



# A reconsideration of the failure of uncovered interest parity for the U.S. dollar<sup>☆</sup>

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## ABSTRACT

We reexamine the time-series evidence on uncovered interest rate parity for the U.S. dollar versus major currencies at short-, medium- and long-horizons. The evidence that interest rate differentials predict foreign exchange returns is not stable over time and disappears altogether when interest rates are near the zero-lower bound. However, we find that year-on-year inflation rate differentials predict excess returns – when the U.S. y.o.y. inflation rate is relatively high, subsequent returns on U.S. deposits tend to be high. We interpret this as consistent with the hypothesis that markets underreact initially to predictable changes in future monetary policy. The predictive power of y.o.y. inflation begins in the mid-1980s when central banks began to target inflation consistently and continues in the post-ZLB period when interest rates lose their primacy as a policy instrument. We attempt to address some econometric problems that might bias the conventional Fama (1984) test.

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## 1. Introduction

Uncovered interest parity (UIP) under rational expectations is the hypothesis that there is no foreign exchange risk premium, or that the expected excess returns on foreign bonds is equal to zero. A very large literature has tested UIP, especially for the U.S. dollar, and a common finding has been that the interest rate differential (U.S. minus foreign interest rate) predicts excess returns, in conflict with the rational expectations/UIP hypothesis.

We reconsider these econometric tests. It has been previously noted that when the hypothesis is tested in subsamples, the model estimates are not stable over time. We examine the power of the interest rate differential to predict excess returns for the U.S. dollar against the major world currencies, with an extended sample. In addition, in assessing the statistical significance of the slope coefficient estimate, we take into account, and correct for, the potential small sample bias first noted by [Stambaugh \(1999\)](#). When we do so, we find that the evidence for a UIP puzzle is weak. The coefficient estimates vary widely over subsamples.

We also consider “medium-run” and “long-run” tests of uncovered interest rate parity, as we define below. These are tests for whether there are predictable returns from an investment strategy that takes positions in short-term deposits in dollars and foreign currencies, and then rolls them over each month.

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Our main contribution is to find that while there has been considerable instability in the relationship between relative interest rates and excess returns, there is a more stable correlation between year-on-year relative inflation rates and subsequent excess returns. We hypothesize that this predictability arises from delayed adjustment to monetary policy changes, though we discuss and cannot rule out the possibility that this relationship arises because of a foreign exchange risk premium.

## 2. Tests of uncovered interest rate parity

Algebraically, the UIP/rational expectations hypothesis is expressed as:

$$E_t s_{t+1} - s_t = i_t - i_t^* \tag{1}$$

where  $s_t$  is the log of the exchange rate expressed as the home currency price of foreign currency,  $i_t$  is the interest rate on a riskless one-period deposit or security in the home country, and  $i_t^*$  is the analogous interest rate in the foreign country.

The uncovered interest parity puzzle arises from an econometric test in Fama (1984) that estimates the slope (and intercept) parameters in the regression<sup>1</sup>:

$$s_{t+1} - s_t = \alpha + \beta(i_t - i_t^*) + u_t \tag{2}$$

where  $\alpha$  and  $\beta$  are parameters to be estimated and  $u_t$  is an error term. One implication of the null hypothesis of uncovered interest parity (hereinafter referred to as UIP) is  $\alpha = 0$  and  $\beta = 1$ .

By a simple algebraic transformation, this regression is equivalent to one that regresses the ex post excess return on foreign bonds on the home minus foreign interest rate differential:

$$s_{t+1} - s_t + i_t^* - i_t = \alpha + \gamma(i_t - i_t^*) + u_t \tag{3}$$

If Eqs. (2) and (3) are estimated by ordinary least squares, the estimates of the intercepts are identical, and the slope coefficient estimates are related as  $\gamma = \beta - 1$ . In this formulation, the null of UIP requires  $\alpha = 0$  and  $\gamma = 0$ .

The ‘‘UIP puzzle’’ or the ‘‘Fama puzzle’’ refers to the finding that the slope coefficient in Eq. (3) is usually found to be negative and often less than  $-1$ . This means when the home interest rate is high, the excess return on the foreign deposit tends to be low.<sup>2</sup>

We reconsider these econometric tests. It has been previously noted that when Eq. (3) is estimated in subsamples, the slope coefficient estimate is not stable over time. For example, Bekaert and Hodrick (2018) present rolling regressions on monthly data, using five-year estimation windows, for the U.S. dollar relative to the euro, the British pound and the Japanese yen. That study finds considerable instability in the slope parameter estimates.<sup>3</sup>

We also undertake such an exercise for the U.S. dollar against the major world currencies, with an extended sample. We find that the slope coefficients from (3) vary widely over subsamples. They tend to be negative in the period before the global financial crisis, but generally are not significantly different from zero using the bias-corrected tests. After 2007, the point estimates of the slope coefficient are positive, but the standard errors are quite large so the null of  $\gamma = 0$  cannot be rejected.

We also consider the correlation of the interest rate differential on the excess return from rolling over short-term bonds. At the ‘‘medium’’ horizon, we estimate:

$$\sum_{j=1}^{12} s_{t+j+1} - s_{t+j} + i_{t+j}^* - i_{t+j} = \alpha_M + \gamma_M(i_t - i_t^*) + u_t^M \tag{4}$$

In this case, we find more evidence of statistical significance, but not stability of the slope coefficient. The estimates of  $\gamma_M$  are found to be negative prior to 2007 and positive afterward.

Then we consider the expected return from rolling over short-term bonds over a long horizon by estimating:

$$E_t \left( \sum_{j=1}^{\infty} s_{t+j+1} - s_{t+j} + i_{t+j}^* - i_{t+j} \right) = \alpha_L + \gamma_L(i_t - i_t^*) + u_t^L \tag{5}$$

The dependent variable in this regression is constructed from vector autoregressions, as described below. Our findings here in sub-samples are fragile, as they depend heavily on the estimated persistence of interest rates. That is, the estimated responsiveness of  $E_t \left( \sum_{j=1}^{\infty} i_{t+j}^* - i_{t+j} \right)$  depends crucially on how persistent the interest differential is measured to be – the more persistent, the

<sup>1</sup> Bilson (1981) is an earlier published paper that performs this test of uncovered interest parity. Fama actually used the forward premium as the independent variable in this regression. During the time period of Fama’s sample, covered interest parity held well, so in this case, as Fama discusses, uncovered interest rate parity and ‘‘forward rate unbiasedness’’ are equivalent. Recent literature has focused on deviations from covered interest parity since 2008. See Du and Schreger (2021) for a survey. We do not address c.i.p. deviations in this paper.

<sup>2</sup> See Engel (1996, 2014) for surveys of these tests of UIP.

<sup>3</sup> Recently, Engel et al. (2019) and Bussière et al. (2018) have noted the ‘‘disappearance’’ of the Fama puzzle in the 2000s. Rossi (2006, 2013) examines the instability in the predictive power for exchange rate changes of the Fama regression.

greater the response. But the persistence is estimated imprecisely in small samples. For the whole sample, we generally find  $\gamma_L$  is not statistically significant, though for one currency it is significantly negative and for another it is significantly positive.<sup>4</sup>

What should we make of this apparent parameter instability? One possibility is suggested by West (2012). Building on the work of Engel and West (2005), that paper demonstrates that the slope coefficient estimate in the Fama regression may be nearly inconsistent if the exchange rate is generated as in a large class of present-value exchange-rate models. In those models, even ones in which uncovered interest parity is posited, the exchange rate is determined by an expected present value of current and future economic fundamental variables. West shows that as the discount factor approaches unity in value, the ordinary least squares estimate of the slope coefficient becomes inconsistent. Intuitively, when the appropriate discount factor is nearly one, the slope coefficient estimates will be unstable over sub-samples, as we find.

An alternative possibility is that the slope coefficient estimates, particularly in samples from the 1980s and 1990s, were reflecting the outcome of some underlying economic process, but that process changed in the 2000s and 2010s. Much of the literature has been devoted to building models of a foreign exchange risk premium to explain the Fama puzzle. However, that literature, in general, does not generally address the parameter instability of the Fama regression. Farhi and Gabaix (2016) is one approach that allows for a change in the slope coefficient, but we shall see that it is not well-suited to explaining the U.S. data. That model posits that one currency has an apparent low expected return during normal times and a high expected return during times of global economic stress. But, for example, while the slope coefficient in regression (3) is estimated to be negative consistently in the 1980s and 1990s, the sign of the regressor changes frequently for most U.S. dollar currency pairs. That implies that it is not the case that one of the currencies consistently offers an ex-ante excess return over this period, in contrast to the implications of Farhi and Gabaix.

Other popular models in this literature that are based on habits or long run risks (e.g. Verdelhan (2010), Bansal and Shaliastovich (2012), and Colacito and Croce (2013)) make different conditional predictions for the joint behavior of interest rates and exchange rate risk premia that could potentially explain the time varying slope coefficient in the Fama regression. Colacito and Croce (2013) note a change in the slope of the Fama regression coefficient before the early 1970s and afterward and attribute the difference to how increased capital mobility influences risk sharing. However, a change in capital mobility is unlikely to be the source of the instability of the coefficient we have seen in the 2000s. Among reduced-form models, Lustig et al. (2014) provides expressions for the regression coefficients as functions of variances of state variables, which could provide clues to the findings we present here. This work is focused on the evolution of the U.S. dollar vis-a-vis a basket of currencies and emphasizes the importance of looking at the dollar for pinpointing the strongest evidence in favor currency return predictability (whether it is driven by risk premia or not). That study suggests that macroeconomic factors may account for the time-variation in the slope coefficient. This idea is further refined in Colacito et al. (2020) which constructs a model based on long-run risks and shows how this might account for empirical findings similar to ours. We discuss this approach in our conclusions section.

Instead, we look at the possibility that a change in how monetary policy is conducted is responsible for the inconsistent sign in regression (3) over sub-samples in rolling regressions. As inflation subsided in the high-income countries, nominal interest rates sank toward zero. At the zero-lower bound, central banks introduced alternative monetary policy tools – “unconventional monetary policy (UMP)” such as forward guidance and quantitative easing. The relationship between interest rates and excess returns may have changed post-2000 because the instruments of monetary policy changed.

One strand of the literature has tried to account for the Fama puzzle by examining the market's reaction to monetary policy changes. Froot and Thaler (1990) and then Eichenbaum and Evans (1995) have suggested that there is a delayed reaction to interest rate changes engineered by monetary policymakers. When money is tightened in the home country, so  $i_t - i_t^*$  rises, the home currency appreciates, so  $s_t$  falls. However, in contrast to the implications of the classic Dornbusch (1976) model that assumes UIP and rational expectations, this approach suggests that the maximum appreciation of the home currency does not occur initially when policy is changed. Instead, there is “delayed overshooting” because markets react slowly to the shock to  $i_t - i_t^*$ . While some market players react quickly and  $s_t$  falls, others adjust their portfolios more slowly, so that the home currency continues to appreciate beyond the initial period of the shock.  $s_{t+1}$  falls more relative to  $s_t$ , which then implies the negative relationship between  $s_{t+1} - s_t$  and  $i_t - i_t^*$  that Fama found empirically. This process is formally modeled in Bacchetta and van Wincoop (2010), with further implications demonstrated in Bacchetta and van Wincoop (2019), in a model in which agents find it costly to adjust their portfolios constantly, and hence the full reaction of the exchange rate to monetary policy changes does not occur immediately.<sup>5</sup>

A related explanation comes from Gourinchas and Tornell (2004), which posits that investors underestimate the persistence of monetary policy changes. For example, when  $i_t - i_t^*$  rises, investors are surprised in period  $t + 1$  that the increase has not dissipated more than it actually does on average. This surprise acts like an unanticipated tightening of monetary policy leading the home currency to be stronger at time  $t + 1$  than it would be under rational expectations of monetary policy. This finding is consistent with the model of expectations formation in Molavi et al. (2021) that posits there is a limit to the complexity of statistical models that agents can assess. Market participants may only be able to assess a model with  $k$  factors driving excess returns. If the

<sup>4</sup> Note that these tests are not the same as those that test uncovered interest parity for long-term bonds. Our test examines whether expected returns from rolling over short-term bonds are predictable. See Chinn and Meredith (2004) and Chinn and Quayyum (2012) for tests of u.i.p. on long-term bonds. Our regressions are also different than Lustig et al. (2019). That paper looks at short-term excess returns on long-term bonds, while we look at long-term excess returns on short-term bonds.

<sup>5</sup> See also the related papers on slow portfolio adjustment of Bacchetta et al. (2020a) and Bacchetta et al. (2020b).

true model is comprised of more than  $k$  factors and the true data generating process decays more slowly than the traders perceive, the slope coefficient in regression (1) will be less than one.

We address the possibility that the UIP puzzle is related to monetary policy by estimating short-run, medium-run and long-run regressions that are analogous to Eqs. (3), (4), and (5), but with year-on-year inflation differences as the regressor:

$$s_{t+1} - s_t + i_t^* - i_t = \alpha + \delta(\pi_t - \pi_t^*) + u_t \quad (6)$$

$$\sum_{j=1}^{12} s_{t+j+1} - s_{t+j} + i_{t+j}^* - i_{t+j} = \alpha_M + \delta_M(\pi_t - \pi_t^*) + u_t^M \quad (7)$$

$$E_t \left( \sum_{j=1}^{\infty} s_{t+j+1} - s_{t+j} + i_{t+j}^* - i_{t+j} \right) = \alpha_L + \delta_L(\pi_t - \pi_t^*) + u_t^L \quad (8)$$

where  $\pi_t - \pi_t^*$  is the home minus foreign inflation rate in the 12 months leading up to period  $t$ . Although all our empirical analysis uses monthly data, we employ the year-on-year inflation rates because they might better measure policymaker's expectations of inflation compared to noisy monthly inflation rates.<sup>6</sup>

We find consistent evidence that the slope coefficients in all these regressions are negative. That finding accords with the stories of delayed reaction by markets, or underestimation of the persistence of monetary policy, when the monetary policy response to inflation is a contraction that causes a currency appreciation. When home inflation rises relative to foreign inflation ( $\pi_t - \pi_t^*$  rises), policymakers may react immediately or with some delay to tighten monetary policy. The home currency should appreciate with a sufficient tightening, but if investors react slowly or do not correctly anticipate the persistence of policy,  $s_{t+1} - s_t$  will fall more than (or not rise as much as) under rational expectations and there will be predictable excess returns on the home currency.<sup>7</sup>

This prediction does not necessarily depend on how monetary policy is implemented – whether interest rates are the instrument of policy or whether the central bank uses unconventional monetary policy. If the central bank can react sufficiently to inflation changes, the exchange rate behavior may be similar under either regime. In line with this, we find that the estimated slope coefficients do not change sign and are consistently negative over time.<sup>8</sup>

In Section 6, we provide a more detailed discussion of this mechanism.

We estimate Eqs. (3)–(8) over four time periods: the entire sample we have for each country (which runs from 1979:06 to 2020:09 for most countries, but starting in 1986:01 for Norway, 1987:01 for Sweden, 1989:01 for Australia and 1997:04 for New Zealand); a sample starting in the mid-1980s that coincides with the inflation-targeting era; a pre-crisis sample, 1987:01–2006:12; and a sample that includes the low interest rate era, 2007:01–2020:09. For the short-term and medium-term regressions, (3), (4), (6), and (7), we also report rolling regression results with 10-year windows using all of the data we have for each exchange rate.<sup>9</sup> In all of our estimates, we use statistics that offer analytical corrections for small-sample bias and serial correlation.

As we shall see, the predictive power of the inflation variable does not become consistently strong until the samples that begin in the mid-1980s. That observation is consistent with the notion that markets have a delayed reaction to monetary policy. Only when central banks began to target inflation more consistently do the markets begin to react to year-on-year inflation as a signal of future monetary policy. As in Clarida and Waldman (2008), bad news about inflation (i.e., high relative inflation) is good news for the currency (i.e., the currency appreciates). Moreover, in the 2000s, as interest rates approached the effective lower bound, central banks found unconventional monetary policy tools to control inflation. During this period, in which the predictive power of interest rates vanishes, y.o.y. inflation differentials continue to have predictive power for excess returns.

All our empirical results relate to the time-series relationship between interest rates and excess returns, or inflation rates and excess returns. We do not, in other words, look at broad cross-sections of returns as in the pioneering work of Lustig and Verdelhan (2007).<sup>10</sup> Hassan and Mano (2019) emphasize the differing implications of cross-section and time-series tests of uncovered interest parity.

Our empirical work is all for U.S. dollar exchange rates, as the failure of UIP in time series has been shown in the literature to be stronger for the dollar than other currencies. We look at the dollar against the G10 currencies, and in our longer samples, with the German mark, French franc and Italian lira in place of the euro.

<sup>6</sup> Engel et al. (2019) find evidence that monthly inflation helps predict returns in the post-2000 era for a subset of currencies considered in this study.

<sup>7</sup> Our finding is reminiscent of the empirical finding of Molodtsova and Papell (2009). That paper finds, during the period of conventional monetary policy, that the estimated Taylor rule forecasts exchange rate changes out of sample, but with the sign opposite of that which would hold under uncovered interest parity.

<sup>8</sup> Chinn and Zhang (2018) previously examined the Fama regression, including a period of a few years (through 2011) in which the major economies approached the ZLB. In that much shorter sample, the deviations from UIP increased, which is the opposite of what we find in our sample that extends to 2020. Ismailov and Rossi (2018) use a sample that goes through 2015, but they emphasize the findings similar to Chinn and Zhang in the sample that goes through only 2011. In fact, their full sample already begins to demonstrate the reversal of the sign of the slope coefficient that we document here.

<sup>9</sup> We also estimated the regression with 5-year rolling windows, and those result in similar conclusions.

<sup>10</sup> See also Lustig et al. (2011, 2014), Verdelhan (2018), Menkhoff et al. (2012a, 2012b), Hassan and Mano (2019) and many others.

Section 3 presents our findings regarding the short-run excess returns. We present results from medium-run tests and long-run tests in Sections 4 and 5, respectively. We offer some conclusions and interpretation in Section 6.

### 3. Short-run

We begin by presenting estimates of the Fama regression, Eq. (2) for the different time periods mentioned above. [Stambaugh \(1999\)](#) showed that in such a regression, if the regressor (the interest rate differential in this case) is serially correlated, and if innovations to the regressor are correlated with the innovations in Eq. (2), the OLS estimate of the slope coefficient as well as the  $t$ -statistic will be biased. In the tables below, we present estimates of  $\beta$  from Eq. (2) and standard errors with the bias corrections derived by [Amihud and Hurvich \(2004\)](#).

[Table 1](#) presents estimates based on the full sample for each currency and includes the sample dates. Recall that a finding that  $\beta \neq 1$  constitutes a rejection of uncovered interest parity, which is equivalent to  $\gamma \neq 0$  in Eq. (3). Moreover, for comparability with the subsequent tests we present as well as with the literature, recall that the literature has tended to find that when  $i_t - i_t^*$  increases, there is a decline in the excess return on foreign bonds,  $s_{t+1} - s_t + i_t^* - i_t$ . That is, the literature usually finds  $\beta < 1$  (and frequently,  $\beta < 0$ ), which, given that  $\gamma = \beta - 1$ , is equivalent to  $\gamma < 0$  ( $\gamma < -1$ ).

The exchange rates we examine are the Australian dollar (AUD), Canadian dollar (CAD), Swiss franc (CHF), German mark (DEM), French franc (FRF), GBP (British pound), Italian lira (ITL), Japanese yen (JPY), Norwegian krone (NOK), New Zealand dollar (NZD), and Swedish krona (SEK). We convert the mark, French franc and lira into euros using the conversion rates at the time of origination of the euro in January 1999.

All the estimated slope coefficients are less than one, consistent with the literature. However, using the bias-corrected estimates, we find that the 95% confidence interval contains unity for all but five of the currencies (CAD, CHF, DEM, GBP, and JPY.) Hence, we find with our full sample that the evidence against UIP is not as strong as has been previously reported in the literature.

The estimates reported in [Table 1](#) use the full sample we have for each currency. Not all currencies have LIBOR rates for the entire period 1979:06–2020:09. Three of these currencies converted into the euro. [Table 2](#) reports the results for the sample that begins in the mid-1980s, 1987:01–2020:09 (except NZD, for which the interest rate data begin in 1997, and AUD which begins in 1989.)

The findings from [Table 2](#) are like those of [Table 1](#). All the estimated slope coefficients are less than one. Only three of the confidence intervals exclude the UIP null – for the Canadian dollar, Swiss franc, and Japanese yen. There are two reasons for the differences between [Tables 1](#) and [2](#). First, [Table 2](#) has no data prior to 1987, and the Fama puzzle was stronger in the data in the 1980s. Second, because the sample is shorter, the standard errors are larger in [Table 2](#).

Included in [Table 2](#) is a fixed-effects panel regression. All the panel estimates we report in the paper omit France and Italy because of the overlap with Germany since the adoption of the euro in January 1999. Furthermore, below we estimate Eq. (6), which uses consumer price inflation to predict excess returns. But those regressions omit Australia and New Zealand because those countries do not report monthly consumer price data. To make the panel regressions comparable, we have omitted Australia and New Zealand in all the panel estimates. We use the [Driscoll and Kraay \(1998\)](#) standard error estimates.

The panel estimate imposes that the slope coefficient is identical across countries (while allowing the intercept to differ) and will be a more powerful test if the restriction is true. As [Table 2](#) shows, the findings against uncovered interest parity are marginal. The estimated slope coefficient is  $-0.14$ , and the null hypothesis ( $\beta = 1$ ) can be rejected with a  $p$ -value of 0.062. This finding is consistent with our country-by-country evidence for the 1987:01–2020:09 period. The null of uncovered interest parity can be weakly rejected.

In fact, when we split the common sample into the pre-Global Financial Crisis (GFC) and the period that includes the GFC, we find the slope coefficient estimates are dramatically different. [Table 3](#) reports estimates of Eq. (2) for the period 1987:01–2006:12, and [Table 4](#) reports the estimates for the 2007:1–2020:09 period. In the earlier period, all the point estimates for the slope coefficient lie below one, though the 95% confidence intervals exclude unity only for CAD, CHF, JPY, and NZD. In other words, the findings are almost identical to the full sample results reported in [Table 1](#).<sup>11</sup> On the other hand, all the slope coefficients in the later period are estimated to be positive, but, importantly, the confidence intervals are very wide, and all contain unity.

The estimates for the panel regressions tell a similar story. On the earlier sample, the slope coefficient estimate is  $-0.371$ , and the null can be weakly rejected, with a  $p$ -value of 0.061. But in the latter period, the point estimate is 4.36, and the standard error of the coefficient estimate is large (2.84), so the null hypothesis of  $\beta = 1$  cannot be rejected.

In [Figs. 1–3](#), we report estimates of the slope coefficient and 95% confidence intervals using 10-year rolling regressions. For space consideration, we display only the results for Germany, the U.K. and Japan. These graphs are representative of the findings for all the countries, and the full set of graphs are presented in the Supplemental Appendix.<sup>12</sup> These figures highlight the instability of the coefficient estimate of the Fama regression. In the graphs, the dates along the horizontal axis mark the beginning of each 10-year sample. The blue shaded areas represent the time periods in which the estimated slope coefficient is significantly less than one at the 5% level. The pink areas are when the estimated coefficient is less than one, but not significantly so. The green

<sup>11</sup> [Table 1](#) reports that the slope coefficient is less than one for the German mark in the full sample, while in the 1989:01–2016:12 sample, the 95% confidence interval does barely contain one.

<sup>12</sup> We also estimate rolling regressions with 5-year windows. They are similar to those we report in the text with 10-year windows, but the confidence intervals are wider.

**Table 1**  
Slope coefficient estimates from Fama regression, full sample.

Currency	Time	Obs.	$b$	95% CI	$p$ -Value
AUD	1989/01–2020/09	381	−0.546	(−2.34, 1.24)	0.09
CAD	1979/06–2020/09	496	−1.031	(−2.18, 0.11)	< 0.01
CHF	1979/06–2020/09	496	−1.467	(−2.90, −0.04)	< 0.01
DEM	1979/06–2020/09	496	−0.919	(−2.51, 0.67)	0.02
FRF	1979/06–2020/09	496	−0.167	(−1.54, 1.20)	0.10
GBP	1979/06–2020/09	496	−1.759	(−3.42, −0.10)	< 0.01
ITL	1979/06–2020/09	496	0.467	(−0.43, 1.37)	0.25
JPY	1979/06–2020/09	496	−1.607	(−3.03, −0.18)	< 0.01
NOK	1986/01–2020/09	417	0.280	(−1.18, 1.74)	0.33
NZD	1997/04–2020/09	282	−0.110	(−3.10, 2.88)	0.46
SEK	1987/01–2020/09	405	0.658	(−1.06, 2.37)	0.70

Notes: This table reports the slope coefficient estimates ( $b$ ) from Eq. (2),  $s_{t+1} - s_t = \alpha + \beta(i_t - i_t^*) + u_t$ , based on the full sample (the longest covers 1979:06–2020:09) for each currency, including Australian dollar (AUD), Canadian dollar (CAD), Swiss franc (CHF), German mark (DEM), French franc (FRF), British pound (GBP), Italian lira (ITL), Japanese yen (JPY), Norwegian krone (NOK), New Zealand dollar (NZD), and Swedish krona (SEK). The exchange rates are against US dollar, and  $s_t$  is the log of the exchange rate expressed as the home currency (dollars) price of foreign currency. Exchange rates of the mark, French franc and lira are converted into euros using the euro conversion rates at the time of origination of the euro in January 1999.  $i_t$  is the interest rate on a riskless one-period deposit or security in the home country (U.S.) and  $i_t^*$  is the analogous interest rate in the foreign country. Sample dates vary across currencies, especially for AUD, NOK, NZD, SEK, as shown in the Time column, due to the data availability of interest rates. The CI column reports the 95% confidence intervals of the slope coefficient, with the bias-corrected standard errors derived by Amihud and Hurvich (2004). The  $p$ -value column reports the  $p$ -value of the two-sided  $t$ -test for the slope coefficient  $H_0: \beta = 1$ .

**Table 2**  
Slope coefficient estimates from Fama regression, longest common sample, 1987:01–2020:09.

Currency	Time	Obs.	$b$	95% CI	$p$ -Value
AUD	1989/01–2020/09	381	−0.546	(−2.34, 1.24)	0.09
CAD	1987/01–2020/09	405	−0.565	(−1.93, 0.80)	0.03
CHF	1987/01–2020/09	405	−1.416	(−3.47, 0.63)	0.02
DEM	1987/01–2020/09	405	−0.628	(−2.51, 1.26)	0.09
FRF	1987/01–2020/09	405	−0.386	(−2.18, 1.41)	0.13
GBP	1987/01–2020/09	405	−0.173	(−2.20, 1.86)	0.26
ITL	1987/01–2020/09	405	0.712	(−1.07, 2.50)	0.75
JPY	1987/01–2020/09	405	−1.061	(−2.86, 0.74)	0.03
NOK	1987/01–2020/09	405	0.461	(−1.05, 1.97)	0.48
NZD	1997/04–2020/09	282	−0.110	(−3.10, 2.88)	0.47
SEK	1987/01–2020/09	405	0.658	(−1.06, 2.37)	0.70
Panel	1987/01–2020/09	2828	−0.138	(−1.33, 1.06)	0.06

Notes: This table reports the slope coefficient estimates ( $b$ ) from Eq. (2),  $s_{t+1} - s_t = \alpha + \beta(i_t - i_t^*) + u_t$ , based on the longest common sample (1987:01–2020:09, except for AUD, NZD, and SEK) of the currencies. See Table 1 for complete notes. The last row reports the slope coefficient estimates ( $b$ ) from fixed-effect Panel regression,  $s_{c,t+1} - s_{c,t} = \alpha_c + \beta(i_{c,t} - i_{c,t}^*) + \tau_c + u_{c,t}$ , with Driscoll-Kraay standard errors. This currency-month panel includes CAD, CHF, DEM, GBP, JPY, NOK, and SEK.

**Table 3**  
Slope coefficient estimates from Fama regression, pre-GFC sample, 1987:01–2006:12.

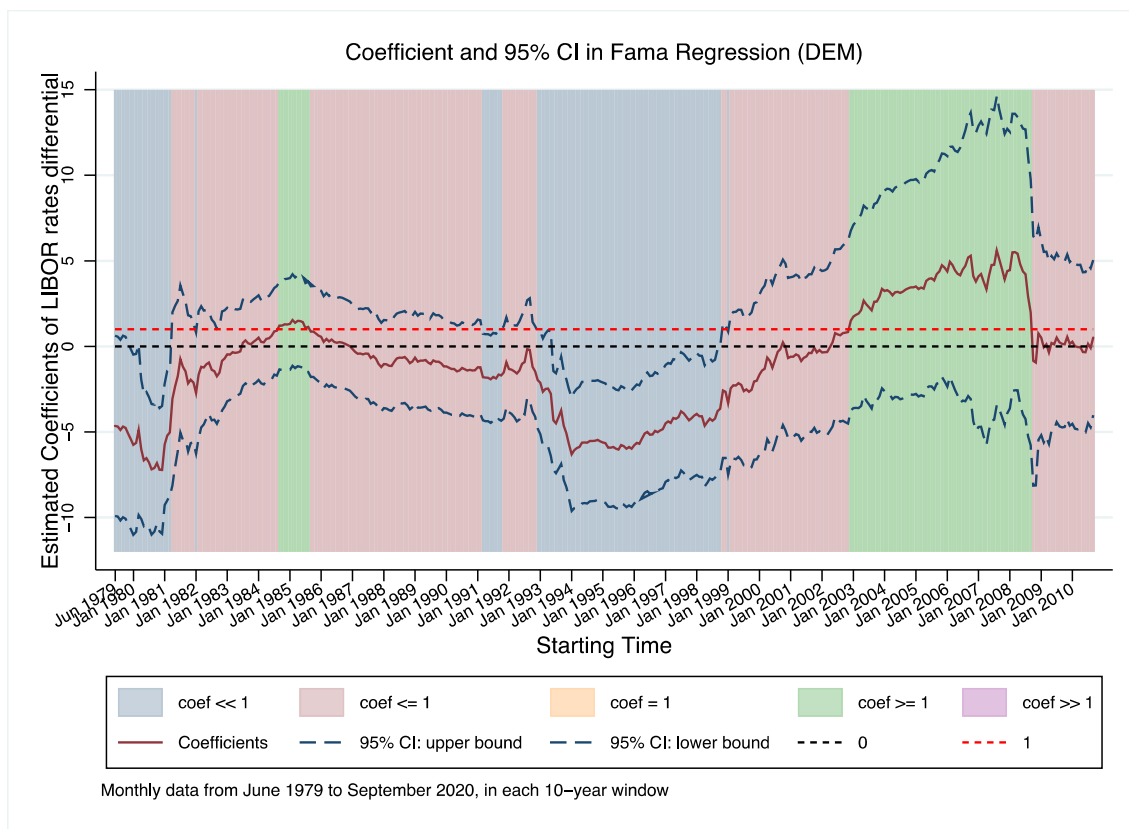
Currency	Time	Obs.	$b$	95% CI	$p$ -Value
AUD	1989/01–2006/12	216	−0.813	(−2.77, 1.14)	0.07
CAD	1987/01–2006/12	240	−0.821	(−2.01, 0.36)	< 0.01
CHF	1987/01–2006/12	240	−1.653	(−3.95, 0.64)	0.02
DEM	1987/01–2006/12	240	−0.978	(−3.07, 1.12)	0.07
FRF	1987/01–2006/12	240	−0.611	(−2.65, 1.43)	0.12
GBP	1987/01–2006/12	240	−0.109	(−2.65, 2.43)	0.39
ITL	1987/01–2006/12	240	0.823	(−1.40, 3.04)	0.88
JPY	1987/01–2006/12	240	−2.448	(−4.60, −0.30)	< 0.01
NOK	1987/01–2006/12	240	0.378	(−1.32, 2.08)	0.47
NZD	1997/04–2006/12	117	−3.714	(−7.38, −0.05)	0.01
SEK	1987/01–2006/12	240	0.823	(−1.14, 2.79)	0.86
Panel	1987/01–2006/12	1680	−0.371	(−1.80, 1.06)	0.06

Notes: This table reports the slope coefficient estimates ( $b$ ) from Eq. (2),  $s_{t+1} - s_t = \alpha + \beta(i_t - i_t^*) + u_t$ , based on the pre-Global Financial Crisis (GFC) common-sample period (1987:01–2006:12, except for AUD and NZD) for each currency and fixed-effect Panel regression. See Tables 1 & 2 for complete notes.

**Table 4**  
Slope coefficient estimates from Fama regression, post-2007:01 sample, 2007:01–2020:09.

Currency	Time	Obs.	<i>b</i>	95% CI	<i>p</i> -Value
AUD	2007:01–2020:09	165	0.022	(−4.13, 4.17)	0.65
CAD	2007:01–2020:09	165	4.56	(−4.82, 13.94)	0.46
CHF	2007:01–2020:09	165	0.885	(−3.38, 5.15)	0.96
EUR	2007:01–2020:09	165	1.507	(−3.11, 6.12)	0.83
GBP	2007:01–2020:09	165	4.564	(−0.86, 9.99)	0.20
JPY	2007:01–2020:09	165	3.609	(0.22, 7.00)	0.13
NOK	2007:01–2020:09	165	2.955	(−2.02, 7.93)	0.44
NZD	2007:01–2020:09	165	3.559	(−0.95, 8.07)	0.27
SEK	2007:01–2020:09	165	0.775	(−3.17, 4.72)	0.91
Panel	2007:01–2020:09	1148	4.36	(−1.21, 9.93)	0.24

Notes: This table reports the slope coefficient estimates (*b*) from Eq. (2),  $s_{t+1} - s_t = \alpha + \beta(i_t - \bar{i}_t) + u_t$ , based on the 2007:01–2020:09 period for each currency and a fixed-effect panel regression. See Tables 1 & 2 for complete notes.



**Fig. 1.** 10-year rolling regressions: DEM.

Notes: This figure presents the estimates of the slope coefficient (*b*) and 95% confident intervals using 10-year rolling Fama regression,  $s_{t+1} - s_t = \alpha + \beta(i_t - \bar{i}_t) + u_t$ , using German mark (DEM) data during 1979:06–2020:09.  $s_t$  is the log of the exchange rate expressed as the home currency (US dollar) price of German mark. The exchange rate is converted into euros using the euro conversion rates at the time of origination of the euro in January 1999.  $i_t$  is the interest rate on a riskless one-period deposit or security in the home country (U.S.) and  $\bar{i}_t$  is the analogous interest rate in Germany. The dates along the horizontal axis mark the beginning of each 10-year sample. The maroon line shows the estimates of the slope coefficient and navy lines draw the 95% confident intervals for each 10-year sample. The blue shaded areas represent the time periods in which the estimated slope coefficient is significantly less than one at the 5% level. The pink areas are when the estimated coefficient is less than one, but not significantly so. The green areas are dates in which the estimated coefficient is greater than one. The purple areas are times in which the slope coefficient is significantly greater than one at the 5% level. The 0 (black dashed line) and 1 (red dashed line) horizontal lines are drawn for reference. The coefficients and 95% confident intervals are bias-corrected following Amihud and Hurvich (2004).

areas are dates in which the estimated coefficient is greater than one. The purple areas are times in which the slope coefficient is significantly greater than one at the 5% level.

The picture that emerges is fairly incoherent, and that is the important lesson. While the Fama puzzle is indicated by the findings that the slope coefficient is significantly less than one, for most of the currencies the time periods over which that is true are

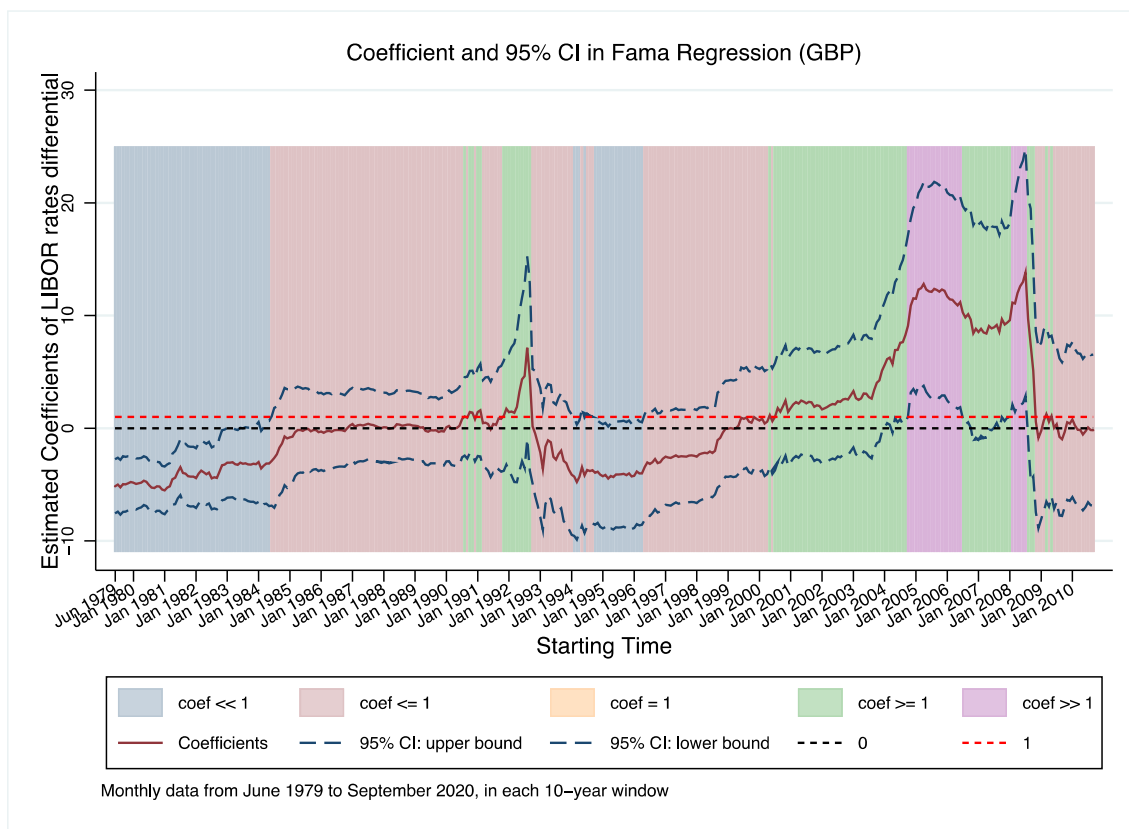


Fig. 2. 10-year rolling regressions: GBP.

Notes: This figure presents the estimates of the slope coefficient ( $b$ ) and 95% confident intervals using 10-year rolling Fama regression,  $s_{t+1} - s_t = \alpha + \beta(i_t - i_t^*) + u_t$ , using British pound (GBP) data during 1979:06–2020:09. See Fig. 1 for complete notes.

relatively short and concentrated in the pre-2000 period. Most of the time, the 95% confidence interval includes unity. For some of the currencies, such as the New Zealand dollar, there are extended periods of time for which the estimated coefficient is greater than one.

In no case is there a sample that starts after January 1999 that has a slope coefficient in the Fama regression that is less than one. Put another way, no sample that ends after the start of either the global financial crisis or the start of the very-low interest rate era evinces the Fama puzzle.

We do not perform joint tests of significance, but it is notable that the estimated slope coefficients rise during the latter part of the sample for all the currencies. There are two possible interpretations of these graphs of rolling estimates of the slope coefficient: The first is that the parameter estimate is very unstable, and there is no true underlying relationship between  $i_t - i_t^*$  and  $s_{t+1} - s_t + i_t^* - i_t$  estimated by regression (2). This possibility is consistent with West's (2012) observation that the parameter estimate is nearly inconsistent as the discount factor in present value models of the exchange rate approaches one. Note that we are not making the case that UIP should be “accepted” for the later periods in which we fail to reject the null of UIP. Indeed, as we shall see when we estimate Eq. (6), inflation differentials have predictive power for excess returns during the 2000s. The standard errors for the slope coefficients reported in Table 4 are quite large, and mechanically that is primarily a result of the fact that the volatility of  $i_t - i_t^*$  is very small during this period. The second possibility is that there is a significant change in the underlying economic relationship between interest rates and excess returns, but we have not detected statistical significance because we have not used a joint test of significance, and because the 10-year window is relatively short.

We performed tests for a structural break in the estimate of the slope coefficient in regression (2), but, unsurprisingly, the results were quite inconclusive. As evidenced from the tables and figures, while the value of the slope coefficient has changed greatly over time, the standard errors of the coefficient estimates in the latter half of the sample are very large. We cannot reject UIP during this period, but also for most currencies, cannot reject the null of no change in the slope coefficient.<sup>13</sup>

<sup>13</sup> We performed Andrews (1993) sup Wald tests for each currency and found significant evidence of a break only for the Japanese yen, Norwegian krone, New Zealand dollar and Swedish krona.



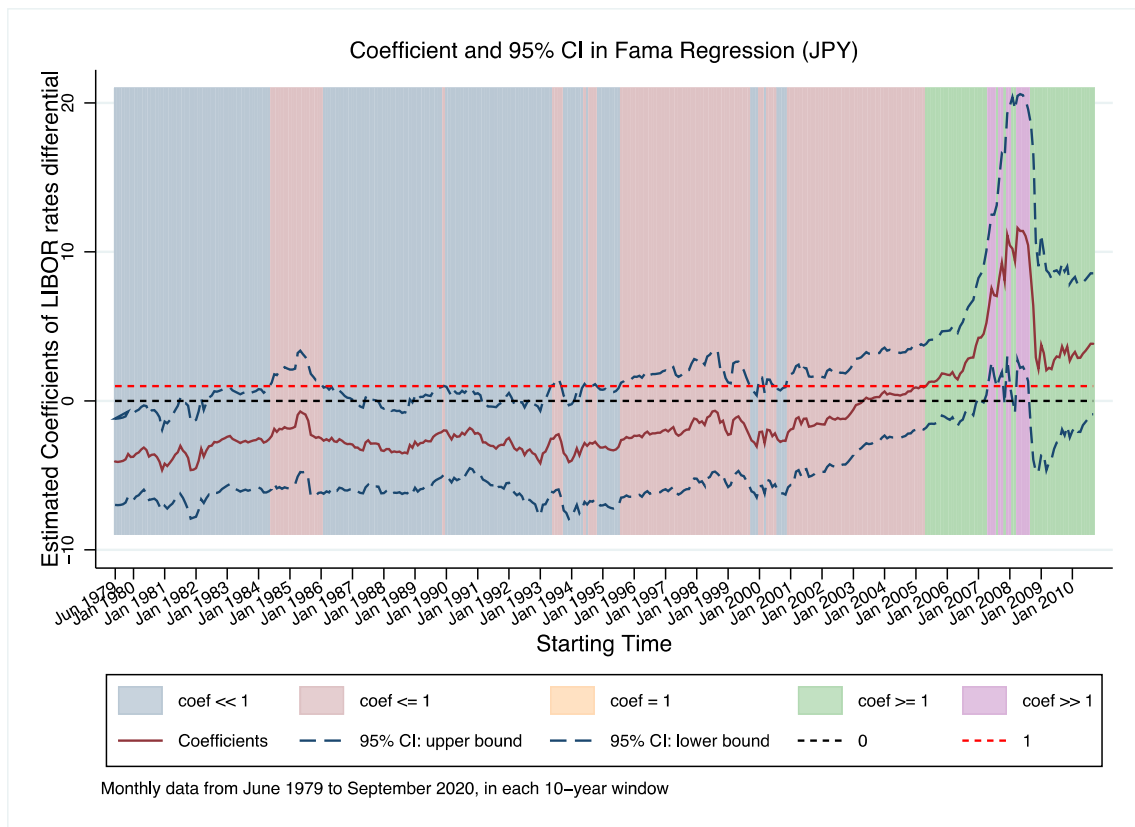


Fig. 3. 10-year rolling regressions: JPY.

Notes: This figure presents the estimates of the slope coefficient ( $b$ ) and 95% confident intervals using 10-year rolling Fama regression,  $s_{t+1} - s_t = \alpha + \beta(i_t - i_t^*) + u_t$ , using Japanese yen (JPY) data during 1979:06–2020:09. See Fig. 1 for complete notes.

We can ask if there are some regularities in the data that can account for this rise in the estimated slope coefficient. Figs. 4–6 plot the estimated slope coefficients from the 10-year rolling Fama regressions and the inverse of the variance of  $i_t - i_t^*$  estimated over the same 10-year time window for Germany, the U.K. and Japan.<sup>14</sup>

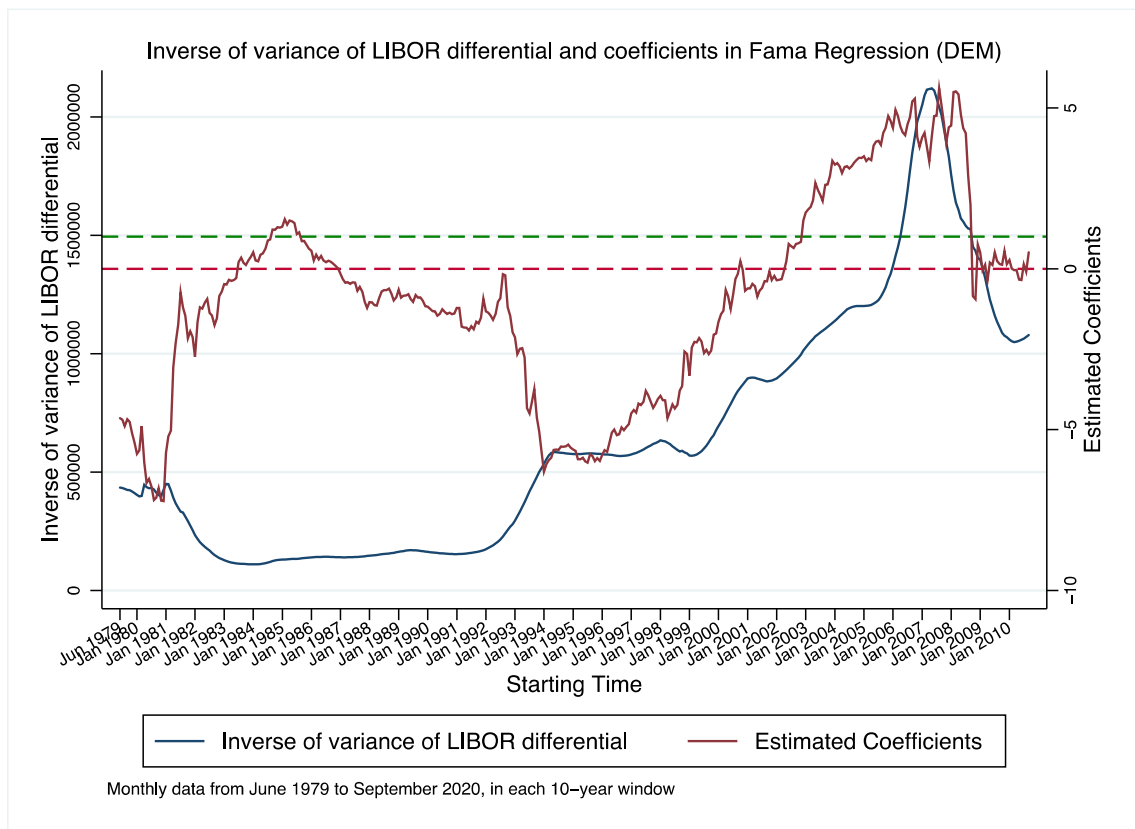
These graphs show a clear and strong positive correlation between the reciprocal of the variance of  $i_t - i_t^*$  and the slope coefficient estimates from rolling regressions of Eq. (2). Table 5 confirms this by reporting estimates of the correlation between the slope coefficient from these rolling regressions beginning with 1989:1, and various measures of the volatility of  $i_t - i_t^*$ , using both 5-year and 10-year windows.

What economic forces lead to this pattern of correlations? One explanation that has been advanced for the puzzling finding of  $\beta < 1$  in the Fama regression is that there is a “peso problem” or the market is incorporating the possibility of a “rare disaster”. Farhi and Gabaix (2016) offer a model of such a phenomenon. That model posits that high-interest rate currencies incorporate a risk premium during “good times” because they are currencies that will depreciate greatly during times of global uncertainty. Thus, the finding of  $\beta < 1$  in the Fama regression arises both because the high-interest-rate currency has a higher expected return than lower-interest-rate currencies, and because of a peso problem in which the sample that contains only good times does not incorporate periods in which the high-interest-rate currency has a large depreciation. During bad times, this same currency has a high interest rate, but the currency depreciates, leading to a slope coefficient less than one in the Fama regression.

The Farhi and Gabaix (2016) model could account for our findings. The model produces a regression coefficient in Eq. (2), the Fama regression, that is always less than one. If the sample over which the model is estimated does not include a “disaster”, the slope coefficient should be negative in their formulation. In the full sample that includes the disaster, the slope increases, but we should still find  $\beta < 1$ , as we do in our full sample. (However, we have noted that the slope is statistically significantly less than one for fewer than half the countries.)

Fig. 7a–c, however, demonstrates some difficulties with this interpretation for the variation in the slope coefficient estimate from Eq. (2) for the U.S. dollar. The figure plots the interest rates for the U.S. and the other countries. The problem, as the graphs illustrate, is that there is not a sustained period in which the U.S. is either the low-interest-rate currency or the high-interest-rate currency in the pre-2007 period. In the Farhi and Gabaix (2016) model, the country with the low interest rate is the less risky

<sup>14</sup> The graphs for all the countries are in the Supplemental Appendix.



**Fig. 4.** Fama slope coefficient and inverse of variance of  $i_t - i_t^*$ : DEM.

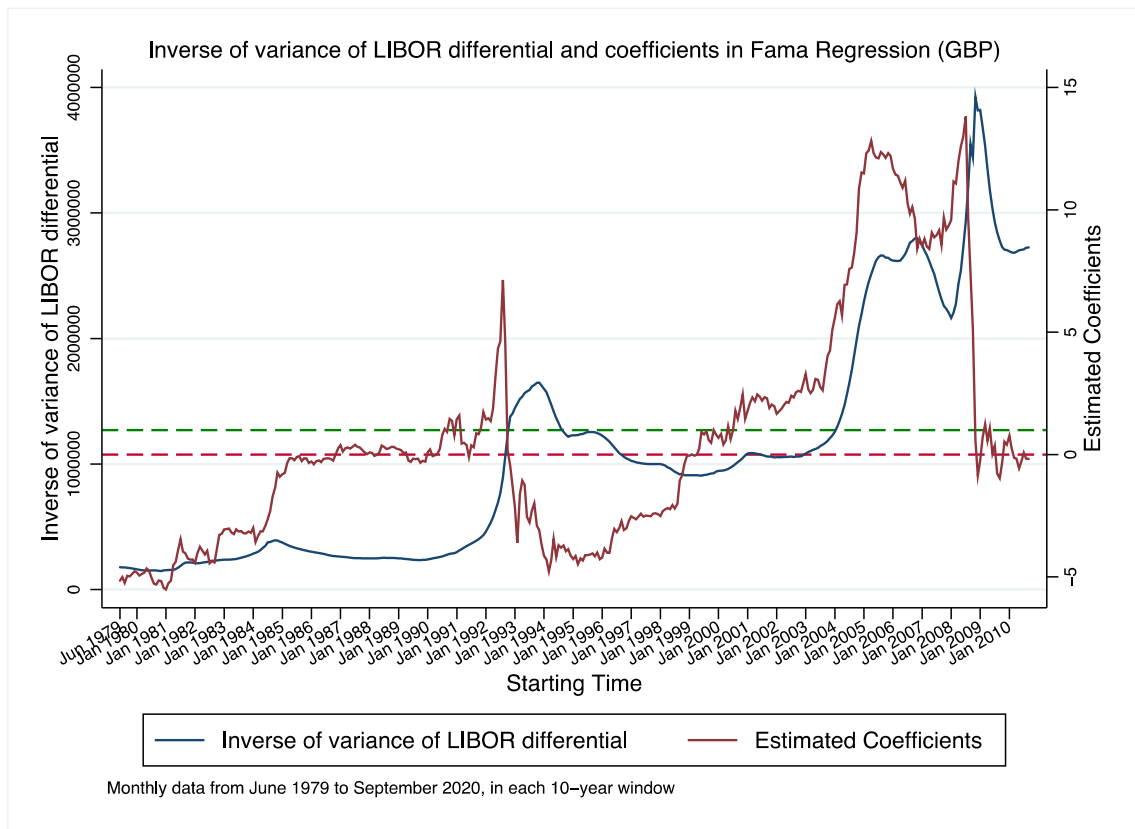
Notes: This figure presents the co-movement of the Fama slope coefficient and the inverse of variance of  $i_t - i_t^*$ . The estimates of the slope coefficient ( $b$ ) using 10-year rolling Fama regression,  $s_{t+1} - s_t = \alpha + \beta(i_t - i_t^*) + u_t$ , based on German mark (DEM) data during 1979:06–2020:09, as in Fig. 1. The maroon line shows the estimates of the slope for each 10-year sample, with starting date in the x-axis, and magnitudes in the right-y-axis. The 0 (red dashed line) and 1 (green dashed line) horizontal lines are drawn for reference of coefficient estimates. The navy line shows the inverse of  $i_t - i_t^*$  in each corresponding 10-year sample, and values are in the left-y-axis.

country, but if the model is correct, the riskiness of the country would have to switch frequently as the sign of the interest rate differential switches in the data. That is, we cannot identify the dollar as, for example, a low-interest-rate currency that is expected to appreciate during times of global uncertainty. Sometimes its interest rate is lower than each of the other countries, and sometimes it is higher. The graphs do reveal that the interest rate of Japan, and to a lesser extent Switzerland, were consistently lower than that of the U.S. pre-2007, but the puzzling behavior of the slope coefficient in regression (2) applies to U.S. dollar regressions, not yen or Swiss franc regressions.

We also can see in Table 6 that the correlation of the slope coefficients from the Fama regression with the market measure of uncertainty, VIX, are not as high as those reported in Table 5. That is, we would expect under the peso problem/rare disaster explanation a strong negative correlation between the Fama coefficient and VIX in the 5-year or 10-year moving average, but that is not the case in fact.

We tentatively offer a different interpretation. As is well known and confirmed by Fig. 7a–c, nominal interest rates in these high-income countries were near zero or below beginning very soon after the onset of the GFC. At such low levels, interest rates are no longer the most useful policy instrument for central banks. Instead, central banks implemented a variety of unconventional policies such as quantitative easing and forward guidance. Our hypothesis is that the change in the slope coefficient in the Fama regression is related to this change in the principle monetary policy instrument. Indeed, post-2007, while the slope coefficient estimates are positive, the confidence intervals are very wide. We think the interest rate differentials may be misleading guides to the relative monetary policy stance in this era. After 2007, U.S. interest rates, while historically close to zero, were higher than in most other countries. At the same time, the Federal Reserve pursued unconventional monetary policies more aggressively than most other countries, and so their overall stance may have been more accommodative.

In other words, prior to the GFC, the finding of  $\beta < 1$  in Eq. (2) could be explained by the models of “delayed reaction” to monetary policy changes. When the Fed, for example, tightened monetary policy,  $i_t$ , and hence  $i_t - i_t^*$  increased, which led to an immediate appreciation of the dollar (a drop in  $s_t$ .) However, perhaps because markets did not perceive how persistent the decline in the interest rate differential would be, or perhaps because expectations are sticky, or perhaps because portfolio adjustment is costly, or perhaps because of balance sheet constraints, the initial appreciation was not the maximum appreciation.



**Fig. 5.** Fama slope coefficient and inverse of variance of  $i_t - i_t^*$ : GBP.  
 Notes: This figure presents the co-movement of the Fama slope coefficient and the inverse of variance of  $i_t - i_t^*$ . The estimates of the slope coefficient ( $b$ ) using 10-year rolling Fama regression,  $s_{t+1} - s_t = \alpha + \beta(i_t - i_t^*) + u_t$ , based on British pound (GBP) data during 1979:06–2020:09, as in Fig. 2. See Fig. 4 for complete notes.

Exchange rates continued to fall, so  $s_{t+1} - s_t + i_t^* - i_t$  fell when  $i_t - i_t^*$  increased. We elaborate on this hypothesis in the concluding section.

That same slow reaction to monetary policy changes may have been at work in the post-2007 era, but monetary policy stance is not well captured by the interest rate differential. To shed some potential light on the subject, we estimated Eq. (6). The regressor in this equation is  $\pi_t - \pi_t^*$ , the difference in the year-on-year inflation rate in the U.S. relative to the foreign country. Our notion is this: year-on-year inflation is a proxy for policymakers' expectations of inflation. When  $\pi_t - \pi_t^*$  rises, the U.S. will very soon tighten monetary policy relative to the foreign country. Then, following the logic of the previous paragraph, delayed reaction by markets leads to high returns on dollar assets relative to the foreign country, implying a negative slope coefficient in this regression.

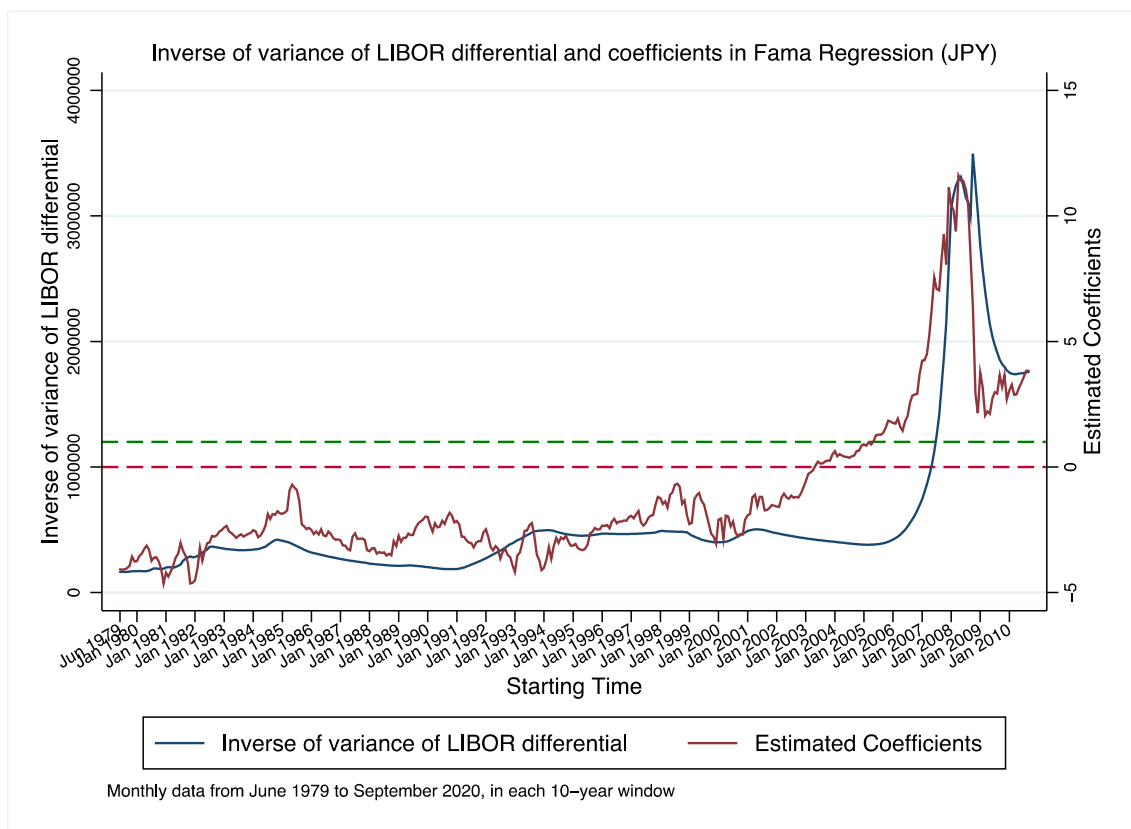
Table 7 offers some evidence that year-on-year inflation is a reasonable proxy for the stance of monetary policy. We do not attempt to estimate a Taylor rule for each country, because measurement of the output gap has been difficult over recent years and is very sensitive to the methodology. Instead, we estimate (as a panel, with fixed effects, using Driscoll-Kraay standard errors), the simple and reproducible relationship:

$$i_{c,t} = \mu_{c,0} + \mu_1 \pi_{c,t} + \mu_2 i_{c,t-1} + u_{c,t}$$

where  $c$  is the country index. As Table 7 shows, over the entire sample, the estimate of  $\mu_1$  is positive and significant.<sup>15</sup> That is also true for the 1987:01–2006:12 sub-period. But in the ZLB period, the estimate of  $\mu_1$  is much smaller and is statistically insignificant. This indicates that year-on-year inflation is a reasonable proxy for monetary policy in the pre-2007 period because interest rates rise in reaction to higher inflation. In the later period, however, interest rates are no longer the main instrument of monetary policy, but we extrapolate from the earlier period to postulate that the unconventional policy tools still respond to year-on-year inflation.

We report estimates for Eq. (6) for the same time periods as we have for the Fama regression. As with the Fama regression, these coefficient estimates and standard errors are bias-corrected using the statistics of Amihud and Hurvich (2004).

<sup>15</sup> Table 7 also reports results for estimating this equation in the form of G10 country relative to the U.S., but the conclusions are identical.

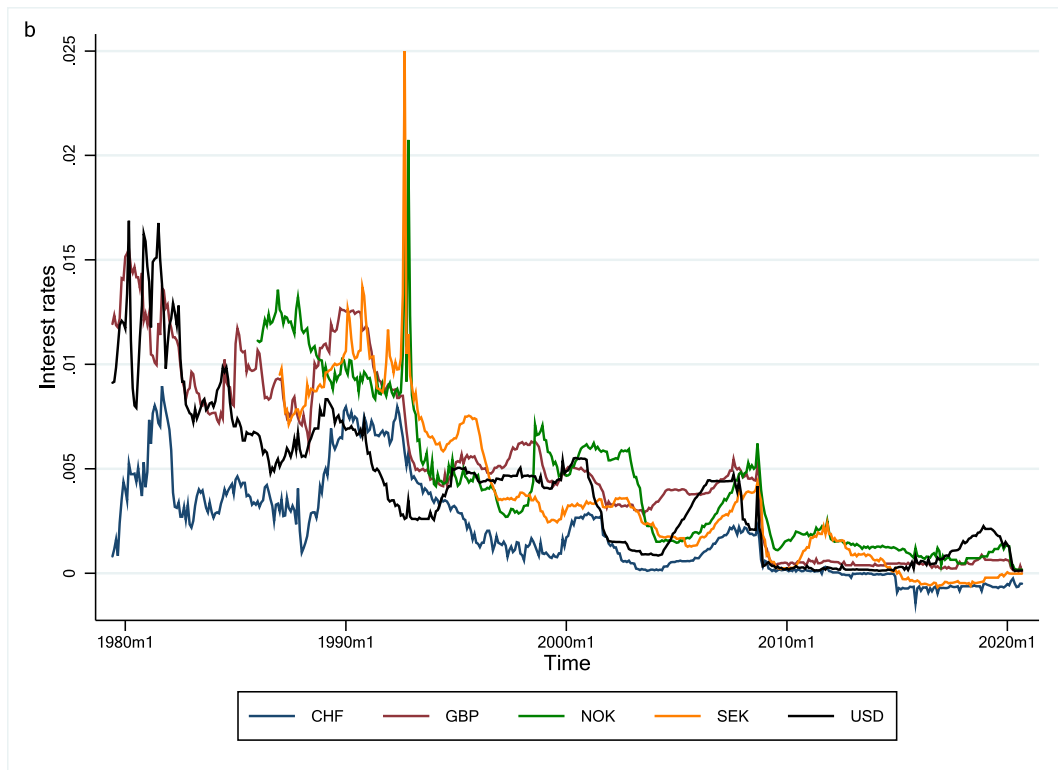
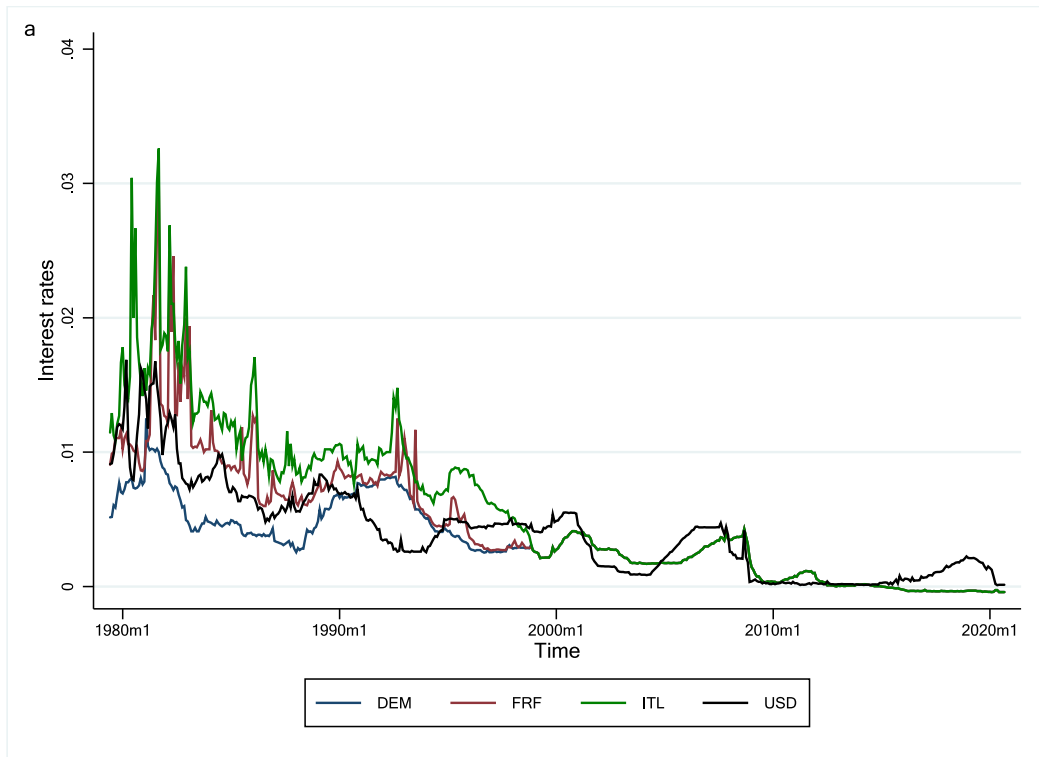


**Fig. 6.** Fama slope coefficient and inverse of variance of  $i_t - i_t^*$ : JPY.  
 Notes: This figure presents the co-movement of the Fama slope coefficient and the inverse of variance of  $i_t - i_t^*$ . The estimates of the slope coefficient ( $b$ ) using 10-year rolling Fama regression,  $s_{t+1} - s_t = \alpha + \beta(i_t - i_t^*) + u_t$ , based on Japanese yen (JPY) data during 1979:06–2020:09, as in Fig. 3. See Fig. 4 for complete notes.

**Table 5**  
 Correlation between Fama coefficient and volatility of  $i_t - i_t^*$ .

Currency	$var(i_t - i_t^*)$		$s. d. (i_t - i_t^*)$		$\frac{1}{var(i_t - i_t^*)}$		$\frac{1}{s.d.(i_t - i_t^*)}$	
	10-Year	5-Year	10-Year	5-Year	10-Year	5-Year	10-Year	5-Year
AUD	-0.223	-0.132	-0.285	-0.176	0.402	0.307	0.377	0.269
CAD	-0.343	-0.282	-0.457	-0.373	0.657	0.331	0.637	0.420
CHF	-0.087	-0.047	-0.148	-0.056	0.432	-0.289	0.341	-0.177
DEM	-0.041	-0.003	-0.148	-0.070	0.543	0.164	0.428	0.189
FRF	-0.038	-0.050	-0.106	-0.134	0.456	0.361	0.340	0.357
GBP	-0.510	-0.385	-0.539	-0.513	0.606	0.824	0.600	0.799
ITL	-0.045	-0.128	-0.082	-0.213	0.354	0.388	0.255	0.394
JPY	-0.668	-0.298	-0.760	-0.436	0.830	0.330	0.846	0.444
NOK	-0.133	-0.084	-0.257	-0.195	0.563	0.154	0.499	0.245
NZD	-0.829	-0.560	-0.855	-0.595	0.906	0.415	0.895	0.526
SEK	0.233	0.182	0.170	0.083	0.170	0.187	0.047	0.140
All	-0.108	-0.095	-0.204	-0.213	0.436	0.377	0.413	0.421

Notes: This table presents the correlation coefficient of the Fama slope coefficient and the volatility of  $i_t - i_t^*$ , for each currency, including each currency and the combined all-country sample. The estimates of the slope coefficient ( $b$ ) using 10-year or 5-year rolling Fama regression,  $s_{t+1} - s_t = \alpha + \beta(i_t - i_t^*) + u_t$ , based on the longest sample for each currency, as in Figs. 1–3. The coefficients are bias-corrected following Amihud and Hurvich (2004). The volatility of  $i_t - i_t^*$  is measured by its variance  $var(i_t - i_t^*)$ , its standard deviation  $s. d. (i_t - i_t^*)$ , the inverse of its variance  $\frac{1}{var(i_t - i_t^*)}$  and the inverse of its standard deviation  $\frac{1}{s.d.(i_t - i_t^*)}$ . For the all-country sample, both the coefficient estimates and volatility measures are demeaned for each country before calculating the correlation.



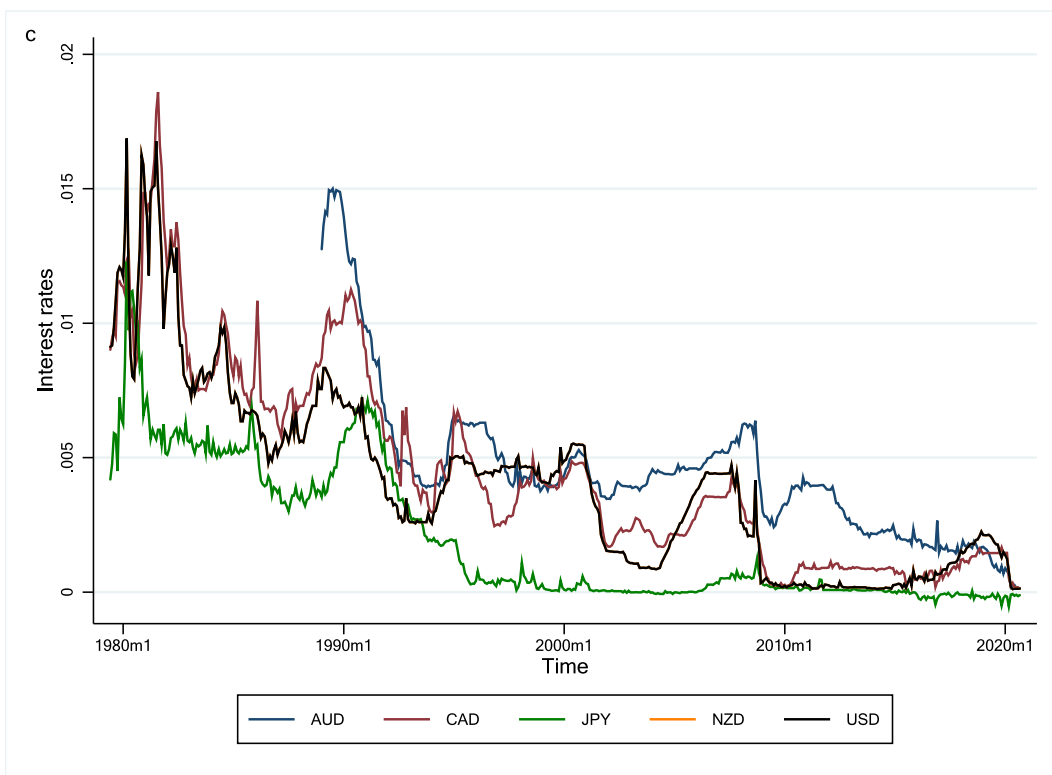


Fig. 7. Interest rates of the U.S. and other countries.

Notes: a: This figure presents the interest rate of the U.S. (black line), Germany (navy line), France (maroon line), and Italy (green line), from 1979:06–2020:09. After the time of origination of the euro in January 1999, all three foreign interest rates are the same, and are the euro interest rates.

b: This figure presents the interest rate of the U.S. (black line), Switzerland (navy line), the U.K. (maroon line), Norway (green line), and Sweden (orange line), from 1979:06–2020:09. Sample length are different for different countries.

c: This figure presents the interest rate of the U.S. (black line), Australia (navy line), Canada (maroon line), Japan (green line), and New Zealand (orange line), from 1979:06–2020:09. Sample length are different for different countries.

Table 6  
Correlation between Fama coefficient and VIX.

	AUD	CAD	CHF	DEM	FRF	GBP	ITL	JPY	NOK	NZD	SEK
10-Year	-0.13	-0.599	-0.33	-0.27	-0.31	-0.13	-0.29	-0.33	-0.27	0.121	-0.365
5-Year	-0.095	-0.44	0.01	-0.07	-0.13	0.04	-0.21	-0.01	0.003	0.065	-0.008

Notes: This table presents the correlation coefficient of the Fama slope coefficient and VIX, for each currency, including AUD, CAD, CHF, DEM, FRF, GBP, ITL, JPY, NOK, NZD, SEK. The estimates of the slope coefficient (b) using 10-year or 5-year rolling Fama regression,  $s_{t+1} - s_t = \alpha + \beta(i_t - i_t^*) + u_t$ , based on the longest sample for each currency, as in Figs. 1–3. The coefficients are bias-corrected following Amihud and Hurvich (2004).

The full-sample results are reported in Table 8. Australia and New Zealand are absent from this table because those countries do not report monthly inflation rates.

We can see that all the estimated slope coefficients are negative, with the exception of France (with a slope of 0.001). Additionally, most are significantly negative at the 5% level (in a two-sided test.) The results are not overwhelmingly strong, allowing for the possibility that these findings are simply noise, but we note that the fact that all slope coefficients save one are negative is evidence of a strong pattern.

Table 9 reports the outcome of estimating Eq. (6) using the longest common sample, 1987:01–2020:09. In this common sample, the findings are quite strong. All the slope coefficient estimates are negative, and all but two (Japan and Norway) are significantly negative at the 5% level. The stronger findings over this time period, which excludes the 1980s, relative to the full sample

**Table 7**  
Response of interest rates to y.o.y. inflation.

Model	Time	Obs.	$\hat{\mu}_1$	95% CI	p-Value
Panel	1987/01–2020/07	4029	0.059	(−0.010, 0.128)	0.044
	1987/01–2006/12	2399	0.099	(−0.009, 0.207)	0.036
	2007/01–2020/07	1630	0.002	(−0.010, 0.014)	0.362
Without ITL and FRF	1987/01–2020/07	3223	0.058	(−0.020, 0.136)	0.074
	1987/01–2006/12	1919	0.104	(−0.027, 0.235)	0.062
	2007/01–2020/07	1304	0.002	(−0.012, 0.016)	0.371
Panel relative to USD	1987/01–2020/07	3626	0.063	(−0.010, 0.136)	0.043
	1987/01–2006/12	2159	0.104	(−0.006, 0.214)	0.032
	2007/01–2020/07	1467	0.006	(−0.006, 0.018)	0.142
Relative, without ITL and FRF	1987/01–2020/07	2820	0.061	(−0.023, 0.145)	0.077
	1987/01–2006/12	1679	0.111	(−0.026, 0.248)	0.058
	2007/01–2020/07	1141	0.008	(−0.006, 0.022)	0.117

Notes: This table reports the slope coefficient estimates ( $\hat{\mu}_1$ ) from fixed-effect Panel regression,  $i_{c,t} = \mu_{c,0} + \mu_1 \pi_{c,t} + \mu_2 i_{c,t-1} + \tau_t + u_{c,t}$ , with Driscoll-Kraay standard errors, based on three samples, 1987/01–2020/07, 1987/01–2006/12 and 2007/01–2020/07. This currency-month panel includes CAD, CHF, DEM, FRF, GBP, ITL, JPY, NOK, SEK and USD in the Panel row. In the without ITL and FRF row, we exclude ITL and FRF. Panel relative to USD and Relative, without ITL and FRF reports estimates from regression  $i_{c,t} - i_t = \mu_{c,0} + \mu_1 (\pi_{c,t} - \pi_t) + \mu_2 (i_{c,t-1} - i_{t-1}) + \tau_t + u_{c,t}$ , based on panel sample including CAD, CHF, DEM, FRF, GBP, ITL, JPY, NOK, SEK and CAD, CHF, DEM, GBP, JPY, NOK, SEK, respectively.  $i_t$  is the interest rate on a riskless one-period deposit or security in the home country (U.S.) and  $i_{c,t}$  is the analogous interest rate in the foreign country  $c$ .  $\pi_t$  is the y.o.y. inflation rate in the U.S. and  $\pi_c$  is the y.o.y. inflation rate in the foreign country  $c$ . The CI column reports the 95% confidence intervals of the slope coefficient. The  $p$ -value column reports the  $p$ -value of the one-sided  $t$ -test for the slope coefficient  $H_0 : \mu_1 = 0, H_1 : \mu_1 > 0$ .

**Table 8**  
Slope coefficient from Eq. (6), full sample.

Currency	Time	Obs.	$d$	95% CI	p-VALUE
CAD	1979/06–2020/09	481	−0.137	(−0.263, −0.0111)	0.03
CHF	1979/06–2020/09	481	−0.357	(−0.574, −0.141)	< 0.01
DEM	1979/06–2020/09	481	−0.326	(−0.521, −0.132)	< 0.01
FRF	1979/06–2020/09	481	0.001	(−0.169, 0.170)	0.995
GBP	1979/06–2020/09	481	−0.2	(−0.398, −0.00222)	0.05
ITL	1979/06–2020/09	481	−0.019	(−0.105, 0.0682)	0.68
JPY	1979/06–2020/09	481	−0.125	(−0.330, 0.0807)	0.24
NOK	1986/01–2020/09	414	−0.137	(−0.293, 0.0189)	0.086
SEK	1987/01–2020/09	402	−0.228	(−0.404, −0.0517)	0.01

Notes: This table reports the slope coefficient estimates ( $d$ ) from Eq. (6),  $s_{t+1} - s_t - (i_t - i_t^*) = \alpha + \delta(\pi_t - \pi_t^*) + u_t$ , based on the full sample (the longest covers 1979:06–2020:09) for each currency, including CAD, CHF, DEM, FRF, GBP, ITL, JPY, NOK, SEK. AUD and NZD are excluded due to their lack of inflation data. Definitions of  $s_t, i_t, i_t^*$  are the same as in Table 1. Sample dates vary across currencies, especially for NOK and SEK, as shown in the Time column, due to the data availability of interest rates. The CI column reports the 95% confidence intervals of the slope coefficient, with the bias-corrected standard errors derived by Amihud and Hurvich (2004). The  $p$ -value column reports the  $p$ -value of the two-sided  $t$ -test for the slope coefficient  $H_0 : \delta = 0$ .

**Table 9**  
Slope coefficient from Eq. (6), 1987:01–2020:09.

Currency	Time	Obs.	$d$	95% CI	p-Value
CAD	1987/01–2020/09	402	−0.319	(−0.570, −0.0682)	0.01
CHF	1987/01–2020/09	402	−0.356	(−0.709, −0.00382)	0.05
DEM	1987/01–2020/09	402	−0.358	(−0.626, −0.0905)	0.01
FRF	1987/01–2020/09	402	−0.805	(−1.211, −0.400)	< 0.01
GBP	1987/01–2020/09	402	−0.285	(−0.521, −0.0494)	0.02
ITL	1987/01–2020/09	402	−0.383	(−0.614, −0.153)	< 0.01
JPY	1987/01–2020/09	402	−0.016	(−0.233, 0.202)	0.89
NOK	1987/01–2020/09	402	−0.116	(−0.280, 0.0468)	0.16
SEK	1987/01–2020/09	402	−0.228	(−0.404, −0.0517)	0.01
Panel	1987/01–2020/09	2821	−0.091	(−0.152, −0.030)	< 0.01

Notes: This table reports the slope coefficient estimates ( $d$ ) from Eq. (6),  $s_{t+1} - s_t - (i_t - i_t^*) = \alpha + \delta(\pi_t - \pi_t^*) + u_t$ , based on the longest common sample (1987:01–2020:09, except for SEK) of the currencies. See Table 8 for complete notes. The last row reports the slope coefficient estimates ( $\delta$ ) from fixed-effect Panel regression,  $s_{c,t+1} - s_{c,t} - (i_{c,t} - i_{c,t}^*) = \alpha_c + \delta(\pi_{c,t} - \pi_{c,t}^*) + \tau_t + u_{c,t}$ , with Driscoll-Kraay standard errors. This currency-month panel includes CAD, CHF, DEM, GBP, JPY, NOK, and SEK.

**Table 10**

Slope coefficient from Eq. (6), 1987:01–2006:12.

Currency	Time	Obs.	$d$	95% CI	$p$ -Value
CAD	1987/01–2006/12	239	−0.291	(−0.484, −0.0982)	< 0.01
CHF	1987/01–2006/12	239	−0.308	(−0.703, 0.0858)	0.13
DEM	1987/01–2006/12	239	−0.273	(−0.575, 0.0287)	0.08
FRF	1987/01–2006/12	239	−0.897	(−1.346, −0.448)	< 0.01
GBP	1987/01–2006/12	239	−0.233	(−0.511, 0.0453)	0.10
ITL	1987/01–2006/12	239	−0.327	(−0.628, −0.0257)	0.03
JPY	1987/01–2006/12	239	−0.213	(−0.601, 0.174)	0.28
NOK	1987/01–2006/12	239	−0.117	(−0.282, 0.0489)	0.17
SEK	1987/01–2006/12	239	−0.183	(−0.363, −0.00161)	0.05
Panel	1987/01–2006/12	1680	−0.112	(−0.181, −0.043)	<0.01

Notes: This table reports the slope coefficient estimates ( $d$ ) from Eq. (6),  $s_{t+1} - s_t - (i_t - i_t^*) = \alpha + \delta(\pi_t - \pi_t^*) + u_t$ , based on the pre-Global Financial Crisis (GFC) common-sample period (1987:01–2006:12) for each currency and fixed-effect panel regression. See Tables 8 & 9 for complete notes.

**Table 11**

Slope coefficient from Eq. (6), 2007:01–2020:09.

Currency	Time	Obs.	$d$	95% CI	$p$ -Value
CAD	2007/01–2020/09	162	−0.471	(−1.159, 0.217)	0.18
CHF	2007/01–2020/09	162	−0.541	(−1.346, 0.264)	0.19
DEM	2007/01–2020/09	162	−1.08	(−1.752, −0.409)	< 0.01
FRF	2007/01–2020/09	162	−0.902	(−1.676, −0.127)	0.02
GBP	2007/01–2020/09	162	−0.549	(−1.040, −0.0583)	0.03
ITL	2007/01–2020/09	162	−0.592	(−1.135, −0.0492)	0.03
JPY	2007/01–2020/09	162	0.187	(−0.0955, 0.470)	0.20
NOK	2007/01–2020/09	162	−0.151	(−0.487, 0.186)	0.38
SEK	2007/01–2020/09	162	−0.683	(−1.345, −0.0211)	0.04
Panel	2007/01–2020/09	1141	−0.031	(−0.203, 0.141)	0.728

Notes: This table reports the slope coefficient estimates ( $d$ ) from Eq. (6),  $s_{t+1} - s_t - (i_t - i_t^*) = \alpha + \delta(\pi_t - \pi_t^*) + u_t$ , based on the 2007:01–2020:09 period for each currency and fixed-effect panel regression. See Tables 8 & 9 for complete notes.

in reported in Table 8 perhaps reflects the stronger commitment of central banks to target inflation post-1987:01. The  $p$ -value for the panel fixed-effects regression is 0.003, indicating strong support for the predictive power of relative y.o.y. inflation.

Importantly, when we look at subsamples of the 1987:01–2020:09 period, the story is not much changed. The number of significant coefficient estimates is reduced because the sample is shorter, but in both the 1987:01–2006:12 period and the 2007:1–2020:09 period, the estimated slope coefficients are all negative (with the sole exception of Japan in the latter period, where the coefficient is slightly positive), as Tables 10 and 11 demonstrate and most are still statistically significant at the 5% level. The notable point is that there is no qualitative change in these regressions between the pre-crisis era and the later time period in which unconventional monetary policies were predominant.

The panel regression also strongly rejects the null of no predictability for the earlier period, though does not for the latter period. The panel regression imposes the restriction that the slope coefficient is the same across all countries, but the estimated coefficient for Japan is positive in the univariate regressions, in contrast to all the other countries. In turn, this is the main contributor to the insignificance of the negative coefficient estimated for the panel.

Figs. 8–10 present the slope coefficient estimates from rolling regressions of Eq. (6) for Germany, the U.K. and Japan.<sup>16</sup> For almost all countries, the coefficient estimates tend to be consistently negative for almost the entire sample after the mid-1980s, with the major exception being Japan which shows a positive coefficient for the most recent time periods. In these figures, the dates along the  $x$ -axis are the starting dates for each 10-year regression window. The areas shaded blue are for a slope coefficient that is significantly negative at the 5% level. Pink areas represent sample periods for which the slope coefficient is negative but not significant at the 5% level. Green areas are for times of positive coefficient estimates – none of which are significant at the 5% level for any time period.

The overwhelming impression from these graphs is that the slope coefficient estimates of Eq. (6) are negative. There are very few periods of positive coefficients. For a few currencies, the estimated coefficient is positive for a short window early in the sample (that is, 10-year estimation windows beginning in the early 1980s.) That window is longer for the Norwegian krone. The latter part of the sample for the Japanese yen also produces some positive coefficient estimates. But overall, the picture is clear that higher  $\pi_t - \pi_t^*$  is associated with lower ex post returns on the foreign currency,  $s_{t+1} - s_t + i_t^* - i_t$ . Even during periods in

<sup>16</sup> The figures for all countries are in the Supplemental Appendix.



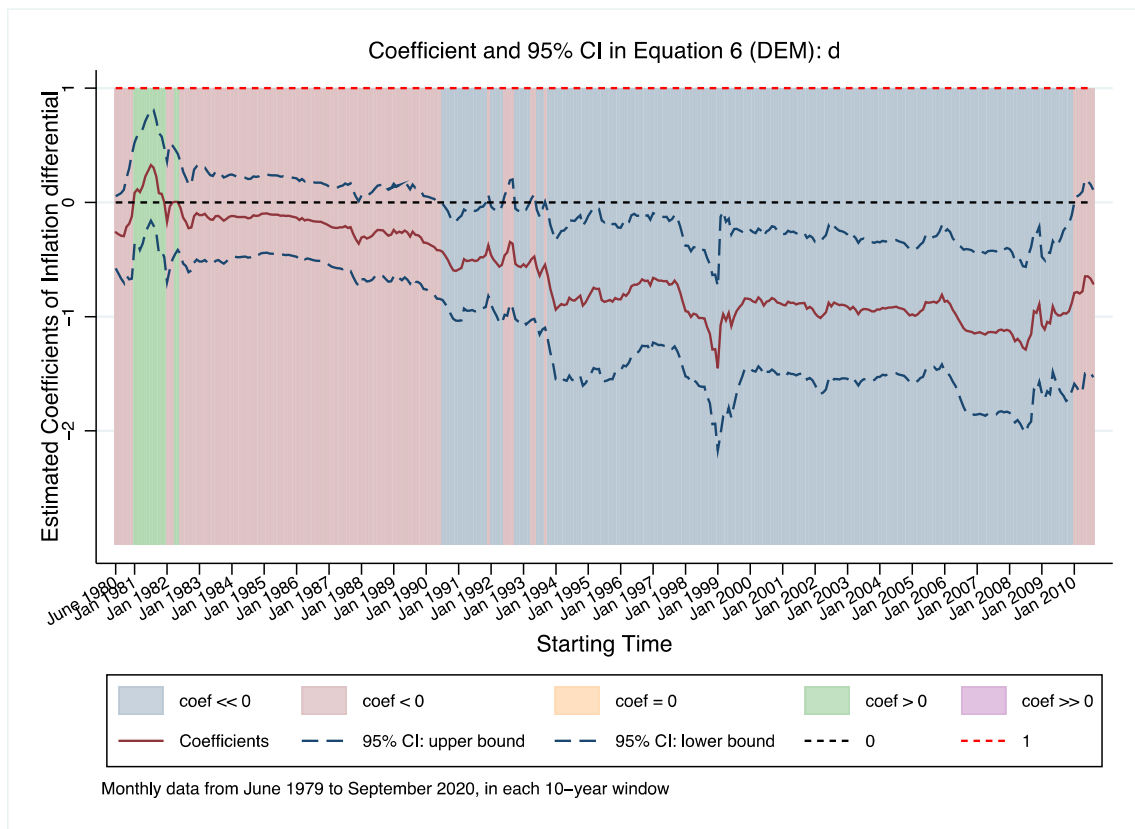


Fig. 8. 10-year rolling regressions, Eq. (6): DEM.

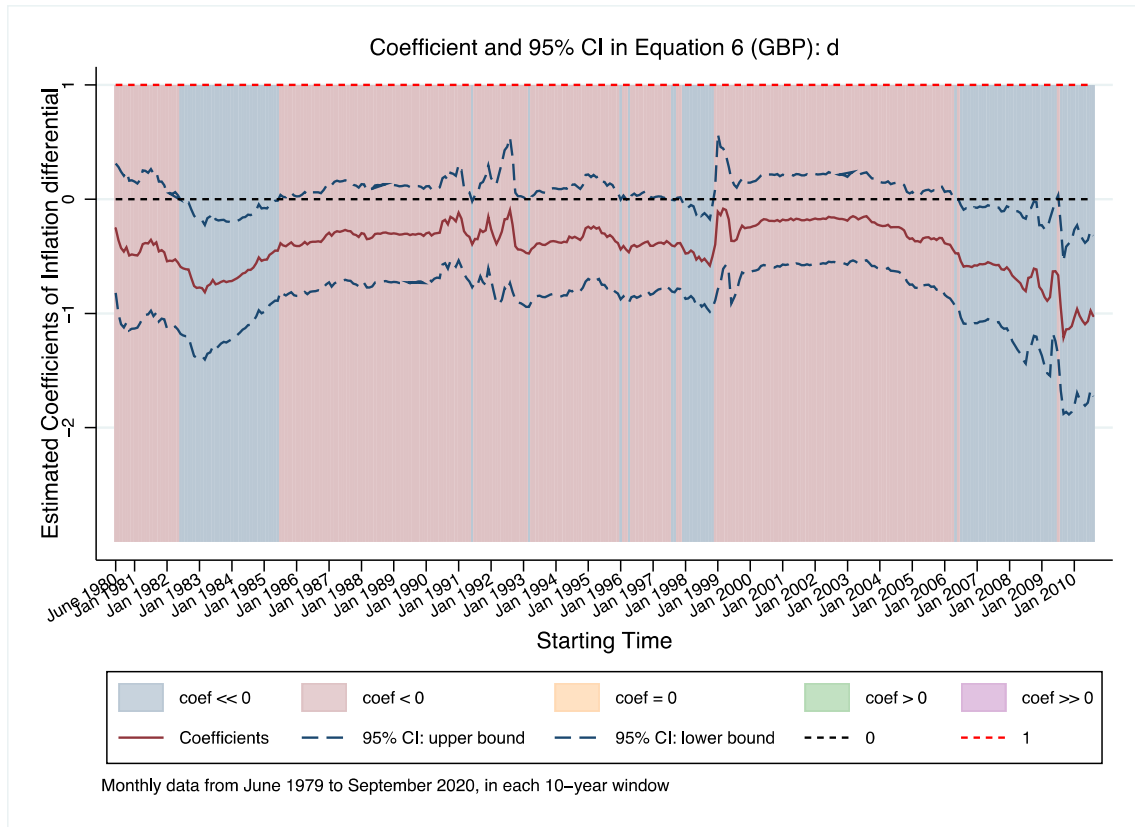
Notes: This figure presents the estimates of the slope coefficient ( $d$ ) and 95% confident intervals using 10-year rolling regression of Eq. (6),  $s_{t+1} - s_t - (i_t - i_t^*) = \alpha + \delta(\pi_t - \pi_t^*) + u_t$ , using German mark (DEM) data during 1979:06–2020:09.  $s_t, i_t, i_t^*$  are the same as in Fig. 1.  $\pi_t$  is the inflation rate in the U.S. and  $\pi_t^*$  is the inflation rate in Germany. Each component of the graph is the same as in Fig. 1, except that we now compare the estimated coefficients with zero instead of one. See Fig. 1 for complete notes.

which the Fama puzzle was seen to hold, prior to 2007, the empirical regularity from this regression is much more consistent. That is, year-on-year inflation is a better predictor of excess returns than the nominal interest rate differential, even prior to 2007. However, for many of the countries, the slope coefficient is not significantly negative at the 5% level over many of the sub-samples, so it is only tentatively that we conclude that an increase in inflation abroad tends to predict higher excess returns on foreign deposits.

The Supplemental Appendix reports estimates from regressions that include both interest rate differentials and year-on-year inflation rate differences:

$$s_{t+1} - s_t + i_t^* - i_t = \alpha + \gamma_1(i_t - i_t^*) + \gamma_2(\pi_t - \pi_t^*) + u_{t+1} \tag{9}$$

The econometrics literature does not provide analytical corrections for bias in the coefficient estimates and standard errors when there is more than one regressor, as is the case here. However, based on the OLS estimates and Newey-West standard errors, we can draw the following inferences. First, the evidence for a negative relationship between  $i_t - i_t^*$  and  $s_{t+1} - s_t + i_t^* - i_t$  is much weaker when the inflation differential is included in the regression, especially in samples that begin in the mid-1980s. On the other hand, the evidence for a negative relationship between  $\pi_t - \pi_t^*$  and  $s_{t+1} - s_t + i_t^* - i_t$  remains strong even when  $i_t - i_t^*$  is included in the regression, except for the three countries of Switzerland, Japan and Norway. This suggests the following interpretation: The usual finding that the slope coefficient is less than one in the Fama regression, (2), really is picking up the reaction to monetary policy. The inflation rate over the previous 12 months,  $\pi_t - \pi_t^*$  is a stronger measure of monetary policy stance than the interest rate differential. When controlling for year-on-year inflation, there is little evidence of additional explanatory power from  $i_t - i_t^*$  except perhaps in the 1979–1988 period for currencies for which we have data. When inflation targeting became more central to monetary policy, both before and after 2007 when interest rates were very low and unconventional monetary policy became more common,  $i_t - i_t^*$  has little explanatory power for excess returns once relative inflation is controlled for.



**Fig. 9.** 10-year rolling regressions, Eq. (6): GBP.  
 Notes: This figure presents the estimates of the slope coefficient ( $d$ ) and 95% confident intervals using 10-year rolling regression of Eq. (6),  $s_{t+1} - s_t - (i_t - i_t^*) = \alpha + \delta(\pi_t - \pi_t^*) + u_t$ , using British pound (GBP) data during 1979:06–2020:09. See Fig. 8 for complete notes.

It is tempting to believe that the significance of the year-on-year inflation differential, along with the insignificance of the interest rate differential in the zero-lower-bound time period, suggests that it is the real interest rate differential that predicts excess returns:

$$s_{t+1} - s_t + i_t^* - i_t = \alpha + \psi(i_t - i_t^* - E_t(\pi_{t+1} - \pi_{t+1}^*)) + u_{t+1} \tag{10}$$

There is immediate reason to be skeptical of this explanation. The puzzle that arises from the estimation of the Fama equation, (3), is that  $\gamma < 0$ . If it is, in fact, the real interest rate differential that matters, we might expect to find  $\psi < 0$  in Eq. (10). But that would mean, holding the nominal interest rate differential constant (as it nearly is when both countries are near the ZLB), the effect of the expected inflation differential on the expected return is positive, since from Eq. (10),  $E_t(\pi_{t+1} - \pi_{t+1}^*)$  enters with a negative sign in predicting excess returns. However, our estimates of the effect of year-on-year inflation on the excess return,  $\delta$ , from Eq. (6), are negative. Unless year-on-year inflation predicts the next month's inflation with a negative sign, the real interest rate specification in (10) does not account for the significance of year-on-year inflation in forecasting excess returns, at least not in a way that is consistent with the Fama puzzle.

To test specification (10), we need a measure of expected inflation in each country. We implemented this by regressing monthly inflation at time  $t + 1$  on the nominal interest rate and y.o.y. inflation at time  $t$ .<sup>17</sup> Using these measures of expected inflation, we construct the real interest rate differential, and estimate Eq. (10). Our findings are very similar to the ones with the nominal interest rate differential as the independent variable, as in the Fama regression. Over the longest sample, starting in 1980, the estimate of  $\psi$  is negative and significant for most countries. In the longest common sample, 1987:01–2020:09, we find  $\psi$  is negative and significant for only Switzerland and Japan. In the pre-ZLB sample, 1987:01–2006:12,  $\psi$  is negative and significant for those two countries and Canada. In the 2007:01–2020:09 period,  $\psi$  is estimated to be positive for all countries, though

<sup>17</sup> In fact, we estimated expected inflation of each country relative to the U.S., using as predictors the country/U.S. relative interest rates and relative year-on-year inflation rates. We also augment this set of predictors with measures of relative output growth and the relative output gap.

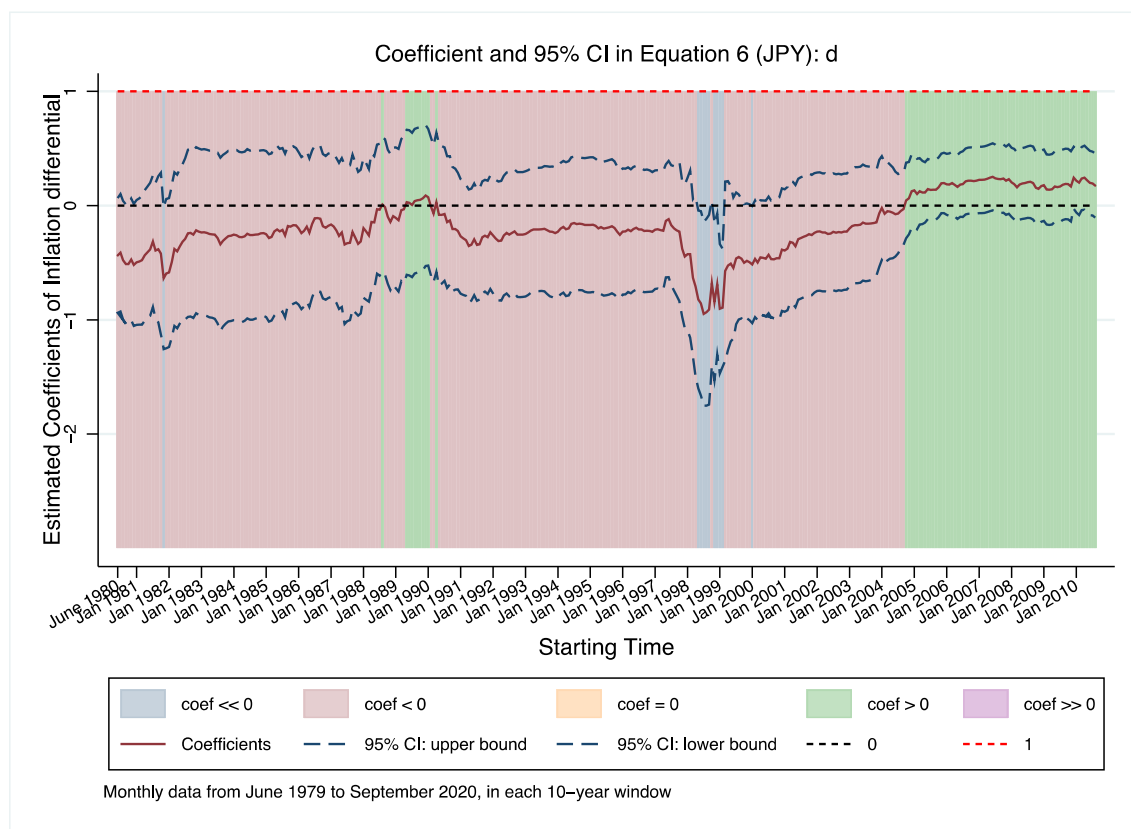


Fig. 10. 10-year rolling regressions, Eq. (6): JPY.

Notes: This figure presents the estimates of the slope coefficient ( $d$ ) and 95% confident intervals using 10-year rolling regression of Eq. (6),  $s_{t+1} - s_t - (i_t - i_t^*) = \alpha + \delta(\pi_t - \pi_t^*) + u_t$ , using Japanese yen (JPY) data during 1979:06–2020:09. See Fig. 8 for complete notes.

insignificant for all but Italy. Hence, we find the real interest-rate differential does not consistently predict excess returns, especially during the ZLB period. These results are reported in detail in the Supplemental Appendix.<sup>18</sup>

#### 4. Medium run

In this section, we consider estimates of Eqs. (4) and (7). The dependent variable in these regressions can be interpreted as the return on an investment strategy of buying foreign exchange, investing in one-month foreign-currency deposits, rolling those deposits over for 12 months then converting the gross investment back into dollars compared to the return from rolling over one-month dollar deposits for 12 months. Our purpose for looking at the medium horizon is to assess whether these predicted excess returns are persistent, which may shed further light on why returns might be predictable.

These regressions are also subject to the bias in parameter estimates and standard errors originally noted by Stambaugh (1999) but are also subject to the problems attendant with “long-horizon” returns regressions. We make use of the bias corrections in the recent study of Boudoukh et al. (2020), using Newey-West standard errors.

We consider the same four estimation periods as we did for our short-run return regressions: the entire sample we have for each country; our longest common sample, 1987:01–2020:09; a pre-crisis common sample, 1987:01–2006:12; and a common sample that includes the low interest rate era, 2007:01–2020:09.<sup>19</sup>

These results are reported in Tables 12–15. For the full sample in Table 12, all the estimated coefficients are negative, except that for the Italian lira. As the short-term interest differential for the U.S. relative to the foreign country increases ( $i_t - i_t^*$ ), the excess return on the foreign investment falls. These coefficients are significantly less than zero at the 5% level for ten of the currencies, and at the 10% level for one more. Only for the Italian lira and Swedish krona do we fail to reject the null of no predictability at the medium horizon. The findings for this relationship are much stronger than what we found over the full sample for the short-run returns in the Fama regression.

<sup>18</sup> The dependent variable in (10) is the “nominal” ex post excess return. The “real” ex post excess return is identical, except for the addition of the forecast errors of inflation. Those forecast errors should not affect the estimates of the slope coefficient in (10) because they are uncorrelated with the regressor in (10), as the nominal interest rates are included in the information set used to make the forecasts of inflation.

<sup>19</sup> Again, we note that New Zealand’s sample does not begin until 1999.

**Table 12**  
Slope coefficient from medium-run regression (4), full sample.

Currency	Time	Obs.	$\hat{\gamma}_M$	CI	p-Value
AUD	1989/01–2020/09	369	−19.787	(−31.870, −7.704)	0.001
CAD	1979/06–2020/09	484	−8.851	(−17.360, −0.342)	0.042
CHF	1979/06–2020/09	484	−19.938	(−27.353, −12.523)	0.000
DEM	1979/06–2020/09	484	−17.274	(−25.271, −9.278)	0.000
FRF	1979/06–2020/09	484	−5.629	(−11.282, 0.025)	0.052
GBP	1979/06–2020/09	484	−16.529	(−24.898, −8.160)	0.000
ITL	1979/06–2020/09	484	1.267	(−3.285, 5.819)	0.586
JPY	1979/06–2020/09	484	−21.122	(−29.625, −12.619)	0.000
NOK	1986/01–2020/09	405	−11.178	(−18.397, −3.959)	0.003
NZD	1997/04–2020/09	270	−28.036	(−48.468, −7.604)	0.008
SEK	1987/01–2020/09	393	−3.689	(−10.915, 3.536)	0.318

Notes: This table reports the slope coefficient estimates ( $\hat{\gamma}_M$ ) from Eq. (4),  $\sum_{j=1}^{12} s_{t+j+1} - s_{t+j} + i_{t+j}^* - i_{t+j} = \alpha_M + \gamma_M(i_t - i_t^*) + u_t^M$  based on the full sample (the longest covers 1979:06–2020:09) for each currency. Definitions of  $s_t$ ,  $i_t$ ,  $i_t^*$  are the same as in Table 1. The CI column reports the 95% confidence intervals of the slope coefficient, with the bias-corrected standard errors derived by Boudoukh et al. (2020). The p-value column reports the p-value of the two-sided t-test for the slope coefficient  $H_0 : \gamma_M = 0$ .

**Table 13**  
Slope coefficient from medium-run regression (4), 1987:01–2020:09.

Currency	Time	Obs.	$\hat{\gamma}_M$	CI	p-Value
AUD	1989/01–2020/09	369	−19.787	(−31.870, −7.704)	0.001
CAD	1987/01–2020/09	393	−8.178	(−19.613, 3.257)	0.162
CHF	1987/01–2020/09	393	−22.346	(−33.727, −10.964)	0.000
DEM	1987/01–2020/09	393	−18.530	(−28.882, −8.178)	0.001
FRF	1987/01–2020/09	393	−14.573	(−23.786, −5.360)	0.002
GBP	1987/01–2020/09	393	−9.341	(−19.476, 0.793)	0.072
ITL	1987/01–2020/09	393	−2.414	(−9.878, 5.050)	0.527
JPY	1987/01–2020/09	393	−24.409	(−35.542, −13.277)	0.000
NOK	1987/01–2020/09	393	−8.159	(−15.568, −0.751)	0.031
NZD	1997/04–2020/09	270	−28.036	(−48.468, −7.604)	0.008
SEK	1987/01–2020/09	393	−3.689	(−10.915, 3.536)	0.318
Panel	1987/01–2020/09	2751	−8.881	(−19.651, 1.889)	0.073

Notes: This table reports the slope coefficient estimates ( $\hat{\gamma}_M$ ) from Eq. (4),  $\sum_{j=1}^{12} s_{t+j+1} - s_{t+j} + i_{t+j}^* - i_{t+j} = \alpha_M + \gamma_M(i_t - i_t^*) + u_t^M$ , based on the longest common sample (1987:01–2020:09, except for AUD, NZD, and SEK) of the currencies. See Table 12 for complete notes. The last row reports the slope coefficient estimates ( $\hat{\gamma}_M$ ) from fixed-effect Panel regression,  $\sum_{j=1}^{12} s_{c,t+j+1} - s_{c,t+j} + i_{c,t+j}^* - i_{c,t+j} = \alpha_{M,c} + \gamma_M(i_{c,t} - i_{c,t}^*) + \tau_c + u_{c,t}^M$ , with Driscoll-Kraay standard errors. This currency-month panel includes CAD, CHF, DEM, GBP, JPY, NOK, and SEK.

**Table 14**  
Slope coefficient from medium-run regression (4), 1987:01–2006:12.

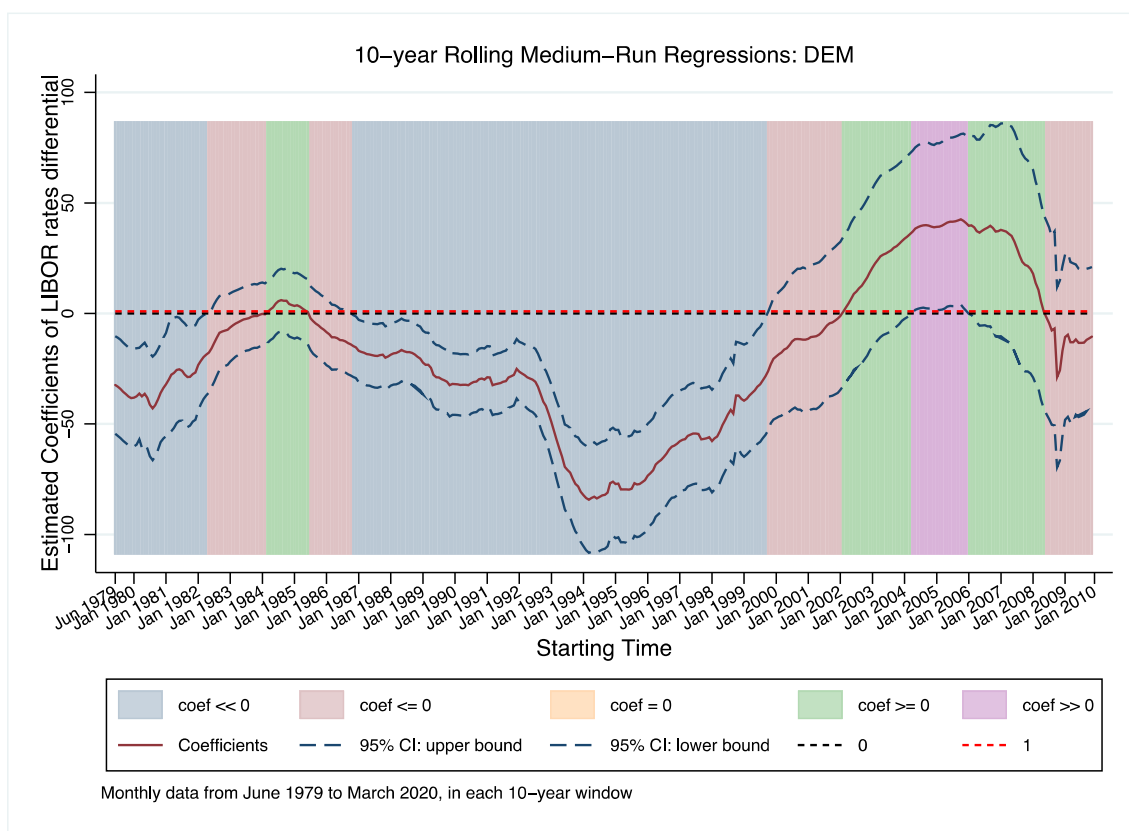
Currency	Time	Obs.	$\hat{\gamma}_M$	CI	p-value
AUD	1989/01–2006/12	216	−21.367	(−33.021, −9.714)	0.000
CAD	1987/01–2006/12	240	−10.180	(−19.230, −1.1303)	0.028
CHF	1987/01–2006/12	240	−24.631	(−37.127, −12.135)	0.000
DEM	1987/01–2006/12	240	−22.294	(−33.510, −11.077)	0.000
FRF	1987/01–2006/12	240	−16.304	(−26.317, −6.291)	0.002
GBP	1987/01–2006/12	240	−5.086	(−18.001, 7.828)	0.441
ITL	1987/01–2006/12	240	−0.169	(−8.887, 8.549)	0.970
JPY	1987/01–2006/12	240	−40.053	(−54.569, −25.537)	0.000
NOK	1987/01–2006/12	240	−6.554	(−13.874, 0.765)	0.081
NZD	1997/04–2006/12	117	−53.738	(−77.247, −30.230)	0.000
SEK	1987/01–2006/12	240	−1.965	(−9.652, 5.722)	0.617
Panel	1987/01–2006/12	1680	−8.272	(−20.183, 3.639)	0.045

Notes: This table reports the slope coefficient estimates ( $\hat{\gamma}_M$ ) from Eq. (4),  $\sum_{j=1}^{12} s_{t+j+1} - s_{t+j} + i_{t+j}^* - i_{t+j} = \alpha_M + \gamma_M(i_t - i_t^*) + u_t^M$ , based on the pre-Global Financial Crisis (GFC) common-sample period (1987:01–2006:12, except for AUD and NZD) for each currency and fixed-effect Panel regression. See Tables 12 & 13 for complete notes.

**Table 15**  
Slope coefficient from medium-run regression (4), 2007:01–2020:09.

Currency	Time	Obs.	$\hat{\gamma}_M$	CI	p-Value
AUD	2007/01–2020/09	153	-17.828	(-46.201, 10.546)	0.220
CAD	2007/01–2020/09	153	51.697	(2.709, 100.684)	0.040
CHF	2007/01–2020/09	153	-1.000	(-31.822, 29.823)	0.949
EUR	2007/01–2020/09	153	5.187	(-22.743, 33.116)	0.716
GBP	2007/01–2020/09	153	36.078	(4.929, 67.226)	0.025
JPY	2007/01–2020/09	153	29.128	(5.357, 52.898)	0.018
NOK	2007/01–2020/09	153	-1.488	(-31.520, 28.544)	0.923
NZD	2007/01–2020/09	153	0.476	(-33.039, 33.991)	0.978
SEK	2007/01–2020/09	153	-0.229	(-25.730, 25.273)	0.986
Panel	2007/01–2020/09	1071	29.155	(1.823, 56.487)	0.045

Notes: This table reports the slope coefficient estimates ( $\hat{\gamma}_M$ ) from Eq. (4),  $\sum_{j=1}^{12} s_{t+j+1} - s_{t+j} + i_{t+j}^* - i_{t+j} = \alpha_M + \gamma_M(i_t - i_t^*) + u_t^M$ , based on the 2007:01–2020:09 period for each currency and fixed-effect Panel regression. See Tables 12 & 13 for complete notes.



**Fig. 11.** 10-year rolling medium-run regressions: DEM.  
Notes: This figure presents the estimates of the slope coefficient ( $\hat{\gamma}_M$ ) and 95% confident intervals using 10-year rolling of Eq. (4),  $\sum_{j=1}^{12} s_{t+j+1} - s_{t+j} + i_{t+j}^* - i_{t+j} = \alpha_M + \gamma_M(i_t - i_t^*) + u_t^M$ , using German mark (DEM) data during 1979:06–2020:09. Definitions of  $s_t$ ,  $i_t$ ,  $i_t^*$  are the same as in Fig. 1. Each component of the graph is the same as in Fig. 8, except that the coefficients and 95% confident intervals are bias-corrected following Boudoukh et al. (2020). See Fig. 8 for complete notes.

The findings are similar when we use our longest common sample, as reported in Table 13. Here all the point estimates of the slope coefficient are negative, and seven are significantly less than zero at the 5% confidence level. The fixed-effects panel regression also finds a negative slope coefficient that is significant at the 10% level. These findings are echoed in Table 14 for the pre-GFC sample of 1989:01–2006:12. All the slope coefficient estimates are negative, and seven are significantly (at the 5% level) negative. However, the panel estimate, while negative, is not statistically significant.

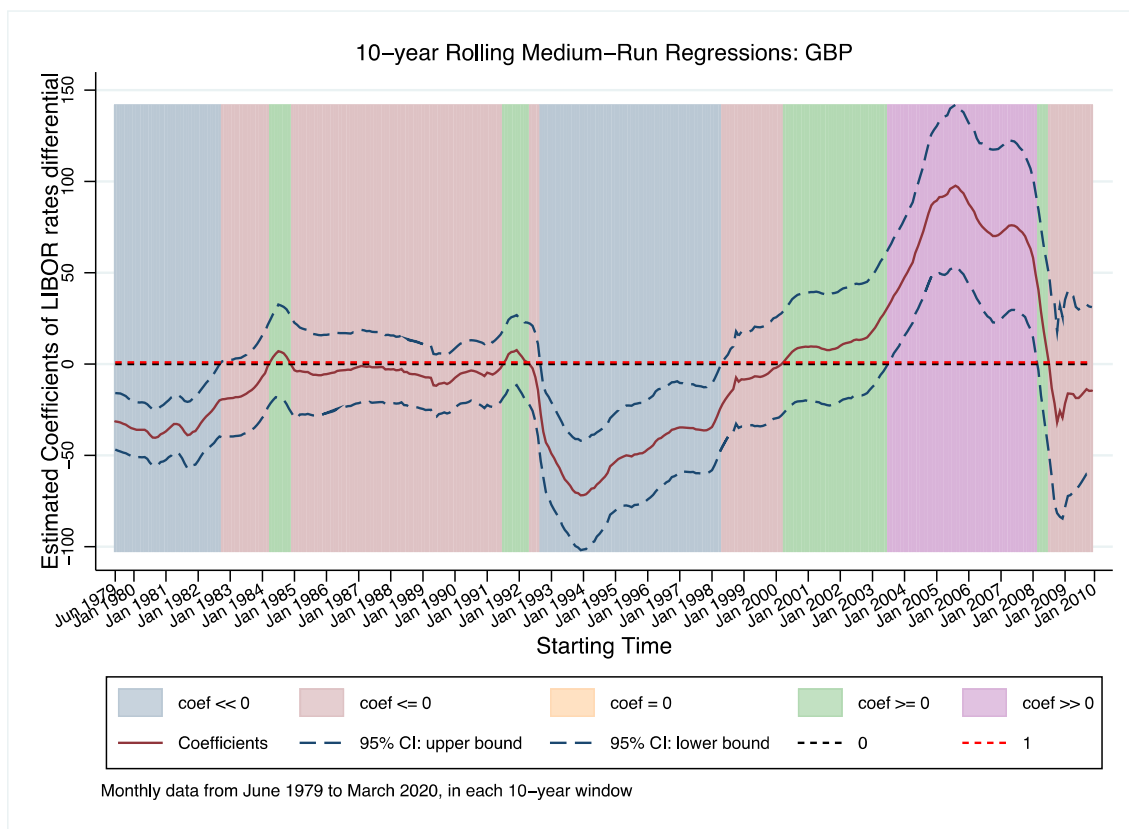


Fig. 12. 10-year rolling medium-run regressions: GBP.

Notes: This figure presents the estimates of the slope coefficient ( $\hat{\gamma}_M$ ) and 95% confident intervals using 10-year rolling of Eq. (4),  $\sum_{j=1}^{12} s_{t+j+1} - s_{t+j} + i_{t+j}^* - i_{t+j} = \alpha_M + \gamma_M (i_t - i_t^*) + u_t^M$ , using British pound (GBP) data during 1979:06–2020:09. See Fig. 11 for complete notes.

The findings are different in the post-2007:01 sample reported in Table 15. There, five of the nine estimated coefficients from regression (4) are positive.<sup>20</sup> Three are significantly positive at the 5% level, and none are significantly negative. The slope coefficient in the fixed-effects panel is positive and significant.

As with the short-run Fama regressions, before 2007, there is strong evidence of a negative slope across the currencies, but after 2007 the evidence is mixed. There is no currency that shows a significantly negative coefficient in either the short-run or medium-run regressions in the post-2007 period.

We also perform rolling regressions, and illustrate the findings in Figs. 11–13 for Germany, the U.K., and Japan, which highlight this sudden shift.<sup>21</sup> The horizontal axis gives the starting date for each 10-year estimation window. The blue-shaded areas represent a slope coefficient significantly negative at the 5% level. Pink areas are for sample periods for which the slope coefficient is negative but insignificant at the 5% level. The area is shaded green for times of positive coefficient estimates that are not significant, and purple when the coefficient is positive and significant.

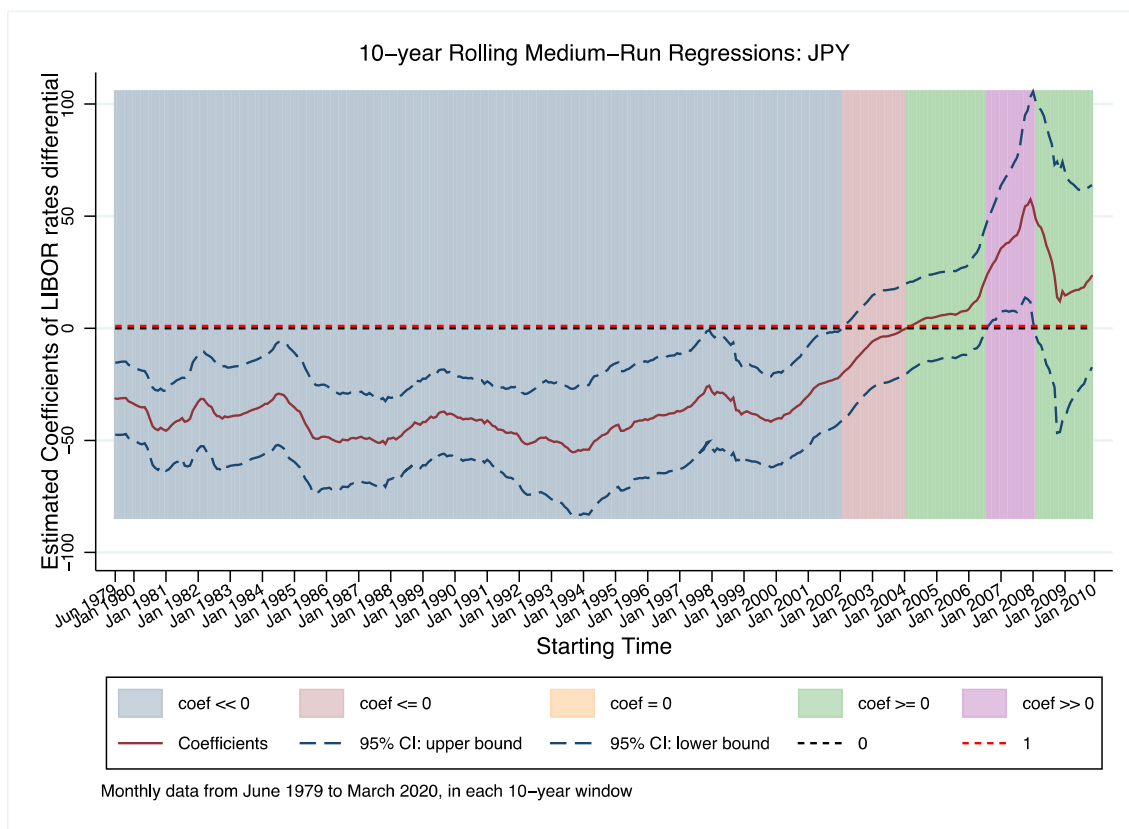
The graphs all show a large swing in the coefficient estimates in the later part of the sample compared to the earlier part. For almost all the currencies, that shift begins with 10-year samples that start in the early 2000s, which coincides with samples in which near-zero interest rates become predominant. In samples that are primarily drawn from the low-interest rate era, the slope coefficients are positive, and usually significantly so. We can conclude that we see the same parameter instability as in the Fama regressions, but with more evidence of a shift in regime from negative to positive coefficients. That is, the finding of the shift in sign of the slope coefficient is more likely to be a genuine shift in the relationship rather than just sampling error.

We turn next to estimates of Eq. (7), in which excess medium-term returns is again the dependent variable, and year-on-year inflation is the regressor. We report the slope parameter estimates and standard errors in Tables 16–19 for the different time periods.

The most striking takeaway from these tables is that all the slope coefficient estimates for all the currencies and time periods are negative, apart from the French franc and British pound for the full sample, and Japan in the last sample, which are marginally

<sup>20</sup> There are only nine currencies in this period because the euro replaced the mark, French franc and lira.

<sup>21</sup> The figures for all countries are available in the Supplemental Appendix.



**Fig. 13.** 10-year rolling medium-run regressions: JPY.

Notes: This figure presents the estimates of the slope coefficient ( $\hat{\gamma}_M$ ) and 95% confident intervals using 10-year rolling of Eq. (4),  $\sum_{j=1}^{12} s_{t+j+1} - s_{t+j} + i_{t+j}^* - i_{t+j} = \alpha_M + \gamma_M(i_t - i_t^*) + u_t^M$ , using Japanese yen (JPY) data during 1979:06–2020:09. See Fig. 11 for complete notes.

positive. The French franc and British pound full sample estimates come from samples that begin in 1979. As we have seen with the one-month excess returns, the empirical evidence that increases in  $\pi_t - \pi_t^*$  predict declines in excess returns on foreign deposits is weaker when the 1980s are included in the sample. If our working hypothesis is correct, this can be explained by the fact that inflation targeting by central banks was not as strongly followed in the 1980s as in later periods. That view is perhaps also bolstered by the case of Japan in recent years that has found its efforts to boost inflation unsuccessful and turned to various non-monetary policies such as fiscal expansion.

In Table 16, which is for the full sample period for each currency, the estimated value of the slope parameter is strongly significantly negative for six of the nine currencies (and insignificant but negative for the lira.) In Table 17, which presents evidence for the 1989:01–2017:01 period, all parameter estimates are negative, and significantly so at the 5% level for six. For the period 1989:01–2006:12, the findings are similar. For the later period when interest rates were near zero, the coefficient estimates are all negative, though fewer are statistically significant.

Figs. 14–16 show the slope estimates from rolling regressions with 10-year windows for Germany, the U.K., and Japan. The figures show that the estimated slope coefficient is negative almost all of the time for all of the currencies, with very few exceptions. For some currencies, windows that start in the early 1980s yield positive parameter values, and the Japanese yen shows a period of positive (but insignificant) parameters at the end of the sample.

Figs. 17–19 (again, for Germany, the U.K., and Japan) offer some perspective into the channel through which inflation is generating expected excess returns.<sup>22</sup> These charts plot the parameter estimates of  $\lambda_k$  from the regressions:

$$s_{t+k} - s_{t-1} = \alpha + \lambda_k(\pi_t - \pi_t^* - (\pi_{t-1} - \pi_{t-1}^*)) + \varepsilon_{t+k}, \quad k = 0, 1, 2, \dots \tag{11}$$

Note that the dependent variable is the change in the exchange rate from time  $t - 1$  to time  $t + k$  (not the change in time  $t$  to  $t + k$ .) For example, when  $k = 0$ ,  $\lambda_0$  gives us the association between the change in the exchange rate between  $t - 1$  and  $t$  and changes at time  $t$  in the inflation rate differential.

<sup>22</sup> The graphs for all countries are in the Supplemental Appendix.

**Table 16**

Slope coefficient from medium-run regression (7), full sample.

Currency	Time	Obs.	$d_M$	CI	p-Value
CAD	1979/06–2020/09	472	−0.868	(−1.713, −0.023)	0.045
CHF	1979/06–2020/09	472	−2.819	(−4.071, −1.567)	<0.001
DEM	1979/06–2020/09	472	−2.154	(−3.300, −1.008)	<0.001
FRF	1979/06–2020/09	472	0.082	(−0.876, 1.040)	0.867
GBP	1979/06–2020/09	472	0.124	(−1.076, 1.324)	0.840
ITL	1979/06–2020/09	472	−0.144	(−0.733, 0.445)	0.632
JPY	1979/06–2020/09	472	−2.493	(−3.744, −1.241)	0.000
NOK	1986/01–2020/09	405	−1.232	(−2.253, −0.210)	0.019
SEK	1987/01–2020/09	393	−2.457	(−3.524, −1.390)	<0.001

Notes: This table reports the slope coefficient estimates ( $d_M$ ) from Eq. (7),  $\sum_{j=1}^{12} s_{t+j+1} - s_{t+j} + i_{t+j}^* - i_{t+j} = \alpha_M + \delta_M(\pi_t - \pi_t^*) + u_t^M$ , based on the full sample (the longest covers 1979:06–2020:09) for each currency, including CAD, CHF, DEM, FRF, GBP, ITL, JPY, NOK, SEK. AUD and NZD are excluded due to their lack of inflation data. Definitions of  $s_t$ ,  $i_t$ ,  $i_t^*$  are the same as in Table 1.  $\pi_t$  is the inflation rate in the U.S. and  $\pi_t^*$  is the inflation rate in the foreign country. Sample dates vary across currencies, especially for NOK and SEK, as shown in the Time column, due to the data availability of interest rates. The CI column reports the 95% confidence intervals of the slope coefficient, with the bias-corrected standard errors derived by Boudoukh et al. (2020). The p-value column reports the p-value of the two-sided t-test for the slope coefficient  $H_0 : \delta_M = 0$ .

**Table 17**

Slope coefficient from medium-run regression (7), 1987:01–2020:09.

Currency	Time	Obs.	$d_M$	CI	p-Value
CAD	1987/01–2020/09	393	−2.116	(−3.505, −0.726)	0.003
CHF	1987/01–2020/09	393	−2.358	(−4.138, −0.577)	0.010
DEM	1987/01–2020/09	393	−2.036	(−3.474, −0.598)	0.006
FRF	1987/01–2020/09	393	−4.051	(−6.326, −1.777)	0.001
GBP	1987/01–2020/09	393	−0.926	(−2.294, 0.443)	0.186
ITL	1987/01–2020/09	393	−2.323	(−3.795, −0.852)	0.002
JPY	1987/01–2020/09	393	−1.386	(−2.791, 0.018)	0.054
NOK	1987/01–2020/09	393	−0.610	(−1.723, 0.504)	0.284
SEK	1987/01–2020/09	393	−2.457	(−3.524, −1.390)	0.000
Panel	1987/01–2020/09	2751	−1.020	(−1.477, −0.563)	<0.001

Notes: This table reports the slope coefficient estimates ( $d_M$ ) from Eq. (7),  $\sum_{j=1}^{12} s_{t+j+1} - s_{t+j} + i_{t+j}^* - i_{t+j} = \alpha_M + \delta_M(\pi_t - \pi_t^*) + u_t^M$ , based on the longest common sample (1987:01–2020:09) of the currencies. See Table 16 for complete notes. The last row reports the slope coefficient estimates ( $\delta_M$ ) from fixed-effect Panel regression,  $\sum_{j=1}^{12} s_{c,t+j+1} - s_{c,t+j} + i_{c,t+j}^* - i_{c,t+j} = \alpha_{M,c} + \delta_M(\pi_{c,t} - \pi_{c,t}^*) + \tau_c + u_{c,t}^M$ , with Driscoll-Kraay standard errors. This currency-month panel includes CAD, CHF, DEM, GBP, JPY, NOK, and SEK.

**Table 18**

Slope coefficient from medium-run regression (7), 1987:01–2006:12.

Currency	Time	Obs.	$d_M$	CI	p-Value
CAD	1987/01–2006/12	240	−2.111	(−3.366, −0.856)	0.001
CHF	1987/01–2006/12	240	−2.504	(−4.663, −0.345)	0.024
DEM	1987/01–2006/12	240	−2.465	(−4.110, −0.819)	0.004
FRF	1987/01–2006/12	240	−6.706	(−9.681, −3.731)	0.000
GBP	1987/01–2006/12	240	−0.361	(−2.012, 1.289)	0.668
ITL	1987/01–2006/12	240	−1.851	(−3.826, 0.125)	0.068
JPY	1987/01–2006/12	240	−4.327	(−6.578, −2.076)	0.000
NOK	1987/01–2006/12	240	−0.804	(−2.120, 0.513)	0.233
SEK	1987/01–2006/12	240	−2.239	(−3.323, −1.155)	0.000
Panel	1987/01–2006/12	1680	−1.165	(−1.688, −0.642)	<0.001

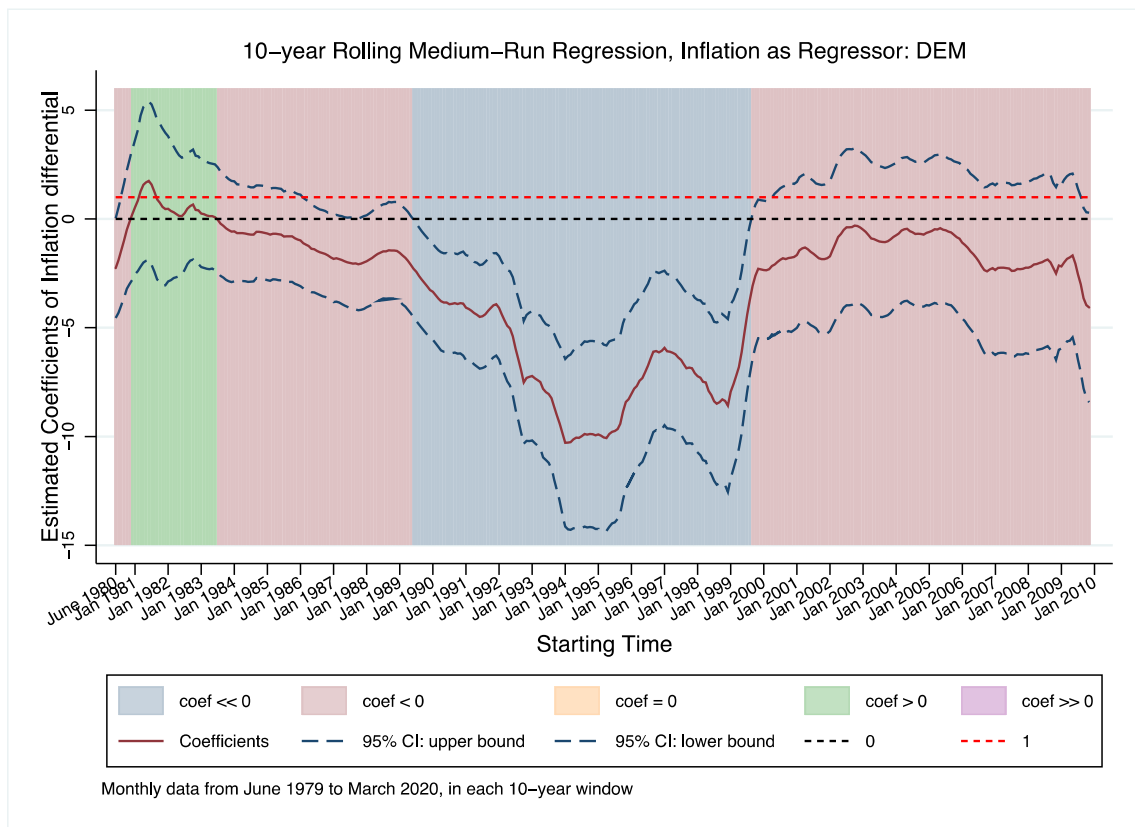
Notes: This table reports the slope coefficient estimates ( $d_M$ ) from Eq. (7),  $\sum_{j=1}^{12} s_{t+j+1} - s_{t+j} + i_{t+j}^* - i_{t+j} = \alpha_M + \delta_M(\pi_t - \pi_t^*) + u_t^M$ , based on the pre-Global Financial Crisis (GFC) common-sample period (1987:01–2006:12) for each currency and fixed-effect Panel regression. See Tables 16 & 17 for complete notes.



**Table 19**  
Slope coefficient from medium-run regression (7), 2007:01–2020:09.

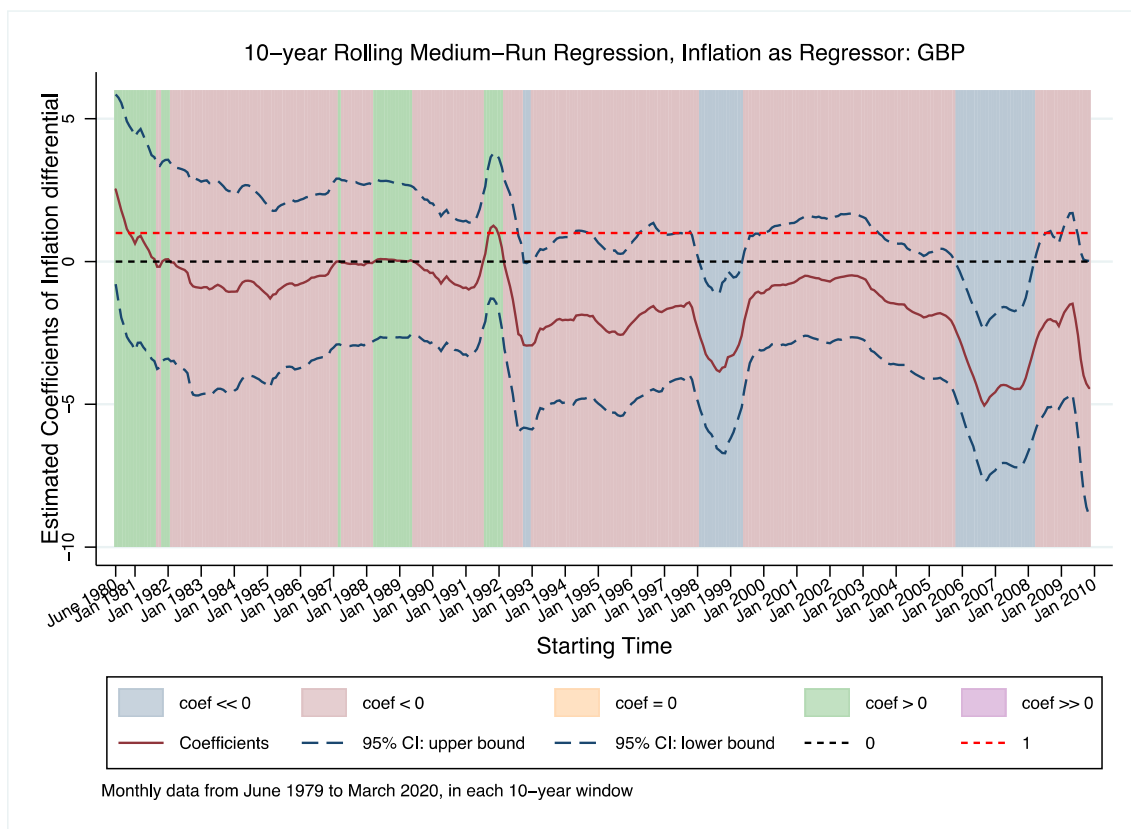
Currency	Time	Obs.	$d_M$	CI	p-Value
CAD	2007/01–2020/09	153	-3.508	(-6.813, -0.203)	0.039
CHF	2007/01–2020/09	153	-2.655	(-5.972, 0.662)	0.119
DEM	2007/01–2020/09	153	-2.252	(-5.716, 1.212)	0.204
FRF	2007/01–2020/09	153	-2.167	(-5.915, 1.580)	0.259
GBP	2007/01–2020/09	153	-4.264	(-6.808, -1.720)	0.001
ITL	2007/01–2020/09	153	-2.567	(-5.360, 0.227)	0.074
JPY	2007/01–2020/09	153	1.386	(-0.473, 3.244)	0.146
NOK	2007/01–2020/09	153	-0.947	(-2.960, 1.066)	0.358
SEK	2007/01–2020/09	153	-3.436	(-7.275, 0.403)	0.081
Panel	2007/01–2020/09	1071	-0.654	(-1.616, 0.308)	0.185

Notes: This table reports the slope coefficient estimates ( $d_M$ ) from Eq. (7),  $\sum_{j=1}^{12} s_{t+j+1} - s_{t+j} + \tilde{i}_{t+j} - i_{t+j} = \alpha_M + \delta_M(\pi_t - \pi_t^e) + u_t^M$ , based on the 2007:01–2020:09 period for each currency and fixed-effect Panel regression. See Tables 16 & 17 for complete notes.



**Fig. 14.** 10-year rolling medium-run regression, inflation as regressor: DEM.  
Notes: This figure presents the estimates of the slope coefficient ( $d_M$ ) and 95% confident intervals using 10-year rolling regression of Eq. (7),  $\sum_{j=1}^{12} s_{t+j+1} - s_{t+j} + \tilde{i}_{t+j} - i_{t+j} = \alpha_M + \delta_M(\pi_t - \pi_t^e) + u_t^M$ , using German mark (DEM) data during 1979:06–2020:09. Definitions of  $s_t$ ,  $i_t$ ,  $\tilde{i}_t$ ,  $\pi_t$ , and  $\pi_t^e$  are the same as in Fig. 8. Each component of the graph is the same as in Fig. 11. See Fig. 11 for complete notes.

We can give a causal explanation to these graphs that is consistent with our hypothesis about delayed reaction to monetary policy, or underestimation of the persistence of monetary policy. For each Figure, the first panel shows estimates of (11) for the pre-GFC period of 1989:01–2006:12, and the second graph for the period 2007:01–2020:09. We find that the  $\lambda_k$  quickly turn negative (though not in all cases immediately) as one would expect if monetary policymakers were targeting inflation, as



**Fig. 15.** 10-year rolling medium-run regression, inflation as regressor: GBP.  
 Notes: This figure presents the estimates of the slope coefficient ( $d_M$ ) and 95% confident intervals using 10-year rolling regression of Eq. (7),  $\sum_{j=1}^{12} s_{t+j+1} - s_{t+j} + i_{t+j}^* - i_{t+j} = \alpha_M + \delta_M(\pi_t - \pi_t^e) + u_t^M$ , using British pound (GBP) data during 1979:06–2020:09. See Fig. 14 for complete notes.

a tighter monetary policy leads to an appreciation.<sup>23</sup> However, the maximum appreciation does not occur immediately, but instead many months later. If investors were adjusting their portfolios continuously to their desired level, and if they had rational expectations of the persistence of monetary policy, the maximum appreciation should occur as soon as the market recognizes that policy will be tightened. Although these graphs are not literally impulse response functions to monetary policy changes or even to news about inflation, they have that flavor. They show us that exchange rates react slowly to changes in inflation in ways that can be anticipated. We know that these predictable exchange rate changes are not mirrored in interest rate changes, and hence there are predictable excess returns.

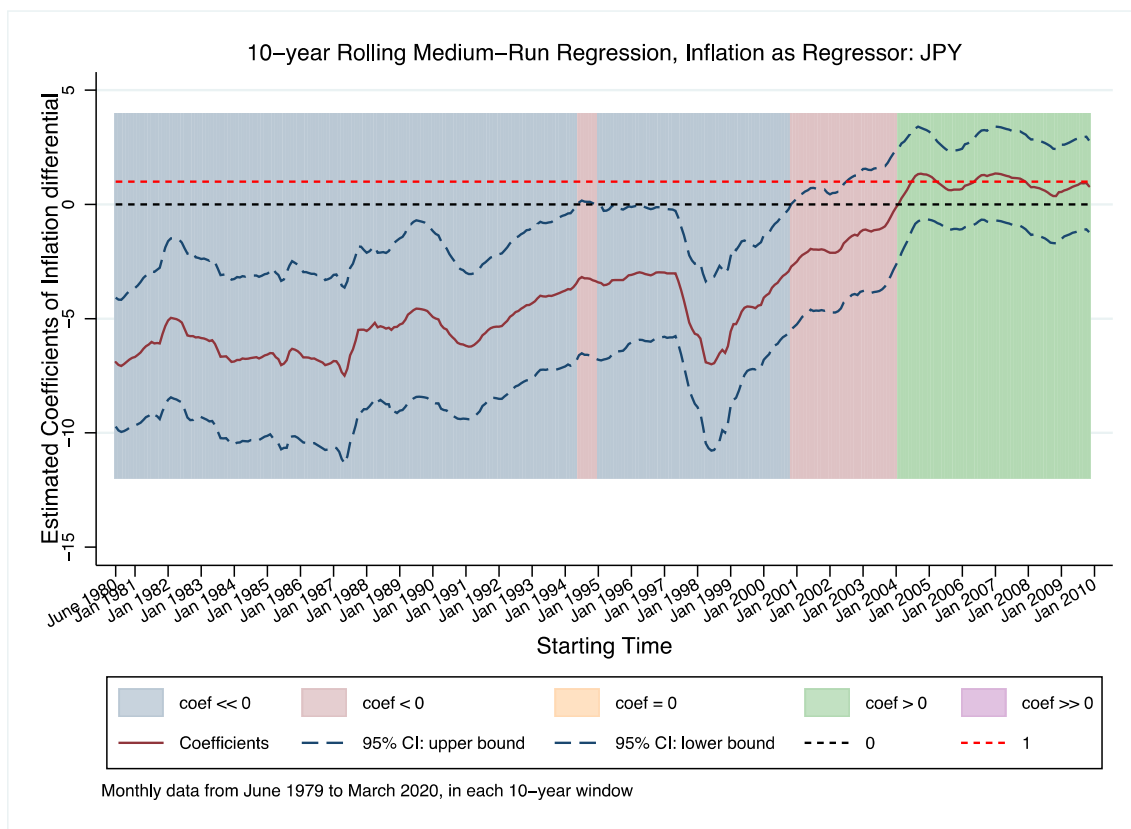
We note that the pattern holds well in both periods, pre- and post-2007:01. Even though the preferred monetary policy instruments changed, the exchange rate reaction to changes in year-on-year inflation is consistent across time.

**5. Long-run**

To understand the regressions with expected long-run returns as the dependent variable, Eqs. (5) and (8), take expectations of the “medium-run” regressions. For example, to motivate Eq. (5), begin by taking the expectation at time  $t$  of the dependent variable in Eq. (4), summing up returns until  $k$  periods in the future:

$$E_t \left( \sum_{j=1}^k s_{t+j+1} - s_{t+j} + i_{t+j}^* - i_{t+j} \right)$$

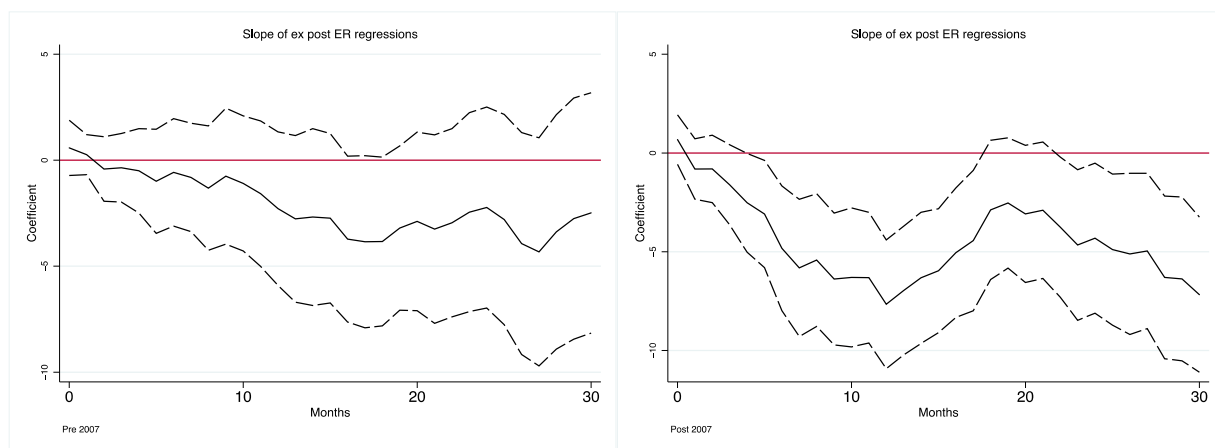
<sup>23</sup> The main exception to this pattern is for Japan, and to a lesser extent for the oil exporters, Canada and Norway.



**Fig. 16.** 10-year rolling medium-run regression, inflation as regressor: JPY.

Notes: This figure presents the estimates of the slope coefficient ( $d_M$ ) and 95% confident intervals using 10-year rolling regression of Eq. (7),

$$\sum_{j=1}^{12} s_{t+j+1} - s_{t+j} + i_{t+j}^* - i_{t+j} = \alpha_M + \delta_M (\pi_t - \pi_t^*) + u_t^M, \text{ using Japanese yen (JPY) data during 1979:06-2020:09. See Fig. 14 for complete notes.}$$



**Fig. 17.** Reaction of exchange rate to inflation differential: DEM.

Notes: This figure presents the estimates of the slope coefficient ( $\hat{\lambda}_k$ ) and 95% confident intervals from Eq. (11),  $s_{t+k} - s_{t-1} = \alpha + \lambda_k (\pi_t - \pi_t^* - (\pi_{t-1} - \pi_{t-1}^*)) + \varepsilon_{t+k}$ ,  $k = 0, 1, 2, \dots, 30$ , using German mark (DEM) data. The left-hand-side figure corresponds to sample period 1987:01-2006:12 and the right-hand-side figure corresponds to 2007:01-2020:07.

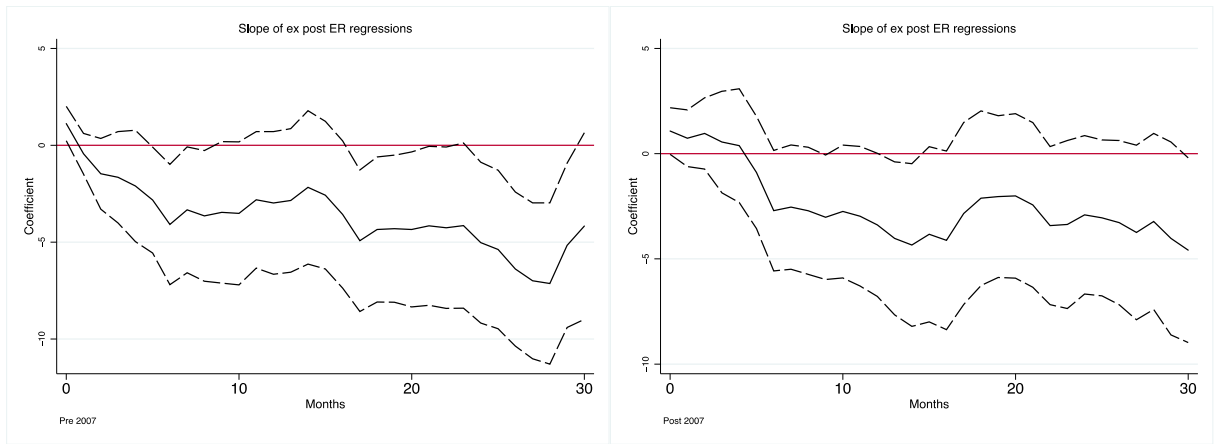


Fig. 18. Reaction of exchange rate to inflation differential: GBP.

Notes: This figure presents the estimates of the slope coefficient ( $\hat{\lambda}_k$ ) and 95% confident intervals from Eq. (11),  $s_{t+k} - s_{t-1} = \alpha + \lambda_k(\pi_t - \pi_t^* - (\pi_{t-1} - \pi_{t-1}^*)) + \varepsilon_{t+k}$ ,  $k = 0, 1, 2, \dots, 30$ , using British pound (GBP) data. The left-hand-side figure corresponds to sample period 1987:01–2006:12 and the right-hand-side figure corresponds to 2007:01–2020:07.

Now subtract the unconditional mean changes in exchange rates, and the unconditional mean interest rate differential:

$$E_t s_{t+k} - s_t - k(\overline{s_{+1} - s}) - E_t \sum_{j=0}^{k-1} (i_{t+j} - i_{t+j}^* - (\overline{i - i^*}))$$

Then take the limit as  $k$  goes to infinity:

$$\lim_{k \rightarrow \infty} [E_t s_{t+k} - s_t - k(\overline{s_{+1} - s})] - E_t \sum_{j=0}^{\infty} (i_{t+j} - i_{t+j}^* - (\overline{i - i^*})) \tag{12}$$

The term  $\lim_{k \rightarrow \infty} [E_t s_{t+k} - s_t - k(\overline{s_{+1} - s})]$  is (minus) the transitory component of the exchange rate in a Beveridge and Nelson (1981) decomposition. Recall any variable with a unit root can be decomposed into a component that is a pure random walk, and a component that is transitory. We shall use a vector autoregression (VAR) to compute a measure of the transitory component of the exchange rate.

The second term on the left-hand-side is the “uncovered interest parity level” of the transitory component of the exchange rate, as defined in Engel (2016). That is, if UIP held, the exchange rate would equal to  $-E_t \sum_{j=0}^{\infty} (i_{t+j} - i_{t+j}^* - (\overline{i - i^*}))$ . This can be seen by rearranging the UIP condition, (1), and iterating forward as in Engel (2016). We can also obtain an estimate of this component from the same VAR mentioned above.

We then regress the measure obtained for the expression in Eq. (12) on  $i_t - i_t^*$  as in Eq. (5), or on  $\pi_t - \pi_t^*$ , as in Eq. (8).

There are two questions we must address that turn out to be very important for the estimates of the quantities in Eq. (12). First, what variables belong in the VAR that we use to produce the dependent variable? Second, how important are small sample considerations in the estimate of the VAR? As it turns out, the two questions are related, and both suggest that results based on shorter samples may not be very reliable.

Since we are interested in the response of exchange rates to interest rate changes and inflation, it seems natural that at a minimum, the VAR that we use should include exchange rates, interest rates and inflation. A key question, though, is whether we model the real exchange rate as stationary or not. Define the log of the real exchange rate as  $q_t \equiv s_t + p_t^* - p_t$ , where  $p_t$  is the log of the consumer price level in the U.S., and  $p_t^*$  is the log of the consumer price level in the foreign country. There is considerable disagreement in the literature about whether the real exchange rate is better modeled as converging or as containing a unit root. In fact, Engel (2000) argues that the question is in essence unresolvable without much longer time series than we use in typical studies. There are plausible reasons why the real exchange rate may be stationary but converging very slowly so that it appears to have a unit root when it does not. Conversely, even if one rejects a unit root using standard tests, there may still be a permanent component that has a small innovation variance and is only detectable in very long samples.

We settle on a VAR that contains the vector of variables  $[q_t \ i_t - i_t^* \ \pi_t - \pi_t^*]$ , but also consider a variant that uses the vector  $[s_t - s_{t-1} \ i_t - i_t^* \ \pi_t - \pi_t^*]$ .<sup>24</sup> Because the interest rate differential and the inflation rate differential are more indisputably

<sup>24</sup> Our estimates for the alternative VAR are reported in the appendix.

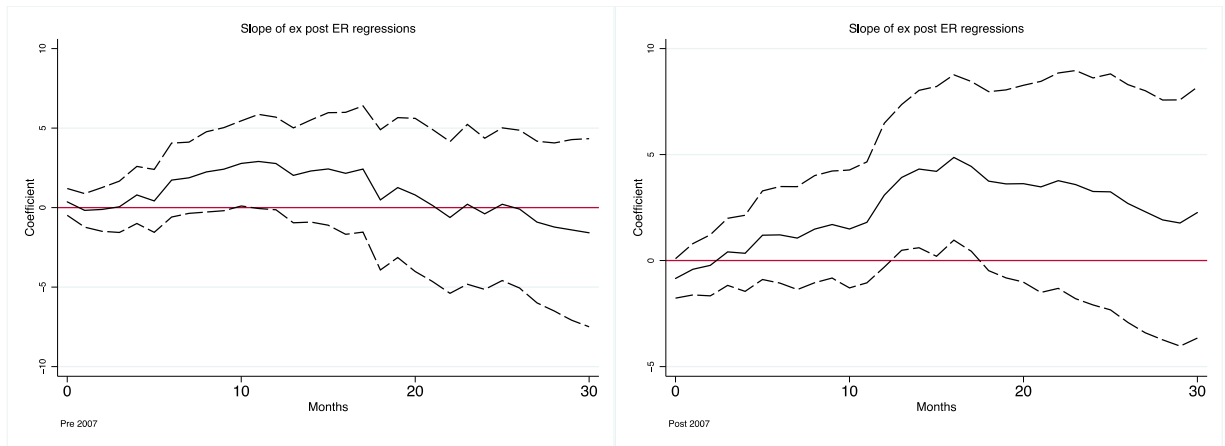


Fig. 19. Reaction of exchange rate to inflation differential: JPY.

Notes: This figure presents the estimates of the slope coefficient ( $\hat{\lambda}_k$ ) and 95% confident intervals from Eq. (11),  $s_{t+k} - s_{t-1} = \alpha + \lambda_k(\pi_t - \pi_t^e - (\pi_{t-1} - \pi_{t-1}^e)) + \varepsilon_{t+k}$ ,  $k = 0, 1, 2, \dots, 30$ , using Japanese yen (JPY) data. The left-hand-side figure corresponds to sample period 1987:01–2006:12 and the right-hand-side figure corresponds to 2007:01–2020:07.

Table 20  
Slope coefficient from long-run regression (13), full sample.

Currency	Time/Obs.	Transitory component	UIP measure	Difference
		$\hat{\eta}_t^T$ [p-val]/95%CI	$\hat{\eta}_t^P$ [p-val] /95%CI	$\hat{\eta}_t^D$ [p-val] /95%CI
CAD	1979/06–2020/09	−53.498 [ $<0.01$ ]	−100.12 [ $<0.01$ ]	46.622 [ $<0.01$ ]
	477	(−67.38, −39.62)	(−118.7, −81.55)	(41.24, 52)
CHF	1979/06–2020/09	−32.333 [ $<0.01$ ]	−22.543 [ $<0.01$ ]	−9.789 [0.079]
	477	(−41.48, −23.19)	(−25.38, −19.71)	(−20.68, 1.098)
DEM	1979/06–2020/09	−38.909 [ $<0.01$ ]	−40.381 [ $<0.01$ ]	1.472 [0.783]
	477	(−48.84, −28.98)	(−41.36, −39.40)	(−9.012, 11.96)
FRF	1979/06–2020/09	−21.947 [ $<0.01$ ]	−26.692 [ $<0.01$ ]	4.744 [0.17]
	477	(−29.14, −14.76)	(−31.77, −21.62)	(−2.018, 11.51)
GBP	1979/06–2020/09	−17.412 [0.011]	−14.112 [ $<0.01$ ]	−3.301 [0.649]
	477	(−30.71, −4.112)	(−15.06, −13.16)	(−17.50, 10.90)
ITL	1979/06–2020/09	−82.55 [ $<0.01$ ]	−87.232 [ $<0.01$ ]	4.683 [0.132]
	477	(−95.31, −69.79)	(−103.7, −70.73)	(−1.403, 10.77)
JPY	1979/06–2020/09	−4.509 [0.474]	−16.29 [ $<0.01$ ]	11.781 [0.143]
	477	(−16.85, 7.829)	(−19.84, −12.74)	(−3.947, 27.51)
NOK	1986/01–2020/09	−29.421 [ $<0.01$ ]	−79.913 [ $<0.01$ ]	50.492 [ $<0.01$ ]
	410	(−36.58, −22.26)	(−85.92, −73.91)	(45.52, 55.47)
SEK	1987/01–2020/09	−109.265 [ $<0.01$ ]	−168.964 [ $<0.01$ ]	59.699 [ $<0.01$ ]
	398	(−144, −74.54)	(−230.2, −107.8)	(32.41, 86.99)

Notes: This table reports the slope coefficient estimates ( $\hat{\eta}_t$ ) from  $s_t^T = \alpha_t^T + \eta_t^T(i_t - i_t^e) + u_t^T$  (Transitory component column),  $s_t^P = \alpha_t^P + \eta_t^P(i_t - i_t^e) + u_t^P$  (UIP measure column), and Eq. (13),  $s_t^T - s_t^P = \alpha_t + \eta_t(i_t - i_t^e) + u_t^D$  (Difference column), based on the full sample (1979:06–2020:09, with variations due to availability of interest rates) for each currency.  $s_t^T = -\lim_{k \rightarrow \infty} [E_t s_{t+k} - s_t - k(\bar{s}_{t+1} - \bar{s})]$  is the transitory component of the exchange rate from the Beveridge-Nelson decomposition, and  $s_t^P = -E_t \sum_{j=0}^{\infty} (i_{t+j} - i_{t+j}^e - (\bar{i} - \bar{i}^e))$  is the value of exchange rate if UIP held. We use a vector autoregression (VAR) to compute the two measures of exchange rate in the stationary case and adopt the small-sample bias correction from West (2016), with which we did not encounter the root greater than one problem. Definitions of  $s_t$ ,  $i_t$ ,  $i_t^e$  are the same as in Table 1. CI in parantheses reports the 95% confidence intervals of the slope coefficient. The  $p$ -value in brackets reports the  $p$ -value of the two-sided  $t$ -test for the slope coefficient  $H_0 : \eta_t = 0$ .

stationary, the former vector imposes that the real exchange rate is stationary, and the latter vector imposes that it is non-stationary. Recent empirical literature has settled on modeling the U.S real exchange rate as stationary (even though there may be a small permanent component), but that is persistent so that the convergence is only detectable in longer samples. We report findings based on the first vector of variables here, and report results under the non-stationarity assumption in the supplemental appendix. We emphasize results estimated from the entire sample, rather than from sub-samples, because of the slow convergence of the real exchange rate.

**Table 21**  
Slope coefficient from long-run regression (14), full sample.

Currency	Time/Obs.	Transitory component	UIP measure	Difference
		$\hat{\zeta}_t^T$ [p-val]/95%CI	$\hat{\zeta}_t^{IP}$ [p-val]/95%CI	$\hat{\zeta}_t$ [p-val]/95%CI
CAD	1979/06–2020/09	-1.822 [ $<0.01$ ]	-4.618 [ $<0.01$ ]	2.796 [ $<0.01$ ]
	477	(-3.147, -0.496)	(-6.450, -2.786)	(2.215, 3.377)
CHF	1979/06–2020/09	0.994 [0.367]	-3.615 [ $<0.01$ ]	4.609 [ $<0.01$ ]
	477	(-1.165, 3.153)	(-4.304, -2.927)	(3.001, 6.217)
DEM	1979/06–2020/09	-2.186 [0.011]	-4.344 [ $<0.01$ ]	2.158 [ $<0.01$ ]
	477	(-3.862, -0.510)	(-5.165, -3.524)	(0.842, 3.474)
FRF	1979/06–2020/09	-1.748 [ $<0.01$ ]	-4.667 [ $<0.01$ ]	2.919 [ $<0.01$ ]
	477	(-2.973, -0.523)	(-5.081, -4.253)	(1.963, 3.874)
GBP	1979/06–2020/09	-0.218 [0.797]	-0.498 [0.017]	0.28 [0.738]
	477	(-1.871, 1.436)	(-0.906, -0.0896)	(-1.360, 1.920)
ITL	1979/06–2020/09	-10.495 [ $<0.01$ ]	-12.071 [ $<0.01$ ]	1.576 [ $<0.01$ ]
	477	(-10.88, -10.11)	(-12.49, -11.66)	(1.158, 1.994)
JPY	1979/06–2020/09	1.89 [0.040]	-2.713 [ $<0.01$ ]	4.603 [ $<0.01$ ]
	477	(0.0907, 3.690)	(-3.201, -2.225)	(2.408, 6.798)
NOK	1986/01–2020/09	-0.2 [0.77]	-4.115 [ $<0.01$ ]	3.915 [ $<0.01$ ]
	410	(-1.539, 1.139)	(-6.129, -2.101)	(2.847, 4.983)
SEK	1987/01–2020/09	-12.124 [ $<0.01$ ]	-23.257 [ $<0.01$ ]	11.133 [ $<0.01$ ]
	398	(-14.90, -9.353)	(-26.84, -19.68)	(10.32, 11.95)

Notes: This table reports the slope coefficient estimates ( $\hat{\zeta}_t$ ) from  $s_t^T = \alpha_t^T + \zeta_t^T(\pi_t - \pi_t^*) + u_t^T$  (Transitory component column),  $s_t^{IP} = \alpha_t^{IP} + \zeta_t^{IP}(\pi_t - \pi_t^*) + u_t^{IP}$  (UIP measure column), and Eq. (14),  $s_t^I - s_t^{IP} = \alpha_t^I + \zeta_t^I(\pi_t - \pi_t^*) + u_t^I$  (Difference column), based on the full sample for each currency. The  $p$ -value in brackets reports the  $p$ -value of the two-sided  $t$ -test  $H_0: \zeta_t = 0$ . See Table 20 for full notes.

The second reason why using longer samples is important is because of small-sample bias in estimating the VAR. This bias is likely to be important in our analysis. From Eq. (12), we can see that the dependent variable depends on the response of  $\lim_{k \rightarrow \infty} [E_t s_{t+k} - s_t - k(\overline{s_{+1}} - \overline{s})]$  relative to the response of  $E_t \sum_{j=0}^{\infty} (i_{t+j} - i_{t+j}^* - (\overline{i} - \overline{i}^*))$  to a change in the regressor. Consider this example. Take the case in which the regressor is the interest rate differential,  $i_t - i_t^*$ , and suppose it turns out that the interest differential follows a first-order autoregression with serial correlation of  $\rho$ . Then our estimate of  $E_t \sum_{j=0}^{\infty} (i_{t+j} - i_{t+j}^* - (\overline{i} - \overline{i}^*))$  is given by  $\frac{1}{1-\hat{\rho}}(i_t - i_t^*)$ , where  $\hat{\rho}$  is the estimate of  $\rho$ . It is well known that the estimate of serial correlation is biased downward. This could have important implications when  $i_t - i_t^*$  is very persistent. If the true value of  $\rho$  is 0.99, but the estimated value is 0.98, the response of  $E_t \sum_{j=0}^{\infty} (i_{t+j} - i_{t+j}^* - (\overline{i} - \overline{i}^*))$  to the interest rate is lowered from the true value of  $100(i_t - i_t^*)$  down to  $50(i_t - i_t^*)$ . Such a large mismeasurement could certainly influence our determination of the sign of the effect of the dependent variable in Eq. (12).

The literature has proposed several ways of dealing with this bias in VAR estimation.<sup>25</sup> We found that when we used some of the corrections from the literature, we were left with VARs that had estimated roots greater than one. Here we have used the small-sample bias correction from West (2016), with which we did not encounter the root greater than one problem. However, especially in the estimates in shorter samples, the slope estimates for regressions (5) and (8) were very sensitive to small differences in the bias correction. Here we report our estimates only for the full sample for each currency because of this problem.

The estimates reported in Tables 20 and 21 for the slope coefficients have the opposite sign of Eqs. (5) and (8). That is, these regressions are of the form:

$$s_t^T - s_t^{IP} = \alpha_L + \eta_L(i_t - i_t^*) + u_t^L \tag{13}$$

$$s_t^T - s_t^{IP} = \alpha_L + \zeta_L(\pi_t - \pi_t^*) + u_t^L \tag{14}$$

where  $s_t^T \equiv -\lim_{k \rightarrow \infty} [E_t s_{t+k} - s_t - k(\overline{s_{+1}} - \overline{s})]$  is the transitory component of the exchange rate from the Beveridge-Nelson decomposition, and  $s_t^{IP} = -E_t \sum_{j=0}^{\infty} (i_{t+j} - i_{t+j}^* - (\overline{i} - \overline{i}^*))$  is the value of the exchange rate if UIP held. In other words, a positive coefficient in these regressions is analogous to a negative coefficient in our previous regressions because they imply that when  $i_t - i_t^*$  or  $\pi_t - \pi_t^*$  increase, expected returns on foreign deposits (over the infinite horizon) rise.

<sup>25</sup> See Bauer et al. (2012), Engsted and Pedersen (2014), Pope (1990).

**Table 22**  
Estimated variances of  $s_t^T$  and  $s_t^{JP}$ , full sample.

Currency	Time	Obs.	$var(s_t^T)$	$var(s_t^{JP})$	F-stat
CAD	1979/06–2020/09	477	0.0219	0.0458	0.4777
CHF	1979/06–2020/09	477	0.0222	0.0031	7.1136
DEM	1979/06–2020/09	477	0.0232	0.0070	3.3279
FRF	1979/06–2020/09	477	0.0184	0.0087	2.1206
GBP	1979/06–2020/09	477	0.0135	0.0008	17.9024
ITL	1979/06–2020/09	477	0.0967	0.1254	0.7710
JPY	1979/06–2020/09	477	0.0247	0.0034	7.2763
NOK	1986/01–2020/09	410	0.0192	0.0432	0.4454
SEK	1987/01–2020/09	398	0.1509	0.3475	0.4341

Notes: This table reports the variance of  $s_t^T$  and  $s_t^{JP}$ . The F-stat column reports the F-statistics of the F-test for the variance difference  $H_0 : var(s_t^T) > var(s_t^{JP})$ , specifically,  $F = var(s_t^T) / var(s_t^{JP})$ . See Table 20 for full notes.

The rightmost column in Table 20 reports the estimates of  $\eta_L$  from Eq. (13). The estimated coefficient has the expected positive sign for seven of the nine currencies. There are three currencies for which the estimated slope is significantly positive at the 1% level for Eq. (13) in which the interest-rate differential is the regressor. Interestingly, these results imply that there is not excess reaction of the exchange rate to changes in either interest rates or inflation rates, in contrast to the findings of Engel (2016). The difference is attributable to the small sample correction that increases the volatility of  $s_t^{JP}$ . That is, when the interest differential is estimated to be more persistent, the UIP value of the exchange rate becomes more volatile.

Table 20 also displays the estimated slope coefficients from regressing  $s_t^T$  and  $s_t^{JP}$  separately on the interest rate differential. As one would expect, all of the estimated slope coefficients are negative. The fact that in most cases the reaction of the transitory exchange rate is smaller (in absolute value) than the reaction of the interest-parity exchange rate demonstrates the finding of no excess volatility. However, the estimates of the response of the interest-parity exchange rate is quite sensitive to the estimated persistence of the interest-rate differential.

Table 21 tells a similar story. The rightmost column reports the estimate of the slope parameter in Eq. (14). The slope estimates are all positive at the 1% level for all the currencies in Eq. (14), which uses the year-on-year inflation rate differential as the regressor, and statistically significant at the 1% level for eight of the nine. The table also reports the slope parameter estimates from regressing  $s_t^T$  and  $s_t^{JP}$  separately on the year-on-year inflation differential. Again, (almost) all of the estimated slope coefficients are negative, as we would expect, but the transitory component of the exchange rate appears to respond less than the interest-parity component.

Indeed, with the correction we have implemented for the small-sample bias in estimating the VAR, there is still mixed evidence of excess volatility of exchange rates, as Table 22 shows. That table reports estimates of the variance of  $s_t^{JP}$  and  $s_t^T$ , and we in only around half of the cases (5 of 9) that  $var(s_t^T) > var(s_t^{JP})$ . In our experience with the estimates, this conclusion is sensitive to the sample period, so the evidence is ambiguous about excess volatility.

The overall picture remains the same – that inflation differentials are a stronger predictor of future excess returns than interest rate differentials, and when the U.S. inflation rate rises, the expected return on foreign deposits declines.

## 6. Conclusions

We find that the famous UIP puzzle is not such a robust puzzle. In fact, the slope coefficient in the Fama regression changes signs in the 2000s. This has been noted previously by, among others, Engel et al. (2019) and Bussière et al. (2018).

One possible explanation for this parameter instability is that, in fact, excess foreign currency returns are unpredictable and the UIP puzzle is merely an illusion. West (2012) has shown that many exchange rate models actually have the implication that UIP holds but that the parameter in the Fama regression will be unstable over time.

Another possible explanation for the parameter instability is that it is related to a genuine change in economic regime. A likely suspect is the change in monetary policy operating procedures. The Fama regression falls apart just when interest rates approach the zero-lower bound. We have offered an explanation for the findings based on the hypotheses of Froot and Thaler (1990), Eichenbaum and Evans (1995), Gourinchas and Tornell (2004), Molavi et al. (2021), and Bacchetta and van Wincoop (2010, 2019). That explanation is supported by empirical examination of the relationship between future excess returns and the current year-on-year inflation rates. Indeed, this relationship emerges in the mid-1980s as inflation targeting became more prevalent and continues in the 2000s even when unconventional monetary policy instruments become more common.

Bussière et al. (2018) also find that the Fama regression no longer produces a negative slope coefficient in the 2000s. However, when the dependent variable (ex post excess returns) is replaced with expected excess returns, where expectations of the future exchange rate are measured from surveys, the slope coefficient is not significantly different than one. That is, uncovered interest parity holds well if the rational expectations hypothesis is jettisoned and replaced with expectations based on surveys of foreign exchange traders (which fail the test of rational expectations.) This is consistent with the possibility that agents do not foresee the

persistence of monetary policy changes, as argued by Gourinchas and Tornell (2004) and Molavi et al. (2021) as we have explained above.

Galí (2020) finds that exchange rates underreact to interest rate changes on long-term bonds but overreact to interest rate changes on short-term bonds. Bacchetta and van Wincoop (2019) demonstrate that this pattern may occur in a model with delayed portfolio adjustment.

### 6.1. Delayed adjustment

The hypothesis we advance is that at time  $t$  agents note that, for example,  $\pi_t - \pi_t^*$  (relative year on year inflation) is below average. If both the home country (the U.S.) and the foreign country target inflation, this implies  $\pi_t - \pi_t^*$  is below their relative targets. Henceforth, for simplicity in this discussion, hold the foreign country variables constant and focus on the U.S. A decline in U.S. inflation signals to the market that U.S. monetary policy will ease – either immediately, or in the very near future.

Suppose the economy were not at the lower bound for interest rates, which we will call the ZLB. This was the situation in the U.S. prior to the early 2000s. If  $i_t$  declines immediately, and uncovered interest parity holds (so there is no risk aversion or constraints to adjusting portfolios), the dollar should depreciate, and its maximum depreciation occurs immediately. After time  $t$ , the dollar is expected to appreciate, so  $E_t s_{t+1} - s_t < 0$ . Under uncovered interest parity, the decline in  $E_t s_{t+1} - s_t$  exactly matches the decline in  $i_t$ .

However, as the previously cited literature has modeled, if portfolio adjustment is costly or if there are constraints on adjusting portfolios, the depreciation of the dollar is not at its maximum at time  $t$ . Some agents shift some of their resources from U.S. interest-bearing deposits to foreign deposits at time  $t$ , but the portfolio rebalancing is insufficient to drive expected returns on U.S. and foreign deposits back into equality, so  $i_t < i_t^* + E_t s_{t+1} - s_t$ . It may take some time for portfolios to fully adjust, so even absent a risk premium, for some extended time of  $j$  periods, we might find  $i_{t+j} < i_{t+j}^* + E_{t+j} s_{t+j+1} - s_{t+j}$ . So, in times when the economies are not at the ZLB, we see the relationship that when  $i_t - i_t^*$  is below average, then also  $i_t - (i_t^* + E_t s_{t+1} - s_t)$  is below average. The interest rate differential,  $i_t - i_t^*$ , forecasts excess returns.

Keeping with the assumption that the economy is not near the ZLB, suppose at time  $t$  agents observe a low outcome for  $\pi_t - \pi_t^*$ , and believe that U.S. monetary policy will ease in period  $t + 1$  rather than in period  $t$ . Under uncovered interest parity, it is still the case that the maximum depreciation would occur at time  $t$ . Then, markets would expect no change in the exchange rate between time  $t$  and  $t + 1$ , ceteris paribus, so  $E_t s_{t+1} - s_t = 0$ . Neither  $i_t$  nor  $E_t s_{t+1} - s_t$  change at time  $t$  if uncovered interest parity holds.

With that example in mind, suppose the economies are at the ZLB. However, central banks can use, for example, forward guidance as an unconventional monetary policy tool. When  $\pi_t - \pi_t^*$  is below average at time  $t$ , the Federal Reserve might credibly announce that at time  $t + k$ , the nominal interest rate,  $i_{t+k}$ , will be lower than markets previously had anticipated. If uncovered interest parity held, the dollar would depreciate immediately at time  $t$ , and then would be expected to remain constant until time  $t + k$ . That is,  $E_t s_{t+k} - s_t$  would equal zero, and the constancy of  $E_t s_{t+k} - s_t$  would match the constancy of short-term interest rates between period  $t$  and period  $t + k - 1$ . Hence, under UIP, the realization of  $\pi_t - \pi_t^*$  would not help forecast excess returns because ex ante excess returns would be zero.

However, suppose portfolio adjustments were costly or were constrained from full adjustment. Then, when  $\pi_t - \pi_t^*$  has a low realization, markets believe that U.S. interest rates will fall in the future. Some agents are able to shift some of their resources from U.S. deposits to foreign deposits, but the rebalancing of portfolios is not sufficient to eliminate the low returns on U.S. assets relative to foreign assets. At time  $t$ , the dollar is expected to depreciate,  $E_t s_{t+1} - s_t > 0$ . Given the constancy of interest rates (at, or very near, the ZLB), we have  $i_t < i_t^* + E_t s_{t+1} - s_t$ , and the foreign deposits will continue to have ex ante excess returns until portfolios fully adjust. We conclude that low values of  $\pi_t - \pi_t^*$  predict high values of  $s_{t+1} - s_t + i_t^* - i_t$ . This negative relationship between  $s_{t+1} - s_t + i_t^* - i_t$  and  $\pi_t - \pi_t^*$  is what we find to hold generally in our data starting in the mid-1980s, as evidence in Tables 8, 9, and 10.

### 6.2. Risk premium

There is another possible interpretation of our finding that the excess returns on U.S. deposits are predictable. Lustig et al. (2014) and Colacito et al. (2020) find that these excess returns are predictable from the U.S. business cycle. If U.S. output growth (measured by industrial production) has positive growth over the past month, or if a measure of the output gap is positive, the excess return on U.S. assets is predictably positive. Colacito et al. further show that this predictability can be motivated by a model of risk averse agents with Epstein-Zin preferences, in which the output gap is a proxy for expected future consumption growth.

In Tables S.15–S.20, we consider whether the output gap or output growth are a better predictor of excess returns than year-on-year inflation. We regress excess returns on the interest rate differential, relative y.o.y. inflation, the growth rate in output, and the output gap, where we use industrial production as a proxy for monthly output. We construct the output gap using the detrending method proposed by Hamilton (2018). The first three tables use only the U.S. output growth and output gap as predictors, while Tables S.18–S.20 take the U.S. measure relative to each other country. For each set of regressions, we estimate over the full 1987:01–2020:09 period as well as the 1987:01–2006:12 and 2007:01–2020:09 sub-periods.



The evidence can be summarized simply that in these regressions, the relative y.o.y. inflation variable retains its predictive power, while the output variables are rarely statistically significant. This is true over the entire sample as well as both sub-samples, using either U.S. output variables alone, or U.S. output measures relative to those of the corresponding country.

These findings do not, however, directly shed light on the risk-premium versus delayed reaction to monetary policy theories. Both the output measures and the inflation variable are proxies for the state of aggregate demand. In the delayed reaction story, the central banks react to high aggregate demand by contracting monetary policy, which leads to appreciation of that country's currency, but with a delayed reaction so that the appreciation is predictable. In the risk premium theory, high aggregate demand predicts high future consumption growth, and therefore a higher stochastic discount factor, which implies a higher risk premium.

These competing hypotheses deserve further investigation and are beyond the scope of this paper. However, we note some preliminary evidence that favors the delayed reaction hypothesis. In the risk premium story, the dollar is predicted to have a positive excess return, for example, when it is riskier. In the zero-lower-bound period, that excess return is reflected as a drop at time  $t$  in  $E_t s_{t+1} - s_t$ , an expected appreciation of the dollar. Because the dollar asset is riskier, this expected appreciation at time  $t$  is generated by a depreciation of the dollar at that time. In other words, if the level of aggregate demand rises at time  $t$  which makes the dollar riskier at that time, then  $s_t - s_{t-1}$  or  $s_t - E_{t-1}(s_t)$  should rise. This is the essence of Engel's (2016) observations about risk-premium models of the Fama puzzle.

In contrast, the delayed reaction explanation suggests very little change or a decline in  $s_t - s_{t-1}$  or  $s_t - E_{t-1}(s_t)$  when aggregate demand is high. We have already seen in Figs. 17–19 and Figs. S.52–S.60 that when an increase in  $\pi_t - \pi_t^e$  predicts a subsequent appreciation of the dollar, the initial change in the exchange rate is muted. This finding is more consistent with the delayed reaction theory. A more careful examination of the different predictions of these models, however, is likely to lead to fruitful insights into the source of excess return predictability.

## 7. Summary

We emphasize that the findings here are not definitive. The economic relationships that determine ex ante excess returns are probably too complicated to be captured in univariate regressions. Our findings do provide evidence that the Fama puzzle is not a stable relationship, and the Fama regression does not generate stable parameter estimates over time. Future research is needed to determine whether, in fact, the data support the hypothesis of delayed portfolio adjustment that may arise from barriers to reallocating assets, from biased expectations, or risk premia.

## Declaration of Competing Interest

Charles Engel has nothing to declare.

Ekaterina Kazakova has nothing to declare.

Mengqi Wang has nothing to declare.

Nan Xiang has nothing to declare.

## Appendix A. Data source

**Table A1**

Data source, exchange rates.

Currency	Data	Time	Source	Specific ticker
AUD	1989/01–2020/09	1989/01–1998/12	FRED	EXUSAL
		1999/01–2020/09	DataStream	AUSTDO\$
CAD	1979/06–2020/09	1979/06–1998/12	DataStream	CNDOLL\$
		1999/01–2020/09	DataStream	CNDOLL\$
CHF	1979/06–2020/09	1979/06–1979/12	DataStream	SWISSF\$
		1989/01–1989/12	FRED	EXSZUS
		1999/01–2017/11	DataStream	SWISSF\$
		2017/12–2020/09	FRED	EXSZUS
DEM	1979/06–2020/09	1979/06–2017/11	DataStream*	USEURSP
		2017/12–2020/02	FRED*	EXUSEU
FRF	1979/06–2020/09	1979/06–2017/11	DataStream**	USEURSP
		2017/12–2020/02	FRED**	EXUSEU
GBP	1979/06–2020/09	1979/06–2017/11	DataStream	USDOLLR
		2017/12–2020/02	FRED	EXUSUK
ITL	1979/06–2020/09	1979/06–2017/11	DataStream***	USEURSP
		2017/12–2020/02	FRED***	EXUSEU
JPY	1979/06–2020/09	1979/06–2017/11	DataStream	JAPAYE\$
		2017/12–2020/02	FRED	EXJPUS
NOK	1986/01–2020/09	1986/01–1998/12	FRED	EXNOUS
		1999/01–2017/11	DataStream	NORKROS

(continued on next page)

**Table A1** (continued)

Currency	Data	Time	Source	Specific ticker
NZD	1997/04–2020/09	1997/04–2020/09	DataStream	NZDOLL\$
SEK	1987/01–2020/09	1987/01–1998/12	FRED	EXSDUS
		1999/01–2017/11	DataStream	SWEKROS\$
		2017/12–2020/09	FRED	EXSDUS

Notes: \*, \*\*, \*\*\* Exchange rates of the German mark, French francs, and Italian lira are converted into euros using the euro conversion rates at the time of origination of the euro in January 1999, which are 1.95583 German mark for 1 euro, 6.55957 French francs for 1 euro, and 1936.27 Italian lira for 1 euro. These conversion rates are from [https://en.wikipedia.org/wiki/Enlargement\\_of\\_the\\_eurozone#Convergence\\_criteria](https://en.wikipedia.org/wiki/Enlargement_of_the_eurozone#Convergence_criteria).

**Table A2**

Data source, 1 M LIBOR rates.

Currency	Data	Time	Source	Specific ticker
AUD	1989/01–2020/09	1989/01–1998/12	FRED	AUD1MTD156N
		1999/01–2020/09	DataStream	ECAUD1M
CAD	1979/06–2020/09	1979/06–2020/09	DataStream	ECCAD1M
CHF	1979/06–2020/09	1979/06–2020/09	DataStream	ECSWF1M
DEM	1979/06–2020/09	1979/06–2020/09	DataStream*	ECEUR1M
FRF	1979/06–2020/09	1979/06–2020/09	DataStream*	ECEUR1M
GBP	1979/06–2020/09	1979/06–2020/09	DataStream	ECUKP1M
ITL	1979/06–2020/09	1979/06–2020/09	DataStream*	ECEUR1M
JPY	1979/06–2020/09	1979/06–2020/09	DataStream	EJJPY1M
NOK	1986/01–2020/09	1986/01–1998/12	Bloomberg	NIBOR1M Index
		1999/01–2017/11	DataStream	ECNOR1M
NZD	1997/04–2020/09	1997/04–2020/09	DataStream	ECNZD1M
SEK	1987/01–2020/09	1987/01–1989/12	Bloomberg	SK0001M Index
		1999/01–2020/09	DataStream	ECSWE1M
USD	1979/06–2020/09	1979/06–2017/11	DataStream	ECUSD1M

Notes: \* use EUR 1 M LIBOR rate. Monthly data uses the last data-available day of each month from the daily data.

**Table A3**

Data source, consumption price index.

Currency	Data	Source	Specific ticker
CAD	1979/06–2020/09	FRED	CANCPALLMINMEI
CHF	1979/06–2020/09	FRED	CHECPALLMINMEI
DEM	1979/06–2020/09	FRED	DEUCPALLMINMEI
FRF	1979/06–2020/09	FRED	FRACPALLMINMEI
GBP	1979/06–2020/09	FRED	GBRCPALLMINMEI
ITL	1979/06–2020/09	FRED	ITACPALLMINMEI
JPY	1979/06–2020/09	FRED	JPNCPALLMINMEI
NOK	1986/01–2020/09	FRED	NORCPALLMINMEI
SEK	1987/01–2020/09	FRED	SWECPALLMINMEI
USD	1979/06–2020/09	FRED	USACPALLMINMEI

## Appendix B. Supplementary Tables and Figures

Supplementary data to this article can be found online at <https://doi.org/10.1016/j.jinteco.2022.103602>.

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