Evidence of Risk Premiums in Foreign Currency Futures Markets

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Weekly data for foreign currency futures prices are examined for evidence of risk premiums. Covariance risks are measured with respect to the excess returns from benchmark portfolios for consumption and wealth. When the parameters representing the prices of the covariance risks are held constant, no risk premiums are detected. However, when these prices are allowed to vary with the conditional expected returns and variances of the benchmark portfolios, possibly reflecting changing investment opportunities, strong evidence of risk premiums is obtained.

Recent examinations of foreign currency futures data have tested the hypothesis that today's futures price is an unbiased predictor of tomorrow's price against an unspecified alternative hypothesis. The accumulated evidence against the unbiasedness hypothesis has not been accompanied by evidence consistent with a particular equilibrium risk premium model.

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Rejecting the unbiasedness hypothesis, Hodrick and Srivastava (1987) argued that positively autocorrelated risk premiums would reconcile the estimates obtained from daily and monthly data. Taylor (1986) did not derive his alternative hypothesis from an equilibrium model; nor did McCurdy and Morgan (1988)) who tested the hypothesis that the risk premium is proportional to the conditional standard deviation of forecast errors, and also illustrated how within-week periodicity contributes to rejection of the unbiasedness hypothesis.

The purpose of this article is to estimate and test a model incorporating time-varying risk premiums for prices of foreign currency futures contracts. The risk premiums arise from covariation of the futures price with consumption and wealth; we measure these risk premiums with respect to the covariances between the futures and weekly returns from benchmark portfolios representing consumption and wealth. Use of benchmark portfolio returns in a multiperiod consumption setting is consistent with a conditional version [Hansen and Richard (1987)] of the capital asset-pricing model (CAPM). Chou, Engle, and Kane (1991) and Harvey (1991) provided evidence against the restriction of a constant ratio of expected excess return to variance of the CAPM benchmark portfolio. We allow the risk premiums to reflect variation, over time, in the conditional covariances and in the prices of the covariance risks (the ratios of the conditional expected excess returns to conditional variances of the benchmark portfolios). We show that risk premiums are detected when the slope coefficients for the conditional covariances of the futures prices with the benchmark portfolio returns are allowed to vary over time, but not otherwise. Some of the other restrictions implied by the models we use are rejected by our tests.

Others have allowed all three components of the risk premiums—the conditional covariance, the conditional expected excess return, and the conditional variance of the benchmark portfolio—to vary individually. These include Mark (1988) for forward markets in foreign currency, McCurdy and Morgan (1991a) for uncovered positions in Eurocurrency deposits, and Harvey (1989) for the stock market; however, none of these considered the covariation with both consumption and wealth. Epstein and Zin (1989) and Giovannini and Weil (1989) have developed theoretical models in which both consumption and wealth are important.

In Section 1, we give a brief summary of the theoretical development leading to our empirical work. Testable versions of the models are formulated in Section 2. In Section 3, we discuss the data, and, in Section 4, we describe the results. We conclude in Section 5.

1. Theoretical Background

1.1 Notation and asset-pricing framework

We assume a single-good, pure-exchange model with a representative consumer. As in Lucas (1982) and Hansen and Hodrick (1983), the model is adapted to price nominal assets. Let M_t be the intertemporal marginal rate of substitution of domestic currency between time t - 1 and time t, F_t be the price at t of a futures contract to deliver one unit of the foreign currency at T, R_t be 1 plus the domestic riskless rate of interest from t - 1 to t, Z_t be 1 plus the foreign riskless rate of interest from t - 1 to t, and R_{it} be 1 plus the rate of return for asset i from t - 1 to t.

First-order conditions for utility maximization lead to Euler conditions,

$$1 = E_{t-1}[M_t R_{it}] = E_{t-1}[M_t R_t] = R_t E_{t-1}[M_t], \tag{1}$$

holding for any asset or portfolio i. M_t is not unique. Epstein and Zin (1989) took preferences to be of the recursive type postulated by Kreps and Porteus (1978). They discussed how different formulations of M_t are obtained from their model by particular choices of parameters reflecting risk aversion and intertemporal substitution. In general, M_t depends on both consumption and wealth, but for particular parameter values one of these two quantities alone is relevant.

1.2 Foreign currency futures positions

Proposition 7 of Cox, Ingersoll, and Ross (1981) established that any equilibrium asset-pricing framework can be used to price futures contracts even though these have zero present value. A futures contract has zero present value when it is initiated at t - 1 because the initial investment outlay is zero. Under the institutional practice of marking to market, the settlement price F_t is the equilibrium price that resets the present value of the contract to zero at t [Black (1976)]. No cash flow beyond that at t need be considered and

$$0 = E_{t-1}[M_t(F_t - F_{t-1})]. \tag{2}$$

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

$$F_{t-1} = E_{t-1}F_t + R_t \operatorname{cov}_{t-1}[M_t, F_t], \tag{3}$$

or, with scaling by the price F_{t-1} known at t - 1,

$$E_{t-1}(F_t/F_{t-1}) - 1 = -R_t \operatorname{cov}_{t-1}[M_t, F_t/F_{t-1}]. \tag{4}$$

Equation (4) defines the conditional nominal risk premium, or expected rate of change of the futures price, in terms of the conditional covariance between the rate of change of futures price and the marginal rate of substitution. Unbiasedness in future prices, $E_{t-1}F_t = F_{t-1}$, does not hold unless the conditional covariance in (4) is zero.

1.3 The benchmark portfolios

One form of M_t and risk premiun in (4) results in the conditional CAPM. For example, Hansen and Hodrick (1983)) following Breeden (1979), postulated the existence of a benchmark portfolio b with nominal return R_{bt} perfectly conditionally correlated with the conditional expectation of the marginal utility of the consumption good. As in Campbell (1987), who used the weaker assumption that the benchmark portfolio was maximally conditionally correlated with M_t, the assumption of perfect correlation establishes that this portfolio must be on the minimum variance frontier. The conditional CAPM is then implied. The unobservable composition of the benchmark portfolio is a problem for empirical work.

Various approaches to the problem of unobservable benchmark portfolio returns have been implemented in recent work. Engle, Ng, and Rothschild (1990) formed time series of excess returns for a set of factor-representing portfolios, with prespecified weights, for monthly Treasury-bill data. Korajczyk and Viallet (1991) estimated the benchmark portfolio returns as a linear combination of monthly excess returns from common stock portfolios from four countries. They used a principal components method to extract the factors and assumed that exchange rates and common stock prices were influenced by the same factors.

A second form of the intertemporal marginal rate of substitution and risk premium in (4) was derived by Epstein and Zin (1989). In their empirical work, Epstein and Zin (1991), Giovannini and Weil (1989), and Giovannini and Jorion (1989) chose similar parametric forms of the Kreps and Porteus recursive preference function. Under the additional assumption that the first differences of the logs of consumption and wealth are jointly normally distributed, the expected log real total return of an asset is a linear function of the covariances with log real total return of wealth and the change in the log of consumption. In other words, the risk of the asset depends on its covariances with wealth and consumption.

For the consumption component, we construct a benchmark portfolio with returns maximally correlated with consumption (MCP). Direct empirical testing with the available consumption data requires a minimum interval of one month between observations. However, stochastic daily interest rates and the daily cash flows generated by marking to market for futures contracts favor the use of futures price data recorded at short intervals. Daily data, however, pose problems for estimating the conditional expected returns from benchmark portfolios in tests of the futures risk premium hypothesis. We compromise by using weekly data, substituting the observable MCP returns for the consumption data. A side benefit of using weekly data is to reduce the impact of the bias induced by trades at the bid or offer price, as discussed by Bossaerts and Hillion (1991).

The weights of the MCP are estimated by relating the rate of growth of monthly consumption to monthly returns from a set of 17 component portfolios, as in Breeden, Gibbons, and Litzenberger (1989). The Appendix describes the steps taken to estimate this portfolio. The MCP weights are applied to daily data for the component portfolios, and the daily returns are aggregated to weekly returns matching the intervals for the futures data.

For the wealth component, we use the Morgan Stanley Capital International world index, a value-weighted stock market portfolio. While this part of our empirical work is open to the critique of Roll (1977), it has the advantage of working with directly observable quantities. Harvey (1991) has discussed the strengths and weaknesses of this index.

2. Test Equations: The Statistical Form of the Model

We analyze each currency separately. We match the series for the rate of change of futures price for a given currency with the two series of excess returns for the benchmark portfolios in a trivariate generalized ARCH model [Engle (1982), Bollerslev (1986)]. In principle, a single multivariate analysis that included all currencies would be preferable to a series of trivariate analyses, but such a model would have a very large number of covariance matrix parameters. It is well known that the conditional variances for rates of change of foreign currency prices vary over time. Evidence for financial data has been surveyed by Bollerslev, Chou, and Kroner (1990). We use the Baba, Engle, Kraft, and Kroner (1989) (BEKK) formulation ensuring positive definiteness in the ARCH model.

Ferson (1989) underlined the role of the short-term interest rate as a determinant of the expected returns that follow from (1). Many researchers since Fama and Schwert (1977) have used short-term interest rates, spreads, or interest rate differentials as instruments for predicting asset or portfolio returns. Recent examples include Campbell (1987), Giovannini and Jorion (1987)) and Harvey (1989). Some variables that have been used as instruments, such as dividend yield, are unsuitable for analysis of weekly data for the world index. We use an interest rate differential computed from the domestic rate and an average foreign rate.

The difference between the realized value and the rational expectation of the scaled payoff from holding the long position in the futures contract is a forecast error, which, by hypothesis, has a conditional mean of zero. Let x_{t-1} be a vector of instruments to predict the returns, in excess of the domestic interest rate, for the market portfolio and the MCP. Let these excess returns be R^*_{mt} and R^*_{ct}, respectively. Let γ_m and γ_c be the corresponding vectors of parameters. Let the error term vector e_t from the trivariate model of futures, market portfolio, and the MCP be conditional normally distributed with covariance matrix H_t. Let A, B, and G_t be symmetric matrices and C be upper triangular.

Initially, the trivariate system of test equations for the rate of change of futures prices is estimated with constant slope coefficients for the conditional covariances h_{fmt} and h_{fct} of the futures prices with the two benchmark portfolios. Let d_m, and d_c be these coefficients. In the Epstein and Zin model they are constants that are functions of risk aversion and intertemporal substitution parameters. The trivariate system is then

$$\frac{F_t}{F_{t-1}} - 1 = \gamma_{0f} + \delta_m h_{fmt} + \delta_c h_{fct} + \epsilon_{ft}, \tag{5}$$

$$R^*_{mt} = \gamma'_m x_{t-1} + \epsilon_{mt}, \tag{6}$$

$$R^*_{ct} = \gamma'_c x_{t-1} + \epsilon_{ct}, \tag{7}$$

$$H_t = C'C + A'\epsilon_{t-1}\epsilon'_{t-1}A + B'H_{t-1}B + G_t. \tag{8}$$

The Epstein and Zin (1991) and Giovannini and Weil (1989) specification implies that the conditionally expected return on the market portfolio is a function of its conditional variance h_{mt} and its covariance with consumption h_{mct}. In our case this leads to the expanded form of (6):

$$R^*_{mt} = \gamma'_m x_{t-1} + \delta_m h_{mt} + \delta_c h_{mct} + \epsilon_{mt}. \tag{9}$$

Subsequently we examine a version of (5) in which the coefficients for the conditional covariances of the futures prices with the market portfolio and the MCP are allowed to vary with the first two conditional moments of the benchmark portfolios. By hypothesis, these changing prices of covariance risk reflect changes in the investment opportunity set. For purposes of testing the null hypothesis of unbiasedness for the futures prices, we introduce an additional pair of multiplicative parameters, μ_m and μ_c, chosen to have a value of zero when the risk premium terms are excluded from the model. The alternative hypothesis in these tests allows μ_m and μ_c to be chosen freely. The futures equation becomes

$$\frac{F_t}{F_{t-1}} - 1 = \gamma_{0f} + \mu_m \gamma_m' x_{t-1} \frac{h_{fmt}}{h_{mt}} + \mu_c \gamma_c' x_{t-1} \frac{h_{fct}}{h_{ct}} + \epsilon_{ft}. \qquad (10)$$

A second null hypothesis maintains both μ_m and μ_c have values of unity, in which case (10) reduces to

$$\frac{F_t}{F_{t-1}} - 1 = \gamma_{0f} + \gamma_m' x_{t-1} \frac{h_{fmt}}{h_{mt}} + \gamma_c' x_{t-1} \frac{h_{fct}}{h_{ct}} + \epsilon_{ft}. \qquad (11)$$

3. Data

Futures prices for the British pound, Canadian dollar, Deutsche mark, Japanese yen, and Swiss franc (BP, CD, DM, JY, and SF, respectively) are taken from the 1985 version of the file provided by the University of Chicago's Center for Research in Futures Markets and updated with data from Reuters Inc. We use futures settlement prices for the contracts with the shortest maturity available at any time up to and including the last Wednesday before the end of the life of the contract. We compute the Wednesday-to-Wednesday rate of change of the futures price. If Wednesday was a holiday, Thursday prices are substituted. The series of 470 observations starts on January 2, 1980, and ends on December 28, 1988, giving an effective sample size of 469, because the first observation is used in the start of the estimation.

In the sample period, the institutionally imposed rules specifying the maximum price change that can occur in one day were relatively unimportant; the limits were not tight for these five currencies, and they were removed entirely on February 22, 1985. For the closest to maturity contract used in this paper, there were no occurrences of limit moves in the BP, four in the CD, eight in the DM, five in the JY, and three in the SF. Of these limit moves, only six occurred on a Wednesday, and we ignored their existence. Kodres (1988) gave a more complete discussion of price limits in empirical work.

The seven-day Eurocurrency interest rates are used to compute the excess returns from the benchmark portfolios and as instruments to predict these returns. When the domestic or any foreign interest rate was unavailable because of holidays, we substitute all rates for the previous day.

4. Empirical Results

We use numerical methods to find the maximum of the log likelihood function of the trivariate system.[1] In Table 1, we summarize the tests of the unbiasedness model against risk premium models in which the coefficients for the conditional covariances of the futures prices with the MCP and the world index are assumed constant. In Table 2, we summarize similar tests of models with and without risk premiums against a general model with no relevant conditional moment parameter restricted. In Table 3, we show the coefficient estimates and standard errors, while, in Table 4, we examine diagnostic statistics based mostly on residuals. In Table 5, we evaluate the models in Table 3 for potentially important omitted variables with outer product of the gradient Lagrange multiplier (OPG-LM) test statistics [Godfrey and Wickens (1982)].

Since many of the time series examined showed evidence that the hypothesis of conditionally normal error terms did not hold, all standard errors are computed to be robust. To do this, we follow a procedure similar to that of Bollerslev and Wooldridge (1988). Let J be the numerical approximation to the matrix of second derivatives with respect to the free variables. Let K be formed by taking the average of the period-by-period outer products of the gradient. The standard errors are computed from the diagonal elements of the matrix $J^{-1}KJ^{-1}$.

The trivariate system of equations for the futures, the market portfolio, and the MCP is estimated for each currency separately. The vector x_{t-1} in (6), (7), and (9) consists of a constant and the interest rate differential variable constructed from the domestic and the average foreign rate. In computing this variable, $R_t/\bar{Z}_t - 1$, the rate for the SF is excluded from the average because of wild rates found toward the end of most months, probably as a result of the reserve requirements in effect throughout most of the sample period.

The diagonal matrix G_t augments the conditional variance of the futures component of the system for the SF analysis and the conditional variance of the wealth component for all the analyses. For the SF the first element of G_t consists of the product of a coefficient f_f and the interest rate differential between the SF rate and the average foreign rate, $Z_t/\bar{Z}_t - 1$. In all analyses the second element on the diagonal is the product of an indicator variable for the week of the October 1987 market crash and its coefficient f_m. Additional variables are included in G_t only for purposes of testing their potential relevance for the conditional covariance specification.

There is one other difference between currencies in the models

[1] We use Numerical Algorithms Group routines E04HBF and E04JBF.

Table 1
Test statistics assuming constant prices of covariance risk

$$\frac{F_t}{F_{t-1}} - 1 = \gamma_{0f} + \delta_m b_{fmt} + \delta_c b_{fct} + \epsilon_{ft}, \quad (5)$$

$$R^*_{mt} = \gamma'_m x_{t-1} + \delta_m b_{mt} + \delta_c b_{mct} + \epsilon_{mt}, \quad (9)$$

$$R^*_{mt} = \gamma'_m x_{t-1} + \epsilon_{mt}, \quad (6)$$

$$R^*_{ct} = \gamma'_c x_{t-1} + \epsilon_{ct}, \quad (7)$$

$$H_t = C'C + A'\epsilon_{t-1}\epsilon'_{t-1}A + B'H_{t-1}B + G_t. \quad (8)$$

BP	CD	DM	JY	SF
5.52	3.33	4.36	2.33	2.78
(0.06)	(0.19)	(0.11)	(0.31)	(0.25)
4.41	1.81	4.57	4.87	1.85
(0.11)	(0.40)	(0.10)	(0.09)	(0.40)

Likelihood ratio tests of the unbiasedness hypothesis corresponding to the constraint $\delta_m = \delta_c = 0$ against the alternative hypothesis that both these parameters are free. In row 1, the excess returns from the market portfolio are given by (9) and the restriction $\delta_m = \delta_c = 0$ in (5) and (9) is tested. *In row 2, the excess returns are given by* (6) *and* $\delta_m = \delta_c = 0$ *in* (5) *is tested.* Below each test statistic the corresponding p value for the χ^2-distribution with 2 degrees of freedom is shown in parentheses.

fitted for the covariance specification. To obtain convergence in the BP analysis, we used the full BEKK structure, instead of setting all off-diagonal terms in the matrices A and B in (8) to zero. This restriction is tested and retained for all other currencies.

4.1 Results for constant prices of covariance risk

Table 1 shows the tests of the unbiasedness model against two versions of a model in which the prices of the covariance risks, or coefficients for the conditional covariances of the futures prices with the excess returns of the benchmark portfolios, are held constant. In the first version, the conditionally expected return from the market portfolio is a function of its conditional variance and its conditional covariance with the MCP, as in (9), as well as a function of the interest rate differential, $R_t/\bar{Z}_t - 1$.

In Table 1 the null hypothesis or no risk premium model, corresponding to the constraint $\delta_m = \delta_c = 0$ in (5) and (9), is tested against the alternative model with both of these parameters free. In the first row of Table 1, the null hypothesis is retained in all currencies. The same inference is also obtained from the second row of Table 1, when the simpler version (6) of the equation for the market portfolio returns is substituted for (9). No evidence has been found, so far, against the null hypothesis of unbiasedness in the futures price. In the next subsection we reevaluate this inference in tests of a model that allows the separate components of the coefficients for the conditional covariances in the futures equation to vary over time.

Table 2
Test statistics allowing prices of covariance risk to change

$$\frac{F_t}{F_{t-1}} - 1 = \gamma_{0f} + \mu_m \gamma_m' x_{t-1} \frac{b_{fmt}}{b_{mt}} + \mu_c \gamma_c' x_{t-1} \frac{b_{fct}}{b_{ct}} + \epsilon_{ft}, \quad (10)$$

$$R^*_{mt} = \gamma_m' x_{t-1} + \epsilon_{mt}, \quad (6)$$

$$R^*_{ct} = \gamma_c' x_{t-1} + \epsilon_{ct}, \quad (7)$$

$$H_t = C'C + A'\epsilon_{t-1}\epsilon_{t-1}' A + B' H_{t-1} B + G_t. \quad (8)$$

H_0	H_1	BP	CD	DM	JY	SF
$\mu_m = \mu_c = 0$	$\mu_m \neq 0$, $\mu_c \neq 0$	11.09 (0.00)	11.09 (0.00)	7.84 (0.02)	12.30 (0.00)	5.16 (0.08)
$\mu_m = \mu_c = 0$	$\mu_m \neq 0$, $\mu_c = 0$	10.68 (0.00)	9.46 (0.00)	5.37 (0.02)	9.75 (0.00)	4.58 (0.03)
$\mu_m = \mu_c = 0$	$\mu_m = 0$, $\mu_c \neq 0$	4.86 (0.03)	8.61 (0.00)	6.70 (0.01)	7.60 (0.01)	4.02 (0.05)
$\mu_m = \mu_c = 1$	$\mu_m \neq 1$, $\mu_c \neq 1$	2.97 (0.23)	3.58 (0.17)	1.59 (0.45)	4.04 (0.13)	0.30 (0.86)

In rows 1, 2, and 3, the unbiasedness hypothesis H_0 is tested against more general models with risk premiums. In row 1, the alternative hypothesis H_1 has risk premiums for both benchmarks. In row 2, the alternative model has risk premiums for wealth but not consumption; in row 3, the alternative model has risk premiums for consumption but not wealth. In row 4, the risk premium model with $\mu_m = \mu_c = 1$ is tested against the more general model with both parameters free. Log likelihood ratio test statistics are shown, and the p values in parentheses are for the χ^2-distribution.

4.2 Results for time-varying prices of covariance risk

Since the unbiasedness hypothesis was retained in the tests with constant coefficients for the conditional covariances in the futures equation, it is important to determine whether this result is due to the auxiliary assumption of constant prices of the covariance risks imposed on the model. In Table 2, these coefficients are replaced by the conditionally expected excess returns from the benchmark portfolios divided by their conditional variances.

The unbiasedness hypothesis, implying $\mu_m = \mu_c = 0$ in (10), is reevaluated in likelihood ratio tests against a more general model as the alternative hypothesis. In the first row of Table 2, both of these parameters are free in the general model, and the test statistics reveal that the null hypothesis is rejected in all currencies except the SF. From this evidence, the risk premiums in the futures prices could be systematically related to wealth or to the MCP or to both. The next two tests are designed to examine these possibilities.

In the second row of Table 2, the unbiasedness hypothesis $\mu_m = \mu_c = 0$ is tested against the alternative model that has μ_m free but $\mu_c = 0$, allowing a risk premium term for wealth but not consumption. The null hypothesis is rejected in favor of a risk premium in all currencies.

In the third row of Table 2, the unbiasedness hypothesis $\mu_m = \mu_c = 0$ is again tested, this time against the alternative model that has

μ_c free but $\mu_m = 0$, allowing a risk premium term for the consumption component but not for wealth. The null hypothesis is again rejected in favor of a risk premium in every currency.

The results of these three tests suggest that, while the risk premiums detected could be attributed to either one of the benchmark portfolios, the futures prices possibly reflect risk premium components from both. The simplest testable hypothesis of risk premiums from both sources involves setting $\mu_m = \mu_c = 1$, as in (11). This new null hypothesis is tested against the more general alternative model with both these parameters free. In this test, reported in the last row of Table 2, the new null hypothesis is retained in all currencies.

Two important results are established in Table 2. First, the unbiasedness hypothesis for currency futures prices is rejected. This result contrasts with the failure to reject the same hypothesis using the more restricted model with constant prices of covariance risk in Table 1. Second, futures prices reflect risk premiums related to their covariances with wealth and consumption. The risk premium model used as null hypothesis in the last row of Table 2 is the model maintained for the rest of the article.

4.3 Detailed examination of the risk premium model

Table 3 shows the estimates of the parameters determining the conditional means in the trivariate system for the model retained in the tests reported in the last row of Table 2.[2] The interest rate differential is successful as an instrument for predicting the conditional mean of the rate of return from the world index and the MCP. Lower conditional expected returns are predicted for both benchmarks when the domestic interest rate is greater than the average foreign rate.

In Table 4, two types of diagnostic tests on the standardized residuals from the maintained model are summarized: tests for remaining persistence and tests for departures from conditional normality. Vectors of standardized residuals u_t are obtained from the raw residuals e_t by setting $u_t = H_t^{-1/2}\epsilon_t$, where the matrix $H_t^{-1/2}$ satisfies $H_t^{-1/2}H_t^{-1/2} = H_t^{-1}$ and is obtained by orthonormal transformation of H_t^{-1}. The tests for persistence in the time series of the standardized residuals are runs tests and autocorrelation function tests and, for the squared residuals, tests for remaining conditional heteroskedasticity. Conditional moment tests examine the standardized residuals for skewness and excess kurtosis.

The main evidence of remaining persistence found in Table 4 is in the results from the portmanteau test for conditional heteroskedasticity in the squared standardized residuals from the world index.

[2]Estimates of the covariance parameters are not shown but are available from the authors.

Table 3
Model estimates allowing prices of covariance risk to change

$$\frac{F_t}{F_{t-1}} - 1 = \gamma_{0f} + \gamma_m' x_{t-1}\frac{b_{fmt}}{b_{mt}} + \gamma_c' x_{t-1}\frac{b_{fct}}{b_{ct}} + \epsilon_{ft}, \quad (11)$$

$$R_{mt}^* = \gamma_m' x_{t-1} + \epsilon_{mt}, \quad (6)$$

$$R_{ct}^* = \gamma_c' x_{t-1} + \epsilon_{ct}, \quad (7)$$

$$H_t = C'C + A'\epsilon_{t-1}\epsilon_{t-1}'A + B'H_{t-1}B + G_t. \quad (8)$$

	Futures	World index		MCP	
	γ_{0f}	γ_{0m}	γ_{1m}	γ_{0c}	γ_{1c}
BP	−0.056	0.372	−0.877	−0.071	−0.300
	(0.065)	(0.086)	(0.203)	(0.060)	(0.146)
CD	0.021	0.202	−0.511	−0.059	−0.391
	(0.054)	(0.042)	(0.101)	(0.057)	(0.126)
DM	−0.080	0.188	−0.436	−0.071	−0.332
	(0.035)	(0.042)	(0.113)	(0.059)	(0.133)
JY	−0.011	0.192	−0.449	−0.064	−0.371
	(0.033)	(0.039)	(0.098)	(0.053)	(0.121)
SF	−0.101	0.185	−0.426	−0.071	−0.323
	(0.040)	(0.038)	(0.102)	(0.053)	(0.120)

The data for the BP, world index, and MCP system were scaled by 100, 100, and 200, respectively. For the CD, the corresponding scaling was 200, 50, 200, and for the DM, JY, and SF, it was 50, 50, 200. The interest rate differential, used to predict the benchmark portfolio returns in (6) and (7), was scaled by 1000. Standard errors in parentheses are robust.

Estimates of their autocorrelation function reveal relatively large contributions to the portmanteau test statistic from lags 5 and 10.

The conditional moment tests detect skewness or excess kurtosis or both in the standardized residuals from the futures equation in all currencies and excess kurtosis in those from the world index equation. We rely on correct specification of the first and second moments and on the use of robust standard errors [Bollerslev and Wooldridge (1988)] for statistical inference in the presence of this conditional nonnormality.

Table 5 shows the OPG-LM tests for variables that might have been inappropriately excluded from the maintained model for the trivariate system [Equations (11), (6), (7), and (8)]. The first group includes the variable representing the relative difference between the domestic and the average foreign interest rate, and the similar variable for the relative difference between the local foreign rate and the average foreign rate. The small p values for the latter in the conditional mean of the BP and, to a lesser extent, the CD futures equations constitute evidence against the model. Similar results were obtained in a bivariate analysis [McCurdy and Morgan (1991b)], and the detailed discussion will not be repeated here. Predictable components of the futures price equations related to the local interest rate differential

Table 4
Diagnostic checks on models in Table 3

	BP	CD	DM	JY	SF
Futures					
R_f	−0.86	0.79	−1.41	−0.36	−2.18
	(0.39)	(0.43)	(0.16)	(0.72)	(0.03)
$Q_f(10)$	6.06	1.61	8.00	16.09	4.86
	(0.81)	(0.99)	(0.63)	(0.10)	(0.90)
$Q_f^2(10)$	5.33	10.87	7.40	8.00	10.56
	(0.87)	(0.37)	(0.69)	(0.63)	(0.39)
S_f	3.74	0.86	16.88	19.05	24.16
	(0.05)	(0.35)	(0.00)	(0.00)	(0.00)
K_f	10.88	13.94	3.18	6.32	1.94
	(0.00)	(0.00)	(0.07)	(0.01)	(0.16)
World index					
R_w	−0.21	−0.42	−0.36	0.54	−0.37
	(0.83)	(0.67)	(0.72)	(0.59)	(0.71)
$Q_w(10)$	8.16	10.20	10.71	8.90	12.09
	(0.61)	(0.42)	(0.38)	(0.54)	(0.28)
$Q_w^2(10)$	25.45	30.16	15.67	19.52	17.26
	(0.00)	(0.00)	(0.11)	(0.03)	(0.07)
S_w	1.46	0.46	1.31	1.61	2.33
	(0.23)	(0.50)	(0.25)	(0.20)	(0.13)
K_w	4.75	5.14	5.68	5.75	5.83
	(0.03)	(0.02)	(0.02)	(0.02)	(0.02)
MCP					
R_c	−1.06	−1.25	−1.06	−0.51	−0.88
	(0.29)	(0.21)	(0.29)	(0.61)	(0.38)
$Q_c(10)$	8.81	9.89	11.12	9.86	10.72
	(0.55)	(0.45)	(0.35)	(0.45)	(0.38)
$Q_c^2(10)$	6.08	9.43	7.55	7.55	7.20
	(0.81)	(0.49)	(0.67)	(0.67)	(0.71)
S_c	1.25	1.15	2.40	1.09	1.90
	(0.26)	(0.28)	(0.12)	(0.30)	(0.12)
K_c	1.01	0.12	0.16	0.06	0.14
	(0.31)	(0.73)	(0.69)	(0.81)	(0.71)

R is the test statistic for runs above the mean, $Q(10)$ and $Q^2(10)$ are the Ljung–Box (1978) form of the portmanteau statistic for the first 10 lags of the autocorrelation functions of the standardized residuals and their squares, and S and K are the Newey (1985) conditional moment statistics for skewness and excess kurtosis. The p values in parentheses are for the χ^2 distribution, except for R, where they are for the unit normal distribution.

may represent risk premiums not captured by the benchmark portfolios. A result obtained by Korajczyk (1985) may be relevant here. With foreign currency forward market data, he inferred that changes in risk premiums in forward markets are correlated with real interest rate differentials.

In the conditional variance, the small p value for the local interest rate differential for the DM parallels the SF where the equivalent variable was not excluded. We decided on the simpler model for the DM because the gains from including the extra variable were smaller than in the SF. The extra conditional covariance terms used in the

Table 5
OPG-LM tests for omitted variables

	BP	CD	DM	JY	SF
Conditional mean:					
Futures					
$R_t/\bar{Z}_t - 1$	2.40	2.26	0.64	0.62	0.34
	(0.12)	(0.13)	(0.42)	(0.43)	(0.56)
$Z_t/\bar{Z}_t - 1$	22.76	4.53	0.85	1.86	0.01
	(0.00)	(0.03)	(0.36)	(0.17)	(0.92)
Conditional variance:					
Futures					
$\lvert Z_t/\bar{Z}_t - 1 \rvert$	0.66	1.66	5.00	0.02	
	(0.42)	(0.20)	(0.03)	(0.89)	
Conditional covariances					
$a_{ij}, b_{ij}, i \neq j$		2.91	9.18	3.42	6.89
		(0.82)	(0.16)	(0.75)	(0.33)
Conditional mean:					
World index					
h_{mt}	0.84	2.02	2.83	1.88	2.61
	(0.36)	(0.16)	(0.09)	(0.17)	(0.11)
h_{mct}	3.26	0.04	0.84	0.01	0.84
	(0.07)	(0.84)	(0.36)	(0.92)	(0.36)
Conditional mean:					
MCP					
h_{ct}	2.29	0.98	0.52	0.02	0.02
	(0.13)	(0.32)	(0.47)	(0.89)	(0.89)

Reference: Godfrey and Wickens (1982).

Tests for the possible relevance of variables not included in various components of the model: interest rate differentials in the case of the equations for the conditional mean and conditional variance of the futures, off-diagonal ARCH covariance terms for all currencies except the BP, and conditional second moments in the equations for the conditional mean for the world index and MCP. p values for the χ^2-distribution (6 degrees of freedom for the conditional covariances) are in parentheses.

BP system but excluded from the other currencies are also tested in these four currencies. These analyses test the need for the off-diagonal terms in the matrices A and B in the BEKK conditional variance specification. The simpler specification is retained.

The panel for the conditional mean of the world index in Table 5 tests the restrictions across equations implied by (9). The test reported in the first row of Table 1 differed from that in the second row by the inclusion of the extra terms contained in (9) but not in (6). In Table 5 it again appears that the extra terms are not sufficiently useful in predicting the conditional returns from the world index and so can be excluded.

Although the estimated models have survived most of the specification tests, the OPG-LM statistics for the local interest rate differentials in the BP and CD, and the significantly negative intercepts for the DM and SF in Table 3, amount to rejection of the model except for the JY. In the other four currencies, there are indications of additional predictable components in futures prices, unrelated to the two benchmark portfolios in the model.

5. Conclusion

Previous analyses of foreign currency futures prices have rejected the unbiasedness hypothesis but not estimated any explicit model of the risk premiums. In this article, we have estimated models with conditional risk premiums in futures prices for five foreign currencies. We have measured covariance risks with respect to benchmark portfolios representing wealth and consumption. In contrast to the empirical work of Giovannini and Weil (l989), Giovannini and Jorion (1989), and Epstein and Zin (1991), our empirical model allows the prices of these covariance risks, hypothesized to reflect the available investment opportunities, to change over time.

When the prices of covariance risk (or slope coefficients for the conditional covariances of the futures prices with wealth and consumption) were held constant, no conditional risk premiums were detected. When the constant slope coefficients were replaced by the conditional expectation of excess returns from the benchmark portfolios and their conditional variances, futures prices for foreign currencies were found to respond to components of risk related to both wealth and consumption.

We also found some evidence of additional predictable components in the rates of change of futures prices for most currencies, so that the particular model maintained for the risk premiums was rejected. A potential limitation of the research is that the trivariate system does not allow a more complete evaluation of the implicit restrictions across equations. Also, our tests use data for a benchmark portfolio maximally correlated with consumption to represent the unobservable weekly consumption and a world equity index to represent wealth or the unobservable market portfolio. Other choices of these benchmark portfolios may be useful.

Appendix: The Portfolio Maximally Correlated with Consumption

We computed nominal consumption per capita from the sum of the monthly data for nondurable goods and services, using Citibase series GMCN82, GMCSN82, and POPCIV for the 108-month period 1980-1988. We obtained nominal daily returns from 16 of the 17 components used by Breeden, Gibbons, and Litzenberger (1989) (BGL), mostly from the University of Chicago's Center for Research in Security Prices. There were some differences between our components

Table 6
Estimates of the MCP from monthly data, 1980–1988

Component portfolio	Estimated coefficient	Standard error	Weight
Constant	0.0048	(0.0015)	
Seven-day Eurocurrency loans	0.156	(0.141)	0.94
High-grade long-term bonds	0.073	(0.028)	0.44
High-yield corporate bonds	−0.059	(0.032)	−0.36
Petroleum	−0.009	(0.008)	−0.05
Banking, finance, real estate	0.022	(0.029)	0.13
Consumer durables	0.014	(0.035)	0.08
Basic industries	0.024	(0.023)	0.14
Agriculture, food, tobacco	−0.035	(0.023)	−0.21
Construction	−0.016	(0.019)	−0.10
Capital goods	0.017	(0.030)	0.10
Transportation	0.008	(0.017)	0.05
Utilities	−0.050	(0.023)	−0.30
Textiles, retailers, wholesalers	−0.014	(0.020)	−0.08
Services	0.004	(0.023)	0.02
Recreation, leisure	−0.009	(0.027)	−0.05
CRSP value-weighted index	0.040	(0.029)	0.24
Total			1.00
Correlations			
Consumption with CRSP	0.265		
Consumption with MCP	0.517		
CRSP with MCP	0.513		

and those of BGL. They used value weighting within each component portfolio, but we used equal weighting. They included two long-term high-grade bond series and also the difference in yields of corporate high-grade and high-yield bonds to give the junk bond yield spread, but we found no series for the corporate high-grade daily returns, and used the returns from the Vanguard High Yield Bond Fund alone. We chose the Vanguard High Grade Bond Fund, which includes government and corporate bonds, for the other long-term bond component portfolio. Where BGL used Treasury bills, we used the seven-day Eurocurrency rates for the U.S. dollar, rolled over in monthly data. In all other details, our component portfolios were the same as theirs.

We computed 108 monthly returns for each component portfolio by compounding the daily returns from mid-month $t-1$ to mid-month t, as recommended by Chen, Roll, and Ross (1986) to correct for the averaging over the month in consumption-data collection. Another procedure, mentioned by BGL, would have been to try to adapt the method introduced by Scholes and Williams (1977) to correct for the noncontemporaneous observation of consumption and returns. This would have involved additional problems when the estimates obtained from monthly data were applied to weekly data for purposes of interpolation. The logarithm of the consumption ratio for consecutive months was then regressed, by ordinary least squares, on a constant and the 16 component portfolio returns. The estimated constant was ignored and the estimated slope coefficients were standardized to sum to unity and applied to the daily returns to obtain the daily and weekly returns from the portfolio maximally correlated with consumption.

Table 6 shows the estimated coefficients and weights from the regression with monthly data, the correlations between the consumption growth rate, the CRSP market portfolio returns, and the returns from the maximally correlated portfolio.

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