, except for Switzerland and Norway (for which we identify only one). However, even when two breakpoints are identified, the second breakpoint is not usually the same.[16] Nonetheless, we decide to use as a common breakpoint those that apply to the euro.

There are several candidate events to associate with the first breakpoint. That time is associated both with the ECB raising rates, and with US *expected* interest

[13] We adopt the standard assumption of no default risk. In general, this is believed to hold, although 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).

[14] Since we are examining one year horizons, the interest rate sample is truncated at 2020M09. Results for three month horizon, reported in Tables 5, 6, and 7, are truncated at 2021M06.

[15] Engel et al. (2021) finds weaker rejection of unbiasedness using a longer sample for the early period, and an alternative estimator for standard errors. They find in a 2007–2020 period, positive coefficients but a general failure to reject unity for the slope coefficient.

[16] We have also conducted the analysis with a first breakpoint at 2006M08, and a second at 2018M01. That breakpoint incorporates exchange rate changes up to 2007M08, which could be considered as the beginning of the Global Financial Crisis, with the turmoil on the US housing market. Using this setup, we again obtain the same pattern of coefficient sign reversals.

rates exceeding actual rates. The second breakpoint is not clearly identified with any given event, although it is about a year and a half after the increase in US policy rates, and underprediction of short term interest rates.[17] Consequently, we separate the sample into early, middle and late periods, with *ex post* exchange rate depreciations ending April 2006, April 2018, and September 2021, respectively. The respective subperiod results are presented in Panels B, C and D of Table 1.

In the pre-crisis (early) period, the coefficients are uniformly negative, and significantly different from unity in all cases. Exchange rate depreciation is strongly—and positively—related to the interest differential. The estimated coefficients range from − 2.1 to − 5.2. The null hypothesis of unity is rejected for all cases. The joint null hypothesis that the constant is zero and the slope coefficient is unity is also uniformly rejected.

Turning the middle period, we obtain drastically different results. The slope coefficient is positive in all instances. In five of eight cases, one can reject the null of a unit coefficient, so even with the positive coefficient, the results are not consistent with the unbiasedness hypothesis. 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, while we obtain this result over a period extending up to 2017.

In the late period, all the slope coefficients save Japan's were negative—ranging from −0.9 to −10.3. The null of a unit coefficient was rejected in all cases. One might think that coefficients should switch back after the zero lower bound is re-attained. It is difficult to determine whether this in fact occurs, given the few observations available especially when using the 12-month horizon. Using 3-month horizons, however, the slope coefficients remain negative, albeit not always significantly so, after February 2020 (see tables in the Appendix).

To highlight the change in how the relationship between interest differentials and *ex post* depreciations change over time, we focus on the Euro in Fig. 5. The stabilization of the interest differential, compared to depreciations, is now obvious. One way to illustrate the contrast pre- and post-crisis, not necessarily evident in Fig. 5, is to show a scatterplot of depreciation against the yield differential. Figure 6 depicts the data for the three periods. In the pre-crisis period, the slope is negative (as in the conventional empirical wisdom), while in the post-crisis period, it is clearly positive. In the late period, the slope is again negative.[18] Another way to illustrate this finding is to show the evolution of the beta coefficients from rolling Fama regressions. Figure 7 shows beta coefficients obtained from regressing the 12 month dollar depreciation against the euro on the US-euro area interest differential, for three year rolling windows. Results confirm the switch of signs of coefficients from negative to positive around the beginning of 2006, and with less certainty a switch to negative again somewhere between mid-2014 and mid-2016.

[17] As indicated by the Survey of Professional Forecasters forecasts of the three month Treasury yield; this is discussed further in Sect. 5.

[18] Figure 11 shows the corresponding graphs for all the currencies.

Table 1 Fama regression results

Coefficient	CAD	CHE	DKR	EUR	JPY	NKR	SKR	GBP
A: Full								
Constant	0.012	0.051	0.017	0.017	0.008	−0.002	0.011	−0.004
	0.010	0.021	0.015	0.016	0.022	0.014	0.017	0.010
Beta	1.310	−1.420***	−1.045***	−1.019***	−0.058	−0.583	−1.084***	−0.108
	1.588	0.872	0.909	0.988	0.755	0.944	0.942	1.109
Adj.R sq.	0.010	0.036	0.018	0.015	−0.004	0.003	0.018	−0.004
F-statistic	3.606	10.684	5.699	4.947	0.033	1.737	5.815	0.054
N	261	261	261	261	261	261	261	261
B: Early								
Constant	0.037	0.137	0.056	0.068	0.086	0.017	0.048	0.006
	0.010	0.022	0.014	0.014	0.023	0.018	0.017	0.021
Beta	−3.793***	−4.888***	−5.180***	−5.213***	−2.419***	−2.158***	−4.141***	−2.136***
	1.227	0.860	1.118	0.956	0.637	0.834	1.022	1.126
Adj.R sq.	0.290	0.438	0.430	0.467	0.274	0.196	0.374	0.104
F-statistic	38.146	72.024	69.601	80.804	35.418	23.252	55.472	11.534
N	92	92	92	92	92	92	92	92
C: Middle								
Constant	0.017	0.006	−0.013	−0.016	−0.037	0.023	−0.015	−0.011
	0.014	0.026	0.017	0.016	0.027	0.022	0.020	0.015
Beta	9.167***	1.520	2.560	3.778***	3.885***	4.127***	2.382	5.331***
	1.751	1.528	1.472	1.337	1.066	1.432	1.290	1.799
Adj.R sq.	0.347	0.017	0.080	0.148	0.195	0.142	0.065	0.219
F-statistic	73.168	3.385	12.790	24.595	33.876	23.451	10.489	39.223
N	137	137	137	137	137	137	137	137
D: Late								
Constant	0.048	0.053	0.113	0.100	−0.036	0.058	0.152	0.100
	0.019	0.030	0.056	0.042	0.012	0.033	0.055	0.016
Beta	−10.10***	−0.865*	−3.986***	−3.797***	2.083**	−10.34***	−6.504***	−7.230***
	2.854	1.021	1.870	1.466	0.533	2.359	1.989	1.308
Adj.R sq.	0.427	0.006	0.252	0.305	0.330	0.458	0.498	0.606
F-statistic	24.061	1.172	11.433	14.628	16.272	27.188	31.792	48.722
N	32	32	32	32	32	32	32	32

Sample period refers to interest rate observations. *(**)[***] denotes significance at the 10%(5%)[1%] marginal significance level for null of unit coefficient. The F-statistic refers to the joint null hypothesis that the intercept is null and slope equal to one

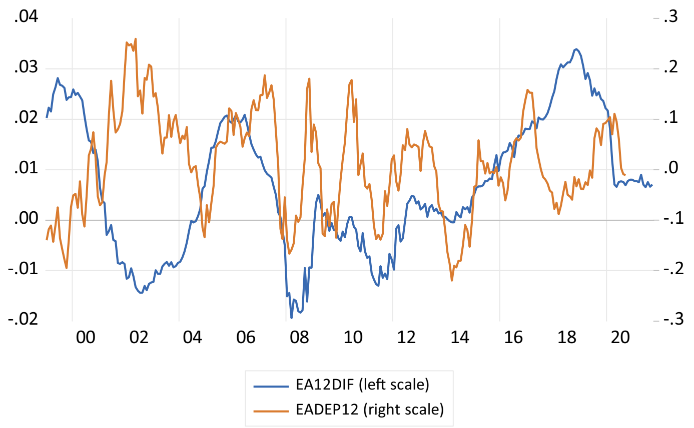


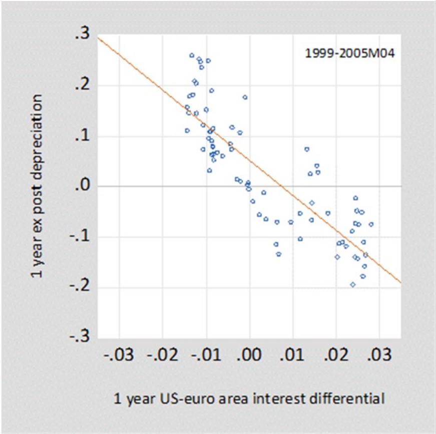
Fig. 5 1 Year Eurocurrency Deposit Rates Differential and 1Y-Ex-Post Depreciation Rate of the US Dollar with respect to euro, end of period (decimal format). *Source* Thomson Reuters Datastream, International Financial Statistics, and authors' calculations

To highlight how the estimated beta coefficients evolve over time for all the currencies, we show in Fig. 8 the coefficients for the corresponding subperiods. In the top panel (early period), the beta coefficients are tightly centered around negative values. In the middle panel (middle period), the coefficients are positive and more widely spread. In the bottom panel (late period), the estimates are mostly negative and very widely dispersed. The switch in slope coefficient signs from the early to middle, and middle to late, holds across currencies with strong regularity, with the sole exception of the Japanese yen. In that particular case, the coefficient switches but once, from the early period to the middle period, and stays constant thereafter (Fig. 9).

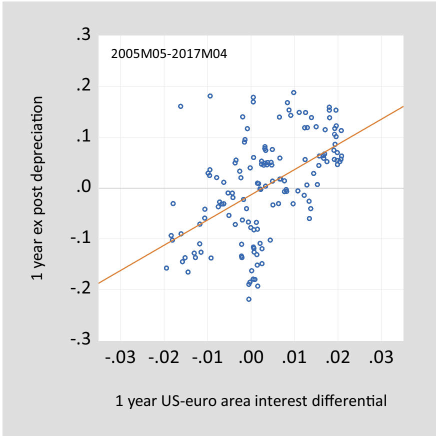
Interestingly, the adjusted R^2 rise substantially from essentially zero in the full sample to values of around 0.2 to 0.8 in the various subsamples. From a statistical perspective, this result is consistent with the conclusion that estimating over the full sample imposes restrictions that are rejected by the data.[19]

One plausible criticism of our finding of sign switches is primarily driven by using the dollar as a base currency. Remarkably, the switch from negative to positive coefficients holds when examining exchange rates using other base currencies (see Table 4). The switch from positive to negative coefficients in 2017 holds for fewer cross rates. Nonetheless, this pattern of results indicates that there is at least one break in the Fama relationship for not just those exchange rates expressed against the US dollar. These results confront the researcher with at least two related questions. The first is the longstanding puzzle of why the bias

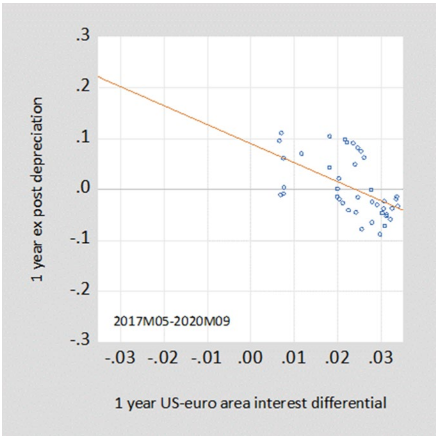
[19] The absolute size of the coefficients is larger after the first period; mechanically, this arises because the regression coefficient is a covariance divided by the variance of the interest differential, and the variance of interest differentials are much smaller post-Crisis, as illustrated in Fig. 2.



Panel (a): Early (pre-crisis)



Panel (b): Middle (Crisis/Post-Crisis)



Panel (c): Late

Fig. 6 Scatterplot of the 1 Year *Ex Post* Depreciation Rate (1 Year Ahead) on 1 Year Eurodeposit Rate Differential of US Dollar with respect to the Euro (decimal format). Regression line in red. Author's calculations based on International Financial Statistics and Thomson Reuters Datastream data

exists; the second is why the correlation changed so much after the crisis, and then again seemingly reverted.

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

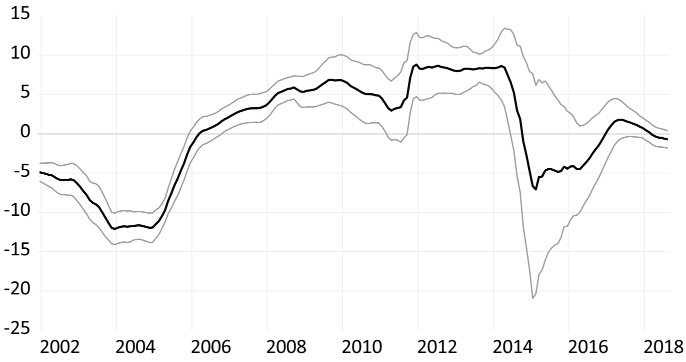


Fig. 7 Estimates of Beta from a 1 Year horizon Fama Regression Euro with respect to the US Dollar 3 Year Rolling Windows (timing refers to interest differentials)**.** Author's calculations based on International Financial Statistics and Thomson Reuters Datastream data

is to include as an additional regressor some variable that proxies for an exchange risk premium, ϵ_t^{rp}. This suggests the following regression equation:[20]

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

where Z is a proxy variable.

We select the VIX as a proxy measure [21]. 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 2018) and for excess returns (Brunnermeier et al. 2008, Habib and Stracca 2012, or Husted *et al.*, 2018).[22]

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 positive in the full sample, though generally non-significant.

[20] 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. Variants 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).

[21] 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 (2019) 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.

[22] See Berg and Mark (2018) for discussion of uncertainty and the risk premium.

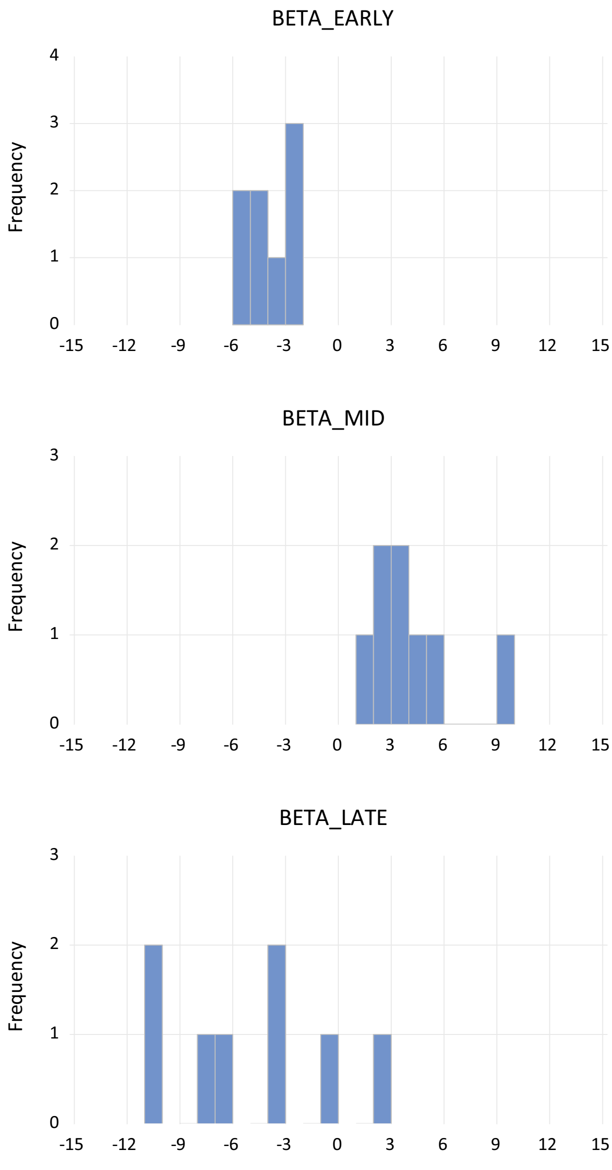


Fig. 8 Estimates of Beta from a 1 Year horizon Fama Regression for Early, Middle, and Late Periods. Author's calculations based on International Financial Statistics and Thomson Reuters Datastream data

Fig. 9 Decomposition of the Deviation from β-1 from 1 Year horizon Fama Regressions with respect to the US Dollar. *Note*: Early subperiod starts at 2003M01. Authors' calculations based on International Financial Statistics and Thomson Reuters Datastream data

We also examined the impact of VIX inclusion in the three subsamples, but overall we do not obtain any significant results. This result suggests that when the slope coefficients switch sign, it's not because of the omission of the VIX.[23]

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 $E_t\left(\epsilon^f_{t+1}\right) = 0$ in Eq. (6). It might be that agents are truly irrational, they use bounded rationality, or 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 Eq. (6) with:

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

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.[24] 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):

$$\hat{s}^M_{t+h} - s_t = \alpha + \beta^{'\left(i_{h,t} - i^*_{h,t}\right)} + u_t \tag{10}$$

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.

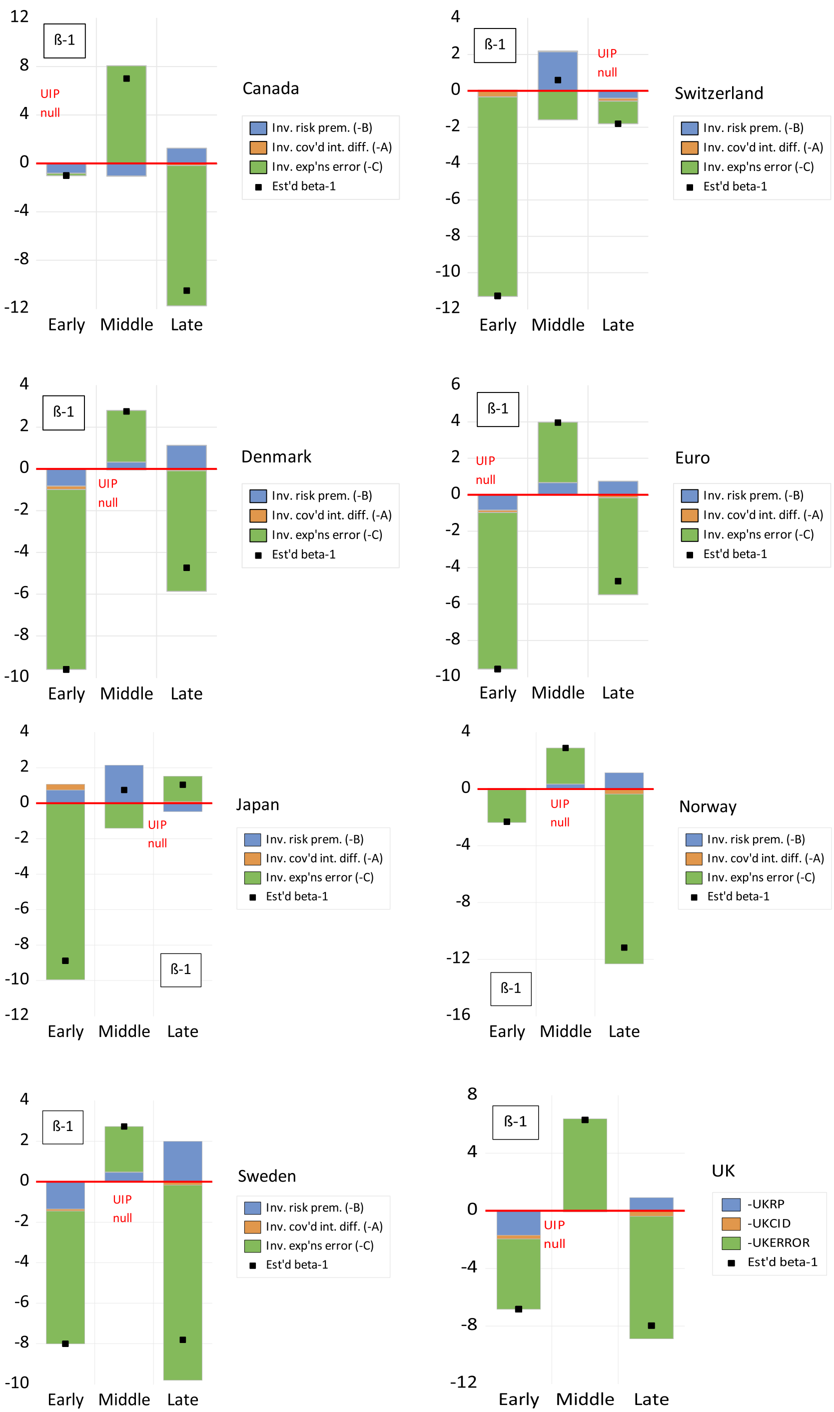
We use as measures of expectations survey data sourced from Consensus Forecasts from 2003M01 to 2021M09. Notice that survey data availability necessitates a change in the sample period.[25]

The results of the regressions are reported in Table 3, using the same format as in Table 1. One of the defining features of the results is (1) the point estimates are

[23] Kalemli-Ozcan and Varela (2021) investigate how the deviation from survey-implied UIP moves with the VIX, as opposed to how *ex post* depreciation moves.

[24] 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.

[25] 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. We have cross checked the results for the euro using Currency Forecasters Digest/FX Forecasts data (as used in Chinn and Frankel 2020). The results are the same when the expected, futures and spot rates are exactly aligned.


Canada
ß-1
UIP null
12 8 4 0 -4 -8 -12
Early Middle Late
Inv. risk prem. (-B)
Inv. cov'd int. diff. (-A)
Inv. exp'ns error (-C)
Est'd beta-1
Switzerland
ß-1
UIP null
4 2 0 -2 -4 -6 -8 -10 -12
Early Middle Late
Inv. risk prem. (-B)
Inv. cov'd int. diff. (-A)
Inv. exp'ns error (-C)
Est'd beta-1
Denmark
ß-1
UIP null
4 2 0 -2 -4 -6 -8 -10
Early Middle Late
Inv. risk prem. (-B)
Inv. cov'd int. diff. (-A)
Inv. exp'ns error (-C)
Est'd beta-1
Euro
ß-1
UIP null
6 4 2 0 -2 -4 -6 -8 -10
Early Middle Late
Inv. risk prem. (-B)
Inv. cov'd int. diff. (-A)
Inv. exp'ns error (-C)
Est'd beta-1
Japan
ß-1
UIP null
4 2 0 -2 -4 -6 -8 -10 -12
Early Middle Late
Inv. risk prem. (-B)
Inv. cov'd int. diff. (-A)
Inv. exp'ns error (-C)
Est'd beta-1
Norway
ß-1
UIP null
4 0 -4 -8 -12 -16
Early Middle Late
Inv. risk prem. (-B)
Inv. cov'd int. diff. (-A)
Inv. exp'ns error (-C)
Est'd beta-1
Sweden
ß-1
UIP null
4 2 0 -2 -4 -6 -8 -10
Early Middle Late
Inv. risk prem. (-B)
Inv. cov'd int. diff. (-A)
Inv. exp'ns error (-C)
Est'd beta-1
UK
ß-1
UIP null
8 4 0 -4 -8 -12
Early Middle Late
-UKRP
-UKCID
-UKERROR
Est'd beta-1

almost uniformly positive (except for the Canadian dollar, in the early period), and (2) coefficients for the Swiss franc in full and middle samples, and Japanese yen are in all samples, are significantly greater than one, confirming that those currencies are considered as safe havens by practitioners. Mechanically, the difference in estimated slope coefficients arises from the fact that *ex ante* and *ex post* measures of depreciation differ substantially, so that the ex ante measures are usually biased predictors. The rejection of the null hypothesis of unit coefficient, despite positive estimates, can in part be attributed to the lower variability of ex ante depreciation, leading to smaller estimated standard errors. These results are consistent with those obtained in previous studies using survey data, including Chinn and Frankel (1993) and Chinn and Frankel (2020)[26].

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. For instance, for the euro dollar exchange rate, the one year *ex ante* changes range from -0.10 to +0.07; *ex post* changes range from -0.22 to +0.26. The corresponding standard deviations are 0.037 and 0.100, respectively. Roughly speaking, *ex post* changes are about three times as large as *ex ante*, for the euro.

Table 3 displays the estimated β' coefficients in the full sample as well as in the three sub-periods. Turning to the full sample results in Panel A, in contrast to the results using *ex post* depreciation, the coefficient on the interest differential is almost always positive. That does not mean that uncovered interest parity holds, as less than half of the cases reject the null of a unit slope coefficient (interestingly, not the euro). And in fact, for all cases save the Canadian dollar the joint null hypothesis of a zero constant and unit slope is resoundingly rejected. Interestingly, the sub-period point estimates (Panels B-D) do not suggest a switch in coefficient signs through the three periods.

Our findings of positive coefficients might be interpreted as an artifact of subsample selection. Applying Bai-Perron tests to the data indicate one or multiple breaks in all cases, even when using a high significance level. However, the estimated slope coefficients for the separate subperiods are all positive.

In sum, our empirical results indicate largely negative correlations between *ex post* depreciation in the early and late periods, and largely positive correlations during the middle period. Inclusion of a conventional risk proxy, the VIX, does not alter these basic results. On the other hand, expected depreciation and the interest differential is almost always positively correlated.

[26] 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 inputs forecasters use.

Table 2 Fama regression augmented with VIX results

Coefficient	CAD	CHE	DKR	EUR	JPY	NKR	SKR	GBP
A: Full								
Constant	−0.057	0.019	0.003	0.000	−0.043	−0.078	−0.058	−0.025
	0.023	0.031	0.038	0.036	0.033	0.033	0.034	0.022
Beta	1.873	−1.173***	−0.885***	−0.844***	0.027	0.341	−0.539	0.060
	1.514	0.921	1.049	1.071	0.780	1.117	0.996	1.114
Gamma	0.002	0.001	0.001	0.001	0.002	0.004	0.003	0.001
	0.001	0.001	0.001	0.001	0.001	0.001	0.001	0.001
Adj.R sq.	0.122	0.047	0.016	0.015	0.038	0.067	0.061	0.001
F-statistic	19.047	7.410	3.107	2.986	6.093	10.348	9.426	1.195
N	261.000	261.000	261.000	261.000	261.000	261.000	261.000	261.000
B: Early								
Constant	0.080	0.162	0.132	0.121	0.021	0.247	0.124	0.024
	0.035	0.061	0.062	0.061	0.045	0.047	0.073	0.055
Beta	−4.310***	−5.018***	−5.775***	−5.522***	−2.381***	−4.780***	−4.637***	−2.164***
	1.217	0.892	1.003	0.916	0.619	0.828	0.905	1.116
Gamma	−0.002	−0.001	−0.004**	−0.002	0.003	−0.012***	−0.004	−0.001
	0.002	0.002	0.002	0.002	0.002	0.002	0.003	0.002
Adj.R sq.	0.319	0.437	0.462	0.481	0.323	0.489	0.398	0.099
F-statistic	22.292	36.338	40.139	43.155	22.736	44.454	31.130	5.999
N	92	92	92	92	92	92	92	92
C: Middle								
Constant	−0.057	0.019	0.003	0.000	−0.043	−0.078	−0.058	−0.025
	0.023	0.031	0.038	0.036	0.033	0.033	0.034	0.022
Beta	1.873	−1.173***	−0.885***	−0.844***	0.027	0.341	−0.539	0.060
	1.514	0.921	1.049	1.071	0.780	1.117	0.996	1.114
Gamma	0.002	0.001	0.001	0.001	0.002	0.004	0.003	0.001
	0.001	0.001	0.001	0.001	0.001	0.001	0.001	0.001
Adj.R sq.	0.122	0.047	0.016	0.015	0.038	0.067	0.061	0.001

Table 2 (continued)

Coefficient	CAD	CHE	DKR	EUR	JPY	NKR	SKR	GBP
F-statistic	19.047	7.410	3.107	2.986	6.093	10.348	9.426	1.195
N	261	261	261	261	261	261	261	261
D: Late								
Constant	−0.050	0.015	0.038	0.047	−0.026	−0.066	0.046	0.014
	0.017	0.042	0.048	0.044	0.019	0.029	0.085	0.034
Beta	−5.022***	−0.386	−2.753***	−2.985***	1.872	−7.239***	−5.070***	−4.896***
	1.920	1.007	1.474	1.311	0.652	1.083	2.161	1.497
Gamma	0.0039	0.0008	0.0019	0.0014	−0.0003	0.0055	0.0031	0.0027
	0.0007	0.0008	0.0007	0.0007	0.0005	0.0011	0.0010	0.0010
Adj.R sq.	0.664	−0.006	0.251	0.289	0.305	0.688	0.479	0.603
f	21.602	2.290	8.328	10.504	1.011	51.273	21.284	27.151
N	41	41	41	41	41	41	41	41

Sample period refers to interest rate observations. *(**)[***] denotes significance at the 10%(5%)[1%] marginal significance level for null of unit coefficient on interest differential, or null of zero coefficient for VIX coefficient. The F-statistic refers to the joint null hypothesis that the intercept is null and slope equal to one

5 Reconciling the Results

The contrasting results obtained using *ex ante* and *ex post* depreciation suggests that understanding the characteristics of exchange rate expectations are critical to solving the puzzle.

To see this point explicitly, consider again the decomposition outlined in Eq. (7), that is:

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

where the relevant interest differential regression coefficients with the covered interest differential, exchange risk, and expectation errors are labelled A, B, and C, respectively. From this decomposition, 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 coefficient switches suggests that the C term, involving forecast errors, is of crucial importance.

Notice that the switch in the risk premium component—the B term—is not particularly central to the switch in the Fama regression slope coefficient for any of the currencies. The foregoing discussion suggests that the reason the puzzle has evolved in going from early to middle period is mainly because of a change in how expectations errors co-move with interest differentials, i.e., the C component.

The sign of the coefficient on the interest differential changes again—from positive to negative—moving from the middle to late period for all the currencies, save the Japanese yen. There the correlation switches, but in a way that is opposite that for the other currencies. The Swiss franc slope coefficient sign switches too, but in this case it's a change in the exchange risk premium correlation which drives the switch. The C component is unchanged in this case. In the other six cases, the switch in how expectations errors move with the interest differential drives the switch in the Fama regression coefficient sign.

What lies behind the change in the C component? For these currencies—save the Japanese yen—the forecast errors as defined in Eq. (6) change from significantly negative in the pre-crisis period to half positive in the middle crisis period. Finally, in the late period, the dollar appreciates more than expected, except with respect to the Swiss France. In fact, the Swiss franc is the only case for which the dollar constantly depreciates against more than expected. The forecast errors—over- or under- prediction—do not correspond to the switches in slope coefficient in the Fama regression.

In words, the overprediction of dollar depreciation is systematically greater, the greater the US-foreign interest differential. One of the characteristics of the 2005-17 period is that for most of the period, US interest rates were consistently expected to rise faster than they actually did. This point is illustrated in Fig. 10, which shows the US three month Treasury yield and the corresponding forecasts for up to one year as

Table 3 Uncovered interest parity regressions

Coefficient	CAD	CHE	DKR	EUR	JPY	NKR	SKR	GBP
A: Full								
Constant	0.000	−0.054	−0.016	−0.017	−0.057	0.033	0.020	0.000
	0.003	0.007	0.005	0.005	0.007	0.004	0.005	0.004
Beta	0.283	2.360***	1.188	1.377	2.987***	1.653***	1.374	0.880
	0.328	0.373	0.290	0.293	0.240	0.258	0.294	0.338
Adj.R sq.	0.000	0.349	0.185	0.217	0.597	0.278	0.206	0.088
F-statistic	0.095	114.882	49.056	59.923	314.866	82.706	55.860	21.520
N	213	213	213	213	213	213	213	213
B: Early								
Constant	−0.002	−0.008	0.013	0.012	−0.019	0.026	0.047	0.005
	0.005	0.014	0.006	0.006	0.014	0.006	0.008	0.007
Beta	−0.394**	1.845	1.141	1.105	2.374***	1.168	0.724	0.465
	0.309	0.510	0.383	0.389	0.347	0.251	0.340	0.316
Adj.R sq.	0.012	0.307	0.199	0.183	0.618	0.268	0.087	0.044
F-statistic	10.720	3.691	2.832	2.613	26.769	10.446	19.421	7.264
N	44	44	44	44	44	44	44	44
C: Middle								
Constant	−0.004	−0.062	−0.027	−0.027	−0.062	0.026	0.008	−0.011
	0.005	0.007	0.004	0.004	0.008	0.005	0.005	0.005
Beta	0.119	2.164***	0.386*	0.673	2.797***	1.352	0.937	0.401
	0.543	0.485	0.360	0.378	0.357	0.364	0.318	0.580
Adj.R sq.	−0.007	0.260	0.012	0.042	0.505	0.150	0.082	0.003
F-statistic	1.350	67.941	49.716	40.936	36.395	16.478	1.266	2.304
N	137	137	137	137	137	137	137	137
D: Late								
Constant	0.011	−0.008	−0.011	−0.002	0.006	0.060	0.002	0.020
	0.005	0.013	0.013	0.013	0.007	0.012	0.013	0.012
Beta	2.450*	0.416	1.536	1.293	0.437*	1.906	2.961***	1.346
	0.735	0.441	0.445	0.457	0.333	0.984	0.499	0.817
Adj.R sq.	0.191	0.019	0.344	0.311	0.013	0.090	0.579	0.076
F-statistic	33.856	28.892	7.474	2.822	1.641	67.109	109.359	8.058
N	32	32	32	32	32	32	32	32

Sample period refers to interest rate observations. *(**)[***] denotes significance at the 10%(5%)[1%] marginal significance level for null of unit coefficient. The F-statistic refers to the joint null hypothesis that the intercept is null and slope equal to one

of the third quarter of each year. To the extent that higher rates are associated with a stronger currency, the fact that rates did not rise in line with expectations meant that the dollar ended up being weaker than anticipated—hence the greater than anticipated dollar depreciation.

This means the reversals in the Fama coefficients is due in part to the larger mistakes in forecasting dollar changes in the post-crisis period, and very little is

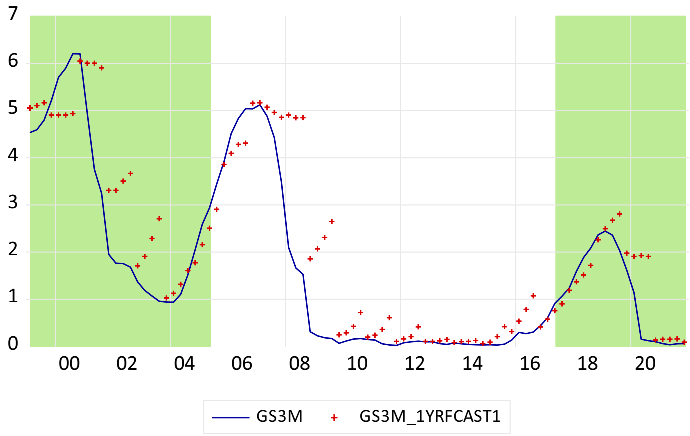


Fig. 10 Three month Treasury yield and Survey of Professional Forecasters mean forecast as of Q3, in %. Green shading denotes early, late periods. *Source*: US Treasury, and Federal Reserve Bank of Philadelphia

attributable to changes in exchange risk co-movements. And still less is associated with covered interest differentials co-movements.[27]

6 Conclusions

Our extensive cross-currency analysis of uncovered interest parity has yielded new empirical results that 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 breaks associated with the global financial crisis and the subsequent period of low interest rates, and the subsequent reversal, are quantitatively and qualitatively much more pronounced. The positive, albeit very large, Fama regression coefficient detected in much of the last decade is not usually 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.

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. 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.

[27] At the three-month horizon, the A component is slightly more important, but remains less significant than the B and C components.

Fourth, a formal decomposition of deviations from the posited value of unity in the Fama regression indicates that the switch in signs from the early to middle period can largely be attributed to the switch in the nature of the co-movement between expectations errors and interest differentials. We find that the switch does not tightly correspond to the period of extended zero lower bound in the US. Rather the break coincides with persistent overprediction of the US short term interest rate and hence overprediction of dollar strength.

From these results, we conclude that the change in the Fama coefficients is one that is primarily driven by systematic expectational errors ruled out in the full information rational expectations framework. Risk—either time varying or time invariant—might be important, but it is not primarily important in driving *ex post* exchange rate changes.

Appendix

See Tables 4, 5, 6, 7, 8 and Fig. 11.

Table 4 Estimated fama coefficients for the various sub-samples for selected base currencies (12 month horizon)

	USD	CAD	CHE	DKR	EUR	JPY	NKR	SKR	GBP
A: Full									
USD		1.310	−1.420	−1.045	−1.019	−0.058	−0.583	−1.084	−0.108
JPY	−0.058	−0.065	−0.911	−0.047	−0.160		−0.146	−0.023	0.725
EUR	−1.019	−0.421	−2.294	0.138		−0.160	1.744	−0.572	0.828
GBP	−0.108	3.122	−0.117	0.612	0.828	0.725	0.960	0.123	
B: Early									
USD		−3.793	−4.888	−5.180	−5.213	−2.419	−2.158	−4.141	−2.136
JPY	−2.419	1.404	−4.260	−1.800	−3.086		0.004	0.261	1.921
EUR	−5.213	−6.628	−6.917	−0.141		−3.086	0.772	−1.908	−3.542
GBP	−2.136	4.262	−2.851	−2.908	−3.542	1.921	−0.398	−3.424	
C: Middle									
USD		9.167	1.520	2.560	3.778	3.885	4.127	2.382	5.331
JPY	3.885	4.181	2.448	2.955	3.246		3.946	2.689	5.350
EUR	3.778	3.011	−0.811	0.125		3.246	6.642	0.754	9.939
GBP	5.331	5.766	5.238	6.351	9.939	5.350	6.248	3.448	
D: Late									
USD		7.142	0.839	1.347	1.830	3.355	2.208	1.368	3.374
JPY	3.355	4.559	2.286	2.762	3.010		4.122	2.555	5.523
EUR	1.830	1.750	−0.318	0.197		3.010	7.194	0.246	7.125
GBP	3.374	4.996	5.428	4.211	7.125	5.523	5.948	2.901	

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

Table 5 Fama regression results (3 month horizon)

coefficient	CAD	CHE	DKR	EUR	JPY	NKR	SKR	GBP
A: Full								
Constant	0.024	0.068	0.029	0.031	0.015	0.016	0.028	0.008
	0.016	0.032	0.022	0.023	0.028	0.021	0.024	0.016
Beta	1.505	−1.965***	−1.447*	−1.723*	0.342	−0.601	−1.656**	−0.145
	2.354	1.320	1.385	1.487	1.076	1.472	1.274	1.684
Adj.R sq.	0.000	0.013	0.007	0.010	−0.003	−0.002	0.011	−0.004
F-statistic	1.150	2.623	2.623	1.678	0.197	1.153	2.177	0.370
N	270	270	270	270	270	270	270	270
B: Early								
Constant	0.045	0.164	0.056	0.069	0.082	0.017	0.055	−0.004
	0.021	0.060	0.031	0.032	0.062	0.037	0.030	0.029
Beta	−3.300*	−5.783***	−5.564***	−5.694***	−1.829	−2.584***	−4.483***	−2.990***
	2.307	2.092	1.933	1.829	1.672	1.624	1.456	1.786
Adj.R sq.	0.036	0.119	0.132	0.146	0.018	0.054	0.150	0.052
F-statistic	3.055	5.276	6.210	6.939	1.473	3.749	8.494	3.950
N	92	92	92	92	92	92	92	92
C: Middle								
Constant	0.034	0.017	0.008	0.004	−0.007	0.077	0.015	0.021
	0.029	0.042	0.028	0.027	0.035	0.027	0.033	0.025
Beta	10.16***	2.297	4.223	5.851***	3.498	6.041**	3.016	8.532**
	4.123	2.424	2.304	2.267	2.097	2.367	2.048	3.403
Adj.R sq.	0.083	0.006	0.053	0.079	0.043	0.083	0.022	0.116
F-statistic	2.498	1.226	1.811	3.256	0.859	4.844	0.841	2.516
N	137	137	137	137	137	137	137	137
D: Late								
Constant	0.026	0.063	0.087	0.097	−0.023	0.048	0.136	0.084
	0.045	0.061	0.075	0.072	0.040	0.076	0.085	0.043
Beta	−4.032	−1.976	−4.333**	−5.211**	1.161	−9.169*	−7.562***	−7.970***
	6.872	2.078	2.602	2.695	1.886	5.904	3.354	3.892
adj.R sq.	−0.011	0.015	0.109	0.149	−0.005	0.040	0.169	0.114
F-statistic	0.307	1.849	3.519	10.892	0.515	5.044	8.096	2.825
N	41	41	41	41	41	41	41	41

Sample period refers to interest rate observations. *(**)[***] denotes significance at the 10%(5%)[1%] marginal significance level for null of unit coefficient on interest differential. The F-statistic refers to the joint null hypothesis that the intercept is null and slope equal to one

Table 6 Fama regression augmented with VIX results (3 month horizon)

Coefficient	CAD	CHE	DKR	EUR	JPY	NKR	SKR	GBP
Full								
Constant	−0.061	0.039	−0.003	0.000	−0.034	−0.078	−0.060	−0.010
	0.048	0.050	0.057	0.053	0.052	0.058	0.071	0.060
Beta	1.879	−1.781***	−1.086	−1.431	0.404	0.529	−1.008	−0.025
	2.342	1.367	1.558	1.576	1.089	1.792	1.311	1.677
Gamma	0.0042*	0.0013	0.0015	0.0014	0.0024	0.0051*	0.0042	0.0009
	0.0025	0.0018	0.0023	0.0022	0.0019	0.0030	0.0036	0.0032
Adj.R sq.	0.035	0.012	0.007	0.010	0.003	0.020	0.028	−0.006
F-statistic	0.837	2.426	1.568	1.703	1.044	1.331	2.197	0.202
N	270.000	270.000	270.000	270.000	270.000	270.000	270.000	270.000
Early								
Constant	−0.022	0.206	0.062	0.062	0.189	0.303	0.022	−0.046
	0.113	0.153	0.137	0.139	0.132	0.169	0.145	0.098
Beta	−5.395***	−7.191***	−7.849***	−7.319***	−1.502	−6.987***	−6.663***	−4.945***
	2.300	2.211	1.979	1.872	1.832	2.338	1.425	1.603
Gamma	0.0021	−0.0013	−0.0012	−0.0004	−0.0049	−0.0170***	0.0003	0.0001
	0.0047	0.0064	0.0058	0.0058	0.0050	0.0080	0.0057	0.0037
Adj.R sq.	0.081	0.148	0.197	0.195	0.019	0.198	0.273	0.128
F-statistic	3.879	7.142	10.020	9.880	1.435	5.869	15.377	6.886
N	76	76	76	76	76	76	76	76
Middle								
Constant	−0.050	−0.064	−0.203	−0.147	−0.125	−0.108	−0.119	−0.043
	0.068	0.066	0.077	0.065	0.075	0.064	0.099	0.081
beta	9.093***	2.456	8.991***	8.242***	2.177	6.375***	4.835**	10.917***
	3.236	1.916	2.620	2.456	1.377	2.297	1.868	3.853
Gamma	0.0041	0.0039	0.0100***	0.0071***	0.0059**	0.0088***	0.0068	0.0040
	0.0041	0.0028	0.0035	0.0030	0.0032	0.0036	0.0056	0.0047
Adj.R sq.	0.116	0.023	0.152	0.138	0.055	0.127	0.082	0.147

Table 6 (continued)

Coefficient	CAD	CHE	DKR	EUR	JPY	NKR	SKR	GBP
F-statistic	3.149	0.482	4.784	12.405	1.393	3.837	2.153	3.405
N	144	144	144	144	144	144	144	144
Late								
Constant	−0.094	0.026	0.043	0.074	−0.023	−0.175	−0.046	0.104
	0.057	0.070	0.103	0.098	0.056	0.101	0.107	0.098
Beta	0.703	−1.498	−3.764*	−4.932***	1.333	−3.583	−4.743*	−8.159*
	6.316	1.992	2.724	2.794	1.828	4.869	3.078	4.864
Gamma	0.0055***	0.0014	0.0025	0.0018	0.0003	0.0107***	0.0073***	−0.0001
	0.0018	0.0018	0.0027	0.0025	0.0018	0.0035	0.0030	0.0030
Adj.R sq.	0.108	−0.002	0.104	0.129	−0.026	0.170	0.213	0.082
F-statistic	2.432	1.419	4.016	4.646	0.092	5.352	7.021	2.144
N	50	50	50	50	50	50	50	50

Sample period refers to interest rate observations. *(**)[***] denotes significance at the 10%(5%)[1%] marginal significance level for null of unit coefficient on interest differential, or null of zero coefficient for VIX coefficient. The F-statistic refers to the joint null hypothesis that the intercept is null and slope equal to one

Table 7 UIP regressions results using survey data on exchange rate expectations (3 month horizon)

coefficient	CAD	CHE	DKR	EUR	JPY	NKR	SKR	GBP
A: Full								
Constant	0.000	−0.054	−0.016	−0.017	−0.057	0.033	0.020	0.000
	0.003	0.007	0.005	0.005	0.007	0.004	0.005	0.004
Beta	0.283**	2.360***	1.188	1.377	2.987***	1.653***	1.374	0.880
	0.328	0.373	0.290	0.293	0.240	0.258	0.294	0.338
Adj.R sq.	0.000	0.349	0.185	0.217	0.597	0.278	0.206	0.088
F-statistic	0.095	114.882	49.056	59.923	314.866	82.706	55.860	21.520
N	213	213	213	213	213	213	213	213
B: Early								
Constant	−0.002	−0.008	0.013	0.012	−0.019	0.026	0.047	0.005
	0.005	0.014	0.006	0.006	0.014	0.006	0.008	0.007
Beta	−0.394**	1.845	1.141	1.105	2.374***	1.168	0.724	0.465*
	0.309	0.510	0.383	0.389	0.347	0.251	0.340	0.316
Adj.R sq.	0.012	0.307	0.199	0.183	0.618	0.268	0.087	0.044
F-statistic	10.720	3.691	2.832	2.613	26.769	10.446	19.421	7.264
N	44	44	44	44	44	44	44	44
C: Middle								
Constant	−0.004	−0.062	−0.027	−0.027	−0.062	0.026	0.008	−0.011
	0.005	0.007	0.004	0.004	0.008	0.005	0.005	0.005
Beta	0.119	2.164***	0.386*	0.673	2.797***	1.352	0.937	0.401
	0.543	0.485	0.360	0.378	0.357	0.364	0.318	0.580
Adj.R sq.	−0.007	0.260	0.012	0.042	0.505	0.150	0.082	0.003
F-statistic	1.350	67.941	49.716	40.936	36.395	16.478	1.266	2.304
N	137	137	137	137	137	137	137	137
D: Late								
Constant	0.011	−0.008	−0.011	−0.002	0.006	0.060	0.002	0.020
	0.005	0.013	0.013	0.013	0.007	0.012	0.013	0.012
Beta	2.450**	0.416	1.536	1.293	0.437	1.906	2.961**	1.346
	0.735	0.441	0.445	0.457	0.333	0.984	0.499	0.817
Adj.R sq.	0.191	0.019	0.344	0.311	0.013	0.090	0.579	0.076
F-statistic	33.856	28.892	7.474	2.822	1.641	67.109	109.359	8.058
N	32	32	32	32	32	32	32	32

Sample period refers to interest rate observations. *(**)[***] denotes significance at the 10%(5%)[1%] marginal significance level for null of unit coefficient. The F-statistic refers to the joint null hypothesis that the intercept is null and slope equal to one

Table 8 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 Forecasts*, Consensus 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
Three month Treasury yield	US Treasury	Quarterly, period average, Start: 1999Q1
Expected 3 month Treasury yield	Federal Reserve Bank of Philadelphia	Quarterly, period average, Start: 1999Q1

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

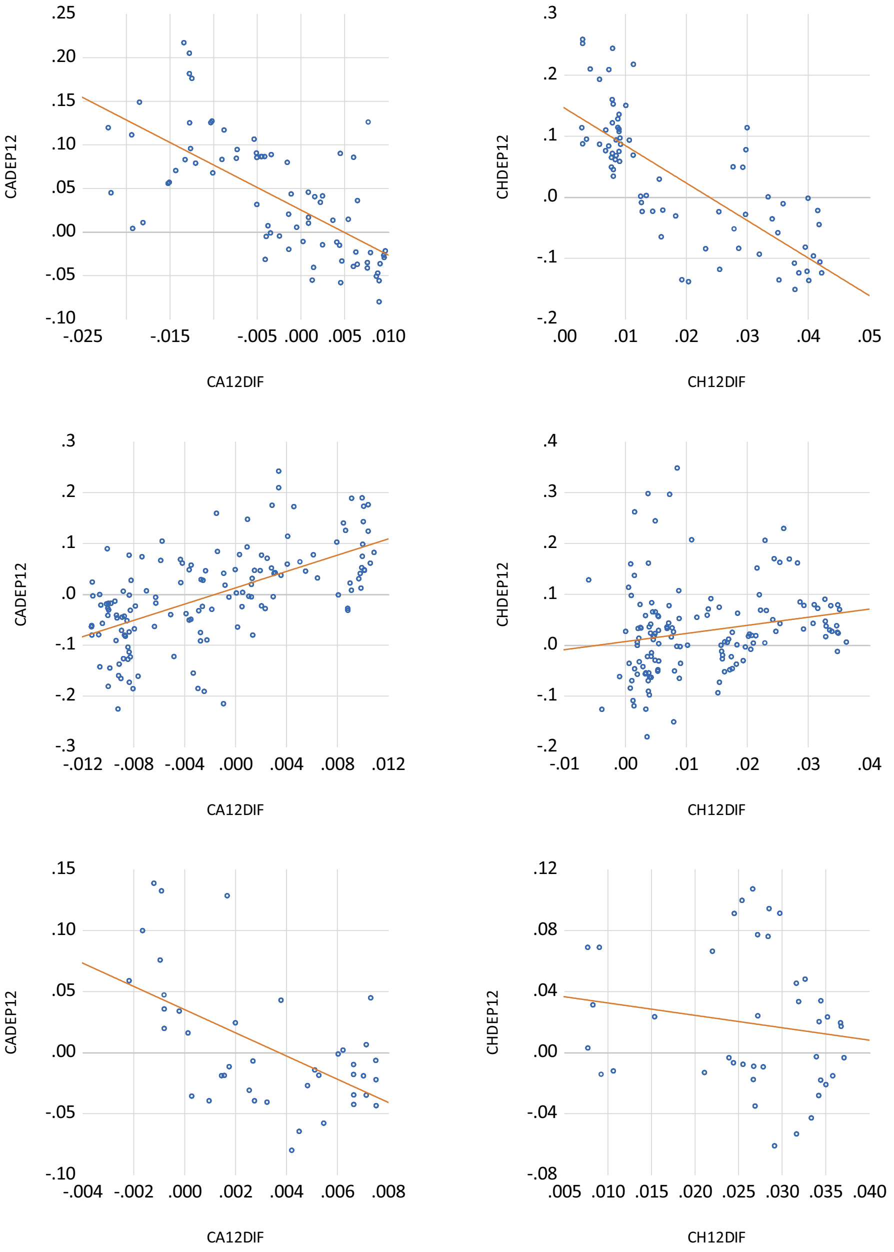


Fig. 11 Scatterplot of the 1 Year Ex-Post Depreciation Rate (1 Year Ahead) on 1 Year Eurodeposit Rate Differential (decimal format). Note: Top graph is Early Period, middle graph is Middle Period, and bottom graph is Late Period. CA denotes Canadian dollar, CH denotes Swiss franc, DK denotes Danish krone, EA denote Euro, JP denotes Japanese yen, NO denote Norwegian krone, SW denotes Swedish krona, and UK denotes British pound. Regression line in red. Authors' calculations based on International Financial Statistics and Thomson Reuters Datastream data

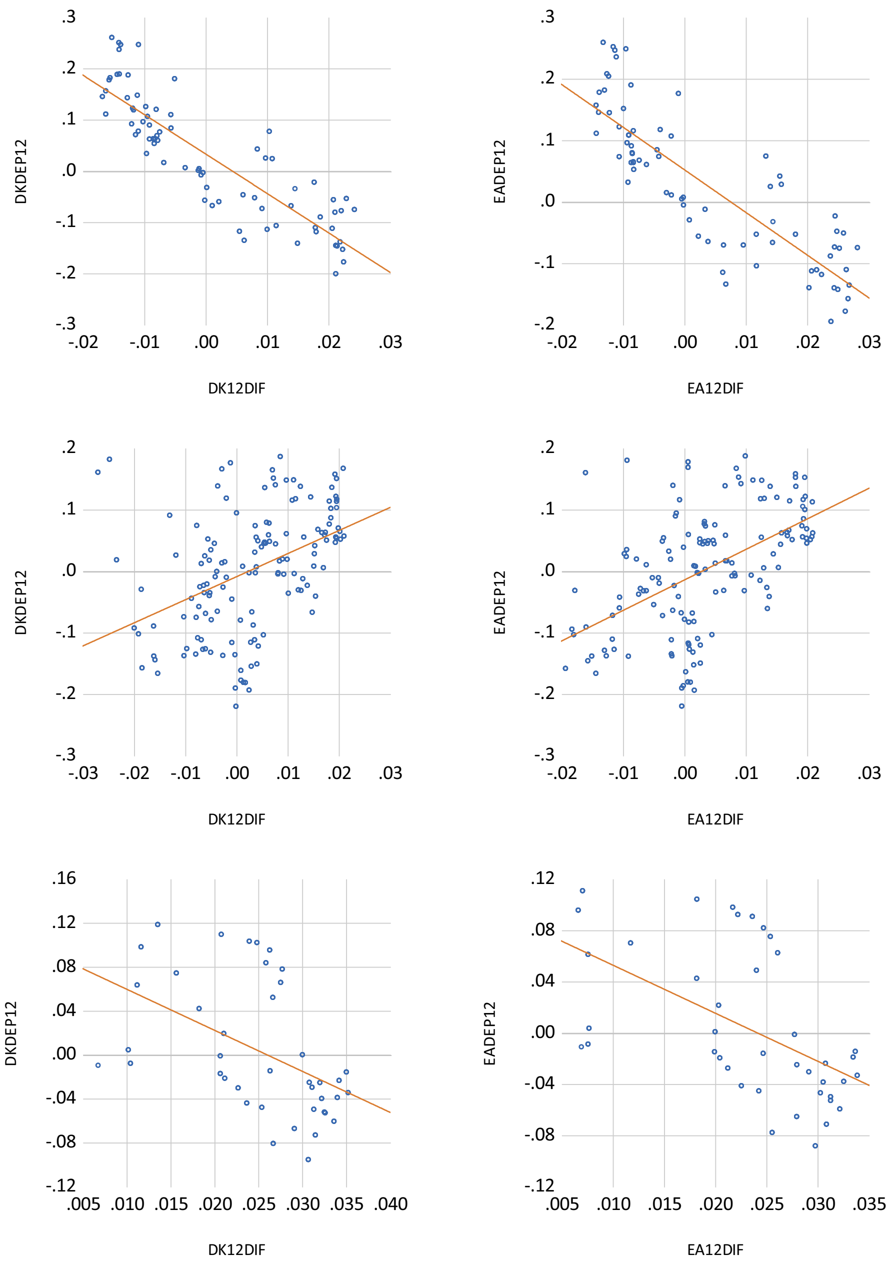


Fig. 11 (continued)

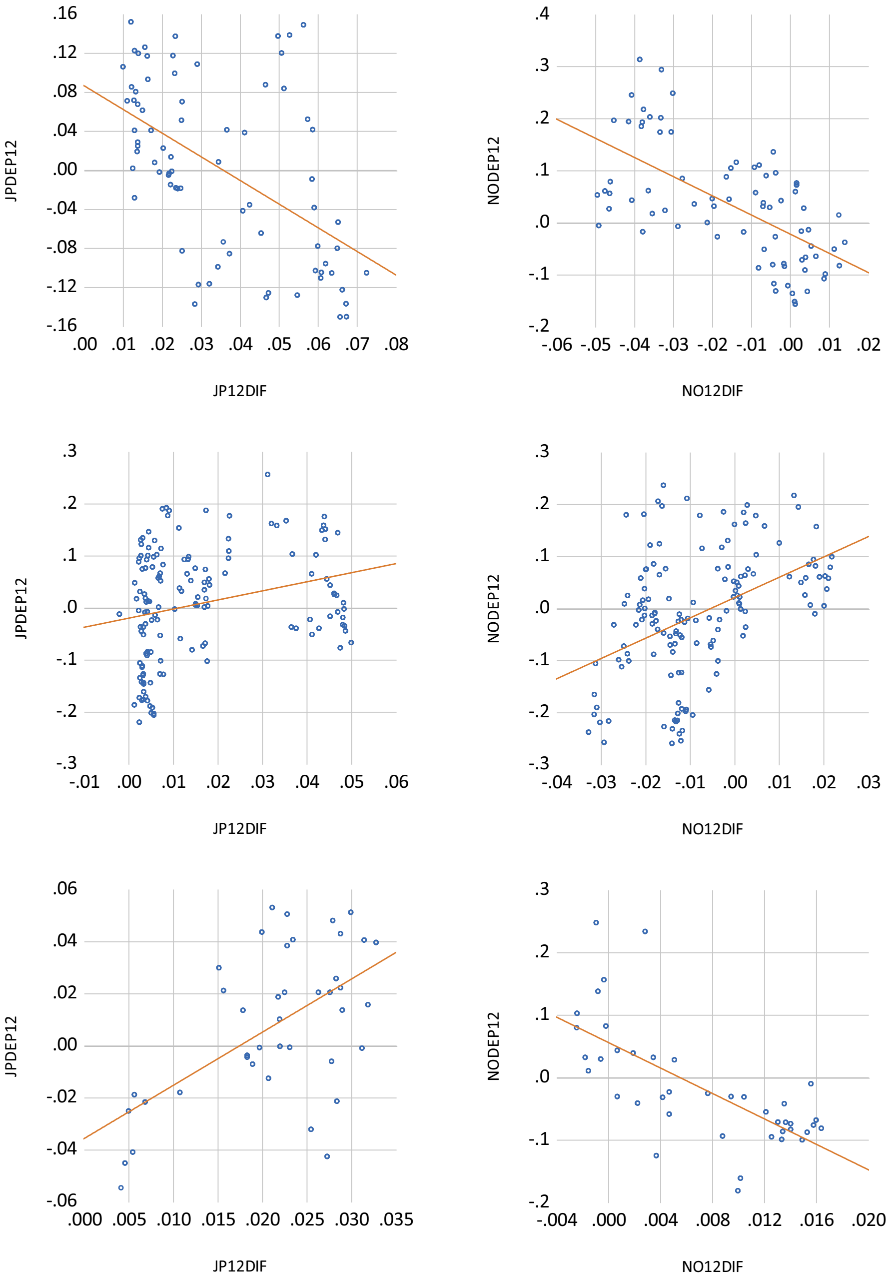


Fig. 11 (continued)

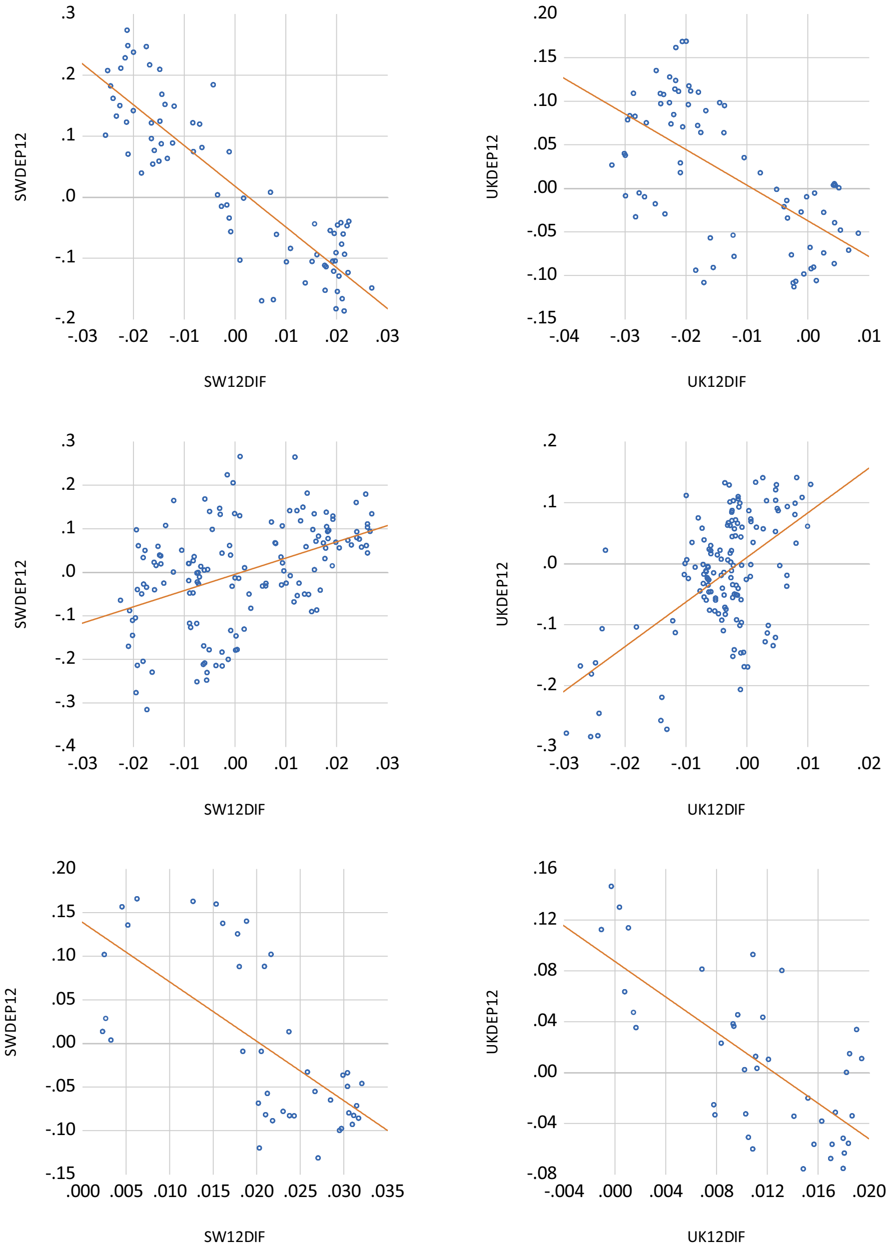


Fig. 11 (continued)

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