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Cross-currency, cross-maturity forward exchange premiums as predictors of spot rate changes: Theory and evidence

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Abstract

This paper shows that, in predicting the spot rate of a currency, the term structures of forward premiums of *other* currencies have incremental information content in addition to the term structure of the currency's own forward premiums. The theoretical model motivating the analysis hinges on the abundant evidence on the co-movements of excess returns from investing in different currencies. The empirical results are obtained through FIML estimation of a vector error correction model for weekly data on three bilateral US dollar exchange rates.

JEL classification: F31; G15

Keywords: Spot exchange rate; Forward exchange premium; Risk premium

1. Introduction

There is a widespread consensus in international finance literature that the so-called unbiasedness efficiency hypothesis, namely that the forward exchange premium is the optimal unbiased predictor of future spot rate changes, does not hold at the empirical level (see, e.g. the surveys by Hodrick (1987) and Engel (1996)). Moreover, a large body of evidence shows that the forward rate tends to mispredict the future path of the spot rate, as the coefficient of a regression of exchange rate

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changes on the forward premium is generally estimated to be negative (Froot and Thaler, 1990). In other words, when the foreign currency is at a forward premium and the level of domestic interest rates is therefore higher than that of the foreign country, an appreciation is likely—the opposite of what is predicted by the uncovered interest parity hypothesis.

Several theoretical explanations have been advanced for the finding of substantial deviations from the unbiasedness hypothesis (see Taylor (1995) for a survey). Under the assumption of rational expectations and non-systematic forecast errors, risk aversion induces market participants to demand a risk premium when holding foreign currency; this risk premium arises as an equilibrium outcome. Alternatively, Froot and Frankel (1989), using survey data, cannot reject the hypothesis that the entire bias in the forward premium is attributable to systematic expectation errors. Another explanation is based on rational systematic forecast errors due, for example, to expectations of a future shift in the distribution of the exchange rate not followed by the materialization of the policy shift motivating such expectations (the "peso problem"; see Lewis (1994) for a survey).

The accumulation of evidence that a rise in the forward exchange rate is likely to be associated with a subsequent fall in the spot rate (an appreciation) has engendered some skepticism over the ability of forward contract markets to convey information about the future evolution of the currency's value (Cumby and Obstfeld, 1984). However, a recent study by Clarida and Taylor (1997) convincingly challenges this view by showing that when the whole term structure of forward rates is considered, the forward premiums at different horizons do contain useful information for predicting the spot rate.

Importantly, all of their hypotheses and conclusions are entirely consistent with the presence of deviations from the unbiasedness hypothesis. More generally, the failure of this hypothesis to stand up to empirical scrutiny is associated with the existence of excess returns from speculation in the foreign currency market. These excess returns are not simply realizations of forecast errors; on the contrary, they are predictable, and their predictability is extensively documented in the literature. [1] Moreover, numerous studies have addressed the issue of whether excess returns on a variety of financial assets in international markets display some co-movements and some degree of proportionality. The main finding is that such commonality is widespread (see, e.g. Campbell and Clarida, 1987; Cumby and Huizinga, 1992; Bekaert, 1995).

Building on this large body of evidence, the present paper examines excess returns on foreign exchange markets across bilateral dollar exchange rates and across the term structure, simultaneously. In particular, it investigates whether the term structure of forward premiums of a currency has information content for predicting the spot rate of *another* currency incremental to that embedded in the latter's own forward premiums. From a simple theoretical model motivating the empirical

[1] See, among others, Campbell and Clarida (1987), Bekaert and Hodrick (1992), Canova and Marrinan (1995) and Bekaert et al. (1997).

analysis, we derive a set of conditions under which information from cross-currency forward premiums at one and several maturities ought to be extracted when predicting the future direction of a spot rate. A key feature of the model is the explicit consideration of the co-movements and proportionality of risk premiums both across foreign currencies and across the maturity spectrum. [2] We explicitly model this commonality in returns using an intertemporal capital asset pricing model, which we adapted to the purposes of our work by considering as excess returns in international markets those arising from uncovered positions in various foreign currencies.

The empirical analysis is performed by estimating a dynamic vector error correction model (VECM) on a system of three bilateral dollar exchange rates. The main issue to investigate is whether the forward premiums of one or more currencies at different maturities have incremental predictive power for the spot rate of another currency, over and above that of the term structure of the latter's own forward premiums.

The paper is organized as follows: Section 2 develops the theoretical framework, motivating the analysis. Section 3 illustrates the empirical methodology, describes the data and presents the estimation results. Section 4 concludes.

2. Theoretical framework

2.1. A simple asset pricing model

As is well known, the forward premium, $(F_{t,j} - S_t)$, can be decomposed into the sum of two unobservable variables: expected depreciation, $E_t(\Delta S_{t+j})$, and deviation from uncovered interest parity, $\Phi_{t,j}$, with $F_{t,j}$ being the j-period forward rate at time t and S_t the spot rate. Assuming that agents are risk averse, the deviation term, $\Phi_{t,j}$, can be interpreted as a time-varying risk premium, originating within the context of an intertemporal capital asset pricing model (ICAPM; see, e.g. Hansen and Hodrick, 1983; Hodrick and Srivastava, 1984).

Consider a representative agent framework with time separable utility over lifetime consumption. We concentrate on three simple investment strategies that the consumer/investor faces (the numeraire currency is assumed to be the dollar). The first strategy consists in investing in a one-period, dollar-denominated deposit that guarantees a riskless nominal gross return ($R^{\mathrm{us}}_{t,1}$) after one period: $\mathrm{Ret}^1_{t,1} = R^{\mathrm{us}}_{t,1}$. The second strategy is an uncovered (open) position in Japanese yen (JPY), investing in a one-period yen-denominated deposit that yields $R^{\mathrm{jpy}}_{t,1}$. The dollar return of the investment at the end of period is $\mathrm{Ret}^2_{t,1} = (S_{t+1}/S_t)R^{\mathrm{jpy}}_{t,1}$, where S_t is the dollar–yen spot exchange rate. The third strategy is an uncovered position in British pounds (GBP),

[2] The issue of excess returns predictability and commonality when allowance is made for different maturities is examined by Huang (1989), Lewis (1991) and Canova and Marrinan (1995), among others.

investing in a one-period pound-denominated deposit that yields $R^{\mathrm{gbp}}_{t,1}$. The dollar return is $\mathrm{Ret}^3_{t,1} = (S^*_{t+1}/S^*_t)R^{\mathrm{gbp}}_{t,1}$, where S^*_t denotes the dollar–sterling spot rate. The first order conditions for expected utility maximization can be summarized as follows:

$$\frac{U'(C_t)}{P_t} = \delta E_t\left[\frac{U'(C_{t+1})}{P_{t+1}}\mathrm{Ret}^i_{t,1}\right] \quad (i = 1,2,3), \tag{1}$$

where $E_t\,\mathrm{Ret}^i_{t,1}$ is the dollar-denominated expected nominal return on the ith investment strategy; $1/P_t$ denotes the purchasing power of the numeraire currency and δ is the discount factor. Dividing both sides of Eq. (1) by $U'(C_t)$, we obtain

$$E_t(\mathrm{Ret}^i_{t,1}Q_{m,t+1}) = 1 \quad (i = 1,2,3), \tag{2}$$

where $Q_{m,t+1}$ denotes the ex-post marginal rate of substitution of the domestic currency between time t and $t+1$: $Q_{m,t+1} = [\delta(U'(C_{t+1})/U'(C_t))(P_t/P_{t+1})]$.

The intuition behind the first order conditions of this problem is the following: an international investor choosing a utility maximizing consumption/saving plan must be indifferent among the following alternatives: (a) use the domestic currency to increase her/his stock of assets denominated in domestic currency; (b) allocate the money to assets denominated in foreign currency (Japanese yen or British pounds) and then convert the proceeds back into dollars; (c) spend the money in present consumption.

In order to obtain an expression for the risk premium, which separates the expected return on the uncovered position, $E_t\mathrm{Ret}^i_{t,1}$, from the certain return on the dollar deposit, Eq. (2) is recalled using the properties of conditional covariance and the fact that, from Eq. (2), the riskless return $R^{\mathrm{us}}_{t,1}$ is equal to $1/E(Q_{m,t+1})$; it follows that

$$E_t\,\mathrm{Ret}^i_{t,1} = \frac{1}{E_t(Q_{m,t+1})}[1 - \mathrm{Cov}_t(\mathrm{Ret}^i_{t,1}, Q_{m,t+1})] \quad (i = 1,2,3). \tag{3}$$

Subtracting $R^{\mathrm{us}}_{t,1}$ from both sides of the above expression, we obtain

$$E_t\,\mathrm{Ret}^i_{t,1} - R^{\mathrm{us}}_{t,1} = -\frac{\mathrm{Cov}_t(\mathrm{Ret}^i_{t,1}, Q_{m,t+1})}{E_tQ_{m,t+1}} \quad (i = 2,3). \tag{4}$$

It follows that the risk premium on the ith investment strategy is inversely related to the conditional covariance between the intertemporal marginal rate of substitution of dollars and the return on the ith investment. The investment whose return has a small covariance with $Q_{m,t+1}$ tends to yield low returns when the investor's marginal propensity to consume is high, that is, when consumption is low. In this case, a higher risk premium is required because the returns are low just when wealth is most valuable, i.e. when consumption is low. Consider now a benchmark investment strategy, defined as one that yields the following nominal return:

$$\mathrm{Ret}^{\mathrm{b}}_{t,1} = Q_{m,t+1}/E_t(Q^2_{m,t+1}). \tag{5}$$

It is straightforward to verify (see Eq. (2)) that $\mathrm{Ret}^{\mathrm{b}}_{t,1}$ is defined so as to satisfy the first order conditions of the maximization problem (Hansen and Hodrick, 1983). Replacing $Q_{m,t+1}$ in Eq. (4) with its expression deriving from Eq. (5), i.e. $Q_{m,t+1} = \mathrm{Ret}^{\mathrm{b}}_{t,1} E_t Q^2_{m,t+1}$, yields the following equation:

$$E_t \mathrm{Ret}^{i}_{t,1} - R^{\mathrm{us}}_{t,1} = -\frac{\mathrm{Cov}_t(\mathrm{Ret}^{i}_{t,1}, \mathrm{Ret}^{\mathrm{b}}_{t,1})}{E_t(\mathrm{Ret}^{\mathrm{b}}_{t,1})} \quad (i = 2, 3). \tag{6}$$

In the case of the benchmark investment $(i = b)$, Eq. (6) reduces to $(E_t \mathrm{Ret}^{\mathrm{b}}_{t,1} - R^{\mathrm{us}}_{t,1}) = -\mathrm{Var}_t(\mathrm{Ret}^{\mathrm{b}}_{t,1})/E_t \mathrm{Ret}^{\mathrm{b}}_{t,1}$. Combining the latter expression with Eq. (6), we have

$$E_t(\mathrm{Ret}^{i}_{t,1} - R^{\mathrm{us}}_{t,1}) = \left[\frac{\mathrm{Cov}_t(\mathrm{Ret}^{i}_{t,1}, \mathrm{Ret}^{\mathrm{b}}_{t,1})}{\mathrm{Var}_t(\mathrm{Ret}^{\mathrm{b}}_{t,1})}\right] E_t(\mathrm{Ret}^{\mathrm{b}}_{t,1} - R^{\mathrm{us}}_{t,1}) \quad (i = 2, 3). \tag{7}$$

Let $\lambda^1_{t,1}$ define the ex-post excess return of the uncovered investment in the yen-denominated deposit over the riskless investment in a dollar-denominated deposit and let $\lambda^2_{t,1}$ denote the ex-post excess return of the uncovered investment in the sterling-denominated deposit over the riskless investment:

$$\lambda^1_{t,1} = (s_{t+1} - s_t + i^{\mathrm{jpy}}_{t,1}) - i^{\mathrm{us}}_{t,1}, \qquad \lambda^2_{t,1} = (s^*_{t+1} - s^*_t + i^{\mathrm{gbp}}_{t,1}) - i^{\mathrm{us}}_{t,1}; \tag{8}$$

all the variables on the right-hand side have been written in lower case after taking logarithms and using the fact that $\log R_{t,1} = \log(1 + i_{t,i}) \approx i_{t,i}$. Taking the conditional expectations of the two excess returns yields

$$E_t(\lambda^1_{t,1}) = E_t \Delta s_{t+1} - (i^{\mathrm{us}}_{t,1} - i^{\mathrm{jpy}}_{t,1}), \qquad E_t(\lambda^2_{t,1}) = E_t \Delta s^*_{t+1} - (i^{\mathrm{us}}_{t,1} - i^{\mathrm{gbp}}_{t,1}). \tag{9}$$

The two terms, $E_t(\lambda^1_{t,1})$ and $E_t(\lambda^2_{t,1})$, are the unobservable risk premiums for the two currencies: call them $\phi^1_{t,1}$ and $\phi^2_{t,1}$. Eq. (7) can be then expressed as follows:

$$\phi^i_{t,1} = E_t \lambda^i_{t,1} = \beta^i E_t(\mathrm{Ret}^{\mathrm{b}}_{t,1} - R^{\mathrm{us}}_{t,1}) \quad (i = 2, 3), \tag{10}$$

where $\beta^i = \mathrm{Cov}_t(\mathrm{Ret}^{i}_{t,1}, \mathrm{Ret}^{\mathrm{b}}_{t,1})/\mathrm{Var}_t \mathrm{Ret}^{\mathrm{b}}_{t,1}$.

The expected return on the benchmark investment in excess of that obtained from the riskless rate, $E_t(\mathrm{Ret}^{\mathrm{b}}_{t,1} - R^{\mathrm{us}}_{t,1})$, represents what is commonly recognized in the literature as a *latent variable*. Eq. (10) captures an important feature of this class of equilibrium models, namely that the risk premiums across different currencies are proportional because they are propelled by a common unobservable factor through a conditional covariance term (the consumption beta, β^i). In order to make the concept of latent variable operational, it is customary in the literature, since the work of Hansen and Hodrick (1983), to project $E_t(\mathrm{Ret}^{\mathrm{b}}_{t,1} - R^{\mathrm{us}}_{t,1})$ on some appropriate observed variables, X_t, of the information set of time t: $E_t(\mathrm{Ret}^{\mathrm{b}}_{t,1} - R^{\mathrm{us}}_{t,1}) = \gamma' X_t + \varepsilon_t$, where ε_t is the linear projection error and γ is a vector of parameters. We believe that natural candidates for information variables are the forward premiums of the two currencies, $(f_{t,1} - s_t)$ and $(f^*_{t,1} - s^*_t)$. By adopting this approach, we use as information variables elements of the expected excess returns over the riskless rate, as defined by Eq. (9). To show that this is appropriate, we first recall that these excess returns are linked to the latent variable, $E_t(\mathrm{Ret}^{\mathrm{b}}_{t,1} - R^{\mathrm{us}}_{t,1})$, by way of Eq. (10). Second, using covered interest parity, we rewrite Eq. (9) as follows:

$$E_t(\lambda^1_{t,1}) = E_t\Delta s_{t+1} - (f_{t,1} - s_t), \quad E_t(\lambda^2_{t,1}) = E_t\Delta s^*_{t+1} - (f^*_{t,1} - s^*_t); \tag{11}$$

that is to say, the forward premiums enter the expressions for the excess returns that we want to estimate jointly. Subsequently, each excess return realized with an open position is projected on a constant and on a common set of information variables given by the *two* forward premiums:

$$\lambda^1_{t,1} = \Delta s_{t+1} - (i^{\mathrm{us}}_{t,1} - i^{ye}_{t,1}) = \alpha_{1,1} + \alpha_{1,2}(f_{t,1} - s_t) + \alpha_{1,3}(f^*_{t,1} - s^*_t) + u^1_{t+1,1}, \tag{12}$$

$$\lambda^2_{t,1} = \Delta s^*_{t+1} - (i^{\mathrm{us}}_{t,1} - i^{uk}_{t,1}) = \alpha_{2,1} + \alpha_{2,2}(f_{t,1} - s_t) + \alpha_{2,3}(f^*_{t,1} - s^*_t) + u^2_{t+1,1}. \tag{13}$$

Each of the coefficients above, $\alpha_{i,j}$, is the product of β_i and γ_j, which have previously been defined. Rational expectations imply that the forecast errors, $u^i_{t+1,1}$, are orthogonal to the information set available at time t. In fact, $u^i_{t+1,1}$ is a linear combination of two disturbance terms: an expectation error, $v_{t+1,1}$, and the error obtained when the latent variable is projected on the vector X_t: $u^i_{t+1,1} = v^i_{t+1,1} + \beta^i\varepsilon_t$ (Hansen and Hodrick, 1983). The above system can be reformulated as

$$\Delta s_{t+1} = \delta_{1,1} + \delta_{1,2}(f_{t,1} - s_t) + \delta_{1,3}(f^*_{t,1} - s^*_t) + u^1_{t+1,1}, \tag{14}$$

$$\Delta s^*_{t+1} = \delta_{2,1} + \delta_{2,2}(f_{t,1} - s_t) + \delta_{2,3}(f^*_{t,1} - s^*_t) + u^2_{t+1,1}, \tag{15}$$

where $\delta_{1,1} = \alpha_{1,1}$, $\delta_{1,2} = (1 + \alpha_{1,2})$, $\delta_{1,3} = \alpha_{1,3}$ and $\delta_{2,1} = \alpha_{2,1}$, $\delta_{2,2} = \alpha_{2,2}$, $\delta_{2,3} = (1 + \alpha_{2,3})$.

2.2. Testable implications

We have provided motivation for including the forward exchange premium of another currency as an information variable. Before testing on empirical grounds whether *both* forward premiums have predictive power for the future movement in the spot rate, it is important to examine the probability limit of the regression coefficients in the equations above and, in particular, of $\alpha_{1,3}$ and $\alpha_{2,2}$, referring to the forward premiums of the *other* currency. Their true values are:

$$\begin{aligned}\alpha_{1,3} = \delta_{1,3} &= \frac{\mathrm{Cov}_t[(f^*_{t,1} - s^*_t), \Delta s_{t+1} - (f_{t,1} - s_t)]}{\mathrm{Var}_t(f^*_{t,1} - s^*_t)},\\ \alpha_{2,2} = \delta_{2,2} &= \frac{\mathrm{Cov}_t[(f_{t,1} - s_t), \Delta s^*_{t+1} - (f^*_{t,1} - s^*_t)]}{\mathrm{Var}_t(f_{t,1} - s_t)}.\end{aligned} \tag{16}$$

If both the covered interest parity condition and rational expectations hold, then the expressions above can be reformulated as

$$\begin{aligned}\delta_{1,3} &= \frac{\mathrm{Cov}_t[E_t\Delta s^*_{t+1} + \phi^*_{t,1}, (s_{t+1} - E_t s_{t+1}) - \phi_{t,1}]}{\mathrm{Var}_t(f^*_{t,1} - s^*_t)},\\ \delta_{2,2} &= \frac{\mathrm{Cov}_t[E_t\Delta s_{t+1} + \phi_{t,1}, (s^*_{t+1} - E_t s^*_{t+1}) - \phi^*_{t,1}]}{\mathrm{Var}_t(f_{t,1} - s_t)}.\end{aligned} \tag{17}$$

Using the fact that the expectation error relative to the spot rate of a currency is orthogonal to both the risk premium and the expected rate of depreciation of the other currency, the expressions above reduce to

$$\begin{aligned}\delta_{1,3} &= \frac{-\mathrm{Cov}_t(\phi_{t,1}, E_t\Delta s^*_{t+1}) - \mathrm{Cov}_t(\phi_{t,1}, \phi^*_{t,1})}{\mathrm{Var}_t(f^*_{t,1} - s^*_t)},\\ \delta_{2,2} &= \frac{-\mathrm{Cov}_t(\phi^*_{t,1}, E_t\Delta s_{t+1}) - \mathrm{Cov}_t(\phi^*_{t,1}, \phi_{t,1})}{\mathrm{Var}_t(f_{t,1} - s_t)}.\end{aligned} \tag{18}$$

Focusing on one of the coefficients, for example $\delta_{1,3}$, we can assert that it is non-trivial, and the forward premium of the other currency $(f^*_{t,1} - s^*_t)$ therefore has incremental information content for Δs_{t+1}, if either the expected depreciation of the dollar against a currency is correlated with the risk premium associated with the other currency or, importantly, the risk premiums themselves display co-movements. This is true, however, only to the extent that: $|\mathrm{Cov}_t(\phi_{t,1},\phi^*_{t,1})| \neq -|\mathrm{Cov}_t(\phi_{t,1}, E_t\Delta s^*_{t+1})|$. The same argument applies, *mutatis mutandis*, to the other parameter, $\delta_{2,2}$. Expressions (18) provide the intuition as to why forward premiums of various currencies might have incremental predictive ability for a spot exchange rate. Indeed, the large body of evidence of co-movements in risk premiums implies that $\mathrm{Cov}_t(\phi_{t,1},\phi^*_{t,1})$ is different from zero. In turn, this is likely to render $\delta_{1,3}$ and $\delta_{2,2}$ different from zero. Therefore, a testable implication of the widely observed commonality in excess returns is that, in predicting the spot rate of a currency, the forward premiums of other currencies have additional information content.

We now explicitly consider the presence of a maturity spectrum of forward rates for the different currencies. In particular, the question arises of whether the theoretical framework examined thus far is confined to a cross-currency but *single* maturity setting or, conversely, it should be extended to allow for forward premiums matching different maturities. Thus, the next step is to investigate how the intertemporal latent variable model accommodates the inclusion of the term structure of forward premiums of the own currency and of another currency. This issue is addressed by expanding our framework to include a larger set of investment strategies.

In particular, for each investment denominated in a particular currency we allow for more than one uncovered position: the number of these depends on the different maturities available in financial market contracts.

Eq. (7) stemming from the first order conditions of the consumer/investor problem is then generalized as follows:

$$E_t(\mathrm{Ret}^i_{t,j} - R^{\mathrm{us}}_{t,j}) = \left[\frac{\mathrm{Cov}_t(\mathrm{Ret}^i_{t,j}, \mathrm{Ret}^{\mathrm{b}}_{t,j})}{\mathrm{Var}_t(\mathrm{Ret}^{\mathrm{b}}_{t,j})}\right] E_t(\mathrm{Ret}^{\mathrm{b}}_{t,j} - R^{\mathrm{us}}_{t,j}), \tag{19}$$

where the index i continues to refer to the currency and j to the holding period of a contract ($j = 1, 2, \ldots, h$). In our framework, the total number of excess returns is now $2h$ and the expected excess returns of the benchmark investments over the corresponding $R^{\mathrm{us}}_{t,j}$ may differ across the contract length j (recall the definition in Eq. (5): i.e. $\mathrm{Ret}^{\mathrm{b}}_{t,j} = Q_{m,t+j}/E_t(Q^2_{m,t+j})$). In other words, there are as many latent variables as the number of different maturities. However, it is possible to conjecture that the h expected excess returns of the benchmark investments over $R^{\mathrm{us}}_{t,j}$ are all projected on the same set of information variables, which is now expanded. [3] Again, the variables entering the expressions for excess returns are selected as predictors. Hence, all forward premiums crossing the maturity spectrum are used, not just those associated with different currencies at a single maturity.

For expositional simplicity, we consider a term structure with only two maturities ($j = 1, 2$). Expressions (12) and (13) become

$$\begin{aligned}\lambda^1_{t,1} &= \Delta s_{t+1} - (i^{\mathrm{us}}_{t,1} - i^{\mathrm{jpy}}_{t,1}) \\ &= \alpha_{1,1} + \alpha_{1,2}(f_{t,1} - s_t) + \alpha_{1,3}(f_{t,2} - s_t) + \alpha_{1,4}(f^*_{t,1} - s^*_t) + \alpha_{1,5}(f^*_{t,2} - s^*_t) \\ &\quad + u^1_{t+1,1},\end{aligned} \tag{20}$$

$$\begin{aligned}\lambda^2_{t,1} &= \Delta s^*_{t+1} - (i^{\mathrm{us}}_{t,1} - i^{\mathrm{gbp}}_{t,1}) \\ &= \alpha_{2,1} + \alpha_{2,2}(f_{t,1} - s_t) + \alpha_{2,3}(f_{t,2} - s_t) + \alpha_{2,4}(f^*_{t,1} - s^*_t) + \alpha_{2,5}(f^*_{t,2} - s^*_t) \\ &\quad + u^2_{t+1,1},\end{aligned} \tag{21}$$

where the generic parameter $\alpha_{i,j}$ is the product of $\beta^{i,j}$, i.e. the covariance–variance ratio term in square brackets of Eq. (19) and the coefficient, γ_j, from the projection of the latent variable onto the set of information variables. The latter equations can be written as

$$\begin{aligned}\Delta s_{t+1} &= \delta_{1,1} + \delta_{1,2}(f_{t,1} - s_t) + \delta_{1,3}(f_{t,2} - s_t) + \delta_{1,4}(f^*_{t,1} - s^*_t) + \delta_{1,5}(f^*_{t,2} - s^*_t) \\ &\quad + u^1_{t+1,1},\end{aligned} \tag{22}$$

[3] Huang (1989) lends some support to this conjecture. However, in his work no attempt is made to estimate different cross-currency and cross-maturity excess returns simultaneously.

$$\Delta s^*_{t+1} = \delta_{2,1} + \delta_{2,2}(f_{t,1} - s_t) + \delta_{2,3}(f_{t,2} - s_t) + \delta_{2,4}(f^*_{t,1} - s^*_t) + \delta_{2,5}(f^*_{t,2} - s^*_t) + u^2_{t+1,1}, \tag{23}$$

with $\delta_{1,3}$, $\delta_{1,5}$, $\delta_{2,3}$, $\delta_{2,5}$ being equal to, respectively, $\alpha_{1,3}$, $\alpha_{1,5}$, $\alpha_{2,3}$, $\alpha_{2,5}$ and $\delta_{1,2}$ and $\delta_{2,4}$ equal to, respectively $(1+\alpha_{1,2})$ and $(1+\alpha_{2,4})$.

The probability limit of the coefficients associated with the cross-currency, cross-maturity premiums, that is, $(f^*_{t,2} - s^*_t)$ in Eq. (22) and $(f_{t,2} - s_t)$ in Eq. (23), are:

$$\begin{aligned} \alpha_{1,5} = \delta_{1,5} &= \frac{\mathrm{Cov}_t[f^*_{t,2} - s^*_t, \Delta s_{t+1} - (f_{t,1} - s_t)]}{\mathrm{Var}_t(f^*_{t,2} - s^*_t)}, \\ \alpha_{2,3} = \delta_{2,3} &= \frac{\mathrm{Cov}_t[f_{t,2} - s_t, \Delta s^*_{t+1} - (f^*_{t,1} - s^*_t)]}{\mathrm{Var}_t(f_{t,2} - s_t)}; \end{aligned} \tag{24}$$

through some algebra of least squares, $\delta_{1,5}$ and $\delta_{2,3}$ reduce to

$$\begin{aligned} \delta_{1,5} &= \frac{-\mathrm{Cov}_t[E_t\Delta s^*_{t+2}, \phi_{t,1}] - \mathrm{Cov}_t[\phi^*_{t,2}, \phi_{t,1}]}{\mathrm{Var}_t(f^*_{t,2} - s^*_t)}, \\ \delta_{2,3} &= \frac{-\mathrm{Cov}_t[E_t\Delta s_{t+2}, \phi^*_{t,1}] - \mathrm{Cov}_t[\phi_{t,2}, \phi^*_{t,1}]}{\mathrm{Var}_t(f_{t,2} - s_t)}; \end{aligned} \tag{25}$$

according to the expression above, the conditions for $\delta_{1,5}$ and $\delta_{2,3}$ to be different from zero are, respectively:

$$\begin{aligned} &|\mathrm{Cov}_t(\phi^*_{t,2}, \phi_{t,1})| \neq -|\mathrm{Cov}_t(E_t\Delta s^*_{t+2}, \phi_{t,1})|, \\ &|\mathrm{Cov}_t(\phi_{t,2,}\phi^*_{t,1})| \neq -|\mathrm{Cov}_t(E_t\Delta s_{t+2,}\phi^*_{t,1})|. \end{aligned} \tag{26}$$

The existence of co-movements across risk premiums matching different currencies and different maturities is likely to induce a non-zero term in the left-hand sides of the above expressions. Again, this justifies the inclusion of the term structure of cross-currency forward premiums as explanatory variables of regressions (22) and (23).

3. Empirical evidence

3.1. Estimation

The conclusions reached in the previous section concerning the predictive content of forward premiums lend themselves to empirical scrutiny. In order to assess the predictive content of forward premiums, a dynamic error correction representation is adopted for the vector, z_t. The latter comprises 15 elements: the spot rates of three currencies against the dollar plus twelve forward rates, each referring to one of the three currencies and one of the four possible maturities. Moreover, lagged forward premiums for each currency and for each maturity are added to the specification. The three currencies are the Japanese yen (JPY), the British pound sterling (GBP) and the German mark (DM). For each of them, the term structure of forward rates comprises four maturities: 4, 13, 26 and 52 weeks. In compact notation, the specification of the differenced VECM is

$$\Delta z_t = \mu + \zeta_1 \Delta z_{t-1} + \cdots + \zeta_{p-1} \Delta z_{t-p+1} + \alpha\beta' z_{t-1} + \varepsilon_t, \tag{27}$$

where $\beta' z_{t-1}$ is the 12-element vector with 12 forward premiums and α is the (15×12) matrix of their associated parameters in each equation.

With regard to the univariate statistical properties of each variable in z_t, there is a widespread consensus in the empirical literature that spot and forward rates are well characterized as unit root processes (see, among others, Hsieh, 1988; Ballie and Bollerslev, 1989). The abundant evidence against rejection of the non-stationarity hypothesis points to the issue of cointegration among the variables included in z_t. The prevailing view in the literature is that spot and forward rates are cointegrated and that each forward premium $(f_{t,j} - s_t)$ is suitably represented as a $I(0)$ process. The numerous studies providing evidence in this direction include the work of Horvath and Watson (1994), Phillips et al. (1996) and Wu and Chen (1998).

Among the 15 equations in the system, three are of special interest for our purpose here, namely, the ones where the dependent variable is the rate of variation in the spot exchange rate: $\Delta s_t^{\mathrm{jpy}}$, $\Delta s_t^{\mathrm{gbp}}$ and Δs_t^{dm}. Assessing whether the term structure of cross-currency forward premiums has incremental information for the future movement of the spot exchange rate is the crucial point to investigate.

The estimation is conducted as follows. First, we provide further evidence to support the prevailing view that spot and forward rates possess a unit root while forward premiums are stationary. Second, to assess the appropriate lag length of the dynamic model, we employ both the Akaike and the Schwarz information criteria. Once the lag length is selected, the VECM in first difference is estimated by OLS. Subsequently, a reduction process of insignificant individual variables is performed in each equation on the basis of a Wald test of exclusion restrictions, conducted at the system level. The restricted dynamic model is then estimated by full information maximum likelihood (FIML).

The empirical investigation is conducted on weekly (annualized) logarithmic observations; the sample period runs from 1977:1 through 1996:52 and the last three and a half years are reserved for out-of-sample-forecasting. The source of the data is the Harris Bank database supported by the Center for Studies in International Finance at the University of Chicago. Observations of exchange rates are recorded on Friday at 2:00 PM (Chicago time) every week. Each observation records the bid price of the currency and in particular the amount of foreign currency that a dealer offers (bids) for one dollar (for any exchange rate except dollar/sterling) and the amount of dollars that a dealer offers for one unit of the foreign currency (for dollar/sterling exchange rate only). Data are then made comparable so that each rate expresses the amount of dollars per unit of the foreign currency. In addition to the spot rates, the Harris dataset contains the annualized value of the percentage forward premium ($(F - S)/S$ in percentage per year); simple manipulations allow us to extract the value of each forward rate.

3.2. Results

The estimation results are presented in this section. With regard to the univariate properties of each variable in the system, in addition to the abundant evidence supporting the unit root characterization of exchange rates and the stationarity of forward premiums, we directly test these hypotheses by conducting augmented Dickey–Fuller tests for the spot and forward rates as well as for the forward premiums. The results confirm the prevailing view. As Table 1 documents, the hypothesis that the log level of each spot and forward rate possesses a unit root is not rejected at standard level of significance. Importantly, the evidence on the forward exchange premiums confirms that these series are, in general, stationary. In order to assess the appropriate lag length of the system, both the Akaike and the Schwarz criteria are applied to the unrestricted first differenced VAR. Table 2 reports the results: the Akaike criterion is minimized when one lag in the differenced system is included, whereas the Schwarz criterion is minimized when no lags appear in the system. The more conservative option of retaining one lag in the differenced system is adopted.

A restricted dynamic VECM is estimated for the system of the three bilateral dollar exchange rates. The emphasis is placed on the three equations where the dependent variable is the spot exchange rate variation: one for the dollar–sterling, one for the dollar–mark and one for the dollar–yen (see Table 3). Exclusion

Table 1
Unit root tests

Spot and forward rates	ADF tests	Forward premiums	ADF tests
(A) Dollar–Yen			
s_t^{jpy}	−2.24	$s_t^{\mathrm{jpy}} - f_{t,4}^{\mathrm{jpy}}$	−26.37
$f_{t,4}^{\mathrm{jpy}}$	−2.26	$s_t^{\mathrm{jpy}} - f_{t,13}^{\mathrm{jpy}}$	−18.86
$f_{t,13}^{\mathrm{jpy}}$	−2.23	$s_t^{\mathrm{jpy}} - f_{t,26}^{\mathrm{jpy}}$	−16.99
$f_{t,26}^{\mathrm{jpy}}$	−2.30	$s_t^{\mathrm{jpy}} - f_{t,52}^{\mathrm{jpy}}$	−13.61
$f_{t,52}^{\mathrm{jpy}}$	−2.41		
(B) Dollar–Sterling			
s_t^{gbp}	−6.43	$s_t^{\mathrm{gbp}} - f_{t,4}^{\mathrm{gbp}}$	−34.78
$f_{t,4}^{\mathrm{gbp}}$	−6.40	$s_t^{\mathrm{gbp}} - f_{t,13}^{\mathrm{gbp}}$	−29.68
$f_{t,13}^{\mathrm{gbp}}$	−6.39	$s_t^{\mathrm{gbp}} - f_{t,26}^{\mathrm{gbp}}$	−25.24
$f_{t,26}^{\mathrm{gbp}}$	−6.40	$s_t^{\mathrm{gbp}} - f_{t,52}^{\mathrm{gbp}}$	−22.14
$f_{t,52}^{\mathrm{gbp}}$	−6.63		
(C) Dollar–Mark			
s_t^{dm}	−2.86	$s_t^{\mathrm{dm}} - f_{t,4}^{\mathrm{dm}}$	−16.49
$f_{t,4}^{\mathrm{dm}}$	−2.88	$s_t^{\mathrm{dm}} - f_{t,13}^{\mathrm{dm}}$	−13.47
$f_{t,13}^{\mathrm{dm}}$	−2.95	$s_t^{\mathrm{dm}} - f_{t,26}^{\mathrm{dm}}$	−10.22
$f_{t,26}^{\mathrm{dm}}$	−3.08	$s_t^{\mathrm{dm}} - f_{t,52}^{\mathrm{dm}}$	−8.34
$f_{t,52}^{\mathrm{dm}}$	−3.31		

The sample is 1977:1–1996:52. The augmented Dickey–Fuller (ADF) Statistics are computed using a lag truncation parameter of 6. The null hypothesis is that the series possesses a single unit root. The 10% and 5% critical values are, respectively, −11.3 and −14.1.

Table 2
Unrestricted vector autoregression on spot and forward rate first differences: selection criteria

VAR order	Akaike criterion	Schwarz criterion
0	6975.3	7010.9
1	6865.6	7436.5
2	6891.5	7997.3

Note: The appropriate lag length for the VAR minimizes the criterion.

restrictions of individual variables from a particular equation are undertaken and a Wald test at the system level for the joint statistical significance of the variables excluded is reported. In particular, the value of the statistic, which is asymptotically distributed as chi-square, indicates that variables' exclusion cannot be rejected (149.1; p-value 0.37).

The FIML estimates lend strong support to the view that, in predicting the spot rate of a currency, the term structures of the other two currencies' forward premiums have incremental information content with respect to that of the own currency's premiums. The hypothesis that the forward premiums of two currencies, say, the dollar–yen and the dollar–mark, do not help to predict changes in the spot rate of the other currency (the dollar–sterling) is strongly rejected. A Wald test at the system level of the joint statistical significance of the forward premiums of the *other* currencies is conducted. The value of the statistic strongly rejects the hypothesis that these variables do not enter the corresponding equations of the system (87.22; p-value 0.001). We also focused on each of the three equations for the spot rate of a currency and test the hypothesis that the term structures of forward premiums of the other two currencies enter significantly in the equation. When the dollar–yen spot rate is to be predicted, the joint hypothesis that the term structures of dollar–sterling and dollar–mark forward premiums do not have predictive power is strongly rejected (Wald test at system level: 51.77; p-value 0.001). Similarly, in the two equations for the dollar–sterling and the dollar–mark spot rates, the term structures of the other two currencies' premiums enter significantly: the Wald tests at the system level are, respectively, 15.72 (with p-value of 0.02) and 28.98 (with a p-value of 0.001). An additional test refers to each currency's term structure individually. For each spot rate equation and for each other currency's term structure, we test the null hypothesis that the other currency's forward premiums at different maturities are jointly not significant: this hypothesis is also largely rejected. For example, in the dollar–yen spot equation, the Wald tests for the term structure of forward premiums is 23.64 for the dollar–sterling currency (p-value: 0.001) and 20.79 for the dollar–mark (p-value: 0.001). [4]

[4] An exception is found in the dollar–mark spot equation, where the Wald test for the joint significance of the dollar–Japanese premiums is 4.05, with a p-value of 0.26. Whilst the null hypothesis is not rejected, the values of the t-statistics suggest that the dollar–yen premiums have some predictive power in the equation (especially the 6- and 12-month forward premiums with t-statistics of 1.65 and −2.01, respectively). Moreover, the joint significance in that equation of the dollar–sterling premiums is very large (Wald test: 26.98 with a p-value of 0.00).

Table 3
FIML estimation of the error correction model for three exchange rates (results for spot rate changes; sample: 1977:01–1993:26)

Model for Δs_t^{gbp} (dollar–sterling)		Model for Δs_t^{dm} (dollar–mark)		Model for Δs_t^{jpy} (dollar–yen)	
Δs_{t-1}^{gbp}	−1.299 (0.762)	$\Delta f_{t-1,4}^{gbp}$	0.821 (0.371)	Δs_{t-1}^{gbp}	−0.294 (0.773)
$\Delta f_{t-1,4}^{gbp}$	2.206 (0.891)	$\Delta f_{t-1,13}^{gbp}$	−0.841 (0.367)	$\Delta f_{t-1,4}^{gbp}$	0.098 (0.832)
$\Delta f_{t-1,26}^{gbp}$	−1.500 (0.423)	Δs_{t-1}^{jpy}	−1.692 (0.660)	$\Delta f_{t-1,26}^{gbp}$	−0.175 (0.093)
$\Delta f_{t-1,52}^{gbp}$	0.591 (0.219)	$\Delta f_{t-1,4}^{jpy}$	2.022 (0.719)	$\Delta f_{t-1,52}^{gbp}$	0.367 (0.139)
Δs_{t-1}^{dm}	0.105 (0.093)	$\Delta f_{t-1,52}^{jpy}$	−0.291 (0.190)	$\Delta f_{t-1,26}^{dm}$	0.150 (0.044)
$\Delta f_{t-1,4}^{dm}$	−0.111 (0.093)	$(s_{t-1}^{gbp} - f_{t-1,4}^{gbp})$	0.146 (0.041)	$\Delta f_{t-1,52}^{dm}$	−0.184 (0.048)
Δs_{t-1}^{jpy}	−0.634 (0.390)	$(s_{t-1}^{gbp} - f_{t-1,13}^{gbp})$	−2.600 (1.115)	Δs_{t-1}^{jpy}	−0.823 (0.754)
$\Delta f_{t-1,13}^{jpy}$	1.128 (0.473)	$(s_{t-1}^{gbp} - f_{t-1,26}^{gbp})$	3.122 (1.227)	$\Delta f_{t-1,4}^{jpy}$	1.541 (0.820)
$\Delta f_{t-1,26}^{jpy}$	0.027 (0.021)	$(s_{t-1}^{gbp} - f_{t-1,52}^{gbp})$	0.214 (0.101)	$\Delta f_{t-1,52}^{jpy}$	−0.643 (0.184)
$\Delta f_{t-1,52}^{jpy}$	−0.461 (0.220)	$(s_{t-1}^{dm} - f_{t-1,4}^{dm})$	2.834 (0.884)	$(s_{t-1}^{gbp} - f_{t-1,4}^{gbp})$	0.207 (0.053)
$(s_{t-1}^{gbp} - f_{t-1,4}^{gbp})$	−0.353 (0.048)	$(s_{t-1}^{dm} - f_{t-1,13}^{dm})$	−2.880 (1.435)	$(s_{t-1}^{gbp} - f_{t-1,13}^{gbp})$	−0.313 (0.149)
$(s_{t-1}^{gbp} - f_{t-1,13}^{gbp})$	0.305 (0.080)	$(s_{t-1}^{dm} - f_{t-1,26}^{dm})$	−2.987 (0.409)	$(s_{t-1}^{gbp} - f_{t-1,26}^{gbp})$	−0.236 (0.175)
$(s_{t-1}^{gbp} - f_{t-1,26}^{gbp})$	0.042 (0.156)	$(s_{t-1}^{dm} - f_{t-1,52}^{dm})$	3.055 (0.890)	$(s_{t-1}^{gbp} - f_{t-1,52}^{gbp})$	0.100 (0.128)
$(s_{t-1}^{gbp} - f_{t-1,52}^{gbp})$	1.324 (0.345)	$(s_{t-1}^{jpy} - f_{t-1,13}^{jpy})$	−0.066 (0.075)	$(s_{t-1}^{dm} - f_{t-1,4}^{dm})$	−0.207 (0.090)
$(s_{t-1}^{dm} - f_{t-1,13}^{dm})$	0.036 (0.078)	$(s_{t-1}^{jpy} - f_{t-1,26}^{jpy})$	0.200 (0.121)	$(s_{t-1}^{dm} - f_{t-1,13}^{dm})$	−0.033 (0.114)

(continued on next page)

Table 3 (*continued*)

Model for $\Delta s_t^{\mathrm{gbp}}$ (dollar–sterling)		Model for Δs_t^{dm} (dollar–mark)		Model for $\Delta s_t^{\mathrm{jpy}}$ (dollar–yen)	
$(s_{t-1}^{\mathrm{dm}} - f_{t-1,26}^{\mathrm{dm}})$	−0.660 (0.251)	$(s_{t-1}^{\mathrm{jpy}} - f_{t-1,52}^{\mathrm{jpy}})$	−0.151 (0.075)	$(s_{t-1}^{\mathrm{dm}} - f_{t-1,26}^{\mathrm{dm}})$	0.003 (0.154)
$(s_{t-1}^{\mathrm{jpy}} - f_{t-1,4}^{\mathrm{jpy}})$	−0.012 (0.039)	Constant	−0.929 (1.198)	$(s_{t-1}^{\mathrm{dm}} - f_{t-1,52}^{\mathrm{dm}})$	−0.043 (0.100)
$(s_{t-1}^{\mathrm{jpy}} - f_{t-1,13}^{\mathrm{jpy}})$	−0.048 (0.106)			$(s_{t-1}^{\mathrm{jpy}} - f_{t-1,4}^{\mathrm{jpy}})$	−0.692 (0.049)
$(s_{t-1}^{\mathrm{jpy}} - f_{t-1,26}^{\mathrm{jpy}})$	0.257 (0.148)			$(s_{t-1}^{\mathrm{jpy}} - f_{t-1,13}^{\mathrm{jpy}})$	0.573 (0.112)
$(s_{t-1}^{\mathrm{jpy}} - f_{t-1,52}^{\mathrm{jpy}})$	−0.228 (0.085)			$(s_{t-1}^{\mathrm{jpy}} - f_{t-1,26}^{\mathrm{jpy}})$	0.403 (0.154)
Constant	−4.354 (1.276)			$(s_{t-1}^{\mathrm{jpy}} - f_{t-1,52}^{\mathrm{jpy}})$	1.056 (0.264)
				Constant	5.179 (1.062)
$Q = 10.06$ (12; p-val. 0.61)		$Q = 8.57$ (12; p-val. 0.74)		$Q = 13.61$ (12; p-val. 0.33)	

Wald Tests for restrictions: (a) forward premiums (fp) of gbp and dm currencies equal 0 in the Δs^{jpy} equation: 51.77 (8; p-val. 0.00); (b) fp's of jpy and gbp currencies equal 0 in the Δs^{dm} equation: 28.98 (7; p-val. 0.00); (c) fp's of jpy and dm currencies equal 0 in the Δs^{gbp} equation: 15.72 (6; p-val. 0.02).
Note: Standard errors are reported in parentheses. Q is the Ljung-Box test at the equation level. Vector multivariate Portmanteau statistic at 1 autocorrelation: 86.98 (100; p-val. 0.82). Wald test for general exclusion restriction: 149.1 (144; p-val. 0.37). Wald test for the restriction that all *other* currencies' premia equal 0: 87.22 (21; p-val. 0.00).

In Table 3, we also provide other evidence on the validity of the specification. For example, we report the value of the Vector Portmanteau statistic, testing for lack of multivariate serial correlation: it is 86.98 with a p-value of 0.82. Moreover, given the length of the sample period analyzed, we also addressed the issue of parameter stability. We identified two episodes that may have potentially provoked a structural break: the Plaza Agreement of September 22, 1985, whereby the finance ministers of the G5 countries agreed to encourage an ''orderly'' depreciation of the dollar, and the Louvre Accord of February 22, 1987, aimed at reducing exchange rate instability. For every week, we computed recursively the system Chow test over a period of three months after each episode. All values of the Chow test lead to non-rejection of the null hypothesis of parameter constancy (0.67 is the lowest p-value obtained).

3.3. Out-of-sample forecasts

In the present paper, we not only investigate the in-sample incremental predictive content of the forward premiums' term structures, as indicated by the theory, but we also examine the dynamic, out-of-sample forecasting performance of our empirical framework. In particular, we estimate the VECM recursively, adding each time a

Table 4
Out-of-sample forecasting performance

Forecasting horizon	RMSE		Theil decomposition of the MSEs					
			VECM			Random walk		
	VECM	Random walk	U^M	U^R	U^D	U^M	U^R	U^D
(a) Dollar–Yen spot rate: s_t^{jpy}								
4 weeks	0.0309	0.0316	0.00	0.78	0.22	0.00	0.79	0.21
13 weeks	0.0661	0.0669	0.02	0.94	0.04	0.01	0.95	0.04
26 weeks	0.0912	0.0956	0.04	0.94	0.02	0.01	0.97	0.02
52 weeks	0.1221	0.1343	0.16	0.83	0.01	0.02	0.97	0.01
(b) Dollar–Mark spot rate: s_t^{dm}								
4 weeks	0.0266	0.0260	0.00	0.75	0.25	0.01	0.73	0.26
13 weeks	0.0474	0.0442	0.01	0.91	0.08	0.02	0.89	0.09
26 weeks	0.0679	0.0619	0.04	0.92	0.04	0.07	0.89	0.04
52 weeks	0.1116	0.1003	0.09	0.90	0.01	0.16	0.82	0.02
(c) Dollar–Sterling spot rate: s_t^{gbp}								
4 weeks	0.0196	0.0172	0.21	0.53	0.26	0.02	0.64	0.34
13 weeks	0.0386	0.0255	0.48	0.46	0.06	0.06	0.80	0.14
26 weeks	0.0644	0.0363	0.56	0.42	0.02	0.07	0.86	0.07
52 weeks	0.1059	0.0464	0.64	0.35	0.01	0.12	0.84	0.04

Note: The forecast period is 1993:27–1996:52. The RMSE statistics use as inputs the dynamic forecasts of the spot rates in (log) level. U^M, U^R and U^D denote, respectively, the bias, the regression and the disturbance proportion in the Theil decomposition of the MSEs.

new observation that was previously left out of the estimation sample (1977:1–1993:26). The recursive procedure yields dynamic forecasts until the end of 1996 that only make use of information available up to the end of the estimation sample. Of course, as the latter shifts recursively, new information becomes available, but it is not used for the forecasting period. Table 4 reports the root mean square errors (RMSE) of dynamic forecasts from the VECM for each spot rate and at various forecasting horizon (4, 13, 26 and 52 weeks). We also compare the out-of-sample forecasting accuracy of the VECM with that of a random walk model. While the in-sample predictive content of the forward premiums was shown to be substantial for all exchange rates, the results from the out-of-sample exercise are, admittedly, less uniform. When the dollar–yen spot rate is predicted, the accuracy of forecasts from the VECM is larger than that from the random walk model, and this holds true for any forecasting horizon. For example, in the one-year horizon, RMSEs of the VECM and of the random walk model are, respectively, 0.122 and 0.134. On the other hand, an opposite pattern emerges in the case of the dollar–sterling and the dollar–mark rates, where the VECM does not improve upon the forecasting performance of a random walk model and the latter seems to have a higher accuracy. [5]

[5] By contrast, in predicting these two currencies' spot rates, the out-of-sample forecasting performance of the VECMs estimated by Clarida and Taylor (1997), based on a single currency set-up, remains superior to that of alternative models, including the random walk.

In light of these discrepancies, for the forecasts of both the VECM and the random walk model, we also provide a Theil decomposition of the MSEs into three components: bias, regression and disturbance. The bias component of the VECM forecasts for the dollar–yen and dollar–mark spot rates is small, suggesting that the magnitude of the MSE is not the consequence of a tendency to estimate too high or too low a level of the spot rate. This is less true, however, for the dollar–sterling spot rate. The disturbance component, i.e. the variance of residuals in the regression of the actual on the predicted spot rate change, is also small for both the VECM and the random walk, with the exception of the four-week horizon, where it is more sizeable. The regression proportion is generally the highest one in each decomposition and captures the extent to which, in the regression of the actual on the predicted spot rate, the estimated parameter differs from one.

It is not obvious why the forecasts of the VECM outperform those of the random walk for the dollar–yen but not for the other two currencies. To some extent, a similar finding is obtained by Mark (1995). In a frequently-cited paper, he characterizes short- and long-horizon changes in the exchange rates in terms of the current exchange rate's deviations from its "fundamental value", where the latter is defined as a linear combination of relative money stocks and relative real income. He shows that when forecasting horizons similar to ours are used (one and four quarters), the out-of-sample regression forecasts beat the random walk in the case of the dollar–yen but not in that of the dollar–dm (the dollar–sterling is not examined). We confirm his result, although, of course, direct comparison with this and other findings must be approached with caution in light of differences in the sample periods and the data used. In our paper the key explanatory factors are forward premiums, which, by covered interest parity, are approximately equal to interest rate differentials. Arguably, the latter variables may well mirror the monetary factors that, especially for the yen, were shown to be important in exchange rate forecasting, even at short-horizon changes.

4. Concluding remarks

The question addressed in the paper is whether, in predicting the spot rate of a currency, the term structures of forward premiums of both this currency and other currencies contain incremental useful information. The motivation underlying the analysis hinges on the abundant evidence on the existence of predictability and co-movements across excess returns from investing in different currencies. We propose a theoretical framework that allows for these co-movements and yields a number of testable implications on the predictive content of forward exchange premiums.

The empirical results, obtained by FIML estimation of a system related to three bilateral dollar exchange rates, strongly support the predictions of our theoretical model. We show that the term structures of other currencies' forward premiums are linearly informative about future movements of a spot rate, in addition to its own currency's term structure. This holds true, by and large, for all exchange rates and all contract lengths. While the evidence in favor of the in-sample predictive power is strong, the results of an out-of-sample forecasting exercise are mixed. The forecasting performance of our model largely improves on that of a random walk for the dollar–yen spot rate but not for the dollar–sterling and the dollar–mark.

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