The Forward Discount Puzzle: Identification of Economic Assumptions

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Abstract

The forward discount puzzle refers to the robust empirical finding that foreign excess returns are predictable. We investigate if expectations errors are the main cause of this predictability using the serial dependence pattern of excess returns implied by economic models as identification device. This approach also allows us to explain why strong predictability of excess returns only occurs during 1980s. Using USD bilateral spot and forward rates from 1975-2009, we show that both the statistically significant positive serial dependence of excess returns in the entire sample and the very weak (mostly insignificant) positive serial dependence in the subsample excluding observations in 1980-87 are consistent with the predictions of the expectations errors explanation. We provide several pieces of new empirical evidence which support the link between the strong predictability in the 1980s and changes in forecasting techniques by foreign exchange market agents.

Keywords: expectations errors, rational expectations risk premium, excess returns, serial dependence, 1980-87.

JEL Classification: F31, G14.

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1 Introduction

The unbiased hypothesis of forward exchange rates states that the forward rate is the optimal predictor of future spot rates and implies that the expected excess return on foreign relative to domestic currency must be zero. However, numerous studies have persistently documented that excess returns are predictable, mainly using the forward discount or past returns as predictors. The forward discount puzzle, one of the most robust puzzles documented in international economics, refers to such findings.[1]

There have been two popular explanations for the forward discount puzzle in the literature: rational expectations risk premia and expectations errors.[2] We reconsider these explanations and investigate the following questions: (i) Is the predictability due to the presence of the time-varying risk premium in an efficient market or evidence of market inefficiency which reflects deviations from strong rationality?[3] (ii) Why does the strong predictability only appear in the 1980s?

Froot and Frankel (1989) first addressed this identification issue using additional data such as survey data. Instead, we look for information directly from the economic model that restricts the relations between the behavior of macro fundamentals and equilibrium expected returns. We pay particular attention to the serial dependence pattern of excess returns over return horizons implied by economic models for the expected excess returns. We argue that empirical evidence on this serial dependence pattern provides a valid criterion for judging the performance of economic models. The main argument is that a mean reverting component in excess returns, representing a violation of the unbiased hypothesis, can generate different serial dependence patterns, depending on the economic model assumptions. On the one hand, a class of rational expectations risk premium models tends to generate 'negative' serial dependence in excess returns mainly because fundamental variables are negatively correlated with the risk premium in these models,

[1] See, Lewis (1995) and Engel (1996) for the survey. For recent contributions, see Chinn (2006).

[2] Several studies also investigate poor finite sample properties of predictive regression tests [see, for example, Evans and Lewis (1995)]. In a recent paper, Bacchetta and van Wincoop (2010) present another explanation based on the model of infrequent portfolio decisions.

[3] For example, Alvarez, Atkeson, and Kehoe (2009) and Verdelhan (2010) are successful for relating the cause of the forward discount puzzle to the rational expectations risk premium. On the other hand, Froot and Frankel (1989), Frankel and Froot (1990a, 1990b), Mark and Wu (1998), and Gourinchas and Tornell (2004) related successfully the cause to expectations errors.

e.g., counter-cyclical exchange risk premia in Verdelhan (2010). Those include a monetary general equilibrium model with time varying risks by Alvarez, Atkeson, and Kehoe (2009) and an external habit-persistence model with time-varying risk aversion by Verdelhan (2010), which are successful in generating highly variable risk premia. On the other hand, a class of models for expectations errors tends to generate 'positive' serial dependence in excess returns. Those include a model of a speculative bubble by Frankel and Froot (1990a, 1990b) and a noise trader model by Mark and Wu (1998).

Our analysis also complements the popular criterion on the performance of economic models in the literature. In order to investigate the influence of the risk premium in foreign exchange markets, studies have focused on whether or not their models can generate a variance of the risk premium higher than that of the expected exchange rate change. This volatility relation was first illustrated by Fama (1984) using the decomposition of the slope coefficient in the regression of the exchange rate change on the forward discount. However, several studies show that the estimated slope coefficient may not be so informative because its distribution is very wide and the magnitude of systematic estimation errors can be large (see, e.g., Baillie and Bollerslev (2000), West (2008), and Moon and Velasco (2010b)). We also present empirical findings which support this conclusion and argue that previous studies may have overemphasized the estimation results from Fama regression. By contrast, our approach is not subject to this shortcoming, identifies the predictability over return horizons, and provides an explicit criterion for the judgement on the model performance.

Using US dollar (USD) bilateral spot and one-, three-, six- and twelve-month forward rates from 1975-2009, we find that excess returns strongly exhibit positive serial dependence over return horizons up to five years, consistent with the prediction of the expectations errors explanation. But the predictability of excess returns becomes very weak in the sample which excludes observations in 1980-87: excess returns are not statistically significant in most currencies, implying that the deviations from the unbiased hypothesis are mainly driven by a particularly influential subset of the data.

This evidence leads us to investigate the second question and to reconsider the speculative bubble hypothesis on the path of the US dollar in the 1980s proposed by

Frankel and Froot (1990a, 1990b). In particular, we link our empirical results with 'changes in forecasting techniques' used by agents in foreign exchange markets in the 1980s. That is, by assuming that the market expectation is a weighted average of expectations of fundamentalists and chartists, we relate the changes in the weights during 1980s to the market's Bayesian responses to the relative forecasting performance of an agent. We also provide several pieces of evidence which support this argument: (i) the very strong predictability of excess returns in 1980-87 is only observed in USD bilateral rates; and (ii) the behavior of the survey expectation is consistent with the market expectation over both subsample periods: when the unbiased hypothesis is decisively rejected and also when it tends to hold.

The organization of the paper follows. Section 2 documents empirical results on the strong predictability of excess returns in the 1980s based on the standard regression tests. Section 3 shows analytically that the two competing economic explanations generate an opposite sign of the serial dependence in excess returns. Section 4 provides empirical results on the serial dependence pattern of excess returns based on the variance ratio test. Section 5 offers an explanation for the appearance of strong predictability in the particular sample period of 1980-87 and Conclusions follow.

2 Strong Predictability of Foreign Excess Returns in the 1980s

We perform the predictability test on one-, three-, six- and twelve-month excess returns based on the following standard regression for currency excess returns,

$$s_{t+k} - f_{t|k} = \alpha_k + \beta_k(f_{t|k} - s_t) + u_{t+k}, \tag{1}$$

where s_t denotes the log of the domestic currency price of foreign currency at time t and $f_{t|k}$ denotes the log of the domestic currency price at time t of foreign currency delivered at $t+k$. Both α_k and β_k are zero for any k under the unbiased hypothesis with the assumptions of rational expectations and risk neutrality. Under these assumptions, the unbiased hypothesis is equivalent to the uncovered interest parity (UIP) hypothesis. From now on, we use both hypotheses interchangeably. Our sample consists of weekly observations of eleven London close mid prices whose description is presented in Table I.

2.1 Standard Results Reproduced

Table III reports the estimated slope coefficients both when regression (1) is estimated for an individual currency and when a single β is estimated for a pooled sample of all currencies and periods of time as well as for two portfolios which consist of five major currencies: the German mark (GDM), the British pound (BRP), the Japanese yen (JPY), the Canadian dollar (CAD), and the Swiss franc (SWF). The first portfolio (Eport) is constructed using an equal weight and the second one (Wport) is constructed based on weights calculated using the data of foreign exchange market turnover by currency pair.

Panel A provides the results from the entire sample period. Consistent with the previous empirical findings, $\widehat{\beta}_k$ are negative (mostly less than -1) and the conventional t-test rejects the UIP hypothesis at a 1% level in almost all cases. For example, the pooled $\widehat{\beta}_k$ are -1.06, -1.04, -0.97, and -0.88, respectively, for the one-, three-, six-, and twelve-month maturities and statistically significant at all the conventional levels. Notoriously, the R^2s in the regressions are very low in magnitude. These results confirm well-known empirical regularities, which refer to the forward discount puzzle.

2.2 Strong Predictability in the 1980s

We now conduct the same predictive test on subsample periods. In particular, we are interested in conducting the test on the sample which excludes observations of 8 years in the period of 1980-87, out of 35-year observations in the entire sample. The separation is motivated by empirical evidence on the predictability of excess returns in the 1980s. For example, studies often find ambiguous results using observations in the 1990s, although salient deviations from UIP are detected based on the data before 1990s [see, e.g., Flood and Rose (2002)]. Evans (1986) and Meese (1986) also provide evidence on the presence of speculative bubbles in foreign exchange markets during the first half period of the 1980s.

Panel B reports the results from regression (1) using the sample excluding observations in 1980-87. From now on, we call it subsample E. In contrast to the results in the entire sample, the UIP hypothesis tends to hold or the rejections become very weak. The t-test does not reject it for most currencies: $\widehat{\beta}_k$ is statistically indistinguishable from zero at a 5% level in most cases. For example, the t-statistics of the pooled estimates are -1.64, -1.51, -1.50, and -1.57, respectively, for the one-, three-, six-, and twelve-month maturities. Both portfolios also produce similar results except for the twelve-month maturity: the estimates are marginally rejected at the 5% level. For individual currencies, the rejections only occur at the 5% level for CAD and the Australian dollar (AUD) for the one-month maturity, and for CAD and the Danish Krone (DAK) for the three-month maturity. For other maturities, the rejection only occurs for JPY for the twelve-month maturity. We find some more rejections at the 10% level for the six- and twelve-month maturities. These results are surprising, compared to those from the entire sample as well as from numerous previous studies.

The results using another subsample which contains observations only in 1980-87 [We call it subsample O] are reported in Panel C and further confirm that the rejections in the entire sample seem to be significantly influenced by this. The absolute values of both the estimated slope coefficients and their t-statistics are much greater than those in the other sample periods. Further, the strong explanatory power of the forward discount, measured by R^2, is noticeable.[4] For example, for BRP, the R^2s are 0.18, 0.43, 0.52, and 0.36 for the one-, three-, six-, and twelve-month maturities, respectively, while they are almost zero in subsample E, indicating that the strong negative correlation between the spot return and the forward discount may be only present in this particular sample period.[5] This pattern appears for most major currencies. Our findings are, in a sense, compatible with numerous previous studies that document the strong predictability of excess returns simply because their samples include observations in the 1980s.

2.2.1 Robustness I: A Parameter Stability Test

For the robustness, we first investigate if the above results are sensitive to the selection of the subsample period. For this, we design a stability test based

[4]For the one-month and three-month maturities, we also calculate R^2 using non-overlapping observations to see if the results are contaminated by overlap of the data. However, they are robust to this change.

[5]Our empirical findings can be related to those of Bansal (1997). He found that deviations from the UIP hypothesis depend systematically on the sign of the interest rate differential across countries using weekly data of GDM and JPY in 1981-95. Specifically, the US dollar was strongly appreciated when the US interest rate was greater than other countries' interest rates but UIP tends to hold when it is less than them. According to our data, the US interest rate was higher than others for most weeks during the time period of 1981-85.

on rolling samples, instead of typical ones using cumulated sample observations.[6] We generate 8-year rolling samples whose size is equal to that of subsample O. The first sample is obtained using weekly observations on each country's spot and one-month forward rates from January 1975 to December 1982. The next sample then contains those from January 1976 to December 1983, that is, observations in 1975 are dropped and those in 1983 are added. And so on. We estimate the slope coefficient in regression (1) for each rolling sample j for $j = 1, \ldots, J$ and obtain its t-statistic. To estimate critical values for the tests on the stability of β_k, we use a bootstrap algorithm specified in detail in Appendix A.

Figure 1 plots the locus of the t-statistics obtained using 8-year rolling samples as well as the 90% uniform test bands with 5 and 95% quantiles from the simulated bootstrap empirical distribution of t-statistics. We reject the null hypothesis that all the slope coefficients in rolling samples of one-month excess returns are zero at the 10% significance level for all five major currencies considered. However, the rejections are mainly due to rolling samples which include some observations in the period of 1980-87. In those samples, the estimates are significantly negative and the corresponding t-statistics are much beyond the lower bound of the 90% band. On the other hand, the t-statistics from rolling samples which do not contain observations in 1980-87 tend to be inside the band. These confirm the empirical results that the predictability of excess returns is very weak in subsample E.

2.2.2 Robustness II: Statistical Issues

We now examine the extent of possible statistical artifacts. In particular, we are concerned about the sensitivity of statistical inference: the t-statistics can change dramatically with the drop/addition of a few observations when the regressor is highly persistent, as discussed in Campbell and Yogo (2006). Therefore, we investigate if the difference in the results between the two samples with/without observations in 1980-87 can be attributable to problems in the specification of the persistency.[7]

[6] We thank Efstathios Paparoditis for the suggestion of this method. Baillie and Bollerslev (2000) also generate rolling samples to investigate the pattern of the slope coefficient in the regression of spot rates on the forward discount.

[7] Similarly, several studies (e.g., Baillie and Bollerslev (2000), and Maynard and Phillips (2001), and references therein) are concerned with the over-rejections of regression tests using conventional critical values, based on empirical evidence on the (near) nonstationary behavior of the forward discount.

For this, we conduct the conditional test developed by Jansson and Moreira (2006) based on the regression model in equation (1) because their test has correct size and desirable power properties in the specification with (near) integrated regressors.[8] For implementing the test, we follow Polk, Thompson, and Vuolteenaho (2006) who apply the same test for stock return predictability.[9] The test is conducted based on the samples with non-overlapping observations since it cannot account for serial dependence in predictive errors. We choose observations at the end of each maturity, for example, at the end of each month for the one-month maturity.

Table IV reports the t-statistics of the estimated slope coefficient in regression (1) and 2.5 and 97.5% quantiles of the conditional t distribution. We find that both the conventional t-test and the conditional test produce quite similar results: both tests agree with the rejections/non-rejections. And the critical values at the two quantiles obtained from the conditional test are quite similar to those from the conventional t-test. For example, both tests decisively reject the null hypothesis in the entire sample for most currencies. On the other hand, they do not reject it for most currencies in subsample E. These results suggest that a statistical phenomenon cannot be the reason for our empirical findings.[10] As Mankiw and Shapiro (1986) illustrate, the conventional t-test over-rejects the null in regression (1) if the contemporaneous correlation between innovations to the regressor and to the dependent variable is large and the regressor is strongly persistent. In our sample, the forward discount is highly persistent [see Table 2], while the contemporaneous correlations are close to zero. For example, the absolute values of the correlations are less than 0.1 for all currencies, suggesting that the absence of statistical distortions may be due, at least partially, to the small contemporaneous correlations.

[8] We also conduct Campbell and Yogo's Q test and find that the results are quite similar between the two tests. For the sake of simplicity, we do not report those from the Q-test.

[9] We use the Matlab codes written by Polk, Thompson, and Vuolteenaho (2006) with a small modification for our study. Their codes are available at
$http://personal.lse.ac.uk/POLK/research/work.htm$.

[10] Our results are consistent with Maynard (2006) who studied the effects of a statistical phenomenon in foreign exchange markets.

2.3 Information in the Estimated Slope Coefficients

In this subsection, we discuss the precision of the estimated slope coefficient as well as the magnitude of the estimation bias. One apparently peculiar phenomenon regarding the results from subsample E is that the sign of the estimated slope coefficients is towards a particular direction, although they are not statistically significant and there is no clear symptom of statistical distortions. Estimates are negative for all cases and less than -1 for many cases. If the distribution of the estimated slope coefficient were normal, then on average we should observe both positive and negative values equally likely.

We argue that this can be explained by the typical present value models. Specifically, Moon and Velasco (2010b) show that it may not be uncommon to observe 'insignificant' estimates with large deviations from the null value and with the same sign, based on the bivariate regression in which both s_t and $f_{t|k}$ are generated under the UIP hypothesis.[11] The key idea can be seen in the decomposition of the contemporaneous correlation, γ,

$$\gamma = \frac{\sigma_{uv}/\sigma_v^2}{\sigma_u/\sigma_v}, \tag{2}$$

where σ_u is the standard deviation of the error term u_{t+1} in regression (1) and σ_v is the standard deviation of the error term in the regression of the forward discount on its lagged value. The scale factor, σ_{uv}/σ_v^2, affects the magnitude of the bias, $E[\widehat{\beta} - \beta]$, in regression (1) [see, e.g., Stambaugh (1999)] and the noise-to-signal ratio, σ_u/σ_v, mainly determines the variance of the estimated slope coefficient, $\widehat{\beta}$. Equation (2) shows that a large magnitude of the bias must be accompanied by a large value of σ_u/σ_v and thus of the standard error of $\widehat{\beta}$, since the absolute value of γ is bounded by one. However, a smaller correlation of γ is not necessarily related to a smaller value of σ_u/σ_v. In an extreme case, σ_u/σ_v can take any value if σ_{uv} is zero. Equation (2) also implies that a large magnitude of bias does not necessarily lead to significant over-rejections of the t-test. Rather, what matters is the ratio between σ_{uv}/σ_v^2 and σ_u/σ_v, which determines the standardized bias of the t-test. Moon and Velasco (2010b) show that, in the present value models, the near unit values of the discount factor and the persistence parameter induce a larger value of σ_u/σ_v as well as a larger magnitude of bias and thus a wider distribution of $\widehat{\beta}$, while

[11] See, also, Baillie and Bollerslev (2000) and West (2008) who reached a similar conclusion on the information content of the estimated slope coefficient.

generating a smaller value of the contemporaneous correlation. Consequently, the estimated slope coefficient becomes less informative and, at the same time, the over-rejections do not occur. Empirical evidence on the large values of the discount factor [see, e.g., Engel and West (2005)] and on strong persistence in the forward discount [see Table 2] supports this argument. Further, the large magnitude of σ_u/σ_v is consistent with the foreign exchange data in that the relative variance between the excess return and the forward discount is in general very large. For example, for the one-month maturity, the relative variances between the two series are in general greater than 150 for most currencies as in Table 2.

This result may challenge the conventional view on Fama (1984)'s volatility relation which has been used as one of the criteria for evaluating the performance of the rational expectations models in terms of generating high volatilities of the risk premium. Since the relation is derived from the magnitude of $\widehat{\beta}$, it is important to understand if the estimates contain valid information. In the next section, we provide an alternative identification strategy which may complement this volatility relation.

3 Identification of Two Competing Explanations: Rational Expectations Risk Premia vs. Expectations Errors

In this section, we consider an implication of general models for the expected excess returns: the serial dependence pattern of excess returns. In particular, we use as an identification device the sign of autocorrelations of excess returns which is critically determined by a feedback from forecasting errors to future expected excess returns. To illustrate this, we present the two well-known explanations for the predictability of excess returns, rational expectations risk premia and expectations errors, and show that they tend to generate opposite signs of the autocorrelation function of those returns.

3.1 An Exchange Rate Model

Following Frankel and Froot (1990b) and Engel and West (2005), we consider a general setup for the determination of the exchange rate

$$s_t = \alpha E_t[\Delta s_{t+1}] + \alpha p_t^e - \alpha p_t + \varpi_t + w_t, \tag{3}$$

where $p_t^e = E_t^m[\Delta s_{t+1}] - E_t[\Delta s_{t+1}]$ is the difference between the market expectation and the rational expectation (expectations errors), p_t is the risk premium, ϖ_t is the log of the real exchange rate, w_t is the linear combination of logs of other fundamental variables such as money and output, and α is constant. ϖ_t represents a deviation from the log of purchasing power parity (PPP), while both p_t^e and p_t represent deviations from the log of UIP,

$$E_t[\Delta s_{t+1}] + p_t^e - p_t = i_t - i_t^* = f_t - s_t \tag{4}$$

where the second equality represents the log of covered interest parity. We assume maturity $k = 1$ in this section and omit the notation for maturity, for example we denote the log of the one-period forward exchange rate by $f_t = \ln F_{t|1}$. In the typical monetary models, equation (3) is derived from home and foreign money demands combined with equation (4) [see, e.g., Engel and West (2005) and Obstfeld and Rogoff (2002) for the rational expectations and Frankel and Froot (1990b) for the expectations errors models].

We assume that the process for w_t follows

$$\Delta w_t = \rho \Delta w_{t-1} + \eta_t, \tag{5}$$

where η_t represents a stochastic disturbance with mean zero and variance σ_η^2. If $\rho = 0$, then the fundamental process follows a random walk. The introduction of more complicated fundamental processes would not change our main results below mainly because the predictability of excess returns is coming from deviations from UIP. Therefore, considering equation (5) may be enough for our objectives.

For modeling the rational expectations risk premium, we begin with considering currency prices in a fairly general setting and then turning to more structured economies. On the other hand, we follow Frankel and Froot (1990b) for modeling expectations errors. Both p_t and p_t^e are assumed to follow a stationary process, respectively, as specified in detail below. With these assumptions and $\rho = 0$, the spot exchange rate can be described by the combination of a random walk fundamental and some persistent stationary components due to the deviations from UIP, mirroring a well-known fads model used for studying the predictability of stock returns in Fama and French (1988) and Poterba and Summers (1988).

3.2 Rational Expectations Risk Premia

In this subsection, we identify the influence of the time-varying rational expectations risk premium on the return predictability. For this, we consider three different risk premium models. One is a general equilibrium sticky-price monetary model [see, e.g., Obstfeld and Rogoff (2002)]; another is a general equilibrium monetary model with an endogenous source of risk variation by Alvarez, Atkeson, and Kehoe (2009); the other is an external habit persistence model by Verdelhan (2010). Considering our objectives, we mainly derive the equilibrium (expected) excess returns from the first model as a benchmark and relegate the presentation of the latter two models to the technical and empirical appendix. Nevertheless, we discuss the key mechanisms and assumptions of all three models to identify the link between the time series behavior of the risk premium and the sign of serial dependence of excess returns. We begin with deriving the risk premium from a currency pricing setting whose definition is exactly applied to all the three models.

Assume that there are no arbitrage opportunities in an economy so that a pricing kernel exists. Let m_{t+1} be the pricing kernel for home (dollar) assets and m^*_{t+1} be the pricing kernel for foreign (euro) assets. These pricing kernels imply that any asset purchased in period t with a dollar excess return of R_{t+1} between periods t and $t+1$ and any asset purchased with a euro return of R^*_{t+1} satisfy the Euler equations, respectively,

$$1 = E_t[m_{t+1}R_{t+1}], \quad 1 = E_t[m^*_{t+1}R^*_{t+1}]. \tag{6}$$

In the complete asset markets, there exist unique m and m^* that satisfy equation (6). Although the pricing kernels are not unique in the incomplete asset markets, they can be chosen to satisfy the same equations [see, for example, Backus, Foresi, and Telmer (2001, Proposition 1)]. So, the pricing kernel for home assets, m_{t+1}, can be related to m^*_{t+1} in the following way,

$$m^*_{t+1} = \frac{m_{t+1}S_{t+1}}{S_t}. \tag{7}$$

Analogously, the log of the price of a one-period dollar-denominated bond and that of a one-period euro-denominated bond can be expressed as

$$i_t = -\log E_t[m_{t+1}], \quad i^*_t = -\log E_t[m^*_{t+1}] \tag{8}$$

where i_t (i^*_t) is the home (foreign) nominal interest rate.

Taking logs of equation (7) and then conditional expectations on a time t information set, and using the expressions for the nominal interest rates as well as the log of covered interest parity, we can define the foreign exchange risk premium, p_t, by[12]

$$p_t = E_t[s_{t+1}] - f_t = (\ln E_t[m_{t+1}] - E_t[\ln m_{t+1}]) - (\ln E_t[m^*_{t+1}] - E_t[\ln m^*_{t+1}]). \quad (9)$$

If both m_{t+1} and m^*_{t+1} follow a conditional log normal distribution, then

$$p_t = \frac{1}{2}(Var_t[\ln m_{t+1}] - Var_t[\ln m^*_{t+1}]). \quad (10)$$

We now consider the reduced-form expression for the foreign exchange risk premium premium, assuming that it follows an univariate AR(1) process,

$$p_t = (1-\varphi)\mu_\epsilon + \varphi p_{t-1} + \epsilon_t, \quad (11)$$

where $0 < \varphi < 1$ and ϵ_t is an *iid* random variable with mean zero and variance σ^2_ϵ. We follow this simplifying assumption on the process for p_t since it mimics the process of the risk premium endogenously driven from the economic models. We also present the reduced form expression for the foreign excess returns, r^e_{t+1},

$$r^e_{t+1} = s_{t+1} - f_t = fd_{t+1} + fp_{t+1} + p_t, \quad (12)$$

where constant terms are omitted for simplicity. The first two terms in the right-hand side of equation (12) are rational forecasting errors which contain two disturbance terms: fd_{t+1} is from the fundamental process and fp_{t+1} is from the risk premium process. For example, fd_{t+1} is the difference between disturbances to home and foreign money growth rates in Obstfeld and Rogoff (2002) and Alvarez, Atkeson, and Kehoe (2009) and the difference between disturbances to home and foreign consumption growth rates in Verdelhan (2010).[13] Equation (12) implies that the forecasting errors between time t and $t+1$ will be correlated with the future values of p_t, which provides an identification device for the sign of autocorrelations of excess returns.

[12] Two popular definitions for the risk premium are interchangeably used in the literature. For example, several studies define the risk premium by $p_t = E_t[s_{t+1}] - f_t$ like us [e.g., Alvarez, Atkeson, and Kehoe (2009) and Verdelhan (2010)], while the other studies define it by $p_t = f_t - E_t[s_{t+1}]$ [e.g., Froot and Frankel (1989), Engel (1996), Backus, Foresi, and Telmer (2001), and Obstfeld and Rogoff (2002)]. However, this would not change the result because the sign of p_t in equation (4) will change accordingly.

[13] Verdelhan abstract from money and inflation and focuses on real risk. So, the currency excess return in equation (12) should be interpreted accordingly.

The covariance of two consecutive ex post excess returns between $t+1$ and $t+2$ is

$$Cov(r^e_{t+1}, r^e_{t+2}) = Cov(fp_{t+1}, p_{t+1}) + \varphi Var(p_t) + \sigma_{\eta\epsilon}, \tag{13}$$

where $\sigma_{\eta\epsilon} = Cov(fd_{t+1}, \epsilon_{t+1})$ captures the relation between shocks to the fundamental value and shocks to the risk premium.[14] Although the sign of $\sigma_{\eta\epsilon}$ is model specific, the three models considered generate either a negative sign or a zero correlation. For example, Alvarez, Atkeson, and Kehoe (2009) built a model which is driven by exogenous shocks to the growth rates of money supply in each country (the difference between home and foreign money supply in their model can be interpreted as the economic fundamental, w, in our setup). They show that a change in the home money growth rate is negatively related to p_t if the money growth rate is persistent and uncorrelated with p_t if the money growth rate is iid, i.e., $\rho = 0$ in our setup [see their equation (39)]. On the other hand, Verdelhan (2010) built an external habit persistent model which is driven by exogenous iid shocks to the growth rates of consumption in each country (the difference between home and foreign consumption in their model can be interpreted as the economic fundamental, w, with $\rho = 0$, in our setup). His model also generates the negative sign of $\sigma_{\eta\epsilon}$, implying counter-cyclical risk premia. As shown in the technical and empirical appendix, the negative sign of $\sigma_{\eta\epsilon}$ is one of the key conditions that generate negative autocorrelations of excess returns in the risk premium models considered. We also find that the sign of $Cov(fp_{t+1}, p_{t+1})$ is either negative or zero in all three models. For example, the negative sign of $Cov(fp_{t+1}, p_{t+1})$ can be easily checked from equations (3) and (4) in the typical monetary models such as Obstfeld and Rogoff (2002). Alvarez, Atkeson, and Kehoe (2009) impose that exchange rate follows a random walk in their calibration, implying that the covariance is zero.

We now examine the behavior of the foreign excess return over long horizons since it also provides information on the importance of the risk premium in exchange rates [see, e.g., Fama and French (1988)]. This can be easily seen in the following covariance between r^e_{t+1} and r^e_{t+1+q} for $q \geq 1$

$$Cov(r^e_{t+1}, r^e_{t+1+q}) = \varphi^{q-1} Cov(r^e_{t+1}, r^e_{t+2}), \tag{14}$$

where φ^{q-1} summarizes the long-horizon predictability of the excess return. Equa-

[14] Campbell (1991) also derives this expression from the present value of stock prices with the expected return process (11).

tion (14) shows that the sign of $Cov(r^e_{t+1}, r^e_{t+1+q})$ is the same as that of $Cov(r^e_{t+1}, r^e_{t+2})$ and the predictability of excess returns will die out eventually. As Fama and French (1988) shows, the accumulation of these autocovariances can induce strong autocorrelations over long horizons, which may not be clearly captured over short horizons. Therefore, examining the pattern of autocorrelations of excess returns over long horizons can take a clearer picture of the importance of a mean reverting component in exchange rates, which is the deviation from UIP.

To see explicitly the relative magnitude among the three quantities in $Cov(r^e_{t+1}, r^e_{t+2})$, we need to resort on a specific economic model. As an example, we calculate the autocorrelation of the excess return based on the model by Obstfeld and Rogoff (2002) with the modification on their assumption of the money supply process, while relegating our analysis on the other models to the technical and empirical appendix. Obstfeld and Rogoff assume that money supply in each country follows a random walk, that is, $\rho = 0$, in our setup. They further assume that the shocks to the money growth rates are time invariant so that the risk premium is constant. To make it time varying, we assume that the distribution of the money growth rates follows an ARCH process so that we can assume (11) for the risk premium process in the spirit of Abel (1988) and Hodrick (1989). In particular, from equations (3)-(5) and (11), we have

$$fd_{t+1} = \eta_{t+1}, \quad fp_{t+1} = -\frac{b}{1-b\varphi}\epsilon_{t+1}. \tag{15}$$

Then, from equations (13) and (15), we have

$$Cov(fd_{t+1}, p_{t+1}) = 0, \quad Cov(fp_{t+1}, p_{t+1}) + \varphi Var(p_t) = \frac{\varphi - b}{(1-b\varphi)(1-\varphi^2)}\sigma^2_\epsilon. \tag{16}$$

Equation (16) shows that the sign of $Cov(fp_{t+1}, p_{t+1}) + \varphi Var(p_t)$ depends only on two parameter values: if the discount factor is less than φ, then it is positive; if the discount factor is greater than φ, then it is negative. Considering the higher values of the discount factor in foreign exchange markets, it is unlikely that its sign is positive.[15] Further, even if the persistence of the risk premium φ is close to or even greater than the discount factor, the relative magnitude would not significantly contribute to the overall sign of autocorrelations of the excess returns since both

[15] Engel and West (2005) show that a reasonable range for the values of the discount factor is between 0.97 and 0.99 for quarterly exchange rates, based both on empirical evidence and on implied values from a theoretical model.

parameters are bounded by one. Of course, if the risk premium follows a unit root process, then excess returns would exhibit positive autocorrelations since its effects would dominate the other components. However, this can be easily checked by looking at the behavior of the foreign excess returns over longer horizons. That is, if $\varphi = 1$, the autocovariance should not decease with q in equation (14), which can be easily checked from the data.

In sum, we predict that the rational expectations risk premium generated from all three models considered in this paper tends to produce a negative autocorrelation in excess returns. However, we emphasize that finding evidence against this prediction does not necessarily imply that the rational expectations hypothesis itself is rejected. Rather, our above analysis suggests that empirical evidence on the serial dependence pattern of excess returns may offer a criterion to judge the performance of economic models and a guidance toward a more plausible model for the data. For example, our analysis shows that the sign of $\sigma_{\eta\epsilon}$ is tightly linked to the sign of serial dependence of excess returns in the risk premium models.

3.3 Expectations Errors

We present a model of expectations errors following Frankel and Froot (1990b). There are three types of risk-neutral agents: one is portfolio managers who participate in currency transactions; the other two are fundamentalists and chartists who are merely issuing their forecasts for the manager.[16] The expectation of the portfolio managers is equal to the market expectation given by

$$E_t^m[s_{t+1}] = (1-\lambda)E_t^f[s_{t+1}] + \lambda E_t^c[s_{t+1}], \tag{17}$$

where $E_t^f[\cdot]$ is the expectation of the fundamentalists, $E_t^c[\cdot]$ is the expectation of the chartists, and $0 \leq \lambda \leq 1$. The weight λ is assumed to be exogenously given and related to the sign of serial dependence of excess returns. The expectation of the fundamentalists is assumed to be regressive,

$$E_t^f[\Delta s_{t+1}] = -\theta(s_t - \overline{s}_t) = -\theta(\varpi_t - \overline{\varpi}), \tag{18}$$

where $\theta > 0$ is the expected adjustment speed of s_t toward to $\overline{s}_t$, ϖ_t is the real exchange rate, and $\overline{s}_t$ is a long-run equilibrium exchange rate level defined by inflation differentials between home and foreign country, $\overline{s}_t = \overline{s}_0 + \ln(P_t/P_0)/(P_t^*/P_0^*)$.

[16] This is different from the noise trader model of Mark and Wu (1998), which is built on the idea of marketplace-aggregation. See, Frankel and Froot (1990b) for the detailed discussion.

This assumption is consistent with Dornbush (1976) and requires that the fundamentalists anticipate future depreciation if the current exchange rate is above the long run equilibrium level. The expectation of the chartists is assumed to be of the form of distributed lags,

$$E_t^c[\Delta s_{t+1}] = -g\Delta s_t, \tag{19}$$

where g is assumed to be greater than and equal to zero.[17] If $g = 0$, then the chartists expect that the exchange rate follows a random walk; if $g > 0$, then the chartists anticipate future depreciation of the currency toward its previous predicted level after observing an currency appreciation. Suppose that the real exchange rate, ϖ, follows an AR(1) process

$$\varpi_t = (1-\psi)\overline{\varpi} + \psi\varpi_t + \nu_t, \tag{20}$$

where $0 \leq \psi < 1$ and ν_t follows an iid Normal distribution with mean zero and variance σ_ν.

Combining equations (3)-(5) and (17)-(20), we can derive the one-period excess return between time t and $t+1$,

$$r_{t+1}^e = \frac{1}{1+\alpha\lambda g}\eta_{t+1} + \frac{1-\alpha(1-\lambda)\theta}{1+\alpha\lambda g}\nu_{t+1} - p_t^e, \tag{21}$$

where we assume $\rho = 0$ in the fundamental process (5), the first two terms in the right-hand side of equation (21) are forecasting errors (which can be defined by $fd_{t+1} + fp_{t+1}$ analogous to the previous subsection), and the expectations error p_t^e is

$$p_t^e = -\left\{\frac{-(1-\psi)+(1-\lambda)\theta(\alpha(1-\psi)+(1+\alpha\lambda g))}{1+\alpha\lambda g}\varpi_t + \frac{\alpha\lambda g + \lambda g(1+\alpha\lambda g)}{1+\alpha\lambda g}\Delta s_t\right\}. \tag{22}$$

Here, we choose the value of θ so that the fundamentalists' expectation can be rational if $\lambda = 0$, that is UIP holds. Under this condition, we find that $Cov(fd_{t+1} + fp_{t+1}, -p_{t+1}^e)$ is positive for a broad range of parameter values. For example, suppose PPP holds so that ϖ_t is constant. Then, it is obvious to show that the covariance is positive. Or, suppose $g = 0$, that is, the chartists believes that the spot exchange rate follows a random walk. Then, the covariance is positive for sufficiently large values of ψ which is consistent with the data.

[17] If $g < 0$, then the chartists have a bandwagon expectation. We rule out this case.

The key assumption for deriving equations (21) and (22) is a fixed value of λ. To relax this assumption, we need to specify how portfolio managers update their weight on the expectation of the chartists. For this, we use a speculative bubble model by Frankel and Froot (1990b) and find that excess returns exhibit positive serial dependence. Although we do not present this result since their model is built in continuous time, it can be easily verified from their equation (24) in Frankel and Froot (p. 108) and from their expressions for the rational expected exchange rate change and the market expectation in Frankel and Froot (p. 112).

We also find that the noise trader model of Mark and Wu (1998), based on De Long, Shleifer, Summers, and Waldmann (1990), tends to generate positive serial dependence in excess returns for a range of parameter values considered in their paper.[18] In sum, we predict that expectations errors generated from these models tend to generate a positive autocorrelation in excess returns.

4 Predictability of Foreign Excess Returns

Our main goal here is to examine the serial dependence pattern of excess returns, which can be used to identify a particular economic explanation presented in Section 3. Our serial dependence test is based on the variance ratio statistic, developed by Lo and McKinlay (1989). Define the population variance ratio $VR(q)$ by

$$VR(q) = \frac{Var(\sum_{i=1}^{q-1} r_{t+i}^e)}{qVar(r_t^e)} = 1 + 2\sum_{i=1}^{q-1}\left(1 - \frac{i}{q}\right)\gamma(i),$$

where q is an aggregation value and $\gamma(i) = Cov(r_t^e, r_{t+i}^e)/Var(r_t^e)$ denotes the autocorrelation of excess returns between t and $t+i$. All autocorrelations must be zero under the null hypothesis of unpredictability. So, $VR(q)$ must be equal to one if excess returns are not serially correlated. If the returns are positively autocorrelated, $VR(q)$ should be greater than one; if the returns are negatively autocorrelated, $VR(q)$ should be less than one.

Variance ratio tests are more appropriate for our objective than other serial dependence tests such as portmanteau methods because they provide direct information on the sign of the serial dependence and have good power properties (see, e.g., Lo and McKinlay (1989) and Moon and Velasco (2010a)).[19] However,

[18] We do not provide our detailed results since its calculation is straightforward. See, for example, equations (12), (22), and (29) in Mark and Wu (1998).

[19] Serial dependence tests based on variance ratios have been used for testing a random walk

if one intends to improve efficiency based on a finer sampling frequency than the maturity, most distributional results on the variance ratio tests are not directly applicable due to the presence of moving average components in excess returns. To take into account this, we follow Moon and Velasco (2010a) where we proposed to split the entire sample into k subsamples so that excess returns are not serially correlated within a subsample under the null hypothesis but are correlated across subsamples.[20] For aggregating information across subsamples, we use the rank-based median variance ratio test which calculates the variance ratio using ranks of returns and obtains the median value of those ratios across subsamples. We relegate the details of the procedure to Appendix B.

We perform the variance ratio test on one-, three-, six- and twelve-month excess returns against the USD using the same weekly data as in Section 2. The tests are conducted at the 5% significant level for both one-sided and two-sided alternatives. Whenever the two-sided test rejects the null hypothesis, the rejection also occurs at the right-tail. Hence, we only report the results for right-tail alternatives for the sake of simplicity. We use aggregation values q of 2, 12, 24, 36, 48, and 60 months relative to a one-month base period for the one-month foreign excess returns, 2, 4, 8, 12, 16, and 20 quarters relative to a one-quarter base period for the three-month returns, 2, 4, 6, 8, and 10 biyears relative to a one-biyear base period for the six-month returns, and 2, 3, 4, and 5 years relative to a one-year base period for the one-year returns. We set the range of aggregation values for each maturity so that the maximum value of q is 5 years. For each maturity, examining statistics at $q = 2$ may be enough if the main objective is to judge the rejection or non-rejection of the UIP hypothesis. In addition, the variance ratio test provides further information helping us to understand the reasons of the rejections against the null hypothesis as discussed in detail below.

of asset prices, in particular for identifying a mean reverting component in asset prices. See, for example, Liu and He (1991) for spot exchange rates, Campbell and Mankiw (1987) and Cochrane (1988) for output, and Poterba and Summers (1988) and Lo and MacKinlay (1988) for stock prices.

[20]To avoid possible confusion, we clarify that the terminology of 'subsample' used in this paragraph as well as in Appendix B is different from that defined using sample periods in the other parts of the paper.

4.1 Results on Serial Dependence of Excess Returns in the Entire Sample

Table V reports the estimated rank-based median variance ratios defined in equation (29) in Appendix B and their right-tail p-values from the empirical distribution of 5,000 bootstrap samples generated using the parametric bootstrap method. For each maturity, the variance ratios are estimated for individual currency as well as for the two portfolios whose construction is explained in Section 2. In particular, the construction of portfolios may alleviate the effects of a country-specific or idiosyncratic noise which makes it difficult to detect the presence of systematic predictable components.[21] Panel A provides the results obtained from the entire sample. Overall, we find that excess returns significantly exhibit positive serial dependence for all maturities, supporting the expectations errors explanation presented in Section 3.

The variance ratio test rejects the null hypothesis for both equally- and turnover-weighted averages of excess returns. The estimated median variance ratios are greater than one and the rejections occur at the right-tail for all aggregation values q ranged from 2 to 60 months. For example, for the one-month turnover-weighted average of excess returns, the estimated median variance ratios, associated with the aggregation values $q = 2$, 12, 24, 36, 48, and 60 months, are 1.1, 1.5, 2.0, 2.2, 2.3, and 2.3, respectively, and their right-tail p-values are all less than 1%. The results for individual currencies are also similar. The test rejects the null at the right tail for all currencies in our sample for the one- and six-month maturity. For the three-month maturity, an exception is AUD. For the one-year maturity, an exception is BRP.[22] The rejection only occurs at $q = 2$ years for JPY.

Regarding the rejections of the UIP hypothesis, these results are consistent with those from the predictive regression using the forward discount as a predictor in Section 2. However, the findings on the strong positive serial dependence of excess returns provide additional information that supports the models of expectations errors. The pattern of the t-statistics of estimated variance ratios over q can be further used to identify a particular explanation. Figure 2 shows how the t-statistics of the estimated variance ratios, defined in equation (28) in Appendix

[21] See, e.g., Fama and French (1988) and Lo and MacKinlay (1988) who used the stock price data.

[22] For the one-year maturity, we only report the results using six currencies such as GDM, BRP, JPY, CAD, SWF, and DAK because of the data availability.

B, evolve with respect to q. The line with asterisks is the locus of t-statistics with q and the two dashed lines correspond to 5 and 95% quantiles from the simulated bootstrap empirical null distribution. In general, the t-statistics tend to increase up to 2 to 3 years and then decrease for all excess returns except CAD and AUD. This hump-shaped pattern of the t-statistics with q is consistent with the behavior of a persistent stationary component in the excess return.[23] As shown in the next subsection, the separation of the entire sample, however, notices that the true cause for the predictability of excess returns seems to be masked in the entire sample.

4.2 Results on Subsamples

As in Section 2, we decompose the entire sample into the two subsamples to further investigate the main cause of the predictability of excess returns. Panel B reports the results from subsample E. In contrast to the results from the entire sample, the predictability of excess returns becomes very weak. Most estimated variance ratios are statistically indistinguishable from one. For example, the variance ratio test does not reject the null for both portfolios except for the turnover-weighted average for the one month-maturity: the rejections occur at the right-tail at the 5% level for some aggregation values q. The pattern of the results are also similar for individual currencies. These results suggest that strong predictability detected in the entire sample may be significantly influenced by the particular sample period

To further see this influence, we conduct the variance ratio test using subsample O and find that excess returns are much strongly predictable [see Table VI]. For example, estimated variance ratios are increasing monotonically with q and the maximum ratio across q is in general greater than 5.[24] Further, the right-tail p-values are close to zero for all aggregation values, suggesting that excess returns display a very strong positive serial dependence in the 1980s. One exception is

[23] Note that, when a stationary component in the excess return is highly persistent, it is hard to detect its departure from the random walk for smaller q. However, the first-order autocorrelation of the increments of the stationary component in excess returns grows as the aggregation value q increases and thus behaves less like random walk increments, leading to increments in the power of the test. On the other hand, the random walk component dominates the effects of the stationary component for larger q, leading to the decline of the power of the test beyond some q.

[24] Bekaert and Hodrick (1992) also show that the implied variance ratios, based on the VAR estimation, are increasing monotonically to above 2.9 using the data of GDM, BRP, and JPY between 1981 and 1989.

CAD: it is not statistically significant. For the robustness, we also construct several subsamples with the same size of 8-years using observations between 1990 and 2009 and conduct the same test. In general, we find that excess returns appear to be unpredictable or the predictability becomes very weak in those samples.

We now discuss the rejection patterns over q from the two subsamples. In contrast to the results from the entire sample, we do not find the hump-shaped pattern of the t-statistics with q in both subsamples. That is, in subsample E, we do not find such pattern in all the one-month excess returns except for JPY [see Figure 3]. This is mainly because the UIP hypothesis tends to hold in this sample. Neither does the pattern of t-statistics look hump-shaped in subsample O [see Figure 4]. The t-statistics are increasing up to about 24-36 months of aggregation values and then decreasing up to about 36-48 months. But they increase again beyond about $q = 48$ months, implying that the rejections of the null hypothesis become stronger as q further increases. A simple model with a persistent stationary component in excess returns does not seem to generate such rejection patterns from both subsamples. We provide a possible explanation for these patterns in the next section, based on the changes in λ in the expectations errors model by Frankel and Froot (1990b).

5 The Speculative Behavior of the US bilateral Rates in 1980-87

Although the evidence on the serial dependence of excess returns against USD supports the expectations errors explanation presented in Section 3, it leaves unanswered the question of why deviations from UIP appear much strongly in the sample period of 1980-87 relative to the other time periods. In this section, we reconsider the model of a speculative bubble by Frankel and Froot (1990a, 1990b), relate this new evidence to 'changes in forecasting techniques', and provide several pieces of empirical evidence which support it.

5.1 Changes in Forecasting Techniques

Recall that the market expectation is defined by the weighted average of the expectations of the two agents in the model of a speculative bubble presented in Section 3. If the market solely considers the forecasts of fundamentalists then the UIP hypothesis holds. Otherwise, the hypothesis does not hold and the deviation becomes stronger as the weight of the expectations of chartists increases. So, the abnormal behavior of USD bilateral rates in the 1980s may be related to significant changes in the weight of the expectations of chartists. In other words, it is possible to observe such behavior if, for some reasons, the market had gradually shifted its attention from the opinions of fundamentalists to chartists in the 1980s but not in the other sample periods.

Frankel and Froot (1990a, 1990b) provided a reason for the significant increase in the weight of chartists over 1981-85. Hence, we only present their key idea: suppose the market expectation had been formed, based exclusively on the expectations of the fundamentalists until 1980. Because the US interest rate had been persistently higher than those of other countries in early 1980s, the US dollar was expected to be depreciated. However, as the value of the dollar increased since 1980 as shown in Figure 5, the forecasts of (relatively more) future dollar depreciation by fundamentalists turned out to be wrong month after month. As a result, the market started to pay less attention to the opinion of fundamentalists but more to that of chartists.[25] This reflects market's Bayesian response to the performance of the fundamentalists in terms of forecasting future exchange rates. Frankel and Froot also explained why weight shifted back to the fundamentalists from the chartists and the collapse of speculative bubbles after 1985 based on the unsustainability of large US current account deficits in the long run (see, also, Krugman (1985)).

5.2 Evidence on Changes in Forecasting Techniques

In the previous subsection, we have related the very strong positive serial dependence of excess returns against USD in the 1980s to the changes in forecasting techniques and provided a reason why weight was gradually shifted from (to) the fundamentalists to (from) the chartists. This speculative behavior explanation on the path of USD in the 1980s fits well with long swings of USD bilateral rates with a large appreciation in 1981-85 and then a large depreciation in 1985-87 [see Figure 5]. Interestingly, we did not observe these swings in other bilateral rates. For example, the time series of JPY bilateral rates against GDM and BRP do not

[25] These changes in the weighted-average forecasts of future spot exchange rates increased demand for dollar and thus the prices of the dollar, which may cause the speculative bubbles in the early 1980s.

exhibit such pattern as shown in Figure 6. Therefore, if our argument is consistent with USD bilateral rates, then we should not expect similar changes in forecasting techniques for other bilateral rates and thus observe different serial dependence patterns of excess returns.

A simple method to investigate this is to conduct the same test using other bilateral rates. For this, we first eliminate the effects of common components in USD bilateral rates using triangular arbitrage in the absence of transaction costs, that is replacing USD with other country's currency as a base currency. Then, we compare the results of the serial dependence tests between these samples: one with USD as a base currency and the other with a currency other than USD. We choose JPY, BRP, or GDM as the base currency since they are major currencies in foreign exchange markets. For the sake of saving the space we only present the results using JPY and relegate those using BRP and GDP to the technical and empirical appendix.[26] Compared to Figure 4 which shows the pattern of t-statistics for the excess returns against USD, we find that excess returns exhibit very different serial dependence patterns. All excess returns against JPY except for CAD are unpredictable over the entire range of q as shown in Figure 7, supporting the argument for changes in forecasting techniques in the 1980s.[27]

Frankel and Froot (1990a) also provided evidence which supports their hypothesis, directly looking at the survey results about the forecasting techniques of foreign exchange forecasting service firms. The survey had been conducted by Euromoney magazine between 1978 and 1988. According to their Table II (p. 184), most forecasting service firms surveyed relied exclusively on economic fundamentals before 1980 for predicting future exchange rates. But they had reversed their position by 1984: none of them were relying exclusively on economic fundamentals in 1984. Since then, the number of forecasting firms that relied on economic fundamentals had increased again.

5.3 Survey Data on the Investors' Expectations

In this subsection we search for a measure of the expectation whose behavior is consistent with equation (17) and thus with our empirical findings on the

[26] The results are quite similar using any one of these currencies as a base currency.

[27] The serial dependence pattern of CAD excess returns against either JPY, BRP, or GDM is very similar to USD excess returns. This may be because CAD behaves very closely with USD.

predictability of excess returns. One possible candidate is survey data on the investors' expectations of future exchange rates which became popular in the literature since first used by Frankel and Froot (1987). Hence, we examine whether or not the time series behavior of the survey data is compatible with that of the market expectation.

The excess return can be decomposed into two parts using the survey expectation:

$$s_{t+k} - f_{t|k} = s_{t+k} - s^m_{t+k} + s^m_{t+k} - f_{t|k}, \tag{23}$$

where s^m_{t+k} is the median survey expectation of the k-period ahead future spot exchange rate. Our analysis below relies on the following assumptions: there is a single market expectation and the measurement error defined by the difference between the market expectation and the survey data is iid. Then, $s_{t+k} - s^m_{t+k}$ represents the sum of rational forecasting errors, expectations errors, and the measurement error, and $s^m_{t+k} - f_{t|k}$ represents the sum of the market risk premium and the measurement error.

Froot and Frankel (1989) used survey data in 1981-85 when the UIP hypothesis was decisively rejected and provided empirical evidence that the deviations are mainly attributable to $s_{t+k} - s^m_{t+k}$ and none to the risk premium, suggesting a close relation between expectations errors and the predictability of excess returns. On the other hand, the sample period for our survey data set is 1988-08 so that we can investigate the behavior of the survey data when the UIP hypothesis tends to hold. Note that $s_{t+k} - s^m_{t+k}$ should not be predictable by anything in the time t information set under the UIP hypothesis if the survey data are consistent with the market expectation. Further, $s^m_{t+k} - f_{t|k}$ should be equal to the measurement error.

We consider two econometric specifications for studying this. The first specification is the OLS regression of the difference between the future spot rate and the survey expectation on the forward discount,

$$s^m_{t+k} - s_{t+k} = a^a_k + \beta^a_k(f_{t|k} - s_t) + u^a_{t+k}, \tag{24}$$

where u^a_{t+k} represents the regression error. If the UIP hypothesis holds, then $\beta^a_k = 0$. If it does not hold and expectations errors are the main cause, then $\beta^a_k \neq 0$. The second specification analyzes the effects of $s^m_{t+k} - f_{t|k}$ on the deviations from UIP. Specifically, we run the following regression of $s^m_{t+k} - s_t$ on the forward discount

$$s^m_{t+k} - s_t = a^b_k + \beta^b_k (f_{t|k} - s_t) + u^b_t, \tag{25}$$

where u^b_t represents a measurement error. $\beta^b_k = 1$ implies that there are no significant effects of the risk premium on the deviations. So, if the survey expectation proxies well the market expectation, we predict that $\beta^a_k = 0$ and $\beta^b_k = 1$ under UIP; on the other hand, we predict that $\beta^a_k \neq 0$ and $\beta^b_k = 1$.

Table VII reports the results from regressions of (24) and (25) using the survey forecasts data obtained from FX4casts. Our sample of the survey data includes three-, six-, and twelve-month prices of USD against GDM, BRP, JPY, CAD, SWF, DAK, FRF, ITL, AUD, and SWK. The data cover the sample periods of 1988:1-2008:12 and the sample frequency is monthly. Accordingly, monthly spot and forward rates from the corresponding sample periods are used. We choose both spot and forward rates at the end of each month. Panel A provides the reproduction of the results from the standard regression of the excess return on the forward discount and Panels B and C report the results from regressions of (24) and (25), respectively. All three regressions are estimated for an individual currency and for a single β_k, β^a_k, and β^b_k using a pooled sample of all currencies and periods of time as well as using the two portfolios. Overall, we find that the behavior of the survey expectation is consistent with the market expectation.

We begin our analysis with the case where the UIP hypothesis does hold. As consistent with the results in Section 2, we do not reject the UIP hypothesis at the 5% level in the pooled sample. Neither do we reject the hypothesis of $\beta^b_k = 1$. But we marginally reject the hypothesis of $\beta^a_k = 0$ for the six- and twelve-month maturities. Regarding individual currencies, there are seven out of ten currencies for which the UIP hypothesis holds. For these currencies, $\widehat{\beta}^a_k$ are not statistically significant from zero. One exception is DAK for the six- and twelve-month maturities: they are significant from zero at the 5% level. Further, $\widehat{\beta}^b_k$ are not significantly different from one for the three-month maturity. However, the results for the six- and twelve-month maturities are mixed in that $\widehat{\beta}^b_k$ are statistically significant at the 5% level for DAK, ITL, AUD, and SWK. We provide a possible explanation for this mixed evidence as well as the marginal rejections of $\beta^a_k = 0$ in the pooled sample in the next subsection.

We now discuss the case where the UIP hypothesis does not hold. The UIP hypothesis is marginally rejected at the 5% level for both portfolios except for the turnover-weighted average of one-month excess returns. Regarding individual currencies, the rejections occur only for four currencies.[28] It occurs for JPY and SWF for all maturities, for the CAD for the three-month, and for GDM for the six-month and one-year. For these portfolios and individual currencies, $\widehat{\beta}_k^a$ is significantly different from zero. An exception is the JPY for the three-month maturity. On the other hand, $\widehat{\beta}_k^b$ is statistically indistinguishable from one for all maturities. Exceptions are JPY for the three-month maturity and GDM for the six-month maturity.[29]

We summarize the results from the two cases above (both where the UIP hypothesis holds and where it does not hold): (i) in general, we do not reject the hypothesis of $\beta_k^b = 1$ in both cases, consistent with the findings by Froot and Frankel as well as with ours on the serial dependence of excess returns in the previous section; (ii) in the case where the UIP hypothesis does not hold, we reject the hypothesis of $\beta_k^a = 0$ consistent again with previous evidence; (iii) $s_{t+k}^m - s_{t+k}$ is not predictable using the forward discount when the UIP hypothesis holds; (iv) the results in (i)-(iii) suggest that the survey expectation behaves consistently with the market expectation defined in equation (17).

5.4 Measurement Error

The key assumption made in regressions (24) and (25) is that measurement error is iid. To investigate the validity of this assumption, we discuss several sources of measurement error that may generate a systematic bias in the regressions and attempt to quantify the magnitude of the bias. First, the survey data were not collected at the same moment as the contemporaneous spot rate was recorded. To see how severely this mismatch affects our results, we conduct the same regressions

[28] Compared to the results from the previous sections, there is a small discrepancy regarding the rejections of the UIP hypothesis. This may be due to using different sample periods and/or different sample frequency. Nevertheless, we confirm that the deviations from UIP become significantly weakened after 1980s.

[29] Our results are consistent with Bacchetta, Mertens, and Wincoop (2008) who also used the same data between August 1986 and July 2005: they found a strong relation between the rejections of the UIP hypothesis and the expectations errors. One apparent difference is that the rejections against the UIP hypothesis in their sample appear somewhat stronger than ours. One reason may be the use of different sample period. In particular, their sample includes part of observations from the period of 1980-87 in which the strongest rejections are occurred.

(24)-(25) by changing the date of selecting forward and spot exchange rates from Monday to Friday of the last week of each month.[30] The results across different selection dates are broadly consistent.[31] Similarly, Bekaert and Hodrick (1993) showed that their findings were not much affected by measurement error due to the failure of taking correctly into account of the delivery structure of the forward contract. One reason for this insensitivity may be related to the strong persistence in the forward discount.

Second, the survey data used in the present paper were constructed from a sample of market participants. Hence, there may exist measurement error if the sample does not capture the opinions of all participants in the market. However, this may not be large enough to overturn the above results since the sample consists of large numbers of wealthy investors and financial institutions who actively participate in the market according to FX4casts. Further, many studies consistently find evidence on the relation between expectation errors and excess returns using survey data in various financial markets and over different sample periods, suggesting that our results may not be entirely attributable to a sample selection bias [see, e.g., Froot and Frankel (1989), Bacchetta, Mertens, and Wincoop (2008), and references therein].

In addition, there may exist other sources of measurement error which we have not been able to take into account. Therefore, our objective below is to identify a systematic pattern between measurement error and the predictor in the regressions without providing any further explanation for the existence of those sources. Suppose that the UIP hypothesis holds. Then, the difference between the estimate and the true value in regressions (24)-(25) can be interpreted as an estimation error if there are no biases. However, if there were a systematic relation between measurement error and the predictor for some reasons, this would generate either a negative or positive bias in the regressions and thus we would consistently observe a tendency of either downward or upward biases of both $\widehat{\beta}_k^a$ and $\widehat{\beta}_k^b$. Since the UIP hypothesis tends to hold for many currencies in our sample during 1988-08, we use those regressions to identify the direction of the bias. That is, if there is a positive bias, then we should consistently observe $\widehat{\beta}_k^a > 0$ and $\widehat{\beta}_k^b > 1$ under

[30] According to FX4casts, the survey was e-mailed on the last Friday of the month. Most of the forecasts come in on Monday of that week and about 10% on Tuesday. About 10% come in the previous Friday.

[31] See the technical and empirical appendix.

the UIP hypothesis; otherwise, we should observe $\widehat{\beta}_k^a < 0$ and $\widehat{\beta}_k^b < 1$ if there is a negative bias.

We find a symptom of a positive bias in those regressions. For all the currencies and maturities for which the UIP hypothesis holds, we find that $\widehat{\beta}_k^b > 1$. On the other hand, $\widehat{\beta}_k^a > 0$ for most cases although they are not statistically significant. These results suggest that there may exist a positive bias. We now attempt to informally estimate to which extent the magnitude of this bias might affect our conclusions. Under the assumption of zero estimation errors on average, we calculate the average of the difference between $\widehat{\beta}_k^b$ and the true value, finding that it is about 0.6. None of them become statistically significant at the 5% level under the UIP hypothesis after the correction of this bias by subtracting the average from both $\widehat{\beta}_k^b$ and $\widehat{\beta}_k^a$. This may explain the mixed evidence for $\widehat{\beta}_k^b$ above. We did the same correction for the case where there are deviations from UIP. However, this does not overturn the significant relation between the expectation error and the forward discount, although the evidence appears somewhat weaker. In sum, we find that measurement error cannot entirely drive our findings, although the positive bias may affect the results on regressions (24)-(25).

6 Conclusions

This paper investigates the predictability of excess returns in foreign exchange markets. We first find that both the statistically significant positive serial dependence of foreign excess returns in the entire sample and the very weak (mostly insignificant) positive serial dependence in Subsample E are consistent with the predictions of the expectations errors explanation. Our analysis also explains why previous studies have consistently documented strong predictability of excess returns: the effects of a particular sample period are masked. We, then, link this new evidence with the changes in forecasting techniques in foreign exchange markets in the 1980s. Finally, we show that the behavior of the survey expectation is compatible with the market expectation over both subsample periods. That is, the difference between the survey expectation and the future exchange rate is not predictable when the UIP hypothesis holds but the expectations errors are closely related to the predictability of excess returns when the UIP hypothesis does not hold. Overall, we find that the expectations errors explanation is consistent with the behavior of excess returns over the entire sample period.

Our results appear robust to the selection of the subsample period, choice of predictors, and the presence of highly persistent regressors. We show that the very strong predictability only appears in the 1980s based on the parameter stability test using rolling samples. Further, we obtain very similar results on the predictability of excess returns either using past returns through the variance ratio test or using the forward discount through the typical regression test. Furthermore, the results from the conditional test by Jasson and Moreira confirm that a statistical phenomenon cannot be the main cause of our findings.

Our new empirical findings on the predictability of excess returns against the USD deserve further attentions. The first concern is related to the dynamic behavior of exchange rates in response to monetary policy shocks. An example is the delayed overshooting phenomenon, first documented by Eichenbaum and Evans (1995), which states the dynamic response of currency depreciation (or excess returns) in response to the change in the interest rate difference between the two countries. We leave as future study for investigating if this phenomenon is also significantly affected by the influential subsample period. The second concern is related to the potential effects of peso problems on the deviations from UIP in 1980-87.[32] Although peso problems can be a natural candidate for explaining the new empirical evidence, we also leave this possibility as future study.

[32]For example, several studies such as Engel and Hamilton (1990), Kaminsky (1993), and Evans and Lewis (1995) related the large swings of the US dollar exchange rates in this time period to peso problems.

Appendix A: The Parameter Stability Test

We describe the computation of bootstrap estimates of the uniform test bands for the stability test under the null of parameter stability $\beta_k(j) = 0$ for all subsample j. The justification of this procedure is straightforward under the assumptions that the regressor is stationary and the error term is iid in equation (1). The procedure follows.

1. Run 8-year rolling regressions based on (1) and obtain residuals only for the last one-year observations in each rolling regression.
2. Obtain a bootstrap resample of these residuals using the block bootstrap method with block size of k.
3. Construct the bootstrap dependent variable using regression (1) with resampled residuals from step 2, the original regressors, and the true parameter values under the null of parameter stability.
4. Run 8-year rolling regressions on the resampled data from step 3 and store the estimated slope coefficients as well as the corresponding t-statistics. Then obtain the minimum and maximum of these sequences of the estimated regression coefficients and the corresponding t-statistics for this resample.
5. Repeat steps 2-4 n times [In the present paper, we set n=10,000.].
6. Calculate 5 and 95 % quantiles from the simulated bootstrap empirical distribution of the minimum and maximum of the estimated rolling regression slope parameters as well as the t-statistics.

Appendix B. The Rank-based Median Variance Ratio Test

We briefly describe the procedure of the rank-based median variance ratio test conducted in Section 4. To take into account the presence of moving average components in excess returns due to using finer sampling frequency, we divide the entire sample $S = \{r_1^e, r_2^e, \ldots, r_T^e\}$ into k subsamples in the following way: $S_j = \{r_j^e, r_{k+j}^e, \ldots, r_{T-k+j}^e\}$, for $j = 1, \cdots, k$. Let the sample variance ratio, $\widehat{VR}_j(q)$, for subsample j, be defined by

$$\widehat{VR}_j(q) = \frac{(qm_k(q))^{-1}\sum_{t=q}^{\frac{T}{k}}\left(r^e_{k(t-1)+j}+\cdots+r^e_{k(t-q)+j}-q\widehat{\mu}_j\right)^2}{(m_k(1))^{-1}\sum_{t=1}^{\frac{T}{k}}\left(r^e_{k(t)+j}-\widehat{\mu}_j\right)^2}, \tag{26}$$

where $\widehat{\mu}_j$ is the mean of S_j and $m_k(q) = (\frac{T}{k} - q + 1)(1 - \frac{kq}{T})$ corrects the biases in the variance estimator.

In order to summarize information across the subsamples, we use the rank-based median variance ratio test from Moon and Velasco (2010a) which pools variance estimates from subsamples of uncorrelated returns under the null. The following is the brief description of the procedures. First, the conventional variance ratio defined in equation (26) is modified using ranks. Let $\xi_j(r^e_t)$ be the rank of r^e_t among the $r^e_j, \ldots, r^e_{T-k+j}$ belonging to S_j. Then, following Wright (2000), a simple linear transformation of the rank $\xi_j(r^e_t)$ is defined by

$$\xi_t = \left(\xi_j(r^e_t) - \frac{T+1}{2}\right) \div \sqrt{\frac{(T-1)(T+1)}{12}},$$

where ξ_t is standardized with sample mean 0 and sample variance 1. The rank-based variance ratio test simply substitutes ξ_t for r^e_t in the definition of q-period sample variance ratio statistic in equation (26):

$$\widehat{R}_j(q) = (qm_k(q))^{-1}\sum_{t=q}^{T/k}(\xi_{k(t-1)+j}+\cdots+\xi_{k(t-q)+j})^2, \quad j = 1, \ldots, k., \tag{27}$$

where the corresponding denominator in $\widehat{VR}_j(q)$ is omitted from the definition of $\widehat{R}_j(q)$ since it is equal to 1 by construction. Second, we obtain the median value of the standardized deviation of the variance ratio statistic from one

$$\widehat{MV}_k(q) = median_{(j=1,\ldots,k)}\frac{\sqrt{T/k}(\widehat{R}_j(q)-1)}{\sqrt{2(q-1)(2q-1)/3q}}. \tag{28}$$

Since both $\sqrt{T/k}$ and the variance of $\widehat{R}_j(q)$ for a given q are constant, we can also use the following statistic without loss of generality,

$$\widehat{MR}_k(q) = median_{(j=1,\ldots,k)}\widehat{R}_j(q). \tag{29}$$

Third, we obtain critical values using a parametric bootstrap procedure in Moon and Velasco (2010a), which leads to better approximations to the actual joint distribution of the variance ratio deviations from one than methods based on asymptotic results.

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Table I Sample Periods

	Starting Date			Ending Date
	One-Month	Three- and six-Month	One-Year	
German deutschmark (GDM)	Jan. 2, 1975	Jan. 2, 1975	Jan. 2, 1975	Dec. 15, 2009
British pound (BRP)	Jan. 2, 1975	Jan. 2, 1975	Jan. 2, 1975	Dec. 15, 2009
Japanese yen (JPY)	Jan. 2, 1975	Jan. 2, 1975	Jan. 2, 1975	Dec. 15, 2009
Canadian dollar (CAD)	Jan. 2, 1975	Jan. 2, 1975	Jan. 2, 1975	Dec. 15, 2009
Swiss franc (SWF)	Jan. 2, 1975	Jan. 2, 1975	Jan. 2, 1975	Dec. 15, 2009
Danish krone (DAK)	July 11, 1979	Jan. 2, 1977	Jan. 2, 1977	Dec. 15, 2009
French franc (FRF)	Jan. 2, 1975	Jan. 2, 1980	Jan. 2, 1980	Dec. 31, 1998
Italian lira (ITL)	Jan. 2, 1975	Jan. 2, 1975	Jan. 2, 1980	Dec. 31, 1998
Dutch guilder (DUG)	Jan. 2, 1980	Jan. 2, 1980	Jan. 2, 1980	Dec. 31, 1998
Australian dollar (AUD)	Oct. 9, 1986	Oct. 9, 1986	Oct. 9, 1986	Dec. 15, 2009
Swedish krona (SWK)	Jan. 2, 1988	Jan. 2, 1988	Jan. 2, 1988	Dec. 15, 2009

The data are simultaneously collected from London close bid and ask prices and obtained from the database of Global Insight. Our sample includes spot prices of USD against the German deutschmark (GDM), the British pound (BRP), the Japanese yen (JPY), the Canadian dollar (CAD), the Swiss franc (SWF), the Danish krone (DAK), the French franc (FRF), the Italian lira (ITL), the Dutch guilder (DUG), and the Australian dollar (AUD) as well as one-, three- , six- and twelve-month prices (forward exchange rates) of USD. Mid prices are used for our empirical study. We use ask prices for the currencies that belong to EMU because of the data availability. We use weekly observations from 1975:1 to 2009:12 and select Wednesday's closing price for forming our sample. If the following Wednesday is missing, then Thursday's price is used (or Tuesday's if Thursday's is also missing).

Table II Summary Statistics

Table IIa One-Month Maturity ($k = 4$)

	GDM	BRP	JPY	CAD	SWF	DAK	FRF	ITL	DUG	AUD
	Panel A Foreign Excess Returns: $s_{t+k} - f_{t\|k}$									
mean	0.16	1.46	0.17	0.83	-0.67	1.72	0.98	2.11	-1.52	4.54
std	38.72	38.25	41.58	22.24	43.08	39.77	37.36	36.93	40.47	42.00
auto(1)	0.08	0.09	0.12	0.00	0.07	0.09	0.10	0.16	0.14	0.02
auto(3)	0.05	-0.01	0.09	0.05	0.02	0.08	0.08	0.07	0.03	0.00
	Panel B forward discount: $f_{t\|k} - s_t$									
mean	1.59	-2.44	3.35	-0.99	3.35	-1.51	-1.99	-6.03	1.56	-2.92
std	3.06	2.90	3.04	1.94	3.53	3.79	3.91	5.49	3.52	2.75
auto(1)	0.96	0.93	0.92	0.92	0.96	0.87	0.68	0.61	0.92	0.95
auto(3)	0.87	0.79	0.82	0.76	0.88	0.68	0.50	0.53	0.82	0.89
T	455	455	455	455	455	396	312	312	247	302

Table IIb Three-Month Maturity ($k = 13$)

	GDM	BRP	JPY	CAD	SWF	DAK	FRF	ITL	DUG	AUD
	Panel A Foreign Excess Returns: $s_{t+k} - f_{t\|k}$									
mean	0.11	1.03	0.40	0.63	-0.41	2.03	0.10	1.98	-1.53	4.20
std	22.95	21.86	24.23	13.14	25.27	22.78	23.69	23.50	24.83	24.22
auto(1)	0.13	0.20	0.14	0.12	0.11	0.16	0.21	0.11	0.22	0.17
auto(3)	0.20	0.07	0.14	0.00	0.15	0.20	0.14	0.05	0.19	-0.17
	Panel B forward discount: $f_{t\|k} - s_t$									
mean	1.63	-2.01	3.12	-0.79	3.10	-1.54	-1.84	-5.90	1.56	-2.68
std	2.89	2.61	2.76	1.63	3.26	3.35	3.41	5.04	3.22	2.60
auto(1)	0.83	0.76	0.82	0.81	0.90	0.78	0.59	0.56	0.84	0.88
auto(3)	0.80	0.47	0.59	0.64	0.81	0.59	0.49	0.32	0.81	0.66
T	140	140	140	140	140	132	76	96	76	93

Table IIc Six-Month Maturity ($k = 26$)

	GDM	BRP	JPY	CAD	SWF	DAK	FRF	ITL	DUG	AUD
	Panel A Foreign Excess Returns: $s_{t+k} - f_{t\|k}$									
mean	-0.08	0.86	0.13	0.41	-0.56	1.67	-0.30	1.42	-1.86	2.77
std	17.88	16.98	18.70	10.40	18.35	17.73	19.30	17.91	20.31	16.55
auto(1)	0.08	0.10	0.11	-0.12	0.07	0.11	0.10	0.07	0.10	0.25
auto(3)	-0.07	0.06	-0.12	0.18	-0.06	0.00	-0.06	-0.15	-0.12	0.01
	Panel B forward discount: $f_{t\|k} - s_t$									
mean	1.62	-1.96	3.20	-0.74	3.04	-1.42	-1.89	-5.71	1.50	-2.68
std	2.79	2.50	2.70	1.50	3.08	3.07	3.15	4.18	3.11	2.66
auto(1)	0.81	0.60	0.65	0.76	0.81	0.67	0.45	0.33	0.79	0.76
auto(3)	0.55	0.30	0.19	0.38	0.61	0.32	0.12	-0.01	0.59	0.46
T	69	69	69	69	69	65	37	47	37	45

Table IId One-Year Maturity ($k = 52$)

	GDM	BRP	JPY	CAD	SWF	DAK	FRF	ITL	DUG	AUD
	Panel A Foreign Excess Returns: $s_{t+k} - f_{t\|k}$									
mean	-0.34	0.67	0.15	0.02	-0.89	1.50	-0.18	0.70	-1.78	3.60
std	12.98	13.61	13.22	7.41	13.15	12.95	14.09	13.25	14.42	12.41
auto(1)	0.28	0.01	0.17	0.09	0.24	0.28	0.40	0.41	0.35	0.11
auto(3)	-0.10	0.00	-0.20	0.03	-0.16	-0.13	-0.09	-0.08	-0.10	0.03
	Panel B forward discount: $f_{t\|k} - s_t$									
mean	1.67	-1.95	3.23	-0.66	3.12	-1.49	-2.02	-4.99	1.44	-2.53
std	2.60	2.22	2.49	1.42	2.89	2.80	2.75	2.58	2.77	2.54
auto(1)	0.75	0.39	0.57	0.61	0.74	0.55	0.35	0.33	0.75	0.52
auto(3)	0.25	0.03	-0.07	0.18	0.30	0.18	-0.22	-0.11	0.13	0.42
T	34	34	34	34	34	32	18	18	18	22

Log foreign excess returns against the US dollar, $s_{t+k} - f_{t|k}$, are measured over a maturity of k weeks, which includes 4 weeks for the one-month maturity, 13 weeks for the three-month, 26 weeks for the six-month, and 52 weeks for the one-year. All log excess returns are annualized in the following way: $(5200/k)(s_{t+k} - f_{t|k})$. Summary statistics are produced using non-overlapping observations in the first subsample out of k-subsamples for each maturity k. 'std' represents the standard deviation of the series, 'auto(n)' denotes n-order autocorrelation, and 'T' is sample size. See the note in Table 1 for the data description.

Table III Results of the predictive regression test

	EPort	WPort	Pool	GDM	BRP	JPY	CAD	SWF	DAK	FRF	ITL	DUG	AUD
Panel A Entire Sample: 1975-2009													
$\widehat{\beta}_4$	-2.11^{**}	-2.10^{**}	-1.06^{**}	-1.62^{**}	-1.93^{**}	-2.50^{**}	-1.62^{**}	-2.19^{**}	-1.28^{**}	-0.58	-0.74	-2.33^{**}	-2.03^{**}
s.e.	0.55	0.57	0.23	0.63	0.67	0.62	0.41	0.59	0.44	0.51	0.38	0.75	0.66
R^2	0.03	0.03	0.02	0.02	0.02	0.04	0.02	0.03	0.02	0.01	0.01	0.04	0.03
$\widehat{\beta}_{13}$	-2.09^{**}	-2.18^{**}	-1.04^{**}	-1.61^{*}	-1.96^{*}	-3.21^{**}	-1.29^{**}	-1.95^{**}	-1.36^{**}	-0.77	-0.58	-2.27^{**}	-1.78^{*}
s.e.	0.63	0.62	0.27	0.65	0.82	0.57	0.48	0.64	0.45	0.82	0.45	0.87	0.80
R^2	0.07	0.07	0.03	0.04	0.05	0.12	0.03	0.06	0.04	0.01	0.02	0.08	0.05
$\widehat{\beta}_{26}$	-2.16^{**}	-2.28^{**}	-0.97^{**}	-1.72^{**}	-1.77^{*}	-3.44^{**}	-1.03	-1.93^{**}	-1.22^{*}	-0.78	-0.38	-2.24^{**}	-1.62
s.e.	0.59	0.56	0.25	0.65	0.86	0.51	0.54	0.55	0.57	0.93	0.48	0.79	0.91
R^2	0.12	0.13	0.05	0.07	0.06	0.21	0.03	0.10	0.04	0.02	0.01	0.12	0.07
$\widehat{\beta}_{52}$	-2.13^{**}	-2.23^{**}	-0.88^{**}	-1.56^{*}	-1.59	-3.57^{**}	-1.01	-1.85^{**}	-0.90	-0.49	1.16	-1.75^{**}	-1.82^{*}
s.e.	0.61	0.56	0.24	0.67	0.82	0.55	0.70	0.50	0.72	0.91	0.68	0.65	0.77
R^2	0.18	0.18	0.05	0.09	0.08	0.35	0.04	0.14	0.03	0.01	0.05	0.11	0.15
Panel B Subsample excluding observations in 1980-1987 (Subsample E)													
$\widehat{\beta}_4$	-1.17	-1.24	-0.62	-1.11	-0.31	-1.44	-1.45^{**}	-1.14	-0.94	-1.27	-0.39	-1.54	-1.67^{*}
s.e.	0.68	0.74	0.38	0.72	0.79	0.90	0.53	0.79	0.50	0.70	0.49	1.03	0.83
R^2	0.01	0.01	0.00	0.01	0.00	0.01	0.01	0.01	0.01	0.01	0.00	0.02	0.01
$\widehat{\beta}_{13}$	-0.99	-1.10	-0.60	-1.06	-0.23	-1.55	-1.26^{*}	-0.79	-1.12^{*}	-0.99	-0.84	-1.52	-1.44
s.e.	0.74	0.82	0.40	0.72	0.93	1.36	0.59	0.84	0.50	0.97	0.56	1.01	1.02
R^2	0.02	0.02	0.01	0.02	0.00	0.01	0.02	0.01	0.03	0.02	0.02	0.04	0.02
$\widehat{\beta}_{26}$	-1.07	-1.26	-0.57	-1.13	-0.04	-2.12	-0.87	-0.81	-1.17	-1.07	-0.81	-1.67	-1.24
s.e.	0.63	0.68	0.38	0.65	0.95	1.09	0.63	0.65	0.66	0.98	0.54	0.89	1.14
R^2	0.03	0.04	0.01	0.04	0.00	0.05	0.02	0.02	0.05	0.04	0.03	0.08	0.03
$\widehat{\beta}_{52}$	-1.33^{*}	-1.52^{*}	-0.66	-1.15	-0.42	-2.64^{**}	-0.69	-0.93	-1.35	-1.07	1.00	-1.60^{*}	-1.53
s.e.	0.67	0.71	0.42	0.68	0.87	0.89	0.77	0.56	0.69	0.93	1.18	0.81	0.94
R^2	0.09	0.11	0.03	0.06	0.01	0.14	0.02	0.04	0.08	0.07	0.06	0.13	0.08
Panel C Subsample with observations in 1980-1987 (Subsample O)													
$\widehat{\beta}_4$	-4.81^{**}	-4.62^{**}	-1.29^{**}	-4.30^{*}	-5.23^{**}	-4.03^{**}	-2.88^{**}	-5.71^{**}	-2.24^{**}	-0.43	-1.58^{*}	-5.54^{**}	
s.e.	1.16	1.24	0.31	1.74	0.99	1.08	0.53	1.34	0.84	0.63	0.70	1.20	
R^2	0.11	0.09	0.03	0.04	0.18	0.10	0.12	0.13	0.03	0.00	0.03	0.10	
$\widehat{\beta}_{13}$	-5.05^{**}	-5.08^{**}	-1.26^{**}	-4.83^{*}	-6.02^{**}	-4.67^{**}	-2.39^{**}	-5.06^{**}	-2.18	-0.82	-0.83	-6.04^{**}	
s.e.	1.46	1.67	0.39	2.27	1.13	1.15	0.76	1.57	1.24	1.23	1.05	1.95	
R^2	0.25	0.22	0.05	0.11	0.43	0.27	0.18	0.21	0.05	0.02	0.01	0.22	
$\widehat{\beta}_{26}$	-5.61^{**}	-5.78^{**}	-1.07^{**}	-6.43^{*}	-5.82^{**}	-5.03^{**}	-2.29^{**}	-5.39^{**}	-0.73	-0.87	0.25	-6.22^{**}	
s.e.	1.49	1.72	0.32	2.52	0.97	1.01	0.88	1.53	1.30	1.52	0.96	2.26	
R^2	0.40	0.35	0.05	0.23	0.52	0.42	0.25	0.31	0.01	0.02	0.00	0.28	
$\widehat{\beta}_{52}$	-5.30^{**}	-5.14^{*}	-0.59^{**}	-4.94	-4.29^{**}	-4.46^{**}	-3.07^{**}	-5.54^{**}	1.81	-0.27	1.62	-4.01	
s.e.	1.65	2.01	0.21	2.62	1.06	1.17	1.11	1.72	1.33	1.68	0.94	3.34	
R^2	0.36	0.27	0.02	0.12	0.36	0.36	0.55	0.35	0.05	0.00	0.05	0.11	

This table reports the results from the OLS regression (1). ‘EPort’ denotes the sample which contains the equally-weighted average of the major five currencies (GDM, BRP, JPY, CAD, and SWF) and ‘WPort’ denotes the sample which contains the foreign exchange turnover-weighted average of the major five currencies. Weights are calculated using the data of foreign exchange market turnover by currency pair from 1992 to 2010 obtained from Bank for International Settlements (BIS). ‘Pool’ denotes the sample which contains all currencies. Newey and West (1987) standard errors denoted by ‘*s.e.*’ are computed based on the lag length of $2(k-1)$. ‘*’ denotes significance at the 5% level and ‘**’ denotes significance at the 1% level.

Table IV Results of the conditional test by Jasson and Moreira

	t-stat	[2.5, 97.5]	T	t-stat	[2.5, 97.5]	T
	Panel A Entire Sample: 1975-2009			Panel B Subsample E: excl 1980-87		
One-month maturity						
EPort	**−3.04**	**[−1.92, 1.99]**	418	**−2.12**	**[−2.05, 1.85]**	322
WPort	**−2.88**	**[−1.91, 1.98]**	418	-1.95	[-2.02 , 1.88]	322
GDM	**−2.46**	**[−1.92, 1.96]**	418	-1.43	[-1.92 , 1.96]	322
BRP	-1.66	[-1.99 , 1.97]	418	-0.21	[-1.98 , 1.97]	322
JPY	**−4.43**	**[−1.92, 1.99]**	418	-1.43	[-1.90 , 1.98]	322
CAD	**−3.14**	**[−2.02, 1.90]**	418	**−2.32**	**[−2.10, 1.73]**	322
SWF	**−2.96**	**[−1.93, 1.97]**	418	-1.15	[-1.97 , 1.92]	322
DAK	**−2.42**	**[−1.94, 2.02]**	364	-1.63	[-2.03 , 1.93]	262
FRF	-1.48	[-1.94 , 2.02]	287	-1.36	[-1.99 , 1.94]	191
ITL	-1.56	[-2.03 , 1.94]	287	-0.83	[-2.02 , 1.91]	191
DUG	**−2.82**	**[−1.92, 1.96]**	227	-1.40	[-1.86 , 2.41]	131
AUD	**−2.90**	**[−2.04, 1.86]**	277	-1.80	[-1.89 , 2.00]	262
Three-month maturity						
EPort	**−2.04**	**[−1.90, 1.99]**	139	-1.65	[-1.92 , 1.96]	107
WPort	-1.95	**[−1.89, 1.99]**	139	-1.47	[-1.89 , 1.99]	107
GDM	-1.76	[-1.89 , 2.00]	139	-1.05	[-1.86 , 2.04]	107
BRP	**−2.21**	**[−1.93, 2.00]**	139	-0.10	[-1.93 , 1.96]	107
JPY	**−4.50**	**[−1.95, 1.99]**	139	-1.15	[-1.92 , 1.99]	107
CAD	**−2.36**	**[−1.98, 2.01]**	139	**−2.22**	**[−2.04, 1.85]**	107
SWF	**−2.32**	**[−1.92, 1.98]**	139	-0.89	[-1.91 , 1.98]	107
DAK	**−2.19**	**[−1.92, 1.99]**	131	-0.85	[-1.79 , 2.11]	87
FRF	-0.51	[-1.96 , 2.01]	75	-0.26	[-1.31 , 2.84]	43
ITL	-1.00	[-2.00 , 2.05]	95	-1.27	[-2.01 , 1.90]	63
DUG	-1.78	[-1.95 , 1.93]	75	-0.97	[-1.45 , 2.93]	43
AUD	**−2.20**	**[−1.99, 1.88]**	92	-1.65	[-1.79 , 2.16]	87
Six-month maturity						
EPort	**−2.18**	**[−1.95, 1.95]**	69	-1.64	[-1.93 , 1.94]	53
WPort	**−2.15**	**[−1.95, 1.95]**	69	-1.58	[-1.94 , 1.93]	53
GDM	**−2.10**	**[−1.96, 1.93]**	69	-1.40	[-1.97 , 1.90]	53
BRP	-1.94	[-1.98 , 1.99]	69	-0.05	[-1.95 , 1.94]	53
JPY	**−4.36**	**[−2.00, 1.96]**	69	-1.90	**[−1.87, 1.99]**	53
CAD	-1.66	[-1.93 , 1.98]	69	-1.45	[-1.92 , 1.97]	53
SWF	**−2.82**	**[−1.92, 1.97]**	69	-1.58	[-2.14 , 1.78]	53
DAK	**−2.46**	**[−2.00, 1.94]**	65	-1.50	[-1.83 , 2.07]	43
FRF	-0.88	[-1.95 , 1.96]	37	-1.12	[-1.44 , 2.68]	21
ITL	-1.22	[-1.99 , 1.99]	47	**−1.99**	**[−1.97, 1.92]**	31
DUG	**−2.05**	**[−1.87, 2.01]**	37	-1.04	[-1.48 , 2.75]	21
AUD	-1.87	[-2.06 , 1.78]	46	-1.04	[-1.77 , 2.17]	43
Twelve-month maturity						
EPort	**−1.99**	**[−1.94, 1.93]**	34	-1.47	[-1.88 , 2.01]	26
WPort	**−1.99**	**[−1.94, 1.93]**	34	-1.47	[-1.89 , 1.99]	26
GDM	-1.91	[-1.97 , 1.90]	34	-1.42	[-1.97 , 1.91]	26
BRP	-1.83	[-1.90 , 2.02]	34	-0.73	[-1.94 , 1.94]	26
JPY	**−5.54**	**[−2.07, 1.83]**	34	**−2.81**	**[−2.12, 1.78]**	26
CAD	-1.59	[-1.95 , 1.93]	34	-1.12	[-1.93 , 1.94]	26
SWF	**−2.88**	**[−2.12, 1.79]**	34	-1.75	[-2.29 , 1.65]	26
DAK	-1.44	[-2.13 , 1.73]	32	-1.42	[-1.94 , 1.94]	21
AUD	**−2.51**	**[−1.88, 2.01]**	23	-1.39	[-1.79 , 2.13]	21

The second and fifth columns report the t-statistic of $\widehat{\beta}_k$ from both the entire sample and subsample E, respectively. The t-statistic in bold means that the null hypothesis is rejected at the

5% level based on the conventional critical value. The third and six columns report the 2.5 and 97.5% quantiles of the t distribution obtained using the conditional test by Jasson and Moreira. The interval of the quantiles in bold means that the null hypothesis is rejected at the 5% level.

Table V Results of the rank-based median variance ratio test

Table Va One-month excess return (k=4)

	Eport	Wport	GDM	BRP	JPY	CAD	SWF	DAK	FRF	ITL	DUG	AUD
	Panel A Entire Sample: 1975-09											
$\widehat{MR}_k(2)$	1.13	1.13	1.10	1.10	1.10	1.04	1.10	1.13	1.10	1.18	1.15	1.09
pvalue	**0.00**	**0.00**	**0.01**	**0.01**	**0.01**	0.21	**0.01**	**0.00**	**0.02**	**0.00**	**0.00**	**0.03**
$\widehat{MR}_k(12)$	1.49	1.54	1.60	1.45	1.59	1.40	1.48	1.72	1.69	1.78	1.90	1.46
pvalue	**0.01**	**0.00**	**0.00**	**0.01**	**0.00**	**0.02**	**0.01**	**0.00**	**0.00**	**0.00**	**0.00**	**0.03**
$\widehat{MR}_k(24)$	1.89	1.99	2.09	1.69	2.04	1.77	1.80	2.23	2.09	2.10	2.42	1.92
pvalue	**0.00**	**0.00**	**0.00**	**0.01**	**0.00**	**0.01**	**0.01**	**0.00**	**0.00**	**0.00**	**0.00**	**0.01**
$\widehat{MR}_k(36)$	2.13	2.24	2.34	1.92	2.16	2.14	1.98	2.44	2.23	2.23	2.62	2.21
pvalue	**0.01**	**0.00**	**0.00**	**0.01**	**0.00**	**0.00**	**0.01**	**0.00**	**0.01**	**0.01**	**0.00**	**0.01**
$\widehat{MR}_k(48)$	2.23	2.30	2.42	2.04	2.13	2.43	1.96	2.53	2.24	2.18	2.66	2.52
pvalue	**0.01**	**0.01**	**0.00**	**0.02**	**0.01**	**0.00**	**0.02**	**0.00**	**0.02**	**0.02**	**0.01**	**0.01**
$\widehat{MR}_k(60)$	2.29	2.31	2.52	1.97	1.95	2.69	2.00	2.59	2.30	2.24	2.77	2.71
pvalue	**0.01**	**0.01**	**0.00**	**0.03**	**0.03**	**0.00**	**0.03**	**0.01**	**0.03**	**0.03**	**0.02**	**0.01**
T	1820	1820	1820	1820	1820	1820	1820	1584	1252	1252	992	1208
	Panel B Subsample excluding observations in 1980-87 (Subsample E)											
$\widehat{MR}_k(2)$	1.11	1.11	1.09	1.06	1.08	1.03	1.10	1.14	1.06	1.15	1.14	1.10
pvalue	**0.01**	**0.01**	**0.02**	0.10	0.05	0.30	**0.02**	**0.00**	0.18	**0.01**	**0.03**	**0.03**
$\widehat{MR}_k(12)$	1.20	1.23	1.26	1.05	1.45	1.34	1.28	1.43	1.33	1.30	1.33	1.42
pvalue	0.17	0.13	0.10	0.37	**0.02**	**0.05**	0.10	**0.04**	0.11	0.13	0.15	**0.03**
$\widehat{MR}_k(24)$	1.47	1.55	1.47	1.00	1.99	1.62	1.43	1.69	1.44	1.12	1.46	1.80
pvalue	0.08	**0.05**	0.07	0.46	**0.01**	**0.03**	0.08	**0.04**	0.13	0.32	0.16	**0.02**
$\widehat{MR}_k(36)$	1.65	1.73	1.65	0.92	2.17	1.89	1.51	1.96	1.19	0.96	1.59	2.09
pvalue	0.06	**0.04**	0.06	0.53	**0.01**	**0.02**	0.10	**0.03**	0.29	0.44	0.15	**0.02**
$\widehat{MR}_k(48)$	1.70	1.74	1.74	0.81	2.05	2.05	1.44	2.18	0.86	0.58	1.54	2.34
pvalue	0.08	0.06	0.07	0.61	**0.03**	**0.02**	0.15	**0.03**	0.49	0.74	0.18	**0.02**
$\widehat{MR}_k(60)$	1.80	1.78	1.83	0.72	1.81	2.20	1.44	2.38	0.79	0.37	1.61	2.50
pvalue	0.08	0.07	0.07	0.67	0.07	**0.03**	0.17	**0.03**	0.52	0.90	0.18	**0.03**
T	1408	1408	1408	1408	1408	1408	1408	1144	832	832	572	1144

Table V reports the estimates of the rank-based median variance ratios and their right-tail p-values. $\widehat{MR}_k(q)$ is the estimate of the rank-based median variance ratio of k-week excess return associated with aggregation value q, defined in equation (29). The right-tail p-probabilities are calculated based on 5000 bootstrap samples generated using the parametric bootstrap method stated in Appendix B. *pvalue* in bold means that the null hypothesis is rejected at the 5% level. We use the same weekly data as Table III.

Table Vb Three-month excess return (k=13)

	Eport	Wport	GDM	BRP	JPY	CAD	SWF	DAK	FRF	ITL	DUG	AUD
Panel A Entire Sample: 1975-09												
$\widehat{MR}_k(2)$	1.14	1.15	1.14	1.12	1.19	1.12	1.13	1.15	1.22	1.17	1.23	1.13
pvalue	**0.00**	**0.02**	**0.03**	0.06	**0.01**	**0.05**	**0.01**	**0.02**	**0.02**	**0.02**	**0.01**	0.10
$\widehat{MR}_k(4)$	1.27	1.30	1.36	1.27	1.34	1.26	1.28	1.41	1.54	1.37	1.54	1.26
pvalue	**0.00**	**0.03**	**0.01**	**0.04**	**0.02**	**0.04**	**0.01**	**0.01**	**0.01**	**0.03**	**0.01**	0.10
$\widehat{MR}_k(8)$	1.64	1.70	1.75	1.43	1.60	1.57	1.54	1.77	1.91	1.68	1.96	1.45
pvalue	**0.00**	**0.01**	**0.01**	0.06	**0.01**	**0.02**	**0.00**	**0.01**	**0.01**	**0.03**	**0.01**	0.09
$\widehat{MR}_k(12)$	1.80	1.86	1.92	1.53	1.66	1.88	1.63	1.93	2.03	1.74	2.09	1.65
pvalue	**0.00**	**0.01**	**0.01**	0.07	**0.04**	**0.01**	**0.01**	**0.01**	**0.03**	0.05	**0.03**	0.07
$\widehat{MR}_k(16)$	1.82	1.81	1.93	1.51	1.52	2.10	1.57	1.95	1.97	1.71	1.96	1.90
pvalue	**0.01**	**0.03**	**0.03**	0.11	0.09	**0.01**	**0.02**	**0.03**	0.06	0.09	0.07	0.06
$\widehat{MR}_k(20)$	1.85	1.78	1.98	1.31	1.37	2.28	1.63	2.03	1.89	1.85	1.86	2.13
pvalue	**0.01**	0.05	**0.04**	0.22	0.17	**0.01**	**0.03**	**0.04**	0.09	0.08	0.10	0.05
Panel B Subsample excluding observations in 1980-87 (Subsample E)												
$\widehat{MR}_k(2)$	1.04	1.05	1.06	1.03	1.12	1.13	1.02	1.07	1.09	1.09	1.04	1.12
pvalue	0.33	0.31	0.25	0.39	0.09	0.07	0.47	0.24	0.30	0.25	0.44	0.12
$\widehat{MR}_k(4)$	1.05	1.08	1.12	1.10	1.24	1.21	1.05	1.22	1.16	1.12	1.13	1.23
pvalue	0.39	0.34	0.24	0.29	0.08	0.12	0.37	0.13	0.28	0.31	0.33	0.12
$\widehat{MR}_k(8)$	1.19	1.27	1.32	1.00	1.58	1.37	1.11	1.38	1.21	0.98	1.20	1.36
pvalue	0.24	0.18	0.14	0.48	**0.03**	0.11	0.32	0.13	0.30	0.48	0.30	0.13
$\widehat{MR}_k(12)$	1.27	1.36	1.41	0.81	1.65	1.56	1.16	1.56	1.26	0.79	1.29	1.62
pvalue	0.22	0.16	0.14	0.67	0.06	0.08	0.29	0.11	0.29	0.63	0.26	0.08
$\widehat{MR}_k(16)$	1.29	1.33	1.45	0.72	1.40	1.59	1.09	1.70	1.10	0.40	1.12	1.86
pvalue	0.22	0.21	0.15	0.71	0.16	0.11	0.36	0.10	0.36	0.92	0.34	0.06
$\widehat{MR}_k(20)$	1.40	1.40	1.63	0.62	1.26	1.64	1.13	1.85	1.34	0.32	1.36	2.11
pvalue	0.19	0.20	0.12	0.76	0.24	0.12	0.32	0.09	0.29	0.94	0.27	**0.05**

Table Vc Six-month excess return (k=13)

	Eport	Wport	GDM	BRP	JPY	CAD	SWF	DAK	FRF	ITL	DUG	AUD
Panel A Entire Sample: 1975-09												
$\widehat{MR}_k(2)$	1.14	1.16	1.20	1.17	1.16	1.20	1.16	1.25	1.30	1.24	1.32	1.17
pvalue	**0.01**	**0.02**	**0.00**	**0.01**	**0.01**	**0.00**	**0.02**	**0.00**	**0.00**	**0.01**	**0.00**	**0.01**
$\widehat{MR}_k(4)$	1.46	1.51	1.53	1.36	1.36	1.56	1.35	1.54	1.62	1.45	1.65	1.37
pvalue	**0.00**	**0.00**	**0.00**	**0.01**	**0.00**	**0.00**	**0.02**	**0.00**	**0.00**	**0.02**	**0.00**	**0.00**
$\widehat{MR}_k(6)$	1.61	1.63	1.64	1.47	1.34	1.81	1.37	1.66	1.76	1.52	1.71	1.47
pvalue	**0.00**	**0.01**	**0.00**	**0.01**	**0.02**	**0.00**	**0.04**	**0.00**	**0.01**	**0.03**	**0.00**	**0.01**
$\widehat{MR}_k(8)$	1.60	1.59	1.62	1.46	1.18	2.01	1.28	1.68	1.70	1.48	1.68	1.58
pvalue	**0.00**	**0.02**	**0.00**	**0.03**	0.13	**0.00**	0.11	**0.01**	**0.03**	0.07	**0.00**	**0.01**
$\widehat{MR}_k(10)$	1.63	1.60	1.69	1.26	1.05	2.19	1.30	1.73	1.78	1.58	1.70	1.71
pvalue	**0.00**	**0.03**	**0.00**	0.13	0.31	**0.00**	0.12	**0.01**	**0.04**	0.07	**0.00**	**0.00**
Panel B Subsample excluding observations in 1980-87 (Subsample E)												
$\widehat{MR}_k(2)$	1.01	1.01	1.08	1.06	1.08	1.13	1.05	1.15	1.18	1.14	1.18	1.12
pvalue	0.53	0.52	0.23	0.37	0.30	0.07	0.39	0.16	0.26	0.27	0.26	**0.05**
$\widehat{MR}_k(4)$	1.12	1.13	1.20	0.99	1.27	1.40	1.07	1.26	1.21	1.07	1.24	1.28
pvalue	0.32	0.31	0.15	0.53	0.16	**0.01**	0.40	0.19	0.32	0.42	0.30	**0.02**
$\widehat{MR}_k(6)$	1.19	1.21	1.25	0.79	1.30	1.57	1.12	1.40	1.24	0.79	1.35	1.41
pvalue	0.27	0.27	0.16	0.74	0.20	**0.01**	0.35	0.16	0.32	0.66	0.27	**0.01**
$\widehat{MR}_k(8)$	1.18	1.16	1.24	0.66	1.11	1.60	1.03	1.53	1.07	0.41	1.23	1.52
pvalue	0.30	0.32	0.21	0.82	0.36	**0.01**	0.44	0.14	0.42	0.93	0.34	**0.01**
$\widehat{MR}_k(10)$	1.29	1.26	1.34	0.56	1.04	1.59	1.10	1.67	1.35	0.34	1.50	1.67
pvalue	0.25	0.27	0.17	0.86	0.41	**0.02**	0.37	0.12	0.34	0.96	0.29	**0.01**

Table Vd One-Year Excess Return (k=52)

	Eport	Wport	GDM	BRP	JPY	CAD	SWF	DAK	Eport	Wport	GDM	BRP	JPY	CAD	SWF	DAK
	Panel A Entire Sample: 1975-09								Panel B Subsample E: excl 1980-87							
$\widehat{MR}_k(2)$	1.30	1.33	1.28	1.15	1.25	1.29	1.26	1.27	1.12	1.11	1.11	0.99	1.13	1.25	1.11	1.11
pvalue	**0.00**	**0.00**	**0.00**	0.07	**0.01**	**0.00**	**0.00**	**0.00**	0.33	0.28	0.24	0.63	0.31	**0.01**	0.14	0.35
$\widehat{MR}_k(3)$	1.48	1.50	1.40	1.19	1.25	1.55	1.33	1.40	1.19	1.23	1.19	0.78	1.19	1.50	1.16	1.25
pvalue	**0.00**	**0.00**	**0.01**	0.08	0.08	**0.00**	**0.00**	**0.00**	0.28	0.14	0.15	0.83	0.29	**0.00**	0.07	0.17
$\widehat{MR}_k(4)$	1.52	1.49	1.42	1.18	1.08	1.73	1.28	1.43	1.19	1.21	1.21	0.67	1.13	1.64	1.07	1.31
pvalue	**0.00**	**0.00**	**0.02**	0.12	0.36	**0.00**	**0.02**	**0.00**	0.31	0.20	0.17	0.86	0.38	**0.00**	0.29	0.17
$\widehat{MR}_k(5)$	1.53	1.46	1.48	1.02	0.96	1.88	1.27	1.51	1.31	1.30	1.27	0.59	1.10	1.66	1.09	1.45
pvalue	**0.00**	**0.01**	**0.03**	0.41	0.56	**0.00**	**0.04**	**0.00**	0.25	0.14	0.13	0.88	0.40	**0.00**	0.22	0.12

Table VI Results of the Rank-Based Median Variance Ratio Test: One-Month Return (k=4) in the sample period of 1980-87

	Eport	Wport	GDM	BRP	JPY	CAD	SWF	DAK	FRF	ITL	DUG
$\widehat{MR}_k(2)$	1.21	1.21	1.14	1.22	1.21	1.02	1.16	1.18	1.13	1.13	1.15
pvalue	**0.01**	**0.01**	0.07	**0.01**	**0.01**	0.43	**0.03**	**0.02**	0.08	0.07	0.05
$\widehat{MR}_k(12)$	2.60	2.69	2.60	2.42	2.36	1.20	2.37	2.65	2.68	2.72	2.76
pvalue	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**	0.28	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**
$\widehat{MR}_k(24)$	3.98	4.14	3.93	3.21	3.50	1.27	3.47	4.27	4.08	4.20	4.17
pvalue	**0.00**	**0.00**	**0.00**	**0.01**	**0.00**	0.26	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**
$\widehat{MR}_k(36)$	4.27	4.38	4.06	4.22	3.60	1.10	3.74	4.55	4.12	4.31	4.45
pvalue	**0.00**	**0.00**	**0.01**	**0.00**	**0.01**	0.34	**0.01**	**0.00**	**0.00**	**0.01**	**0.00**
$\widehat{MR}_k(48)$	4.47	4.56	4.02	4.86	4.00	1.00	3.60	4.92	4.38	4.54	4.47
pvalue	**0.01**	**0.00**	**0.01**	**0.00**	**0.01**	0.40	**0.01**	**0.00**	**0.00**	**0.01**	**0.01**
$\widehat{MR}_k(60)$	6.05	6.10	5.30	6.67	4.99	1.51	4.79	6.27	5.86	5.93	5.75
pvalue	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**	0.22	**0.00**	**0.00**	**0.00**	**0.00**	**0.00**
T	416	416	416	416	416	416	416	416	416	416	416

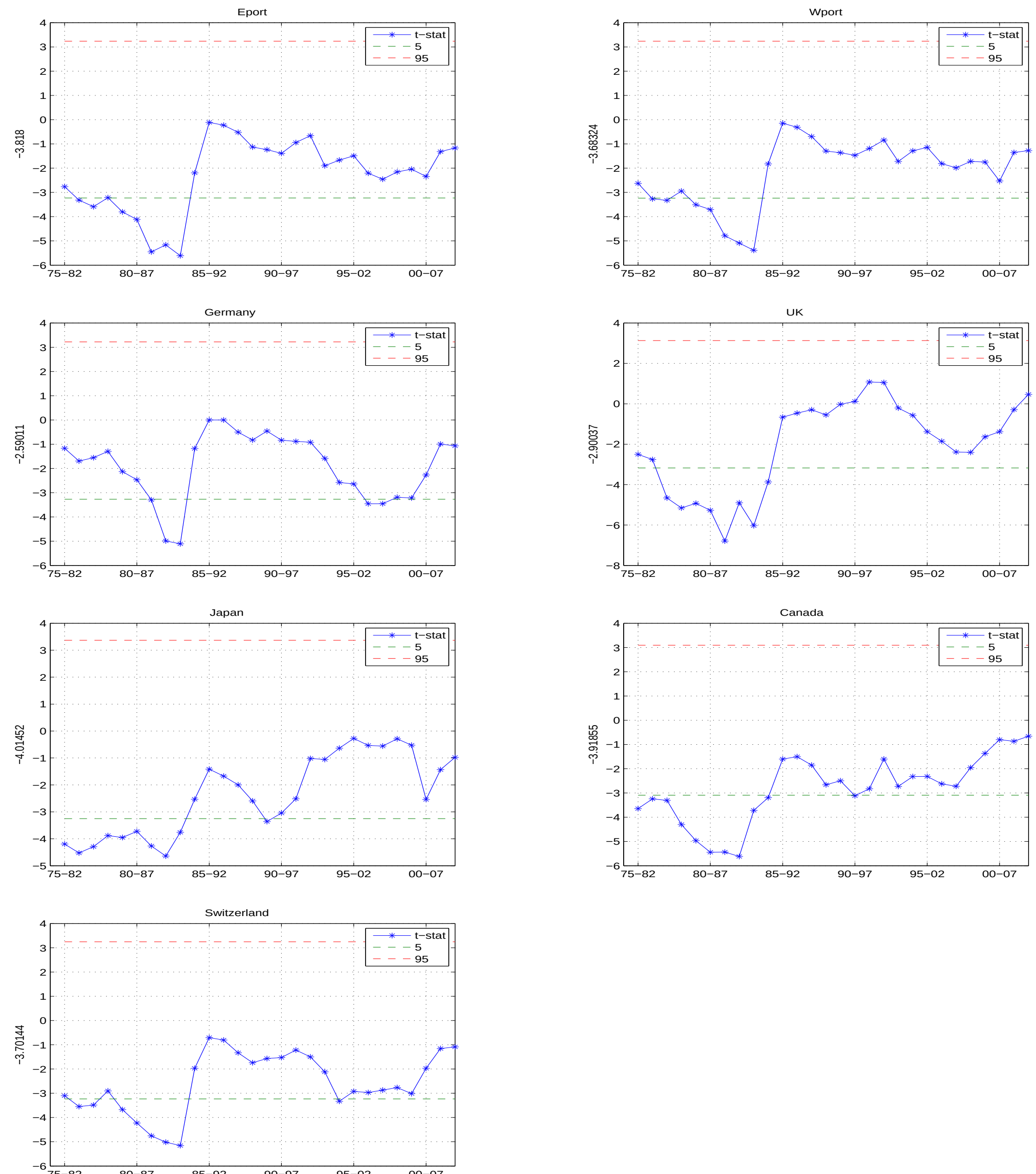
Note: refer to Table V.

Table VII Results on the Behavior of the Survey Expectations

	EPort	WPort	Pool	GDM	BRP	JPY	CAD	SWF	DAK	FRF	ITL	AUD	SWK
Table VIIa Three-Month Maturity													
Panel A UIP hypo test: OLS regressions of $s_{t+k} - f_{t\|k}$ on $f_{t\|k} - s_t$													
$\widehat{\beta}_3$	-1.86^{*}	-1.76	-0.95	-1.83	-0.65	-3.09^{**}	-1.53^{**}	-2.44^{*}	-1.17	-0.94	1.89	-1.52	-0.25
se	0.91	1.04	0.67	1.00	1.50	0.95	0.57	1.13	0.79	1.08	1.68	1.00	1.40
Panel B Excessive speculation test: OLS regressions of $s^m_{t+k} - s_{t+k}$ on $f_{t\|k} - s_t$													
$\widehat{\beta}^a_3$	1.91	2.30^{*}	0.85	3.39^{**}	0.20	1.64	1.27	2.96^{*}	1.52	1.27	-0.89	2.16	0.65
se	0.99	1.14	0.67	1.14	1.51	1.13	0.65	1.25	0.85	1.10	1.92	1.11	1.44
Panel C: OLS regressions of $s^m_{t+k} - s_t$ on $f_{t\|k} - s_t$													
$\widehat{\beta}^b_3$	1.05	1.54	0.89	2.56^{*}	0.55	-0.45^{**}	0.74	1.51	1.35	1.33	2.00^{*}	1.64^{*}	1.40
se	0.43	0.47	0.24	0.61	0.64	0.49	0.18	0.43	0.28	0.41	0.41	0.27	0.33
Table VIIb Six-Month Maturity													
Panel A UIP hypo test: OLS regressions of $s_{t+k} - f_{t\|k}$ on $f_{t\|k} - s_t$													
$\widehat{\beta}_6$	-1.86^{*}	-1.76^{*}	-0.98	-1.91^{*}	-0.26	-3.35^{**}	-1.18	-2.35^{*}	-1.27	-1.02	1.38	-1.32	-0.38
se	0.73	0.84	0.61	0.92	1.40	0.74	0.65	0.95	0.90	1.03	1.32	1.10	1.23
Panel B Excessive speculation test: OLS regressions of $s^m_{t+k} - s_{t+k}$ on $f_{t\|k} - s_t$													
$\widehat{\beta}^a_6$	2.33^{**}	2.54^{**}	1.02^{*}	3.06^{**}	0.76	2.37^{*}	1.10	3.06^{**}	2.17^{*}	1.31	-0.67	2.27	1.31
se	0.70	0.83	0.52	0.84	1.40	0.97	0.74	0.94	0.97	0.97	1.43	1.19	1.26
Panel C: OLS regressions of $s^m_{t+k} - s_t$ on $f_{t\|k} - s_t$													
$\widehat{\beta}^b_6$	1.47	1.77	1.04	2.15^{*}	1.50	0.02	0.91	1.71	1.90^{*}	1.29	1.71	1.95^{**}	1.93^{**}
se	0.46	0.52	0.31	0.55	0.64	0.51	0.24	0.53	0.39	0.42	0.37	0.30	0.33
Table VIIc One-Year Maturity													
Panel A UIP hypo test: OLS regressions of $s_{t+k} - s_t$ on $f_{t\|k} - s_t$													
$\widehat{\beta}_{12}$	-1.92^{*}	-1.83^{*}	-1.00	-1.95^{*}	-0.28	-3.62^{**}	-1.02	-2.29^{*}	-1.37	-1.09	1.03	-1.58	-0.41
se	0.82	0.90	0.60	0.97	1.21	0.72	0.84	0.94	0.93	0.97	1.25	0.96	1.15
Panel B Excessive speculation test: OLS regressions of $s^m_{t+k} - s_{t+k}$ on $f_{t\|k} - s_t$													
$\widehat{\beta}^a_{12}$	2.41^{**}	2.47^{**}	0.86^{*}	2.51^{**}	1.07	3.03^{**}	0.83	2.65^{**}	2.18^{*}	1.04	-0.99	2.35^{*}	1.23
se	0.68	0.76	0.42	0.85	1.18	0.84	0.91	0.78	0.99	0.91	1.29	0.99	1.23
Panel C: OLS regressions of $s^m_{t+k} - s_t$ on $f_{t\|k} - s_t$													
$\widehat{\beta}^b_{12}$	1.49	1.64	0.85	1.56	1.79	0.41	0.82	1.36	1.81^{*}	0.95	1.04	1.77^{**}	1.82^{**}
se	0.37	0.42	0.26	0.51	0.45	0.43	0.26	0.53	0.36	0.12	0.12	0.26	0.27

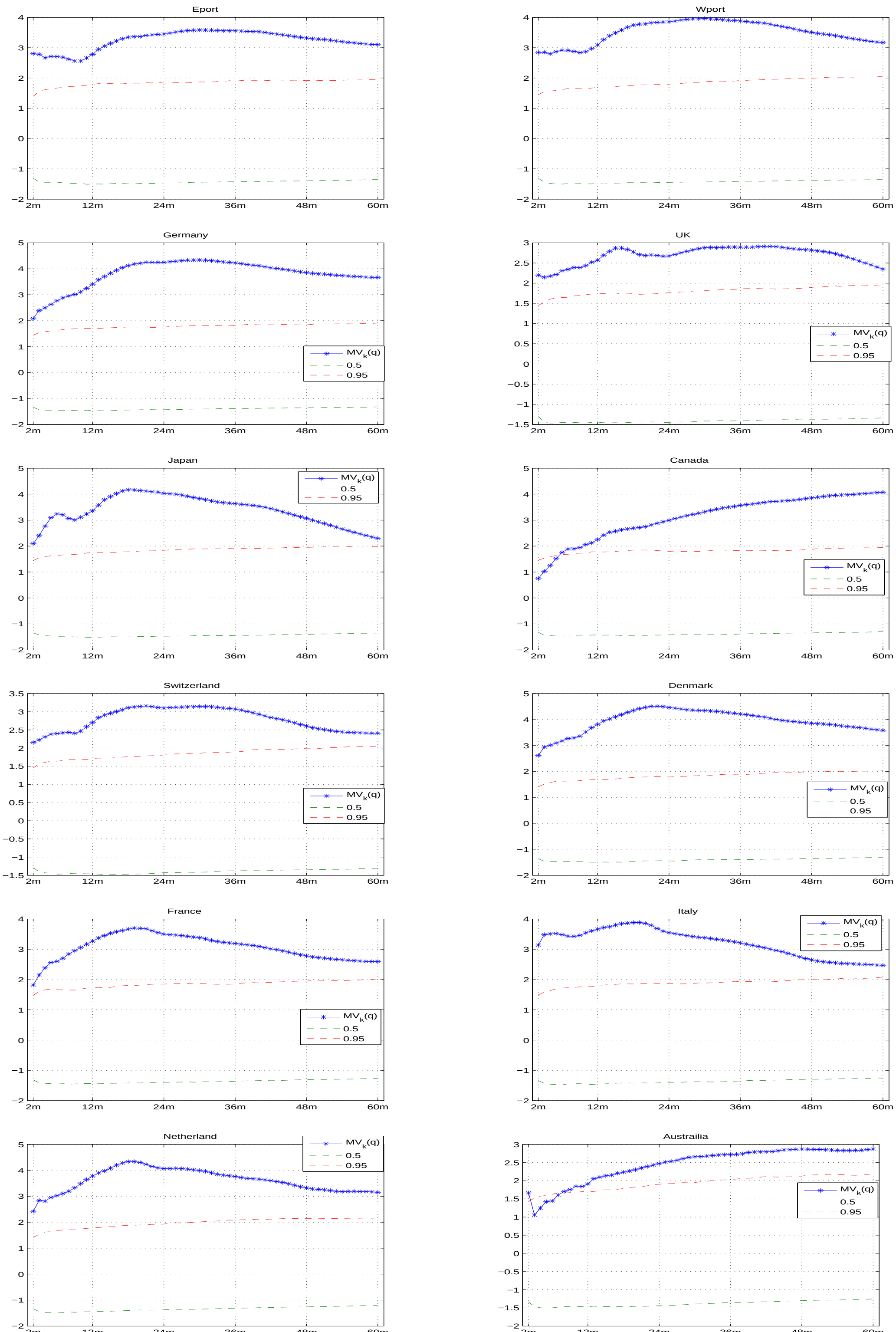
Data frequency is monthly and spot and forward rates are chosen at the end of each month. But we use the same notation k for a maturity as the previous tables. Accordingly, $k = 3$ in this table should be interpreted as the three-month maturity. The sample period is 1988:1-2008:12. Newey and West (1987) standard errors denoted by '*s.e.*' are computed based on the lag length of $2(k-1)$. '*' represents the 5% significant level and '**' represents the 1% level. See also note in Table III.

Figure 1: Parameter Stability Tests Based on 8-year Rolling Samples: 1975-2009



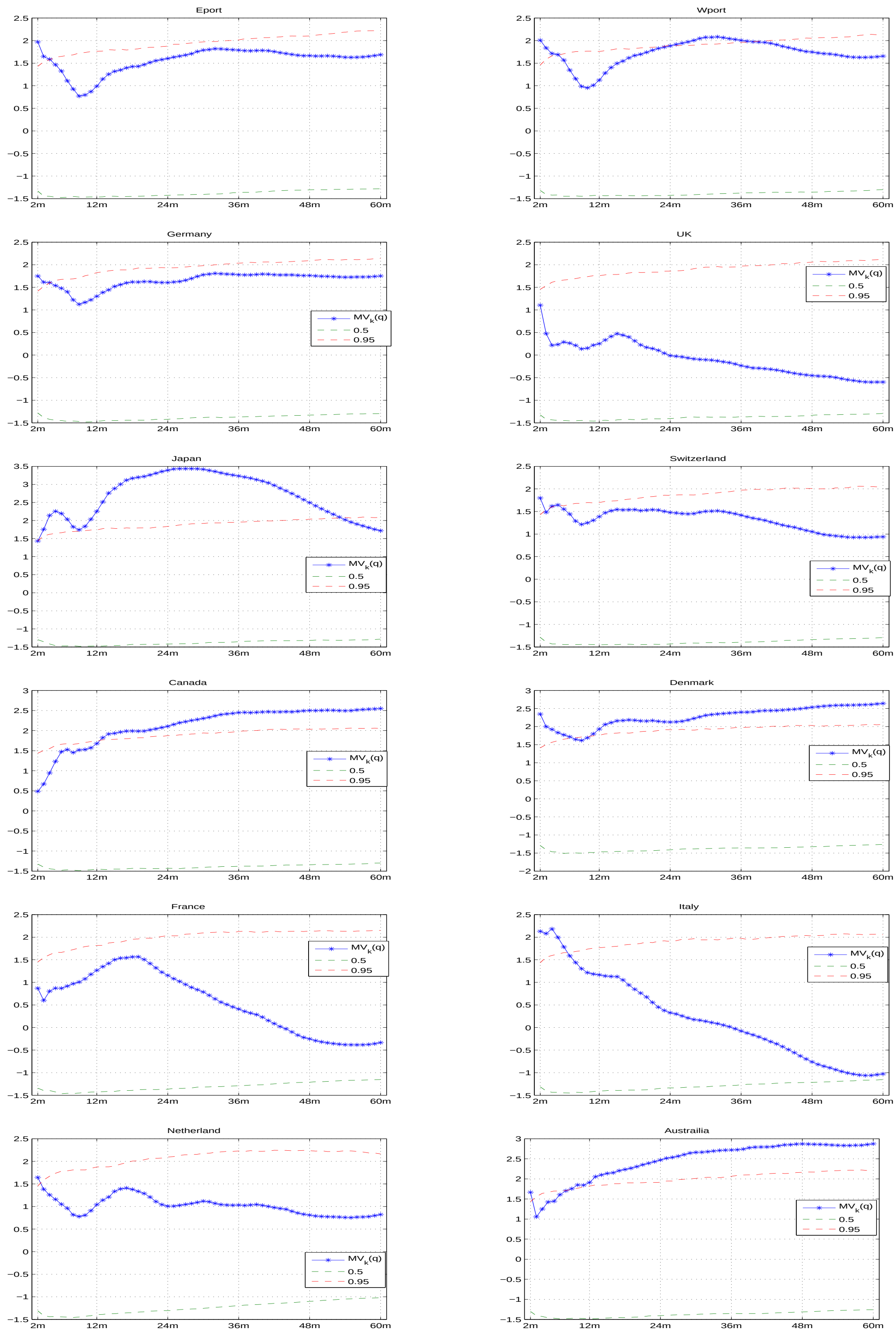
We construct eight-year rolling samples. The t-statistic for each rolling sample is obtained from regression (1). The line with asterisks is the locus of the t-statistics across rolling samples and the two dashed lines denote 5 and 95% quantiles. Construction of the uniform bands is specified in Appendix A. The number on y-axis is the t-statistic from the entire sample.

Figure 2: Patterns of the t-statistics ($\widehat{MV}_k(q)$) with q for one-month excess returns ($s_{t+4} - f_{t|4}$): 1975-2009



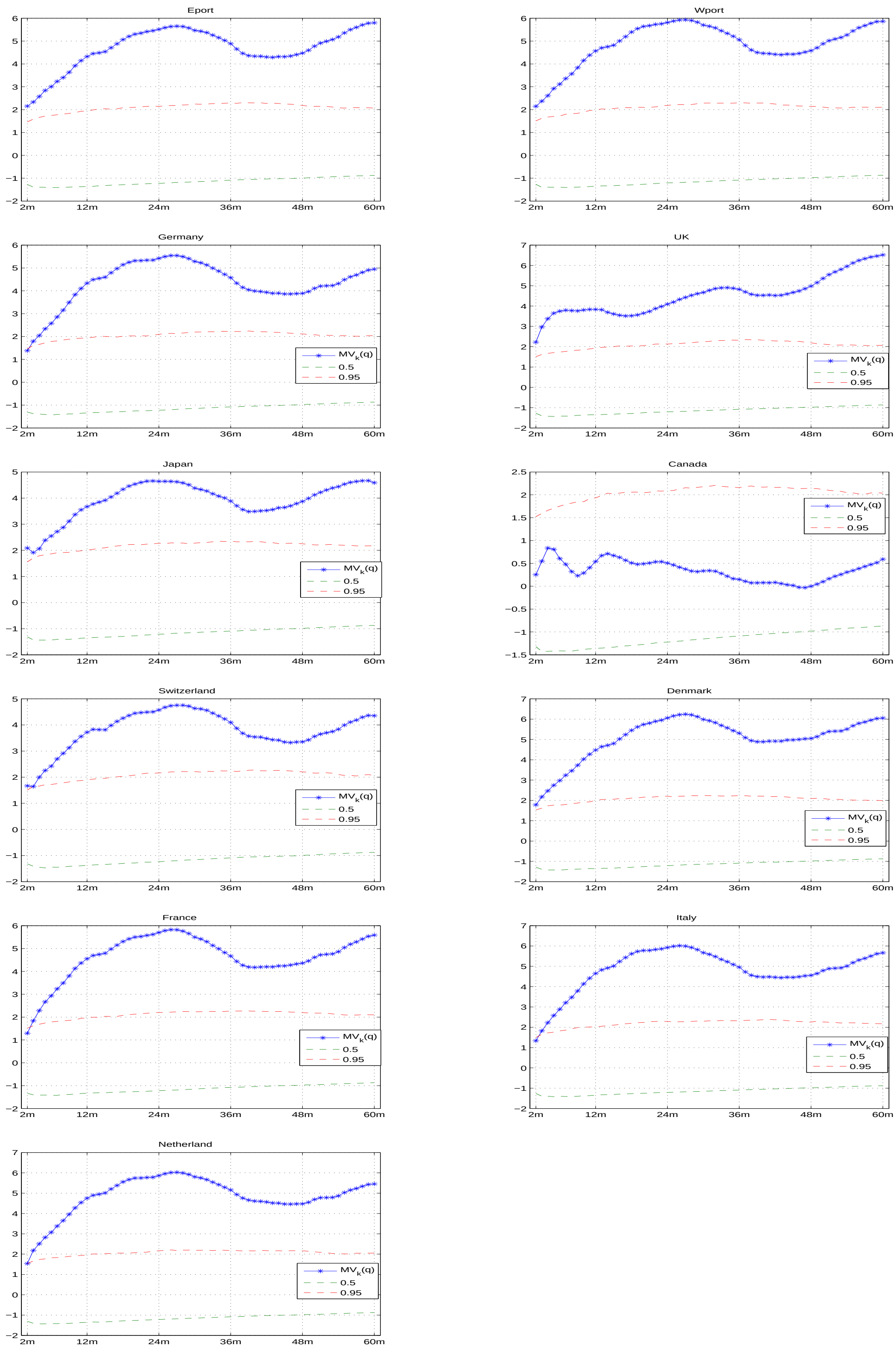
The line with asterisks is the locus of the t-statistics of the estimated median variance ratios with q, $\widehat{MV}_k(q)$, and the two dashed lines correspond to 5 and 95% quantiles from the simulated bootstrap empirical distribution obtained from 5000 bootstrap samples following the method in Appendix B.

Figure 3: Patterns of the t-statistics $(\widehat{MV}_k(q))$ with q for one-month excess returns $(s_{t+4} - f_{t|4})$: excluding observations in 1980-87



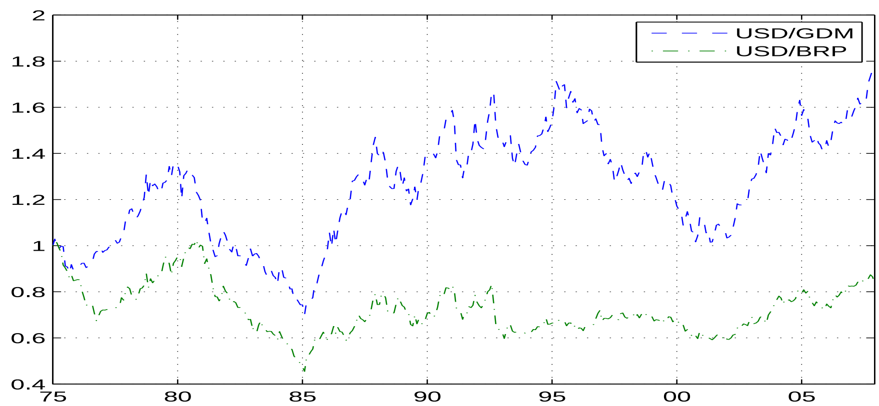
Note: refer to Figure 2.

Figure 4: Patterns of the t-statistics ($\widehat{MV}_k(q)$) with q for one-month excess returns ($s_{t+4} - f_{t|4}$): 1980-1987



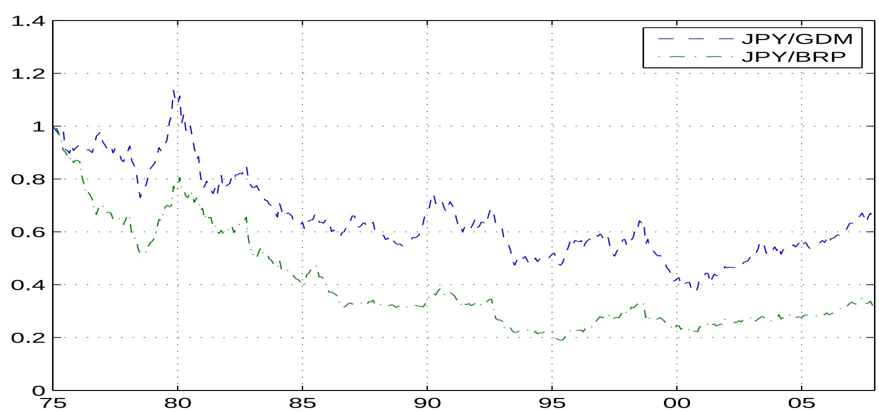
Note: refer to Figure 2.

Figure 5: Time series of the USD bilateral rates against BRP and GDM



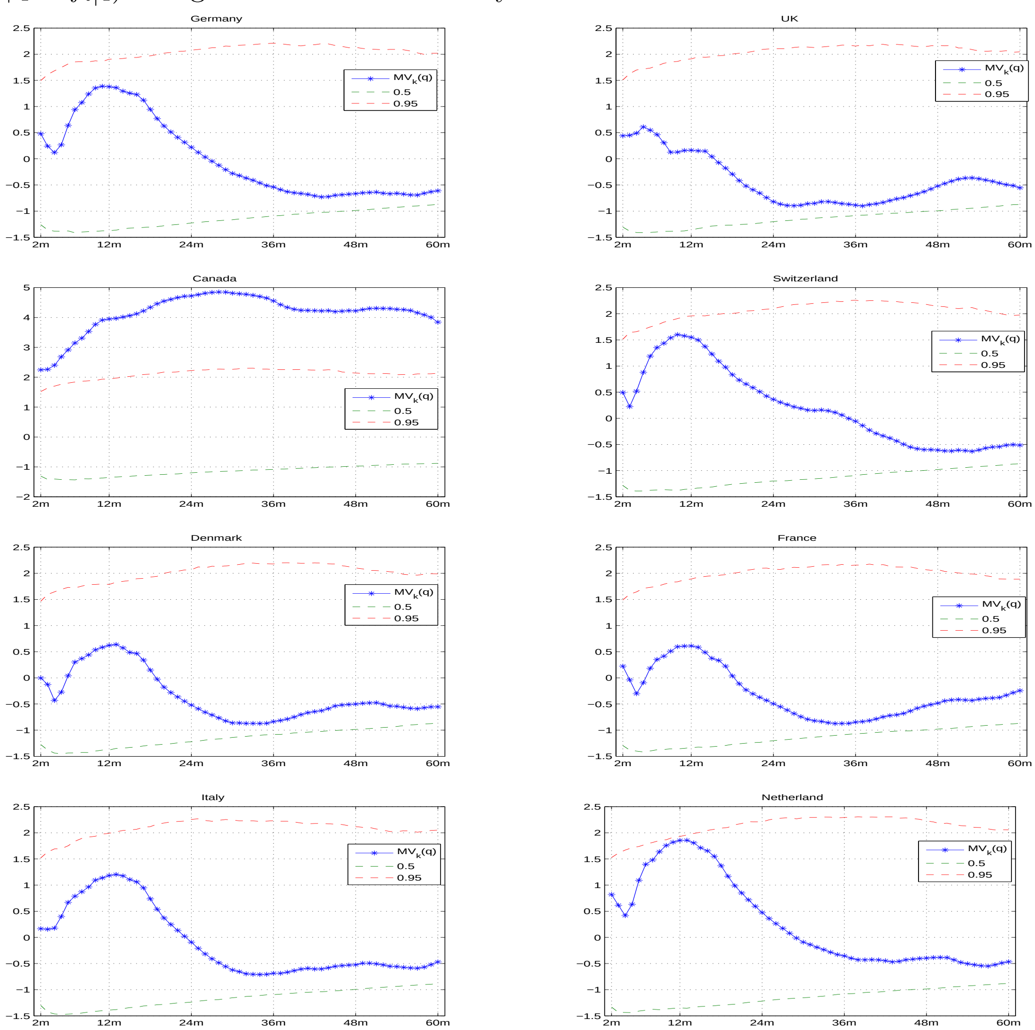
Note: refer to Figure 6

Figure 6: Time series of the JPY bilateral rates against BRP and GDM



Data consists of monthly spot rates between 1975:1 and 2009:11. Observations at the end of each month are selected. The series for each currency are normalized using the first observation.

Figure 7: Patterns of the t-statistics $(\widehat{MV}_k(q))$ with q for one-month excess returns $(s_{t+4} - f_{t|4})$ using JPY as base currency: 1980-1987



Note: refer to Figure 2. We replace the USD with the JPY as a base currency using triangular arbitrage.