# Intertemporal risk in foreign currency markets

by
Thomas H. McCurdy and Ieuan Morgan
Queen's University, Kingston, Ontario, Canada K7L 3N6

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## Introduction and Background

The concept of an efficient capital market (attributed to Fama, 1965) is one in which all available and relevant information is "fully reflected" in asset prices. Essentially, efficient markets process all available and relevant information about future market developments such that asset prices provide appropriate signals for resource allocation.

As emphasized by Fama (1991, p. 1575), "market efficiency per se is not testable." That is, market efficiency can only be evaluated jointly with a particular pricing model that is postulated to generate the equilibrium expected returns. Models of market equilibrium imply particular specifications of preferences, production technology, endowments, information structure and market structure, as well as a specification of expectations formation. The concept of rational expectations implies that no relevant information will be ignored.

In empirical work, the efficient-market hypothesis has often been equated with the martingale hypothesis. This proposition implies that expected returns are a fair game, that is, increments in value (price adjusted for distributions such as dividends or interest payments) are unpredictable.

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Economic models that are consistent with a martingale process for asset prices require rather strong assumptions. As discussed in LeRoy (1989), sufficient assumptions include risk-neutral agents with common and constant discount rates, common information and rational expectations — as well as perfect markets with no transactions costs or costs of information. Relaxing any of these assumptions will lead to more general models — many of which, unlike the martingale hypothesis, will involve returns that are partly predictable. These more general models can then be evaluated with respect to their success at explaining observed predictable returns.

For example, relaxing the assumption of risk-neutrality will generally destroy the correspondence between the martingale hypothesis and informationally efficient markets since non-zero expected excess returns can exist in equilibrium as (potentially time-varying) rewards for bearing risk. Alternatively, relaxing the assumption about uniform and freely available information can lead to problems reconciling the Fama (1965) concept of market efficiency with competitive equilibrium (Grossman and Stiglitz, 1980) — as does sequential trading. The latter introduces the possibility of an infinity of equilibria (for example, Gourieroux and Laffont, 1982) and thus a potential indeterminacy of relevant information. A more detailed modelling of the economic environment (such as including additional details concerning the distribution of information, the structure of markets, and the heterogeneity of agents) may be needed to determine a particular equilibrium. Nevertheless, each of these more general models entails some model-specific concepts of informational and allocative efficiency.

In summary, evidence of predictable returns does not necessarily imply that the market is informationally inefficient. Tests always involve a joint hypothesis of informational efficiency and the particular pricing model that is postulated to generate the equilibrium expected returns. If the null hypothesis is the martingale, rejection implies predictable returns but does not lend support to any particular alternative hypothesis concerning the source of those returns unless the martingale is tested against a specific pricing model as the alternative hypothesis. Examples of such alternative models allowing predictable returns include time-varying risk premiums and rational learning about stochastic regime switches. Some of the



alternatives to the martingale hypothesis and the hypothesis that forward exchange rates are unbiased predictors of future spot rates, are presented or reviewed in Baillie and McMahon (1989), Boothe and Longworth (1986), Engel and Hamilton (1990), Frankel and Froot (1990), Hodrick (1987), LeRoy (1989), Lewis (1989), Meese (1989) and Obstfeld (1989).

## 1. Scope and Relation to Previous Research

In this paper, we focus on the alternative to the martingale hypothesis provided by the intertemporal asset-pricing model (IAPM). The intertemporal asset-pricing model was extended to price nominal assets by Hodrick (1981), Stulz (1981), Lucas (1982), Hansen and Hodrick (1983) and others. Most early applications of the IAPM were directed at financial forward markets, but the theory applicable to commodity futures contracts (Cox, Ingersoll and Ross, 1981 and Richard and Sundaresan, 1981) was also applied to financial foreign currency futures by Hodrick and Srivastava (1987) and McCurdy and Morgan (1987, 1988). These early applications to futures data rejected the martingale hypothesis, but did not test any specific model of equilibrium as the alternative hypothesis.

According to the IAPM, 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. With monthly consumption data, it is possible to test for the particular form of the risk premium implied by the consumption-based IAPM (for example, Hodrick, 1989 and Kaminsky and Peruga, 1990). However, in addition to the usual problems involved in measuring consumption, even at monthly intervals (see, for example, Breeden, Gibbons and Litzenberger, 1989), analyses of futures data favour the use of data recorded at short intervals because of daily marking-to-market.

An alternative empirical approach has been to use benchmark portfolio returns to capture time variation in the intertemporal marginal rate of substitution. The consumption-based asset-pricing model maintains that a single state variable, aggregate consumption, is adequate for determining marginal utility. The one-period capital-asset-pricing model (CAPM) uses aggregate wealth or the market return as a single state variable, whereas the intertemporal version proposed by Merton (1973) hypothesizes that a vector of states may be required to determine marginal utility. Use of benchmark portfolio returns in a multi-period consumption setting is consistent with a conditional version of the CAPM. This link between a consumption-based intertemporal asset-pricing model and the conditional CAPM has been described in Hansen and Hodrick (1983), Campbell (1987) and Hansen and Richard (1987).

There have been several approaches to measuring the benchmark portfolio in empirical implementations of the conditional CAPM. One approach has been to treat the benchmark portfolio as unobservable and derive testable conditions by imposing further assumptions. For example, Campbell (1987), Campbell and Clarida (1987), Giovannini and Jorion (1987), Cumby (1990) and Lewis (1990) assume that the ratio, across assets, of covariances with the benchmark portfolio return is constant over time. Others have replaced the excess returns associated with the benchmark portfolio with excess returns for a set of factor-representing portfolios (Engle, Ng and Rothschild, 1990 and Korajczyk and Viallet, 1991).

In another approach, an observable benchmark portfolio is assumed to be on the conditional mean-variance frontier. These conditional beta models have generally allowed the quantity of risk, the covariance with the benchmark return, to vary intertemporally but have assumed that the market price of that risk is constant (for example: Bollerslev, Engle and Wooldridge, 1988; Engel and Rodrigues, 1989; and Giovannini and Jorion, 1989).

For the most part, models of risk have had limited empirical success in explaining predictable returns in foreign currency markets. However, recent work (e.g., Mark, 1988 and McCurdy and Morgan, 1991) has shown that one can explain at least some part of the predictable component of exchange rate returns as compensation for bearing risk if both the quantity of risk and the market price of risk are allowed to vary. Chou, Engle and Kane (1992) and Harvey (1991) also provide evidence against the restriction of a constant market price of risk when the latter is measured as the ratio of expected excess return to variance for the CAPM benchmark portfolio. When prices of risk vary — owing, perhaps, to consumption



or investment opportunity sets differing across countries (Stulz, 1992), or attitudes towards risk varying with wealth or states of the world — investors will be concerned about which currency their portfolio is held in and may want to protect themselves from or exploit some part of the fluctuations in currency prices.

In this paper, we review and update the analyses in McCurdy and Morgan (1991, 1992b). In particular, we investigate whether some proportion of predictable excess returns associated with open positions in foreign currencies can be attributed to time-varying risk when risk is modelled using a single-beta, conditional capital-asset-pricing model. Taking a broadly diversified international portfolio of equities as a benchmark, we test for time-varying risk premiums associated with both foreign currency futures and spot (Eurocurrency deposit) positions.

In order to evaluate the intertemporal covariance risk associated with open foreign currency spot positions, we use foreign currency spot prices and Eurocurrency interest rates to construct a time series of weekly excess returns on Eurocurrency deposits — rather than using actual forward prices. Since covered interest rate parity holds, forward premiums and Eurocurrency interest rate differentials must be equal. Nevertheless, with weekly data there are some econometric advantages in using excess returns on Eurocurrency deposits. For example, the disadvantage of weekly sampling of prices for 30-day forward contracts is the moving-average form of the error-term process induced by the overlapping information content of the price series.

We find that interest rate differentials (differences between the nominal interest rates of the United States and foreign countries) have predictive power for the excess returns on the world equity portfolio. 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) and Fama and French (1990), have all identified some function of U.S. interest rates — in particular, term premiums and default premiums (junk bond spreads) — as variables that account, predictively, for some proportion of the variation of monthly returns on the stock market. Again, using monthly data, Harvey (1991) found that U.S. term and default premiums are also useful for predicting the world equity portfolio return. In our work with weekly data, when the difference between the U.S. and the average foreign Eurocurrency interest rate is included, the improved estimate of the expected excess return on wealth enhances the ability of the conditional CAPM to detect predictable components in the foreign currency returns, strengthening the evidence for systematic risk associated with uncovered positions in foreign currencies.

Our econometric specification involves joint quasi-maximum-likelihood (QML) estimation of the first and second conditional moments of a system consisting of the payoffs to the foreign currency position and the excess returns on the benchmark portfolio. Using a multivariate generalized ARCH process to parameterize the evolution over time of the conditional second moments, allowing both the market price and the quantity of risk associated with the foreign currency positions to vary intertemporally, and using the difference between the U.S. and the average foreign interest rate as an instrumental variable for the expected excess return from the benchmark portfolio, we detect time-varying risk premiums for all five currencies examined. The time-series properties of the estimated risk premiums associated with futures and spot positions for a particular currency are very similar, suggesting that hedging will be effective.

Our estimated risk premiums do not explain a large proportion of the realized positive and negative payoffs from the uncovered futures and spot positions. This is not unusual since most asset excess returns are difficult to predict. With monthly data for the 1980-90 period, Campbell and Hamao (1992) explain only 1.0 per cent of the variation of U.S. stock market returns with five instrumental variables. With intervals shorter than a month, some instrumental variables may be even less successful because their own variation over time is slow. Rejection of the martingale hypothesis indicates that some part of the weekly excess returns associated with foreign currency positions can be systematically predicted using available information. The empirical success of our risk-premium model (postulated as an alternative to the martingale hypothesis) should be evaluated on the basis of the proportion of those *predictable* excess returns that that model of risk can explain.

A single currency exhibits a significant test statistic associated with inappropriate exclusion of a term for its own conditional standard deviation. Tests for omitted variables also show that the relative difference between the individual



foreign currency interest rate 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 organization of the paper is as follows. Section 2 outlines the theoretical framework, including the equilibrium valuation of expected returns and the specification of the benchmark portfolio. Section 3 explains the test equations and their testable restrictions, while Section 4 summarizes the empirical results. Conclusions are then presented in Section 5.

## 2. Theoretical Framework

### 2.1 Equilibrium valuation

In this section, we apply an intertemporal asset-pricing model to payoffs from open positions in foreign currencies. In particular, we evaluate expected excess returns from Eurocurrency deposits and the expected change in foreign currency futures prices.

All prices are expressed in terms of units of domestic currency (assumed to be the U.S. dollar) and the interest rates are nominal. Let

$C_t$ = the number of units of the single consumption good consumed at time $t$;
$P_t$ = the price per unit of the (consumption) good at $t$;
$F_t$ = the futures price at $t$ for a contract that expires at $T$;
$S_t$ = the domestic currency spot price for a unit of foreign currency;
$R_t$ = one plus the domestic (U.S.) rate of interest from $t$ to $t+1$;
$Z_t$ = one plus the foreign interest rate from $t$ to $t+1$;
$R_{st}$ = $(S_t/S_{t-1})Z_{t-1}$ = return, at $t$, in dollars from an investment of one dollar at $t-1$ in the foreign asset;
$M_t$ = the generalized discount or nominal benchmark variable for the period $t-1$ to $t$;
$R_{Bt}$ = return on a benchmark portfolio that is postulated to be conditionally mean-variance efficient; and
$R_{wt}$ = return on the Morgan Stanley Capital International (MSCI) world index of equities.

With the assumption of complete markets, the no-arbitrage condition uniquely defines a generalized discount or nominal benchmark variable $M_t$ for equilibrium asset-pricing. In general, $M_t$ in conjunction with $E_{t-1}$, the mathematical expectations operator conditional on information available at $t-1$, is a present-value operator (Richard and Sundaresan, 1981) used to convert expected payoffs at $t$ to equilibrium present values at $t-1$. For example, a one-dollar investment in a one-period domestic asset with a payoff of $R_{t-1}$ has a present value of one dollar, that is,

$$1 = E_{t-1}[M_t R_{t-1}]. \quad (1)$$

Asset-pricing paradigms specify the nominal benchmark variable $M_t$ in different ways. In utility-based valuation theories, equation (1) is a particular application of the fundamental valuation equation (Constantinides, 1989) that equates the price of a claim to the expectation of the product of the future payoff and the marginal rate of substitution of the representative investor. In other words, it is the stochastic Euler condition derived from first-order conditions associated with utility maximization. With nominal payoffs, $M_t$ is the intertemporal marginal rate of substitution of domestic currency. For example, for the special case of a time-additive utility function with constant time discount factor $\delta$,

$$M_t = \delta \frac{u'_t}{u'_{t-1}} \frac{P_{t-1}}{P_t} \quad (2)$$

in which $u'$ is the marginal utility of real payoffs and $(P_{t-1}/P_t)$ reflects the change in the purchasing power of domestic currency.

When discussing empirical implementation below, we expand on various ways of measuring the nominal benchmark variable $M_t$. At this point, we proceed by applying the fundamental valuation equation to the payoffs associated with open positions in foreign currencies.



### 2.1.1 *Spot positions*

Application of the fundamental valuation equation to the one-period expected cash flow associated with a one-dollar investment in the foreign asset implies that

$$1 = E_{t-1}[M_t R_{st}]. \quad (3)$$

Combining equations (1) and (3), and using the definition of covariance, we obtain

$$E_{t-1}[R_{st}] - R_{t-1} = -R_{t-1}\, cov_{t-1}[M_t; R_{st}]. \quad (4)$$

That is, the expected excess returns associated with a foreign currency spot position are proportional to the conditional covariance between the payoff on the foreign currency position and the nominal benchmark variable $M_t$. Under the null hypothesis of our maintained model, this expected excess return is the equilibrium compensation for bearing risk. This nominal risk premium would be zero in equilibrium if the conditional covariance in equation (4) were zero. Risk-neutrality and a deterministic price level are sufficient conditions for that to be the case. However, stochastic changes in the purchasing power of domestic currency could result in a non-zero conditional covariance between the nominal benchmark variable $M_t$ and the nominal payoff $R_{st}$ — even under risk-neutrality.[1]

As indicated in equation (2), in a utility-based model, $M_t$ will be a function of marginal utility and the purchasing power of domestic currency. According to equation (4), the risk premium (the expected profit from an uncovered long position in the foreign asset) will be positive if $cov_{t-1}[M_t; R_{st}] < 0$ — for example, if the position has a high payoff when the marginal utility from a dollar's worth of consumption, $u'(C_t)/P_t$, is low. This would be the case if consumption is high

1. Engel (1992) provides a more detailed analysis of the relationship between nominal excess returns and risk premiums in real terms.

(marginal utility is low) and/or if the purchasing power of domestic currency is low (the domestic price level is high). This consumption-based interpretation of intertemporal risk involves equilibrium compensation for holding particular assets to the extent that they provide a hedge against adverse consumption outcomes. Analogous interpretations exist for other asset-pricing paradigms. That is, the conditional covariance of a payoff with the benchmark variable $M_t$ is a general measure of intertemporal risk.

### 2.1.2 *Futures positions*

Even though a futures contract has zero present value when it is initiated at $t-1$, it is possible to apply the conditional CAPM to the change in the futures price. Although the futures price is not by itself the value of an asset, "Proposition 7" of Cox, Ingersoll and Ross (1981, p. 327) equates the futures price with the present value of a series of payments corresponding to an asset. Application of an equilibrium model to the asset defined by Cox, Ingersoll and Ross, must then extend to the futures price itself. Under marking-to-market, the settlement price $F_t$ will be the equilibrium price that resets the value of the contract to zero at $t$. Therefore, applying the fundamental valuation equation to the payoffs from a long position in foreign currency futures gives

$$0 = E_{t-1}[M_t(F_t - F_{t-1})]. \qquad (5)$$

If we use equations (1) and (5) and the definition of covariance,

$$F_{t-1} = E_{t-1}F_t + R_{t-1}\ cov_{t-1}[M_t; F_t], \qquad (6)$$

and scale by the price $F_{t-1}$ which is known at $t-1$,

$$E_{t-1}(F_t/F_{t-1}) - 1 = -R_{t-1}\ cov_{t-1}[M_t; F_t/F_{t-1}]. \qquad (7)$$



Equation (7) defines the expected rate of change of the futures price in terms of the conditional covariance between the nominal benchmark variable $M_t$ and the rate of change of the futures price. In our model, this is the conditional nominal risk premium associated with a long position in foreign currency futures; unbiasedness of the futures price does not hold unless the conditional covariance is zero.

## 2.2 Empirical specification of the benchmark variable

In the conditional CAPM, a benchmark portfolio is hypothesized to be on the conditional mean-variance frontier. If there is an asset or portfolio with return $R_m$ perfectly conditionally correlated with $M_t$, then a benchmark portfolio giving returns $R_B$ that are a linear combination of $R_m$ and the risk-free return will be conditionally mean-variance efficient. The equilibrium expected excess return on any asset $j$, will then depend on the quantity of risk, its conditional covariance with the benchmark portfolio return, and on the market price of the covariance risk. That is,

$$E_{t-1}[R_{jt}] - R_{t-1} = cov_{t-1}[R_{jt}; R_{Bt}]\frac{(E_{t-1}[R_{Bt}] - R_{t-1})}{var_{t-1}[R_{Bt}]}. \qquad (8)$$

As noted in Section 1.2, several approaches to measuring such a benchmark portfolio have been proposed. One approach has been to treat the benchmark portfolio as unobservable and use a latent variable model to explain returns. An alternative approach has been to assume that a particular observable portfolio used as a benchmark is on the conditional mean-variance frontier.

Since we wish to capture time variation in the intertemporal marginal rate of substitution, $M_t$, using the conditional moments of the returns from a benchmark portfolio, the assumption of a return that is perfectly conditionally correlated with $M_t$ is stronger than necessary. Breeden, Gibbons and Litzenberger (1989) proceed by constructing a portfolio that has returns that are *maximally* correlated with the growth rate of consumption. McCurdy and Morgan (1992a) use benchmark portfolios for both consumption (a maximum-correlation portfolio) and wealth (a world-equity portfolio) in an empirical implementation motivated by non-expected utility explanations of asset prices.

In this paper, we use the return on the Morgan Stanley Capital International (MSCI) world equity index, $R_w$, as a measure for the benchmark-portfolio return, $R_B$, in (8). Choosing a market index, as in Mark (1988) and Harvey (1991), is open to the Roll (1977) critique. Nevertheless, the MSCI world index represents extensive international diversification. The issue of whether this choice for the benchmark portfolio in the single-beta formulation of the conditional pricing relations is adequate to price all the relevant risk associated with foreign currency spot and futures prices has been addressed in some detail in McCurdy and Morgan (1991, 1992b) using various tests for misspecification.

## 3. Test Equations

Let $y_{st} = (S_t/S_{t-1}) - (R_{t-1}/Z_{t-1})$ be the (scaled) excess return on the foreign currency spot position, and $y_{ft} = (F_t/F_{t-1}) - 1$ be the rate of change in the futures price.

Using equation (8) and the observable world portfolio return $R_w$, we can re-express (4) and (7) in terms of the single-beta conditional capital-asset-pricing relations,

$$E_{t-1}[y_{st}] = cov_{t-1}[y_{st};R^*_{wt}]\frac{E_{t-1}[R^*_{wt}]}{var_{t-1}[R^*_{wt}]}, \tag{9}$$

$$E_{t-1}[y_{ft}] = cov_{t-1}[y_{ft};R^*_{wt}]\frac{E_{t-1}[R^*_{wt}]}{var_{t-1}[R^*_{wt}]}. \tag{10}$$

The conditional asset-pricing relations (9) and (10) specify that excess returns on a Eurocurrency deposit and the rate of change of a futures price are expected to be proportional to their conditional covariances with the excess returns on the world equity portfolio, $R^*_{wt}$. The factor of proportionality is the market price



of covariance risk measured as the conditionally expected excess return on the benchmark portfolio divided by its conditional variance. Note that both the quantity of risk (the conditional covariance) and the market price of that covariance risk are allowed to vary. A bivariate model jointly estimates the first and second conditional moments of $y_{jt}$ and $R^*_{wt}$ for $j=s$ or $f$.

Under rational expectations, the realized value of the scaled payoffs in (9) and (10) are equal to the conditional expectation plus an error term, $\varepsilon_{jt}$, that is uncorrelated with past information. The corresponding error term for the benchmark portfolio is $\varepsilon_{wt}$. Let $h_{jt}$ be the conditional variance of $y_{jt}$ (either the excess return on the Eurocurrency deposit, $j=s$, or the rate of change in the futures price, $j=f$ ), $h_{jw,t}$ the conditional covariance between $y_{jt}$ and $R^*_{wt}$, $h_{wt}$ the conditional variance of $R^*_{wt}$, $x_{w,t-1}$ a vector of instruments (an intercept, the lagged excess return on the benchmark portfolio, and the difference between the U.S. and an average non-U.S. interest rate) used to predict the benchmark portfolio return, and $\mu$ a parameter that can be restricted to zero to exclude the risk premium from the model. Pairing the time series of the excess returns on the Eurocurrency deposit $(j = s)$, or the rate of change of the futures price $(j = f)$, for a particular currency with the excess returns from the benchmark portfolio gives the bivariate intertemporal risk test equation system

$$y_{jt} = \gamma_{oj} + \mu_j h_{jw,t}\frac{[\gamma'_w x_{w,t-1}]}{h_{wt}} + \varepsilon_{jt}, \tag{11}$$

$$R^*_{wt} = \gamma'_w x_{w,t-1} + \varepsilon_{wt}. \tag{12}$$

Time variation in the conditional second moments of financial data has been extensively documented (for references, see the survey by Bollerslev, Chou and Kroner, 1992). We parameterize the conditional covariance matrix associated with the system (11) and (12) using the Baba, Engle, Kraft and Kroner (1991) positive definite form for generalized ARCH (Engle, 1982 and Bollerslev, 1986). We also allow for asymmetric ARCH effects (Nelson, 1991 and Glosten, Jagannathan and Runkle, 1989) by adding a component for $u_{jt} = min\{0, \varepsilon_{jt}\}$ and $u_{wt} = min\{0, \varepsilon_{wt}\}$ to the conditional covariance matrix.

Let $C$ be an upper triangular matrix of parameters and $A$, $D$ and $B$ be symmetric matrices of parameters. The vector of error terms, $\varepsilon_t' \equiv [\varepsilon_{jt}, \varepsilon_{wt}]$, is assumed to have a conditional bivariate normal distribution with zero mean and conditional covariance matrix,

$$H_{jt} = C_j'C_j + A_j'\varepsilon_{j,t-1}e_{j,t-1}'A_j + D_j'u_{j,t-1}u_{j,t-1}'D_j + B_j'H_{j,t-1}B_j + G_{j,t-1}. \quad (13)$$

For example, when $D$ is set to zero for purposes of testing symmetric generalized ARCH, we can express (13) as

$$\begin{bmatrix} h_{jt} & h_{jwt} \\ h_{jwt} & h_{wt} \end{bmatrix} = \begin{bmatrix} c_j & 0 \\ c_{jw} & c_w \end{bmatrix} \begin{bmatrix} c_j & c_{jw} \\ 0 & c_w \end{bmatrix} + \begin{bmatrix} a_j & a_{jw} \\ a_{jw} & a_w \end{bmatrix} \begin{bmatrix} \varepsilon_{j,t-1}^2 & \varepsilon_{j,t-1}\varepsilon_{w,t-1} \\ \varepsilon_{j,t-1}\varepsilon_{w,t-1} & e_{w,t-1}^2 \end{bmatrix} \begin{bmatrix} a_j & a_{jw} \\ a_{jw} & a_w \end{bmatrix} + \begin{bmatrix} b_j & b_{jw} \\ b_{jw} & b_w \end{bmatrix} \begin{bmatrix} h_{j,t-1} & h_{jw,t-1} \\ h_{jw,t-1} & h_{w,t-1} \end{bmatrix} \begin{bmatrix} b_j & b_{jw} \\ b_{jw} & b_w \end{bmatrix} + \begin{bmatrix} \phi'_j g_{j,t-1} & \phi'_{jw} g_{jw,t-1} \\ \phi'_{jw} g_{jw,t-1} & \phi'_w g_{w,t-1} \end{bmatrix}. \quad (14)$$

Note that $g_{j,t-1}$, $g_{w,t-1}$ and $g_{jw,t-1}$ are vectors of instruments known at time $t-1$. These vectors are used to add explanatory variables to the variances and the covariance, and also to evaluate the specification of $H_t$ using Lagrange multiplier (LM) omitted-variable tests. Our maintained model includes an indicator variable in $g_{w,t-1}$ that allows the conditional variance of the benchmark portfolio return to differ for the week of the market crash (the third week of October, 1987).



The test-equation systems are designed to evaluate separately the intertemporal covariance risk associated with the foreign currency futures and spot positions. Any significant differences in the size or variation of the risk premiums associated with the two positions will provide indirect evidence as to the effectiveness of a hedge.

## 4. Data and Empirical Results

### 4.1 The data

Futures prices for the British pound, Canadian dollar, Deutsche mark, Japanese yen and Swiss franc (BP, CD, DM, JY, 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, recorded at 1:16 to 1:26 p.m. CST, 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 626 observations starts on 2 January 1980 and ends on 26 December 1991, giving an effective sample size of 625, because the first observation is used in the start of the estimation.

In the sample period, the institutionally imposed rules for futures 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 22 February 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 made no adjustment for them.

Foreign currency spot prices are the average of bid and ask prices from the interbank market recorded at 2:00 p.m. EST.

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

## 4.2 The results

Quasi-maximum-likelihood (QML) estimation is implemented with standard errors computed to allow robust inference in the presence of departures from conditional normality (Bollerslev and Wooldridge, 1992).[2] Tables 1 to 4 summarize the main results for the various positions in the different currencies. Table 1 reports output from estimation of the bivariate intertemporal-risk model (11), (12) and (13), applied to excess returns on Eurocurrency deposits for each currency. Table 2 evaluates the model in Table 1 by reporting results from some residual-based diagnostic tests and Lagrange multiplier tests of various restrictions such as omitted variables. Tables 3 and 4 do the same for the rate of change of the futures price for each currency. Figures 1 to 4 plot the estimated risk premiums associated with the futures and spot positions for the CD and the JY.

Table 1 reports empirical evidence concerning our application of the intertemporal risk-premium model to foreign currency spot positions. As indicated in (9), the conditional capital-asset-pricing implementation of the intertemporal-risk model stipulates that $\gamma_{0s} = 0$ and $\mu_s = 1$ in (11). The first panel of Table 1 includes estimates of these parameters along with the associated robust standard errors. The intercept is insignificantly different from zero in all cases — except perhaps the SF for which the estimated intercept has a robust $t$-value of 2.01. The estimates of $\mu_s$ are all positive and, with the possible exception of the JY, insignificantly different from 1. These results support the maintained model.

Panel A in Table 1 also reports the parameter and robust standard error estimates associated with our parameterization of the conditional mean of the excess return on the benchmark (world-equity) portfolio, as in equation (12). Note that the predictable component of (12) enters into (11) through the market price of covariance risk. The coefficient, $\hat{\gamma}_{1w}$, for the lagged excess return, $R^*_{w,t-1}$, is significant in the case of the CD and the JY, and the excess return on the world-

2. 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 product of the gradient. Engle and Gonzalez-Rivera (1990) discuss the potential loss of efficiency associated with QML estimation and propose a more efficient semi-parametric approach.



**Table 1**

**Bivariate Model of Spot Returns**

$$y_{st} = \gamma_{0s} + \mu_s \frac{h_{swt}}{h_{wt}} \gamma'_w x_{w,t-1} + \varepsilon_{st},$$
$$R^*_{wt} = \gamma'_w x_{w,t-1} + \varepsilon_{wt}.$$

**Panel A: Estimates**

| | BP | CD | DM | JY | SF |
|---|---|---|---|---|---|
| $\gamma_{0s}$ | -0.045 | 0.026 | -0.099 | -0.076 | -0.143 |
| | (0.070) | (0.047) | (0.054) | (0.073) | (0.071) |
| $\mu_s$ | 2.07 | 3.33 | 2.14 | 2.76 | 1.93 |
| | (1.03) | (1.95) | (0.92) | (0.86) | (0.95) |
| $\gamma_{0w}$ | 0.152 | 0.131 | 0.146 | 0.146 | 0.137 |
| | (0.068) | (0.078) | (0.057) | (0.067) | (0.073) |
| $\gamma_{1w}$ | 0.070 | 0.093 | 0.081 | 0.096 | 0.088 |
| | (0.067) | (0.042) | (0.051) | (0.039) | (0.053) |
| $\gamma_{2w}$ | -0.478 | -0.464 | -0.448 | -0.370 | -0.422 |
| | (0.134) | (0.166) | (0.118) | (0.147) | (0.157) |

**Panel B: Test of $\mu_s = 0$**

| | BP | CD | DM | JY | SF |
|---|---|---|---|---|---|
| *LR* | 7.47 | 10.15 | 7.89 | 15.89 | 5.88 |
| | (0.01) | (0.00) | (0.01) | (0.00) | (0.01) |
| *t* | 2.56 | 2.47 | 2.85 | 3.29 | 2.54 |
| | (0.01) | (0.01) | (0.00) | (0.00) | (0.01) |
| **Robust *t*** | 2.01 | 1.71 | 2.33 | 3.21 | 2.03 |
| | (0.04) | (0.09) | (0.02) | (0.00) | (0.04) |

**Note:**

Panel A: Standard errors in parentheses are robust. $\gamma_{1w}$ is the autoregressive term coefficient for the benchmark portfolio, $\gamma_{2w}$ the coefficient for the interest rate differential.

Panel B: *LR* is the likelihood-ratio test statistic and the $p$-values in parentheses are for the $\chi^2_1$ distribution. For the $t$ tests, the $p$-values are for the unit-normal distribution.

**Table 2**

**Diagnostics and Tests of Restrictions: Spot Model**

| | BP | CD | DM | JY | SF |
|---|---|---|---|---|---|
| **Panel A: Diagnostic checks** | | | | | |
| $R_s$ | -0.03 | -0.46 | -0.75 | -0.08 | -2.67 |
| | (0.98) | (0.65) | (0.45) | (0.94) | (0.01) |
| $Q_s(10)$ | 9.92 | 4.02 | 14.39 | 14.88 | 13.00 |
| | (0.45) | (0.95) | (0.16) | (0.14) | (0.22) |
| $Q_s^2(10)$ | 10.05 | 3.95 | 16.22 | 11.78 | 11.32 |
| | (0.44) | (0.93) | (0.09) | (0.30) | (0.33) |
| $Q_{sm}(10)$ | 9.23 | 3.08 | 12.89 | 11.48 | 21.34 |
| | (0.51) | (0.98) | (0.23) | (0.32) | (0.02) |
| $S_s$ | 3.13 | 0.01 | 8.60 | 21.70 | 22.20 |
| | (0.08) | (0.92) | (0.00) | (0.00) | (0.00) |
| $K_s$ | 8.75 | 17.02 | 30.87 | 71.44 | 39.15 |
| | (0.00) | (0.00) | (0.00) | (0.00) | (0.00) |
| **Panel B: Tests of restrictions** | | | | | |
| $\varepsilon_{t-1}$ | 0.65 | 0.97 | 0.62 | 0.01 | 0.52 |
| | (0.42) | (0.33) | (0.43) | (0.92) | (0.47) |
| $h_t^{1/2}$ | 0.08 | 1.28 | 0.72 | 1.61 | 5.49 |
| | (0.78) | (0.26) | (0.40) | (0.21) | (0.02) |
| $z_{t-1}/\bar{z}_{t-1}$ | 16.81 | 11.94 | 1.81 | 1.42 | 0.91 |
| | (0.00) | (0.00) | (0.18) | (0.23) | (0.34) |
| $D$ | 30.76 | 17.73 | 15.73 | 10.86 | 12.54 |
| | (0.00) | (0.00) | (0.00) | (0.01) | (0.01) |

Note: $R_s$ is the test statistic for runs above the mean and the $p$-values in parentheses are for the unit-normal distribution. $Q_s(10)$, $Q_s^2(10)$ and $Q_{sm}(10)$ are the Ljung-Box (1978) test statistics for the first 10 lags of the autocorrelation function for the standardized residuals for the spot returns, their squares, and for the cross-products of the standardized residuals for the spot returns and the world portfolio returns, respectively. The $p$-values are for the $\chi_{10}^2$ distribution. $S_s$ and $K_s$ are the Newey (1985) conditional moment test statistics for skewness and kurtosis, respectively. For these and for the tests of restrictions in the mean, the $p$-values are for the $\chi_1^2$ distribution. $D$ is the likelihood-ratio test statistic for a restricted model with the constraint $d_{11} = d_{12} = d_{22} = 0$.



equity portfolio is significantly lower when U.S. rates are higher than the average of the non-U.S. rates ($\hat{\gamma}_{2w}$ is significantly negative).

Panel B of Table 1 reports the results of tests of the hypothesis that intertemporal risk premiums are zero when risk is modelled using the single-beta conditional CAPM. The robust $t$-test of $\mu_s = 0$ rejects the hypothesis of zero risk premiums for all currencies except the CD.

As noted in footnote 2, QML estimation may involve some loss of efficiency. Therefore, we also report the $t$-statistics calculated with conventional maximum-likelihood standard errors computed from the diagonal values of the inverse of the numerical approximation to the matrix of second derivatives. Panel B also provides likelihood-ratio (LR) tests of the restriction $\mu_s = 0$. These two tests reject the zero risk premium hypothesis more convincingly and for all currencies but, unlike the robust $t$-test, they are not robust to departures from conditional normality.

A comparison of Tables 3 and 1 reveals that the results for uncovered spot positions in foreign Eurocurrency deposits are very similar to those for the futures positions. The robust $t$-tests in Table 3 show that the estimates of $\mu_f$ are not significantly different from unity in any currency, consistent with the conditional capital-asset-pricing model. Furthermore, with one possible exception (the SF), the $\mu_f$ estimates are significantly different from zero indicating that intertemporal risk premiums can explain at least some proportion of the predictable returns.

Panel A of Tables 2 and 4 reports some diagnostic tests on standardized residuals from the model (11), (12) and (13) applied to the spot and futures data, respectively. These diagnostic tests generally do not reveal any inadequacy of the estimated models except for some remaining serial correlation for the SF, skewness in the standardized residuals associated with some currency positions and excess kurtosis for all currencies. For inferences in the presence of these departures from conditional normality, we rely on consistency of the QML estimates of the parameters and robust standard errors (Bollerslev and Wooldridge, 1992).

Panel B of Tables 2 and 4 shows LM tests for variables that may have been inappropriately excluded from (11). These omitted variables include an MA(1) term, the conditional standard deviation associated with (11), and the relative

**Table 3**

**Bivariate Model of Changes in Futures Prices**

$$y_{ft} = \gamma_{0f} + \mu_f \frac{h_{fwt}}{h_{wt}} \gamma'_w x_{w,t-1} + \varepsilon_{ft},$$

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

**Panel A: Estimates**

| | BP | CD | DM | JY | SF |
|---|---|---|---|---|---|
| $\gamma_{0f}$ | -0.048 | 0.031 | -0.118 | -0.086 | -0.148 |
| | (0.053) | (0.052) | (0.069) | (0.075) | (0.070) |
| $\mu_f$ | 1.95 | 3.75 | 2.30 | 2.71 | 1.84 |
| | (0.70) | (1.80) | (1.02) | (0.86) | (1.04) |
| $\gamma_{0w}$ | 0.156 | 0.135 | 0.156 | 0.153 | 0.154 |
| | (0.064) | (0.068) | (0.067) | (0.066) | (0.067) |
| $\gamma_{1w}$ | 0.071 | 0.088 | 0.072 | 0.090 | 0.085 |
| | (0.045) | (0.048) | (0.056) | (0.042) | (0.058) |
| $\gamma_{2w}$ | -0.478 | -0.494 | -0.473 | -0.400 | -0.435 |
| | (0.135) | (0.165) | (0.123) | (0.142) | (0.157) |

**Panel B: Test of $\mu_f = 0$**

| | BP | CD | DM | JY | SF |
|---|---|---|---|---|---|
| *LR* | 4.68 | 12.10 | 8.00 | 15.00 | 4.83 |
| | (0.03) | (0.00) | (0.01) | (0.00) | (0.03) |
| *t* | 3.10 | 2.62 | 2.88 | 3.27 | 2.36 |
| | (0.00) | (0.01) | (0.00) | (0.00) | (0.02) |
| Robust *t* | 2.79 | 2.08 | 2.25 | 3.15 | 1.77 |
| | (0.01) | (0.04) | (0.02) | (0.00) | (0.08) |

Note:

Panel A: Standard errors in parentheses are robust. $\gamma_{1w}$ is the autoregressive term coefficient for the benchmark portfolio, $\gamma_{2w}$ the coefficient for the interest rate differential.

Panel B: *LR* is the likelihood-ratio test statistic and the $p$-values in parentheses are for the $\chi^2_1$ distribution. For the $t$-tests, the $p$-values are for the unit-normal distribution.



**Table 4**

**Diagnostics and Tests of Restrictions: Futures Model**

**Panel A: Diagnostic checks**

| | BP | CD | DM | JY | SF |
|---|---|---|---|---|---|
| $R_f$ | -0.67 | 0.70 | -0.23 | 1.30 | -2.49 |
| | (0.50) | (0.48) | (0.82) | (0.19) | (0.01) |
| $Q_f(10)$ | 10.63 | 4.34 | 13.62 | 15.30 | 11.18 |
| | (0.39) | (0.93) | (0.19) | (0.12) | (0.34) |
| $Q^2_f(10)$ | 8.69 | 6.32 | 16.35 | 11.25 | 13.82 |
| | (0.56) | (0.79) | (0.09) | (0.34) | (0.18) |
| $Q_{fm}(10)$ | 12.08 | 4.76 | 16.04 | 13.22 | 24.22 |
| | (0.28) | (0.91) | (0.10) | (0.21) | (0.01) |
| $S_f$ | 3.39 | 0.15 | 8.39 | 25.11 | 21.26 |
| | (0.07) | (0.70) | (0.00) | (0.00) | (0.00) |
| $K_f$ | 7.92 | 7.01 | 31.84 | 76.05 | 34.96 |
| | (0.00) | (0.01) | (0.00) | (0.00) | (0.00) |

**Panel B: Tests of restrictions**

| | BP | CD | DM | JY | SF |
|---|---|---|---|---|---|
| $\varepsilon_{t-1}$ | 1.78 | 0.01 | 1.56 | 0.89 | 1.84 |
| | (0.18) | (0.92) | (0.21) | (0.35) | (0.18) |
| $h^{1/2}_t$ | 0.25 | 0.92 | 0.79 | 1.65 | 5.47 |
| | (0.62) | (0.34) | (0.37) | (0.20) | (0.02) |
| $z_{t-1}/\bar{z}_{t-1}$ | 13.33 | 10.48 | 2.20 | 1.74 | 1.48 |
| | (0.00) | (0.00) | (0.14) | (0.19) | (0.22) |
| $D$ | 26.12 | 16.87 | 14.88 | 11.13 | 10.74 |
| | (0.00) | (0.00) | (0.00) | (0.01) | (0.01) |

Note: $R_f$ is the test statistic for runs above the mean and the $p$-values in parentheses are for the unit-normal distribution. $Q_f(10)$, $Q^2_f(10)$ and $Q_{fm}(10)$ are the Ljung-Box (1978) test statistics for the first 10 lags of the autocorrelation function for the standardized residuals for the futures, their squares, and for the cross-products of the standardized residuals for the futures and the world portfolio, respectively. The $p$-values are for the $\chi^2_{10}$ distribution. $S_f$ and $K_f$ are the Newey (1985) conditional moment test statistics for skewness and kurtosis, respectively. For these and for the tests of restrictions in the mean, the $p$-values are for the $\chi^2_1$ distribution. $D$ is the likelihood-ratio test statistic for a restricted model with the constraint $d_{11} = d_{12} = d_{22} = 0$.

difference between the local and the average foreign interest rate. The SF exhibits a significant LM test statistic ($p$-value of 0.02) associated with inappropriate exclusion of a term for its own conditional standard deviation. Tests for omitted variables also show that the relative difference between the individual foreign currency interest rate and an average foreign currency interest rate would add explanatory power to the equation for the conditional mean of two currencies (BP and CD). These results could indicate sources of risk additional to those captured by the single-beta formulation of the conditional capital-asset-pricing model.

The last row of Panel B reports an LR test for symmetric versus asymmetric ARCH. This LR test reveals the importance of the asymmetric form of ARCH used in the estimated models. Much of the relevance of the asymmetric form stems from the equity-portfolio component, and may be attributed to leverage.

Taken together, the results shown in Tables 1 and 3 indicate that an uncovered foreign currency spot (Eurocurrency deposit) position has significant time-varying covariance risk associated with it, as does an uncovered position in foregin currency futures. The similarity of the estimates for the spot and futures positions suggests that opposite positions in these markets should lead to an effective hedge.

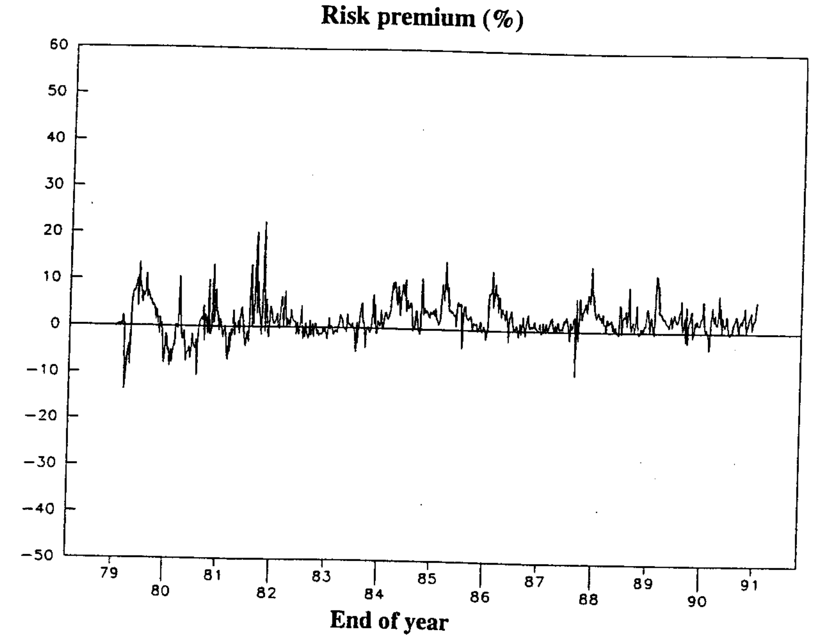
Figures 1 to 4 illustrate the above result for the CD and the JY. Plots 1 and 2 show the estimated risk premiums for the two CD positions and plots 3 and 4 do the same for the JY. The time-series behaviour of the plots are very similar for the two positions indicating that both the quantity of covariance risk and the required return per unit of risk are roughly equal for the two positions in a given currency.

The sample averages for the estimated risk premiums are positive for three (the BP, CD and JY) and negative for the other two of the five currencies for this sample. If one uses standard errors that are robust to heteroscedasticity and autocorrelation, these average risk premiums are significantly different from zero for just two of the currencies. However, our analyses estimate conditional risk premiums and, as summarized in Tables 1 to 4, the conditional risk premiums are significantly different from zero in almost all cases.

The estimated risk premiums do not explain a very large proportion of the *realized* positive and negative payoffs from the uncovered futures and spot



**Figure 1**
**CD Spot**

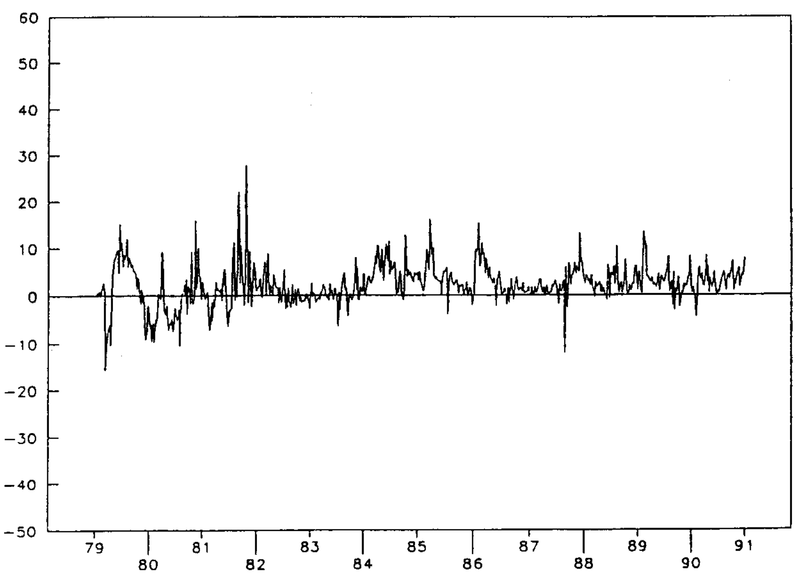



positions. The proportion of *ex post* variation that is explained by the *ex ante* risk premiums ranges from 0.7 to 2.6 per cent for the various currency positions. As noted in the Introduction, this degree of predictability compares favourably with results for other asset returns — especially given that our analysis uses higher frequency returns (weekly instead of monthly) that are generally more difficult to predict.

## 5. Concluding Comments

Uncovered positions in foreign currency spot and futures markets are risky. Estimated risk premiums change sign and exhibit considerable variation over time

**Figure 2**
**CD Futures**

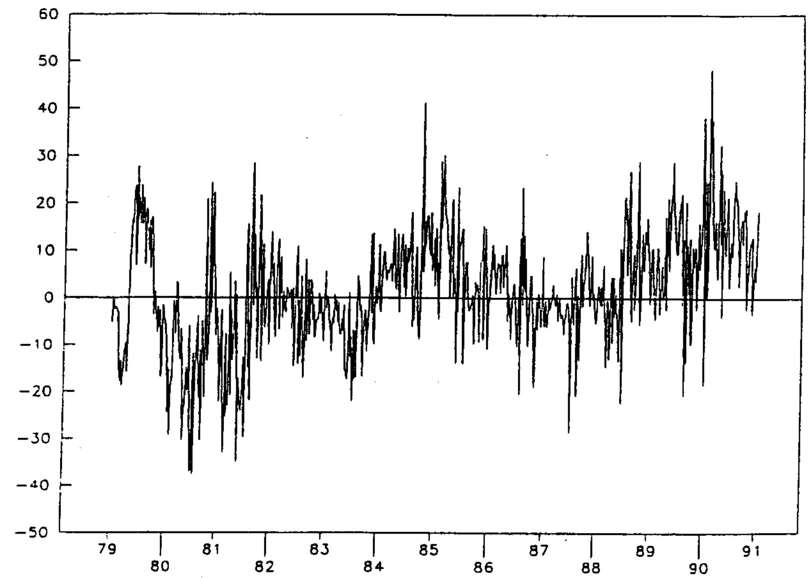



and across currencies. Although our estimated conditional risk premiums do not explain a large proportion of the realized positive and negative payoffs from the uncovered futures and spot positions, they are in almost all cases significantly different from zero.

The fact that our estimated risk premiums do not explain a very high proportion of the *ex post* or realized variation in payoffs is not unusual or surprising. Most asset *excess* returns are difficult to predict — especially at high frequencies. The empirical success of our risk-premium model should be assessed on the basis of whether or not it explains *predictable* excess returns. The empirical results summarized in Tables 1 to 4 generally support our maintained model. However,

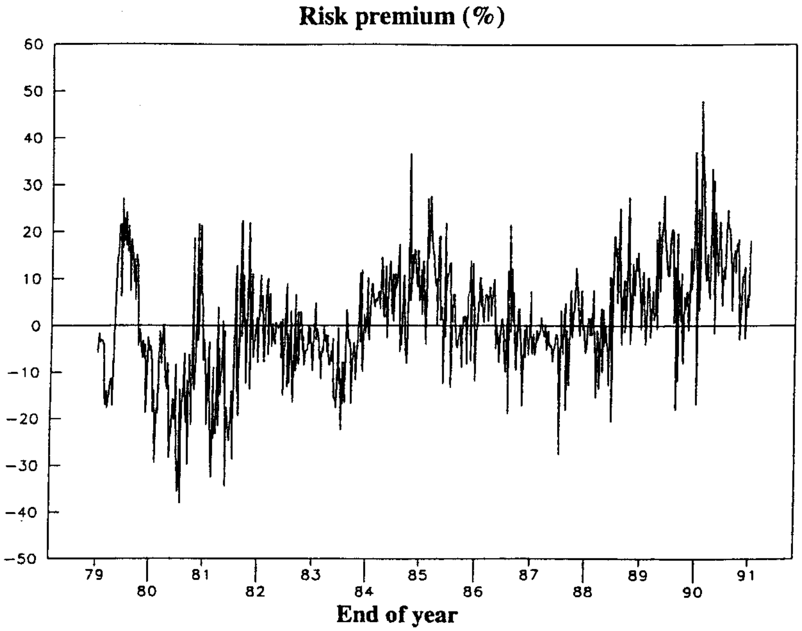


**Figure 3**
**JY Spot**




based on the specification tests in Tables 2 and 4, there was some evidence of predictable returns that were not explained by the model. This could be because our particular pricing model did not capture all the risk present in the market, or it could be due to failure of some other component of the joint hypothesis. In order to evaluate such alternatives, one needs to model them explicitly and derive the implied structure for expected returns — for example, as predicted by a multi-beta, as opposed to a single-beta, risk model.

**Figure 4**
**JY Futures**

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