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# Tests for a Systematic Risk Component in Deviations From Uncovered Interest Rate Parity

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In the intertemporal asset pricing model, investments in spot foreign currencies involve time-varying risk proportional to the conditional covariance of the value of the position with the intertemporal marginal rate of substitution of domestic currency. We detect such risk premia in deviations from uncovered interest rate parity using weekly spot currency prices and Eurocurrency interest rates. Our tests use the conditional capital asset pricing model with a world equity index as benchmark to represent aggregate wealth.

## 1. INTRODUCTION

Traditional structural models of spot exchange rate determination have had difficulty out-performing a martingale model for forecasting nominal exchange rates (see, for example, Meese and Rogoff (1983)). A recent alternative to the martingale process is provided by the intertemporal asset pricing model (IAPM). According to this model, domestic investors who hold open positions in foreign currencies may face time-varying risk that is proportional to the conditional covariance of the value of the position with the intertemporal marginal rate of substitution of domestic currency.

Although the hypothesis that forward exchange rates are unbiased predictors of future spot rates has usually been rejected, most existing models of time-varying risk premia in foreign currency markets have met with limited empirical success. Since the unbiasedness hypothesis jointly maintains no risk premium and the particular model or asset-pricing paradigm used to generate the expected future spot rates, its rejection could be explained by a time-varying risk premium or by several other hypotheses, such as rational learning about stochastic regime switches, speculative bubbles, the 'peso problem', or the failure of the maintained rational expectations assumption. Recent work has concentrated on generalizing the model that produced the hypothesis of unbiasedness to allow direct empirical testing of one or more of the other hypotheses.[1]

In this paper, we focus on a time-varying risk alternative hypothesis. Examples of models of risk premia that have been applied to foreign currency data include: those using international financial relations such as interest rate parity, purchasing power parity and the Fisher real interest equation to relate differences in expected real interest rates to risk premia in forward markets (Korajczyk (1985)); constant beta models (Cosset (1984)); international portfolio balance models (Frankel (1982)), non-parametric approaches (Meese and Rose (1991), Wickens and Thomas (1989)); and representative agent intertemporal asset pricing models (IAPM).

1. For details, see Boothe and Longworth (1986), Hodrick (1987), Baillie and McMahon (1989), Lewis (1989), Meese (1989), Obstfeld (1989), and Engel and Hamilton (1990).

The IAPM was extended to price nominal assets by Stulz (1981), Hodrick (1981), Lucas (1982), Hansen and Hodrick (1983), and others. Many applications of the model to financial forward markets followed. Some have used the conditional variance (Domowitz and Hakkio (1985)) or the conditional variance-covariance matrix (Baillie and Bollerslev (1990)) of forward rate forecast errors to measure the time-varying risk premia. The international portfolio models with time-varying conditional variances and covariances (for example, Diebold and Pauly (1988), Engel and Rodrigues (1989) and Giovannini and Jorion (1989)) are also derived from a dynamic setting. Analogously, conditional beta intertemporal asset pricing models have been estimated using forward (Mark (1988)) and futures (McCurdy and Morgan (1990)) data. Others have used a specific structure on preferences and the distribution of either the exogenous or the endogenous processes to test the particular form of the risk premium implied by the consumption-based IAPM (for example, Mark (1985), Cumby (1988), Hodrick (1989), Backus, Gregory and Telmer (1990), Kaminsky and Peruga (1990)).

In this paper, we use foreign currency spot prices and Eurocurrency interest rates to construct a time series of weekly excess returns on an uncovered foreign currency position which we refer to as deviations from uncovered interest rate parity (UIRP).[2] The conditional capital asset-pricing model is used to test for any non-diversifiable component of that series. Under certain assumptions, there is a direct relationship between the conditional covariance risk derived from the consumption-based IAPM and the conditional systematic risk associated with the conditional beta formulation of the IAPM. The theory predicts interdependence between the conditional mean of deviations from uncovered interest rate parity and the relevant conditional covariance with the benchmark portfolio. Our econometric specification allows the conditional risk premia to reflect time variation in the covariance components of the currency betas and in the conditional variance and expected excess return from the benchmark portfolio. For example, unlike Bollerslev, Engle and Wooldridge (1987), Engel and Rodrigues (1989) and Giovannini and Jorion (1989), we do not impose the restriction that the price of covariance risk is constant over time. Like Mark (1988), we choose an internationally diversified equity portfolio as benchmark but our work differs from his by the use of instruments to predict the return from this portfolio and by the GARCH formulation (Engle (1982), Bollerslev (1986)), instead of ARCH in specifying the conditional moments.

We find that interest rate differentials (differences between the nominal interest rates of the U.S. and foreign countries) have predictive power for the excess returns on the world equity index. This result parallels those found by several authors for the U.S. market in isolation. Fama and Schwert (1977), Keim and Stambaugh (1986), Campbell (1987, 1990), and Fama and French (1989), have all identified some function of US interest rates, in particular, term premia, default premia (junk bond spreads) and a relative bill rate (the 30-day Treasury Bill rate relative to its previous 1-year moving average) as variables that account, predictively, for some proportion of the variation of stock market monthly returns. Harvey (1990, Table III) found, using monthly data, that U.S. term and default premia are also useful for predicting the world equity index return. In our case, when the difference between the U.S. and the average foreign interest rate is included,

2. In other words, we construct "synthetic" forward positions, rather than using actual forward market data. Some examples of earlier analyses of deviations from uncovered interest rate parity include Cumby and Obstfeld (1981), Hsieh (1984), Gregory (1987) and Diebold and Pauly (1988). Goodhart (1988, p. 438) has argued that "speculation in the foward exchange market in particular is of negligible importance; there is much more speculation in the form of the adoption of open spot positions". As long as covered interest rate parity obtains, forward premia and interest rate differentials should be equalized.

the improved estimate of the expected excess return on wealth enhances the ability of the conditional beta model to detect ex ante deviations from uncovered interest rate parity, strengthening the evidence for systematic risk associated with uncovered positions in foreign currencies.

Using an average interest rate differential as an instrument for predicting the benchmark portfolio excess return, we find evidence in favour of the conditional beta measure of time-varying risk premia for all five currencies. Two currencies exhibit a significant test statistic associated with inappropriate exclusion of a term for their conditional variances, as in the GARCH-M model of Bollerslev, Engle and Wooldridge (1987). Tests for omitted variables also show that the difference between the U.S. and the individual foreign currency interest rate, or the difference between the latter and an average foreign currency interest rate, would add explanatory power to the equation for the conditional mean of two currencies. These results could indicate sources of risk additional to those captured by the single beta formulation of the conditional capital asset-pricing model.

The paper is organized as follows. Section 2 reviews the intertemporal nominal asset pricing paradigm used to value an open foreign currency position. Section 3 presents the test equations for risk premia associated with those positions. Section 4 discusses the data and the results and Section 5 offers some concluding comments.

## 2. THEORETICAL BACKGROUND

### 2.1. *Evaluation of positions in foreign currencies*

Let

$C_t =$ the number of units of the good consumed at $t$,

$p_t =$ the price per unit of the consumption good at $t$,

$M_{t-1,t} =$ the intertemporal marginal rate of substitution of domestic currency between time $t-1$ and time $t$,

$S_t =$ the spot price at $t$ of one unit of the foreign currency,

$R_{t-1} =$ one plus the US riskless nominal rate of interest from $t-1$ to $t$,

$Z_{t-1} =$ one plus the foreign riskless nominal rate of interest from $t-1$ to $t$,

$R_{st} =$ $\dfrac{S_t}{S_{t-1}} Z_{t-1} =$ rate of return in dollars from an investment in the foreign asset,

$R^*_{st} =$ $R_{st} - R_{t-1} =$ excess return on the uncovered foreign currency position or deviations from UIRP,

$R_{Bt} =$ rate of return on a benchmark portfolio which is postulated to be conditionally mean variance efficient,

$R_{wt} =$ rate of return on the Morgan Stanley Capital International (MSCI) world equity index.

First-order conditions for maximum expected utility lead to stochastic Euler conditions, in which $E_{t-1}$ refers to expectations conditional on information $I_{t-1}$,

$$1 = R_{t-1} E_{t-1}[M_{t-1,t}], \tag{1}$$

for the present value of the one-period cash flow generated from a dollar invested in a nominally riskless domestic asset, and

$$1 = E_{t-1}[M_{t-1,t}R_{st}], \tag{2}$$

for that generated by a foreign asset.

Sufficient conditions for (1) and (2) to hold include an additive time-separable multi-period utility function with a constant time-preference factor, $\delta$. In this case, the intertemporal marginal rate of substitution is

$$M_{t-1,t} \equiv \delta \frac{u'(C_t)}{u'(C_{t-1})} \frac{p_{t-1}}{p_t}.$$

However, as has been noted by several authors, for example, Dunn and Singleton (1986), these assumptions are stronger than necessary.

From (1) and the definition of covariance applied to (2),

$$E_{t-1}[R_{st}] - R_{t-1} = -R_{t-1} cov_{t-1}[M_{t-1,t}, R_{st}]. \tag{3}$$

The conditionally expected excess return or nominal risk premium associated with an uncovered position in the foreign currency is proportional to the conditional covariance of the spot price with the intertemporal marginal rate of substitution of domestic currency. This would be zero under risk neutrality and a deterministic price level. Stochastic changes in the purchasing power of the domestic currency imply that risk neutrality is not sufficient for the conditional covariance in (3) to be zero.[3]

Under risk aversion, the risk premium (or expected profit from an uncovered long position) will be positive when the conditional covariance between $M_{t-1,t}$ and $S_t$ is negative. In the time-separable case, the conditional covariance will be negative if, for example, the position has a high payoff when the marginal utility from a dollar's worth of consumption, $u'(C_t)/p_t$, is low—either because consumption is high (marginal utility is low) or the purchasing power of the dollar is low (domestic price level is high). Conversely, the risk premium will be negative if the conditional covariance in (3) is positive. In this case, the expected return on the foreign investment is lower in equilibrium than that on the domestic investment because fluctuations in the spot price of foreign currency are such that the foreign investment provides a hedge against adverse consumption outcomes.

### 2.2. *A testable form of the model*

Empirical implementation of asset pricing relations such as (3) has taken several forms. Since measurement of consumption and price-level data is problematic for data observed at intervals as short as a day or a week, it is useful to re-express the consumption-based (3) in terms of a conditional capital asset pricing model for which all component returns are available at the desired frequency.[4] Following Hansen and Richard (1987), the asset pricing relation is expressed in terms of a benchmark portfolio on the conditional mean

3. Engel (1990) analysed the relationship between risk premia in real terms and nominal excess returns.

4. Another reason why such models might receive more empirical support than the corresponding consumption-based model is provided by Epstein and Zin (1989) and Giovannini and Weil (1989). With a specification of preferences allowing identification of separate parameters relating to risk aversion and intertemporal substitution, a unit elasticity of intertemporal substitution implies myopic choice of consumption and savings while relative risk aversion of unity results in myopic portfolio allocation. Giovannini and Weil (1989) claimed that the empirical evidence is more consistent with the latter, supporting the (conditional) CAPM-type specification.

variance frontier. Breeden (1979) and Hansen and Hodrick (1983) postulated the existence of a portfolio with nominal return $R_{mt}$ perfectly conditionally correlated with $M_{t-1,t}$. Portfolios with returns $R_{Bt}$ that are linear combinations of $R_{mt}$ and the riskless rate will be conditionally mean-variance efficient. The equilibrium expected return on any asset is a function of its conditional beta with that benchmark portfolio. We can then re-express (3) as a conditional beta asset-pricing relation for which the conditional covariance in (3) is replaced by a quantity that is more easily measured in data measured at short intervals, obtaining

$$E_{t-1}[R^*_{st}] = \frac{cov_{t-1}[R^*_{Bt}, R^*_{st}]}{var_{t-1}[R^*_{Bt}]} E_{t-1} R^*_{Bt}, \tag{4}$$

in which $R^*$ indicates a return in excess of the riskless rate.

Since we wish to capture time variation in the intertemporal marginal rate of substitution $M_{t-1,t}$ using the conditional moments of the returns from a benchmark portfolio, the assumption of a return which is perfectly conditionally correlated with $M_{t-1,t}$ is stronger than necessary. Breeden, Gibbons and Litzenberger (1989) and McCurdy and Morgan (1990) construct benchmark portfolios whose returns are maximally correlated with the growth rate of consumption. Others, for example, Campbell (1987), Giovannini and Jorion (1987), Korajczyk and Viallet (1990), and Engle, Ng and Rothschild (1990) treat the benchmark return as unobservable and use either a latent variable approach or factor representing portfolios to estimate the benchmark portfolio returns. In this paper, we use the return on the Morgan Stanley Capital International (MSCI) world equity index as the benchmark portfolio, replacing $R^*_{Bt}$ in (4) by $R^*_{wt}$. Choosing an observable equity index, as in Mark (1988) and Harvey (1990), is clearly subject to the critique of Roll (1977). However, our use of the MSCI world equity index is intended to recognize extensive international diversification as important for positions denominated in foreign currencies. Nevertheless, it is possible that systematic risk with respect to the world equity benchmark portfolio will not price all the relevant risk. Our tests for missing variables are conducted, in part, to detect whether additional factors are present in the data.

## 3. TEST EQUATIONS

The risk premium in (4) is the product of the conditional beta of deviations from UIRP and the conditional expected excess return from the benchmark portfolio. The conditional beta consists of the conditional covariance of the deviations from UIRP with the excess return on the benchmark portfolio divided by the conditional variance of the latter. The test equations must allow those time-varying second moments, as well as the conditional expected return on the benchmark portfolio, to influence deviations from UIRP. To accomplish this, we pair the deviations from UIRP series for a given currency with the benchmark portfolio excess return series and use a GARCH model for the conditional heteroscedasticity in the innovations. Following Baba, Engle, Kraft and Kroner (1989), we use a positive-definite parameterization of the bivariate GARCH model.

To obtain a test equation from (4), we replace rational expectations by the realised values minus forecast errors. The rational expectations assumption implies that the forecast errors have a conditional mean of zero. We specify a moving-average process for the test equation residuals of the benchmark portfolio excess return series to capture any potential non-synchronized trading effects associated with the index.

Using the notation $h_{st}$ for the conditional variance of $R^*_{st}$, $h_{wt}$ for the conditional variance of $R^*_{wt}$, $h_{swt}$ for the conditional covariance between the two excess returns, $\mu$

for a multiplicative parameter that can be chosen to have a value of zero to exclude the risk premium term from the model, and $x_{s,t-1}$, $x_{w,t-1}$, $g_{s,t-1}$, $g_{w,t-1}$ and $g_{sw,t-1}$ for vectors of explanatory variables known at time $t-1$, the system of test equations is

$$R^*_{st} = \gamma'_s x_{s,t-1} + \mu \frac{h_{swt}}{h_{wt}} (\gamma'_w x_{w,t-1} + \psi_w \varepsilon_{w,t-1}) + \varepsilon_{st}, \tag{5}$$

$$R^*_{wt} = \gamma'_w x_{w,t-1} + \psi_w \varepsilon_{w,t-1} + \varepsilon_{wt}, \tag{6}$$

$$\varepsilon_t \mid I_{t-1} \sim N(0, H_t).$$

In (6), for symmetric matrices $C$, $A$, $B$ and $\Phi$

$$H_t = C + A'\varepsilon_{t-1}\varepsilon'_{t-1}A + B'H_{t-1}B + \Phi_{t-1}, \tag{7}$$

$$\begin{bmatrix} h_{st} & h_{wst} \\ h_{wst} & h_{wt} \end{bmatrix} = \begin{bmatrix} c_s & c_{ws} \\ c_{ws} & c_w \end{bmatrix} + \begin{bmatrix} a_s & a_{ws} \\ a_{ws} & a_w \end{bmatrix} \begin{bmatrix} \varepsilon^2_{s,t-1} & \varepsilon_{w,t-1}\varepsilon_{s,t-1} \\ \varepsilon_{w,t-1}\varepsilon_{s,t-1} & \varepsilon^2_{w,t-1} \end{bmatrix} \begin{bmatrix} a_s & a_{ws} \\ a_{ws} & a_w \end{bmatrix}$$

$$+ \begin{bmatrix} b_s & b_{ws} \\ b_{ws} & b_w \end{bmatrix} \begin{bmatrix} h_{s,t-1} & h_{ws,t-1} \\ h_{ws,t-1} & h_{w,t-1} \end{bmatrix} \begin{bmatrix} b_s & b_{ws} \\ b_{ws} & b_w \end{bmatrix} + \begin{bmatrix} \phi'_s g_{s,t-1} & \phi'_{sw} g_{sw,t-1} \\ \phi'_{sw} g_{sw,t-1} & \phi'_w g_{w,t-1} \end{bmatrix}.$$

The dependent variable in (5) is the ex post excess return on an uncovered foreign currency position which, under the null hypothesis of the model, should equal the conditional nominal risk premium plus a rational expectations forecast error. The conditional risk premium is the product of the conditional beta and the conditional expected return on the benchmark portfolio. The vector of explanatory variables, $x_{s,t-1}$ includes an intercept in the estimated version of the maintained model. Otherwise, $x_{s,t-1}$ is used in the tests for omitted variables which might have explanatory power under alternative specifications of the time-varying risk premium model.

Our first specification of an empirical model for $R^*_{wt}$ included an intercept and $h_{wt}$, the conditional variance of the world equity return, as indicated by the CAPM.[5] However, as demonstrated in the empirical results below, modelling $R^*_{wt}$ as a function of $h_{wt}$ was dominated in our sample by the structure given by (6). This specification includes an MA (1) term to capture the effects of non-synchronized trades of the components of the index, and $x_{w,t-1}$ consisting of an intercept and the interest rate differential computed from the average of the foreign rates, $(R_{t-1}/\bar{Z}_{t-1}) - 1$.

The vectors $g_{w,t-1}$, $g_{s,t-1}$ and $g_{sw,t-1}$, which can be used to augment the basic structure for the conditional variances and covariances $h_{wt}$, $h_{st}$ and $h_{swt}$, are also used to test for omitted variables. The estimated version of the maintained model for $h_{wt}$ includes an indicator variable $C_t$ which takes the value 1 for the week of the October 1987 market crash and zero otherwise.

In sum, in this specification the risk premia vary as a result of variation in the expected excess returns from the benchmark portfolio and in the variance and covariance components of the currency betas. Maximum likelihood estimation is implemented with a Numerical Algorithms Group optimization routine. Joint estimation of the first and

5. Rearranging (4) and using (3) and (1) to substitute out $E_{t-1}[R^*_{st}]$, it is possible to express the expected excess return on the benchmark portfolio as a function of its conditional variance and $E_{t-1}[M_{t-1,t}]$. (See Campbell (1987)). Variables in addition to $h_{wt}$ may be useful in capturing the time-variation in $R^*_{wt}$.

second conditional moments is particularly attractive for this application because of interdependence between the conditional moments. We avoid a possible cost of using MLE, in terms of non-robustness, by using quasi-maximum likelihood estimation (White (1982), Weiss (1986), Bollerslev and Wooldridge (1988)) which allows inference in the presence of departures from conditional normality. Robust standard errors are computed from the diagonal elements of the matrix $J^{-1}KJ^{-1}$ where $J$ is the numerical approximation to the matrix of second derivatives with respect to the free variables and $K$ is the numerical estimate of the information matrix, formed by taking the average of the period-by-period outer products of the gradient. All standard errors quoted in the paper are robust.

## 4. DATA AND EMPIRICAL RESULTS

Daily observations on foreign currency spot prices and Eurocurrency 7-day interest rates were used to construct a time-series of weekly deviations from uncovered interest rate parity. The spot prices and the 7-day Eurocurrency interest rates are the arithmetic average of the London close bid and ask rates obtained from Reuters. Spot prices are expressed as the U.S. dollar price of a unit of foreign currency. Wednesday-to-Wednesday rates of change of price were computed when both prices were available. If the Wednesday price was missing in a given week, the Thursday price was substituted for the final price for that week and for the initial price for the following week. The 7-day Eurocurrency interest rates were converted from annualized rates to rates applicable to 7 days. When any interest rate was unavailable because of holidays, we substituted all interest rates for the previous day. The weekly benchmark portfolio returns were computed from the MSCI daily world equity index valued in U.S. dollars. No attempt was made to incorporate dividends in the returns because the available dividend data are accumulated over monthly intervals instead of being recorded for each day. The MSCI index was constructed from closing values of the nineteen country component indices. Our primary reason for sampling weekly instead of daily was to lessen the complications arising from the closing of stock markets in these countries at different times in the global day.

The sample covers 1980 to 1988 inclusive. The observations for the first week were used for construction of regressors for the conditional means and variances so that the sample size is 469 weeks. The interest rate differential was scaled by multiplication by 1000. All other variables, except the intercepts, were scaled by multiplication by 10.

The empirical results are summarized in Tables 1 to 4. Table 1 presents the coefficient estimates with the associated robust standard errors. Table 2 gives the results of a battery of diagnostic tests on the standardized residuals from the model reported in Table 1. Table 3 evaluates the model for potentially important omitted variables. Finally, Table 4 provides the statistical evidence from several tests investigating the significance of the estimated conditional risk premia.

A model that allows the risk premia to change signs is likely to be particularly relevant for foreign currencies. The attractive feature of the Baba, Engle, Kraft and Kroner (1989) specification for the conditional variances and covariances is that it allows the conditional covariances to change sign over time while preserving positive definiteness. Joint estimates of the bivariate system of equations (5) to (7) were obtained for each currency individually in five separate analyses. This procedure has the potential disadvantage that, in principle, there may be efficiency gains from treating the set of five currencies and the benchmark portfolio as one system. However, such a model would have a very large number of parameters and be extremely difficult to estimate.

TABLE 1

*Estimates for deviations from uncovered interest rate parity and the benchmark portfolio*

Deviations from uncovered interest rate parity

$$R^*_{st} = \gamma_{0s} + \mu \frac{h_{swt}}{h_{wt}}(\gamma_{0w} + \gamma_{1w}(R_{t-1}/\bar{Z}_{t-1} - 1) + \psi_{1w}\varepsilon_{w,t-1}) + \varepsilon_{st}$$

Benchmark

$$R^*_{wt} = \gamma_{0w} + \gamma_{1w}(R_{t-1}/\bar{Z}_{t-1} - 1) + \psi_{1w}\varepsilon_{w,t-1} + \varepsilon_{wt}$$

| | $\gamma_{0s}$ | $\mu$ | $R^2$ | $\gamma_{0w}$ | $\gamma_{1w}$ | $\psi_{1w}$ | $R^2$ |
|---|---|---|---|---|---|---|---|
| BP | −0·011 | 1·956 | 0·030 | 0·037 | −0·093 | 0·040 | 0·042 |
| | (0·007) | (0·512) | | (0·009) | (0·019) | (0·047) | |
| CD | −0·001 | 1·638 | 0·010 | 0·042 | −0·096 | 0·027 | 0·041 |
| | (0·003) | (0·692) | | (0·009) | (0·021) | (0·051) | |
| DM | −0·021 | 1·434 | 0·018 | 0·035 | −0·092 | 0·082 | 0·042 |
| | (0·006) | (0·445) | | (0·009) | (0·020) | (0·053) | |
| JY | −0·009 | 2·177 | 0·039 | 0·036 | −0·084 | 0·101 | 0·041 |
| | (0·008) | (0·656) | | (0·010) | (0·025) | (0·048) | |
| SF | −0·024 | 1·207 | 0·014 | 0·035 | −0·089 | 0·071 | 0·042 |
| | (0·007) | (0·428) | | (0·009) | (0·021) | (0·055) | |

$$h_{st} = c_s + a_s^2\varepsilon^2_{s,t-1} + 2a_s a_{sw}\varepsilon_{s,t-1}\varepsilon_{w,t-1} + a^2_{sw}\varepsilon^2_{w,t-1} + b_s^2 h_{s,t-1} + 2b_s b_{sw} h_{sw,t-1} + b^2_{sw} h_{w,t-1}$$

$$h_{swt} = c_{sw} + a_s a_{sw}\varepsilon^2_{s,t-1} + (a_s a_w + a^2_{sw})\varepsilon_{s,t-1}\varepsilon_{w,t-1} + a_w a_{sw}\varepsilon^2_{w,t-1} + b_s b_{sw} h_{s,t-1} + (b_s b_w + b^2_{sw}) h_{sw,t-1} + b_w b_{sw} h_{w,t-1}$$

$$h_{wt} = c_w + a_w^2\varepsilon^2_{w,t-1} + 2a_w a_{sw}\varepsilon_{s,t-1}\varepsilon_{w,t-1} + a^2_{sw}\varepsilon^2_{s,t-1} + b_w^2 h_{w,t-1} + 2b_w b_{sw} h_{sw,t-1} + b^2_{sw} h_{s,t-1} + \phi_{1w} C_1$$

| | $c_s$ | $a_s$ | $b_s$ | $c_{sw}$ | $a_{sw}$ | $b_{sw}$ | $c_w$ | $a_w$ | $b_w$ | $\phi_{1w}$ |
|---|---|---|---|---|---|---|---|---|---|---|
| BP | 0·0008 | 0·262 | 0·947 | 0·0011 | −0·0545 | 0·0112 | 0·0042 | 0·353 | 0·861 | 0·0042 |
| | (0·0004) | (0·037) | (0·011) | (0·0003) | (0·0260) | (0·0678) | (0·0020) | (0·038) | (0·041) | (0·0026) |
| CD | 0·0006 | 0·429 | 0·817 | 0·0002 | −0·0029 | 0·0092 | 0·0054 | 0·291 | 0·856 | 0·0049 |
| | (0·0004) | (0·099) | (0·098) | (0·0002) | (0·0174) | (0·0144) | (0·0026) | (0·071) | (0·038) | (0·0025) |
| DM | 0·0050 | 0·414 | 0·801 | 0·0017 | −0·0592 | 0·0219 | 0·0037 | 0·358 | 0·864 | 0·0040 |
| | (0·0020) | (0·064) | (0·059) | (0·0005) | (0·0369) | (0·0836) | (0·0016) | (0·036) | (0·033) | (0·0028) |
| JY | 0·0024 | 0·295 | 0·885 | 0·0009 | −0·0692 | 0·0341 | 0·0037 | 0·375 | 0·854 | 0·0039 |
| | (0·0047) | (0·105) | (0·148) | (0·0007) | (0·0471) | (0·1845) | (0·0017) | (0·049) | (0·038) | (0·0028) |
| SF | 0·0032 | 0·336 | 0·885 | 0·0014 | −0·0596 | 0·0225 | 0·0040 | 0·359 | 0·857 | 0·0039 |
| | (0·0039) | (0·090) | (0·099) | (0·0006) | (0·0376) | (0·0974) | (0·0017) | (0·040) | (0·035) | (0·0025) |

*Note.* Standard errors shown in parenthesis are robust. $R^2$ is the proportion of variation explained for the excess returns $R^*_{st}$ and $R^*_{wt}$. $C_t$ is an indicator variable taking the value 1 on the week of the market crash.

TABLE 2

*Diagnostic checks on models in Table 1*

| | $r_s$ | $Q_s(10)$ | $Q_s^2(10)$ | $S_s$ | $K_s$ | $P_s$ | $r_w$ | $Q_w(10)$ | $Q_w^2(10)$ | $S_w$ | $K_w$ | $P_w$ | $Q_{sw}(10)$ | $P_{sw}$ |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| BP | −0·81 | 8·45 | 4·20 | 0·44 | 7·42 | 0·17 | −0·78 | 7·56 | 13·88 | 0·03 | 1·33 | 0·34 | 18·94 | −0·19 |
| | (0·38) | (0·58) | (0·94) | (0·51) | (0·01) | (0·87) | (0·44) | (0·67) | (0·18) | (0·86) | (0·25) | (0·75) | (0·04) | (0·85) |
| CD | 2·46 | 4·93 | 5·33 | 0·72 | 11·98 | −1·85 | −0·78 | 7·80 | 22·10 | 0·39 | 1·50 | 0·72 | 7·47 | −2·88 |
| | (0·01) | (0·90) | (0·87) | (0·40) | (0·00) | (0·06) | (0·44) | (0·65) | (0·01) | (0·53) | (0·22) | (0·48) | (0·59) | (0·00) |
| DM | −1·86 | 8·00 | 12·85 | 8·13 | 1·19 | −1·64 | −0·69 | 7·65 | 15·52 | 0·01 | 0·87 | 0·28 | 14·67 | −0·35 |
| | (0·06) | (0·63) | (0·23) | (0·00) | (0·28) | (0·10) | (0·49) | (0·66) | (0·11) | (0·92) | (0·35) | (0·78) | (0·14) | (0·75) |
| JY | 0·28 | 21·24 | 5·78 | 15·29 | 8·69 | −0·23 | 0·51 | 8·03 | 18·83 | 0·08 | 0·89 | 0·29 | 5·89 | 0·23 |
| | (0·78) | (0·02) | (0·83) | (0·00) | (0·00) | (0·82) | (0·61) | (0·63) | (0·04) | (0·78) | (0·35) | (0·77) | (0·82) | (0·82) |
| SF | −1·38 | 9·31 | 8·39 | 13·01 | 0·00 | −0·55 | −0·23 | 7·40 | 18·50 | 0·02 | 1·02 | 0·35 | 23·60 | −0·26 |
| | (0·17) | (0·50) | (0·59) | (0·00) | (0·99) | (0·58) | (0·82) | (0·69) | (0·05) | (0·89) | (0·31) | (0·73) | (0·01) | (0·80) |

*Note.* $r$ is the test statistic for runs above the mean; $Q(10)$ the Ljung–Box (1978) form of the portmanteau stastic for autocorrelation in the first 10 lags of the standardized residuals, $Q^2(10)$ the same for the squared standardized residuals, and $Q_{sw}(10)$ the same for the cross products of the standardized residuals; $S$ and $K$ the (Newey (1985), Tauchen (1985)) conditional moment test statstics for skewness and kurtosis, respectively, and $P$ is the Pagan and Sabau (1987) test statistic computed from robust standard errors. The $p$-values, shown in parentheses, are for the chi-square distribution except for $r$ and $P$, where they are for the unit normal distribution.

TABLE 3

*OPG-LM tests for omitted variables*

| | BP | CD | DM | JY | SF |
|---|---|---|---|---|---|
| In the conditional mean for $R^*_{st}$ | | | | | |
| $R^*_{s,t-1}$ | 0·21 | 1·46 | 0·09 | 0·23 | 0·06 |
| | (0·65) | (0·23) | (0·76) | (0·63) | (0·81) |
| $R^*_{w,t-1}$ | 0·02 | 3·51 | 0·07 | 0·73 | 0·06 |
| | (0·89) | (0·06) | (0·79) | (0·39) | (0·81) |
| $(R_{t-1}/\bar{Z}_{t-1})-1$ | 1·25 | 1·80 | 0·21 | 6·93 | 0·10 |
| | (0·26) | (0·18) | (0·65) | (0·01) | (0·75) |
| $(R_{t-1}/Z_{t-1})-1$ | 7·87 | 16·15 | 0·11 | 1·43 | 0·05 |
| | (0·01) | (0·00) | (0·74) | (0·23) | (0·82) |
| $(Z_{t-1}/\bar{Z}_{t-1})-1$ | 16·79 | 9·91 | 0·09 | 0·09 | 0·03 |
| | (0·00) | (0·00) | (0·76) | (0·76) | (0·86) |
| $h_{st}$ | 0·88 | 3·80 | 12·17 | 1·32 | 9·95 |
| | (0·25) | (0·05) | (0·00) | (0·25) | (0·00) |
| In the conditional mean for $R^*_{wt}$: | | | | | |
| $h_{wt}$ | 1·35 | 1·03 | 0·99 | 0·06 | 0·30 |
| | (0·25) | (0·31) | (0·32) | (0·81) | (0·58) |
| In the conditional variance $h_{st}$ | | | | | |
| $C_t$ | 0·62 | 0·52 | 0·57 | 1·04 | 1·18 |
| | (0·43) | (0·47) | (0·45) | (0·31) | (0·28) |
| $(R^*_{s,t-1})^2$ | 1·79 | 6·57 | 1·32 | 0·66 | 1·78 |
| | (0·18) | (0·01) | (0·25) | (0·42) | (0·18) |
| $(R^*_{w,t-1})^2$ | 0·57 | 3·76 | 2·99 | 6·34 | 3·25 |
| | (0·45) | (0·05) | (0·08) | (0·01) | (0·05) |
| $\lvert(Z_{t-1}/\bar{Z}_{t-1})-1\rvert$ | 0·99 | 1·79 | 0·87 | 0·44 | 4·09 |
| | (0·32) | (0·18) | (0·35) | (0·51) | (0·04) |
| In the conditional covariance $h_{swt}$ | | | | | |
| $C_t$ | 1·31 | 6·40 | 2·67 | 0·07 | 2·98 |
| | (0·25) | (0·01) | (0·10) | (0·79) | (0·08) |
| $(Z_{t-1}/\bar{Z}_{t-1})-1$ | 0·16 | 0·74 | 0·08 | 1·21 | 0·26 |
| | (0·69) | (0·39) | (0·78) | (0·27) | (0·61) |

*Note.* $p$-values, for the chi-square distribution with one degree of freedom are shown in parentheses.

TABLE 4

*Evidence concerning risk premia*

| | BP | CD | DM | JY | SF |
|---|---|---|---|---|---|
| $c_{sw}=a_{sw}=b_{sw}=0$ | 22·86 | 12·02 | 24·80 | 29·62 | 21·70 |
| | (0·000) | (0·007) | (0·000) | (0·000) | (0·000) |
| $\mu=0$ | 13·40 | 5·88 | 10·12 | 17·14 | 7·24 |
| | (0·000) | (0·015) | (0·001) | (0·000) | (0·007) |
| $\gamma_{0s}=0,\ \mu=1$ | 5·14 | 0·96 | 9·12 | 6·42 | 9·66 |
| | (0·077) | (0·619) | (0·010) | (0·040) | (0·010) |
| Average risk premium | −0·54 | 0·93 | −7·16 | 2·89 | −8·55 |
| Standard deviation | (13·31) | (3·59) | (12·59) | (16·10) | (11·60) |
| Average realized $R^*_{st}$ | −0·42 | 0·59 | −3·33 | 4·31 | −4·37 |
| Standard deviation | (84·67) | (33·10) | (84·07) | (77·83) | (92·22) |

*Note.* The first three rows report LR test statistics with $p$-values for the chi-square distribution in parentheses. The average excess returns reported in the last two rows are expressed as annual percentages.

Coefficient estimates and standard errors are given in Table 1. For presentation purposes, the conditional mean and conditional variance-covariance panels of Table 1 group the estimates for the five currencies. The conditional mean for $R^*_{st}$ includes an intercept and the conditional risk premium as indicated by (5). In theory, the parameter $\mu$ should be unity. From a $t$-test, this cannot be rejected for any currency. The intercept was negative and significantly different from zero for the DM and the SF. This is evidence against the maintained model for those cases. Since ex ante deviations from UIRP represent expected excess profits, the intercept could be capturing the sample mean of additional risk factors not captured by the conditional beta model of time-varying risk. Alternatively, the intercept could reflect differences in taxation across the domestic and foreign assets or other imperfections outside the model.

The conditional mean of the weekly excess return on the benchmark portfolio, includes an intercept, a first-order moving average term, and the average interest rate differential, $(R_{t-1}/\bar{Z}_{t-1})-1$, where $\bar{Z}$ is the average interest rate for the BP, CD, DM and JY.[6] Empirically, the coefficient estimate for this average interest rate differential is significantly negative, as will be discussed below. The MA (1) coefficient is not significant in every case but was included to capture any persistance associated with non-synchronized trading (for similar evidence, see Chou (1988)).[7]

Evidence of conditional heteroscedasticity in the excess returns for all currencies is shown by the $c$, $a$ and $b$ estimates in Table 1. Although the cross-equation restrictions implied by (7) make it difficult to link the persistence in a particular component of the conditional variance-covariance matrix to particular parameters, it is clear that the strong statistical significance of the estimates of $a_s$, $b_s$, $a_w$ and $b_w$ reflects strong persistence in the conditional variances. The statistical importance of the estimates of $c_{sw}$, $a_{sw}$ and $b_{sw}$, associated with the conditional covariance is convincingly demonstrated in the first row of Table 4. Suppression of these three parameters would reduce the conditional covariances considerably. Finally, the estimate of the parameter $\phi_{1w}$, associated with the indicator variable for the week of the market crash included in the conditional variance of the excess return for the benchmark portfolio, is positive and about 1·5 standard errors in size.

The bivariate models reported in Table 1 are evaluated with diagnostic tests and Lagrange multiplier tests for potentially important omitted variables. The results are summarized in Tables 2 and 3 respectively.

The Pagan-Sabau test is designed to test whether the estimated conditional variance models capture the pattern of heteroscedasticity observed in the squared residuals. In Table 2, the test statistics, $P_s$ and $P_w$, are based on the $t$-statistic, computed from robust standard errors, for the slope coefficient in a regression of the difference between the squared value of the raw residual and the conditional variance estimate on an intercept and the conditional variance estimate itself. This test does not reject the null hypothesis of zero slope in any currency, nor does it reject the specification of the conditional variance for the benchmark portfolio. We also apply this test to the conditional covariance estimates. These test statistics, given by $P_{sw}$, suggest that the CD covariance specification underpredicts outliers.

6. Adding the SF interest rate differential to the average resulted in a slightly inferior instrument for predicting the benchmark portfolio excess return. This may be due to the abnormally high SF interest rates towards the end of each calendar month.

7. Since $h_{wt}$ is also a function of $\varepsilon_{t-1}$, it is possible that the MA process replaces, to some extent, the conditional variance term implied by the static CAPM. Just including an intercept in $x_{w,t-1}$ in (6) and conducting LM tests for inappropriate exclusion of $h_{wt}$ resulted in $p$-values of 0·16, 0·41, 0·45, 0·67 and 0·65 for the five currency models.

The remaining tests reported in Table 2 evaluate the standardized residuals.[8] Evidence of remaining serial dependence includes small $p$-values for the runs test for the CD, the portmanteau test for the first ten lags of the autocorrelation function of the squared standardized residuals, in the JY and for the similar test with the cross products of the standardized residuals in the SF. These portmanteau tests may be affected by the presence of predetermined regressors and any remaining time variation in higher order moments (Cumby and Huizinga (1988)). From the conditional moment tests (Newey (1985), Tauchen (1985)) there is strong evidence of skewness in three currencies (DM, JY and SF). One source of the skewness for the SF case may be the large outliers in the 7-day Eurocurrency interest rates during the last week of some months. There is also evidence of excess kurtosis in the standardized residuals for the BP, CD and the JY. We rely on correct specification of the first two moments and our use of robust standard errors for statistical inference in the presence of this conditional non-normality.

Table 3 evaluates the models reported in Table 1 for potentially important omitted variables using Lagrange multiplier tests based on the outer product of the gradient (Godfrey and Wickens (1982), Davidson and MacKinnon (1990)). In particular, the vectors $x_{s,t-1}$ and $g_{s,t-1}$ are used to include the variables tested in the conditional mean and the conditional variance of $R_s^*$. In addition, the vector $g_{sw,t-1}$ is used to test for variables which may have been inappropriately excluded from the conditional covariance $h_{swt}$. We focus on testing for lagged components of the conditional means, various interest rate differentials, GARCH-M characteristics and effects of the October 1987 market crash.

We begin with the results of the OPG-LM tests for variables omitted from the conditional mean of $R_{st}^*$. Evidence of a missing variable in the mean would be important because it would indicate a possible inadequacy of the conditional beta risk premium in (4) as a measure for the conditional covariance term in (3). There is some evidence against the models estimated for the conditional means, $R_{st}^*$, of the deviations from UIRP. This is indicated by the high test statistics for the interest rate differentials $(R_{t-1}/Z_{t-1})-1$ and $(Z_{t-1}/\bar{Z}_{t-1})-1$ in the BP and CD, and to some extent for the average interest rate differential, $(R_{t-1}/\bar{Z}_{t-1})-1$, in the JY. Note that this is in addition to the effect of the average interest rate differential which enters the conditional risk premium through its importance in predicting $R_{wt}^*$.

Previous work on forward price data (Fama (1984), Bean (1985), Hodrick and Srivastava (1986), Backus, Gregory and Telmer (1989)) has often stressed the result that negative slope coefficient estimates are usually obtained in regressions of realized changes in spot prices on forward premia. In our case, the model presented in Section 2 would specify a coefficient of unity on the relative interest rate, $R_{t-1}/Z_{t-1}$, if the dependent variable was $S_t/S_{t-1}$ but for deviations from UIRP the relative interest rate is already incorporated in the composite dependent variable in (5). The results reported in Table 3 for the conditional means of the CD and the BP could be interpreted as evidence against this theoretical restriction on the interest rate differential.

Table 3 also reports LM test statistics evaluating whether the conditional variances of the two excess returns have been inappropriately excluded from their own conditional means. The capital asset pricing model implies that the conditional mean of the benchmark portfolio is proportional to its conditional variance. However, Table 3 demonstrates that the inclusion of $h_{wt}$ (with a constant coefficient) would not improve the fit of the model. On the other hand, Table 3 also shows that adding $h_{st}$ to (5) would result in more of the

8. Let $W_t$ be the matrix such that $W_tW_t=H_t^{-1}$, then the standardized residuals are obtained from the vector of raw residuals $\varepsilon_t$ as $u_t=W_t\varepsilon_t$.

persistence of the $R^*_{st}$ process being captured in the DM and SF. Whether or not this result is interpreted as evidence that unsystematic risk is being priced for those cases, it suggests the presence of additional sources of risk not captured by the single conditional beta model.

The specification reported in Table 1 is retained in almost all the tests for variables omitted from the conditional variance of $R^*_{st}$. The only low $p$-values are 0·01 for the square of the lagged deviation from UIRP, $R^*_{s,t-1}$, added to the conditional variance of the CD excess return and 0·01 for the square of the lagged excess return on the benchmark portfolio, $R^*_{w,t-1}$, in the JY conditional variance. Since our OPG-LM test statistics are not robust to departures from non-normality so that they will reject the null hypothesis too often (see Bollerslev and Wooldridge (1988)), these $p$-values are unlikely to be low enough to indicate a substantial improvement in the specification if the model were expanded to include those omitted variables.

Finally, Table 3 reports the results of some OPG-LM tests for variables omitted from the conditional covariance $h_{swt}$ between the benchmark excess return and the excess return from the currency position. Adding an indicator variable for the October 1987 market crash to the conditional covariance in the CD would have improved the fit, as indicated by the $p$-value of 0·01. However, adding the difference between the local currency interest rate and the average foreign currency interest rate, $(Z_{t-1}/\bar{Z}_{t-1})-1$, to the conditional covariances would not have improved the model for any of the currencies. Thus, unlike the application reported in Giovannini and Jorion (1987), our function of relative interest rates contributes to the price of covariance risk rather than the covariance itself.

In Table 4 we provide evidence related to the statistical significance of our conditional risk premia estimates. The covariation of the ex ante deviations from UIRP and the expected excess returns on the benchmark portfolio suggests a common source for the variation in excess returns on foreign currency spot positions and on the benchmark portfolio. The conditional covariances, and consequently the conditional betas, are positive on average and exhibit considerable time-variation. We now evaluate whether this covariation results in a statistically significant time-varying risk premium as specified by the maintained model in (4).

A likelihood-ratio (LR) test for restricting $\mu=0$, so that the conditional mean of $R^*_{st}$ does not include the conditional risk premium, is given in row 2 of Table 4. This restriction is clearly incompatible with the data for all the currencies although it is marginal for the CD. This is evidence of (time-varying) systematic risk associated with uncovered positions in foreign currencies.

The single beta conditional asset pricing model specifies that the intercept, $\gamma_{0s}$, in (5) should be zero and that $\mu$ should be unity. When these restrictions were evaluated individually in Table 1 with robust $t$-tests the null hypothesis that $\mu=1$ was retained in every currency but the estimated intercept was significantly negative for both the DM and the SF. The third row of Table 4 shows the LR test statistics for the joint restrictions on these parameters. The joint hypothesis is also rejected in the DM and SF.

In summary, conditional beta risk premia have been detected for all currencies but we have also found evidence against our particular specification of the time-varying risk. Significantly negative intercepts for the DM and SF (Tables 1 and 4), the GARCH-M feature of the conditional variance in the conditional mean of $R^*_{st}$ for the same two currencies (Table 3), and the additional explanatory power of interest rate differentials for the BP, CD and to some extent the JY (Table 3), all provide evidence of additional predictable components in the conditional mean of $R^*_s$. For example, changes in interest rates (in this case, interest rate differentials) might act as proxies for stochastic changes in investment opportunities and as such reflect the "hedging demands" inherent in the multi-beta model of Merton (1973). This could be the case in our analyses if the conditional single beta empirical implementation fails to capture adequately all the intertemporal sources of risk impled by the consumption-based IAPM which generated (3).

## 5. CONCLUDING COMMENTS

This paper has evaluated weekly deviations from uncovered interest rate parity, constructed using foreign currency spot prices and Eurocurrency interest rates, for the presence of a non-diversifiable component. Empirical implementation of the intertemporal asset pricing model is achieved with a conditional beta formulation which uses the Morgan Stanley Capital International world equity index as the benchmark portfolio. This benchmark portfolio represents extensive international diversification and provides directly observable data from which to compute returns for relatively short intervals of time.

The results indicate significant conditional systematic risk, with respect to the world equity benchmark portfolio, for the weekly deviations from uncovered interest rate parity. Detection of the conditional beta risk premia depended, in part, on the power of the interest rate differential to predict the conditional expected excess return on the world equity portfolio.

There is evidence that the conditional beta model of the time-varying risk premia did not capture all the predictable components in the excess returns associated with foreign currency positions. Although other interpretations of the evidence are certainly possible, it seems that additional risk factors should be investigated in an expanded model as well as alternative empirical reference portfolios in the single beta model.

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