NBER WORKING PAPER SERIES

IS IT RISK?
EXPLAINING DEVIATIONS FROM UNCOVERED INTEREST PARITY

Robert E. Cumby

Working Paper No. 2380

NATIONAL BUREAU OF ECONOMIC RESEARCH
1050 Massachusetts Avenue
Cambridge, MA 02138
September 1987

Comments from seminar participants at the University of Pennsylvania, the University of Chicago, Northwestern University, Yale University, and the University of Michigan, along with helpful conversations with Bill Greene, Bob Flood, Joel Hasbrouck, Robert Hodrick, John Huizinga, Bob Korajczyk, Maury Obstfeld, and Rene Schultz, and the research assistance of Kew-Chul Chee are acknowledged with thanks. The research reported here is part of the NBER's research program in International Studies. Any opinions expressed are those of the author and not those of the National Bureau of Economic Research.

NBER Working Paper #2380
September 1987

Is it Risk?
Explaining Deviations from Uncovered Interest Parity

ABSTRACT

This paper analyzes ex-ante returns to forward speculation and asks if these returns can be explained by models of a foreign exchange risk premium. After presenting evidence that both nominal and real expected speculative profits are non-zero, the paper examines if real returns to forward speculation are consistent with consumption-based models of risk premia. Estimates of the conditional covariance between real speculative returns and real consumption growth are presented and, like ex-ante returns to forward speculation, they exhibit statistically significant fluctuations over time and often change sign.

Robert E. Cumby
GSBA
New York University
100 Trinity Place
New York, NY 10006

I. Introduction

It is by now widely accepted that forward exchange rates are not unbiased predictors of future spot exchange rates and that therefore there are predictable nonzero returns to forward speculation. Several authors have pointed out that if forward speculation involves systematic risk, speculative returns should be nonzero. This observation is consistent with numerous theoretical models of a foreign exchange risk premium that have appeared in the literature. However, implementing empirical models of time-varying risk premia has proven to be very difficult in general and previous attempts have not been successful in explaining returns to forward speculation.[1]

This paper analyzes ex-ante returns to forward speculation and seeks to determine if these returns can be explained by models of a foreign exchange risk premium. After providing measures of ex-ante returns to forward speculation, the paper considers whether the returns are consistent with risk neutrality of investors and arise only due to a covariance between nominal speculative returns and the future purchasing

----

[1] Examples of unsuccessful attempts to model these ex-ante returns as a risk premium include Frankel (1982), who finds that in a portfolio balance model based on mean-variance optimization, the hypothesis of risk neutrality cannot be rejected, and Frankel and Engel (1984) who find that the data are inconsistent with the constraints implied by period-by-period mean-variance optimization. While Hansen and Hodrick (1983) are successful in modeling the risk premium in a CAPM-type model with an unobserved benchmark portfolio, Hodrick and Srivastava (1984) find that when more data are added, the constraints implied by the model are rejected. Domowitz and Hakkio (1985) find that the data are not consistent with a risk premium that depends on the conditional variance of exchange rate forecast errors. Korajczyk (1985) meets with some success in modeling forward forecast errors as time-varying risk premia, but his tests assume that real exchange rate changes are unpredictable. This assumption is rejected by Cumby and Obstfeld (1984).

power of money.[2] If this is the case, ex-ante real profits will be zero even though ex-ante nominal profits are nonzero. Since the evidence is found to be clearly inconsistent with the hypothesis of zero expected real profits, the paper then examines if real returns to forward speculation are consistent with consumption-based models of risk premia. Tests such as those suggested by Hansen and Hodrick (1983) and Gibbons and Ferson (1985) are considered first. These tests require that we assume that the conditional covariances of real returns to forward speculation and the rate of change of real consumption move together over time for all currencies. The restrictions implied by the consumption-based model and the proportionality of the conditional covariances are rejected by the data. Next, estimates of these conditional covariances are presented and the evidence shows that, like the ex-ante returns to forward speculation, they exhibit statistically significant fluctuations over time and often change sign. The comovements of the estimated ex-ante returns to forward speculation and the conditional covariances suggest that, on a qualitative level, ex-ante returns behave in a way that is consistent with consumption-based models of the foreign exchange risk premium. Importantly, a large decrease in speculative returns in all currencies that is found in 1981 is accompanied by a similar change in the conditional covariances. The assumption of proportional conditional covariances is also tested. A final set of tests is carried out in which conditional covariances are allowed to change over time and the results indicate that observed real speculative returns cannot be explained fully by consumption-based models.

----

[2] Frenkel and Razin (1980) and Engel (1984) suggest that this may be the case.

II. Measuring Ex-Ante Returns to Forward Speculation

In this section we discuss measuring ex-ante returns to forward foreign exchange speculation and examine the behavior of these ex-ante returns over time. Let S_t^j be the spot exchange rate for currency j at time t and let $F_{t,k}^j$ be the k-period forward exchange rate for currency j at time t. Both S_t^j and $F_{t,k}^j$ are expressed as units of domestic currency (dollars) per unit of foreign currency. An investor taking a long forward position in the foreign currency at time t will buy the foreign currency forward, agreeing to pay $F_{t,k}^j$ dollars at time t+k for each unit of currency j purchased. When the forward contract matures at t+k, the investor sells the foreign currency in the spot market, receiving S_{t+k}^j dollars for each unit of foreign currency sold. It will prove useful in the empirical work that follows to define a normalized return by dividing the dollar return to a long forward position in currency j by S_t^j.[3]

$$(1) \qquad \pi_{t,k}^j = (S_{t+k}^j - F_{t,k}^j)/S_t^j$$

The uncovered interest parity hypothesis states that expected profits to forward speculation are zero. That is $E_t(\pi_{t,k}^j) = 0$, where $E_t(.)$ denotes a conditional expectation given information available at time t. Tests of uncovered interest parity can be carried out by noting that if expected profits are zero, in a large sample realized profits should be unforecastable. Suppose that the econometrician observes some data X_t that is included in the information set of agents at time t. These observable data should not help predict realized profits if the uncovered interest parity hypothesis is true, so that a test of interest parity may be carried out by regressing realized profits on X_t and testing if the coefficients are jointly zero.

----

[3] Meese and Singleton (1982) find that normalization is needed to induce stationarity.

(2) $\pi^j_{t,k} = X_t\lambda_j + \mu^j_{t,k}$

Several studies using a framework like (2) have found strong evidence that the coefficients are not jointly zero and have concluded that the nonzero ex-ante returns to forward speculation are consistent with a time-varying risk premium in the foreign exchange market.[4] The (nominal normalized) risk premium may be defined as the expected return to forward speculation, $\rho^j_{t,k} = E_t(\pi^j_{t,k})$.[5]

Perhaps as a result of the failure to present an empirically tractable model of the risk premium that is consistent with the data, the risk premium is widely considered to be small. Recently, this conventional wisdom has been challenged by Fama (1984) and Hodrick and Srivastava (1986) who find that, while the risk premium may be small relative to exchange rate forecast errors, the variance of the risk premium has been approximately as large as the variance of expected exchange rate changes. Although the evidence presented in these papers shows that time variation in the risk premium has been important in explaining movements in the forward premium, it does not provide information about the magnitude of the risk premium or how it moves over time.

Fortunately, other techniques are available that allow us to examine the movements of ex-ante speculative returns over time. Since the econometrician observes a set of variables X_t, a reasonable choice of an

----

[4] A review of the literature can be found in Hodrick (1987).

[5] If the price level is uncertain, nonzero expected nominal returns to forward speculation may be due to a covariance between the future purchasing power of money and the nominal return to forward speculation even if individuals are risk neutral, so that equating the nominal return to forward speculation with a risk premium is not strictly correct. See Frenkel and Razin (1980), Engel (1984), and section III below. However, since it has become standard practice in the literature to do so, and since the evidence in section III shows that such a covariance is not the explanation for nonzero expected nominal speculative returns, we will go ahead with this definition.

estimator for ex-ante returns is the best linear predictor of $\rho^j_{t,k}$ given X_t, that is the projection,

$$(3) \quad \rho^j_{t,j} = X_t\lambda_j + u^j_{t,k}$$

The projection error, $u^j_{t,k}$, is orthogonal to X_t by construction.

Since ex-ante profits are unobservable, we work with observable realized profits, which can be decomposed into expected profits and a forecast error.

$$(4) \qquad \pi^j_{t,k} = E_t(\pi^j_{t,k}) + \varepsilon^j_{t,k}$$

The key assumption behind the econometric methodology that will allow us to make inferences about the behavior of ex-ante returns based on the observed behavior of realized returns is the assumption that expectations are rational. That is we assume that forecast errors are unforecastable given information available at the time the forecast is made, so that $\varepsilon^j_{t,k}$ is orthogonal to X_t. Combining (3) and (4) we obtain the regression equation,

$$(5) \qquad \pi^j_{t,k} = X_t\lambda_j + u^j_{t,k} + \varepsilon^j_{t,k} = X_t\lambda_j + \mu^j_{t,k}$$

Because $\pi^j_{t,k}$ and X_t are observable, this equation can be estimated and we can use $\hat{\rho}^j_{t,k}$, the fitted values from (5), as our estimates of ex-ante returns for each currency.[6] It is clear from examining (5) that the regression used to examine the behavior of ex-ante returns is the same as (2), the regression used to test uncovered interest parity.

Consistent estimates of the covariance matrix of λ_j are calculated with methods outlined by Hansen (1982) and Cumby, Huizinga and Obstfeld (1983) that allow for serial correlation and conditional heteroscedasticity in the residuals. The first of these is important since monthly

----

[6] Since both components of $\mu^j_{t,k}$ are orthogonal to X_t, OLS will be consistent provided X_t and $\mu^j_{t,k}$ are stationary and ergodic. The methods used here are the ones proposed in Mishkin (1981) to examine the behavior of ex-ante real interest rates.

observations on quarterly returns (k=3) are used in the empirical work carried out below. As a result, $\varepsilon^j_{t,k}$ will follow a second-order moving-average process. In addition, the projection error need not be serially uncorrelated. The second feature is also important since the assumption of conditional homoscedasticity of returns to forward speculation has been rejected by Cumby and Obstfeld (1984), Hodrick and Srivastava (1984), Domowitz and Hakkio (1985), and Giovannini and Jorion (1987).

There is an important difference between using equation (2) or (5) as a test of interest parity or as a model to examine the movement over time in ex-ante returns. Interest parity will be rejected if <u>any</u> information useful in predicting speculative returns is included in X_t. To construct a precise measure of ex-ante speculative returns using the methodology described here, however, most of the relevant information for predicting speculative returns must be included in X_t, so that $u^j_{t,3}$ is small. When this is the case, $\mu^j_{t,3}$ will have essentially the time series properties of the forecast error, $\varepsilon^j_{t,3}$. Thus a diagnostic check of the adequacy of the specification of X_t is to see whether the residuals from (5) have the properties of an MA(2). Since the projection error need not be serially uncorrelated, and may even resemble an MA(2), finding that the residuals exhibit the properties of an MA(2) is not conclusive evidence that the projection error is unimportant. However, finding that the residuals do not have the properties of an MA(2) is evidence that the projection error is important.

We now turn to the results from estimating the ex-ante returns to forward speculation. We consider the U.S. dollar relative to five currencies, the U.K. pound sterling, Deutsche mark, Canadian dollar, Swiss franc, and French franc, over the period January 1974 to December 1986. In all cases we adopt a three-month holding period so as hopefully to minimize the problems arising from the inexact timing of monthly price data.[7]

In estimating the ex-ante nominal return to a long forward position in the jth currency, the X_t variables are chosen following Hodrick and Srivastava (1984) to be a constant, the forward premia, and the squared forward premia for each of the 5 currencies.[8] Before looking at the estimated ex-ante returns, we should first examine the autocorrelation of residuals, reported in Table 1A, in order to see if they behave like an MA(2). Under the null hypothesis that the residuals are serially uncorrelated, the autocorrelations have asymptotic standard errors of approximately $1/\sqrt{T}$, or 0.08. In all cases but the second autocorrelation for the Canadian dollar, the first two autocorrelations are significantly different from zero and of the expected magnitude. The other autocorrelations are small, indicating that the residuals reasonably approximate a second-order moving average. Table 1B contains the χ^2 statistics to test the hypothesis that all coefficients but the constant are zero. As can be seen, there is strong evidence against uncovered interest parity in all five of the currencies as well as against the hypothesis that expected profits are constant. Table 1 also reports the results of tests of uncovered interest parity in which the U.K. pound and the Deutsche mark are used as the base currency in place of the U.S. dollar. These tests also strongly reject uncovered interest parity.

Figures 1 and 2 display the estimated profits to a long forward position in the DM and the U.K. pound respectively. The solid lines in

----

[7] The data appendix contains a detailed description of the data and sources.

[8] All tests reported in the paper have also been carried out lagging X_t one period to accommodate possible reporting lags. In no instance does this affect the results of the hypothesis tests.

the center show the estimated values of the ex-ante profits. The dashed lines provide 95% confidence intervals.[9] Similar plots of the ex-ante return to speculation in the other three currencies are not presented to conserve space. The other three plots are very similar to those presented here, except that the magnitude of Canadian dollar returns is smaller than the others. The ex-ante returns move considerably over time and frequently change sign. While the estimates of the returns are somewhat imprecise as is evidenced by the sometimes wide confidence intervals, periods when the ex-ante returns are significantly negative can be identified in all currencies and periods in which they are significantly positive can be identified in most of the currencies. The estimated magnitude of the ex-ante return is frequently quite large, and perhaps disconcertingly so. It is interesting to note, however, that the magnitudes estimated by Domowitz and Hakkio (1985) using very different methods are similarly large. In addition, the large magnitudes are generally accompanied by large standard errors, making the large magnitudes less bothersome.

An interesting feature of all of the speculative returns is the sharp decline in the expected profits in all of the currencies that occurs in mid-1981. Among the explanations of the dollar's rise in this period is a portfolio shift in favor of dollar-denominated assets. If this explanation is correct, we would expect to see a decrease in the expected return to dollar-denominated assets during this period. Instead, the expected return to dollar-denominated investments rises. The evidence

----

[9] The standard errors are calculated according to the formula in Mishkin (1981), assuming that the variances of the forecast errors are large relative to the variances of the projection errors. As was noted earlier, this assumption on the relative importance of the two components of the composite error term in equation (5) is not contradicted by the correlogram of the residuals in Table 1.

presented here is inconsistent with the claim that a portfolio shift was behind the strong dollar in 1981.[10]

Figure 3 presents the ex-ante return to a forward position that is long in French francs relative to the DM. As is the case with the other figures, periods of significantly positive and significantly negative returns can be discerned. Along with the results in Table 1B, this indicates that finding significant speculative returns is not specific to the choice of the U.S. dollar as a base currency.

III. Modeling Ex-Ante Returns as a Risk Premium

This section uses models of utility-maximizing representative agents to explain the returns observed in section II. The first explanation investigated is that agents are risk neutral and that nominal profits arise only due to a covariance of nominal profits with the future price level. After rejecting this explanation, consumption-beta models of a risk premium are considered and tests such as those suggested by Hansen and Hodrick (1983) and Gibbons and Ferson (1985) are implemented.

Models of intertemporal asset pricing assume that consumer-investors maximize the utility of consumption over their lifetime subject to a sequence of budget constraints. The optimality condition of a representative domestic consumer-investor has been used to examine the pricing of forward foreign exchange contracts by Stockman (1978), Frenkel and Razin (1980), Hansen and Hodrick (1983), and Mark (1985), among others. Their work shows that if a representative consumer-investor is at an interior optimum,

$$(6) \quad E_t[(U'(c_{t+k})/U'(c_t))\pi^j_{t,k}/(p_{t+k}/p_t)] = 0.$$

----

[10] Obstfeld (1985) also presents evidence that a portfolio shift is not behind the rise of the dollar during this period.

where c_t is real consumption and p_t is the price level.

Frenkel and Razin (1980) and Engel (1984) use this condition to point out that even if investors are risk neutral (so that the marginal utility of consumption is constant), expected nominal profits to forward speculation may be nonzero but expected real profits should be zero. Thus rejection of uncovered interest parity does not provide evidence of a risk premium in the forward foreign exchange market. However, finding ex-ante real profits would provide such evidence. Define the real return to forward speculation in currency j, $r^j_{t,k}$, as

(7) $$r^j_{t,k} = [(S^j_{t+k} - F^j_{t,k})/S^j_t]/(p_{t+k}/p_t) = \pi^j_{t,k}/(p_{t+k}/p_t).$$

The hypothesis that expected real returns are zero may then be tested in a manner analogous to the tests of interest parity described above. Again, we assume the econometrician observes some data X_t that are included in the time-t information set of agents, and consider the projection of expected real profits onto X_t.

(8) $$E_t(r^j_{t,k}) = X_t\alpha_j + u^j_{t,k}$$

Next, decomposing realized real profits, $r^j_{t,k}$ into their conditional expectation and a forecast error, δ, we have,

(8') $$r^j_{t,k} = X_t\alpha_j + u^j_{t,k} + \delta^j_{t,k} = X_t\alpha_j + \eta^j_{t,k}.$$

If expected real profits are zero, then in a large sample, they should be unforecastable given information available when the speculative position is taken. The hypothesis of no expected real profits can be tested by testing that the coefficients in (8') are zero.

In testing for nonzero expected real profits we need to consider what X_t variables to use in the regressions. In principle any variables in the information set are reasonable candidates. In the first tests carried out, we use the forward premia and the squared forward premia as was done above in order to determine if the same data that prove useful for predicting nominal profits are also useful for predicting real profits. Next, several "real" variables are considered. The X_t used in the second set of tests are the forward premia in each of the five currencies, U.S. inflation $(p_{t+3}/p_t - 1)$ lagged 3 and 12 months, the rate of change of consumption $(c_{t+3}/c_t - 1)$ lagged 3 months, U.S. industrial production growth $(IP_{t+3}/IP_t - 1)$ lagged three months, and the U.S. terms of trade (TOT_t). These data are chosen since forward premia have proven useful in predicting nominal returns. Consumption and industrial production growth and the terms of trade are employed since various models suggest that these should affect savings and investment decisions and therefore affect equilibrium expected real returns.

Table 2 contains the results from regressions of realized real speculative profits on both sets of X_t variables along with the χ^2 statistics testing the hypothesis that expected real profits are constant. Table 2A contains the results obtained when the forward premia and squared forward premia are used, while Table 2B contains the results obtained from the second set of X_t variables. As can be seen from the tables, the hypotheses that expected real profits are constant is rejected at standard significance levels in all cases. Thus the evidence clearly shows that, contrary to the suggestions of Frenkel and Razin (1980) and Engel (1984), the finding of nonzero nominal profits cannot be attributed solely to a covariance between nominal returns and an uncertain future price level. The evidence is clearly inconsistent with the hypothesis that investors are risk neutral.

Stulz (1981,1984), following Breeden (1979), derives a consumption-based international asset pricing model. The assumption that trading takes place continuously allows him to move to the limit of continuous time and derive an equilibrium relationship between asset returns.[11] A discrete-time conditional consumption-based asset pricing model for a representative domestic consumer-investor can be obtained using the results in Hansen and Richard (1987), where a "generic" conditional asset pricing model is examined.[12] While restrictions on equilibrium returns may be obtained in this way, the means by which Stulz's results on aggregation across countries can be obtained in a conditional discrete-time framework remain unknown. All that we require here, however, is that a consumption-based capital asset pricing model exist for a representative domestic consumer-investor. In Stulz's model, expected real returns to a long forward position in currency j will satisfy,

$$(10)\quad E_t(r^j_{t,k}) = \beta_{j,t}E_t[r^p_{t,k} - r^z_{t,k}],$$

where $r^p_{t,k}$ is the real return on the benchmark portfolio, $r^z_{t,k}$ is the real rate of return on a portfolio whose real return is conditionally uncorrelated with domestic real consumption, and $\beta_{j,t}$ is the "consumption beta" of forward speculation in the jth currency from the point of view of a domestic investor.

$$\beta_{j,t} = \mathrm{cov}_t(r^j_{t,k},\Delta c_{t+k}/c_t)/\mathrm{cov}_t(r^p_{t,k},\Delta c_{t+k}/c_t),$$

where $\Delta c_{t+k} = c_{t+k} - c_t$.

Since (10) must hold for all assets, we can divide $E_t(r^j_{t,k})$ by the expected return on forward speculation in an arbitrarily chosen

----

[11] Grossman and Shiller (1982) show how a consumption-beta model can be derived from the first-order conditions of a representative consumer-investor such as (6), by taking the limit as the trading interval goes to zero. They point out that distributional assumptions or assumptions about the functional form of the utility function are alternatives to the use of continuous-time analysis.

[12] As Hansen and Richard (1987) point out, the consumption-based capital asset pricing model implies that the benchmark return in their analysis is the return on the aggregate consumption portfolio.

reference currency, currency 1.[13] Doing so we obtain,

$$(11)\quad E_t(r^j_{t,k}) = (\beta_{j,t}/\beta_{1,t})E_t(r^1_{t,k})$$

If we combine (11) with the projection equations, (8) and (8′), we obtain,

$$r^j_{t,k} = (\beta_{j,t}/\beta_{1,t})(X_t\alpha_1) + \eta^1_{t,k}$$

Thus $\alpha_j = (\beta_{j,t}/\beta_{1,t})\alpha_1$. Gibbons and Ferson (1985) show that when the ratio of the consumption betas is constant over time (or, equivalently, the conditional covariances between asset returns and the rate of change of real consumption are proportional across currencies), a test of the asset pricing model (10) can be carried out by estimating a system of projection equations and testing the hypothesis that the coefficients in each equation are proportional to the coefficients in the first equation.[14] If there are N assets and k regressors in each of the projection equations, there will be Nk regressors in the system but only k + (N-1) parameters when the proportionality restrictions are imposed. There are thus Nk - (k+N-1) parameter restrictions that can be tested. If the model is correct and if the auxiliary assumptions concerning the constancy of the relative consumption betas and the rationality of expectations are correct, these parameter restrictions should be satisfied by the data.

Estimation of the restricted system of equations and testing of the parameter restrictions can be carried out using Hansen's (1982) generalized method of moments (GMM) procedure. Hansen and Hodrick (1983) show how tests to proportionality restrictions can be carried out using the

----

[13] Gibbons and Ferson (1985) propose the tests currently described as a means of testing the Sharpe-Lintner version of the CAPM. The tests may be thought of as tests of any single beta asset pricing model.

[14] The assumption that the conditional covariances are proportional across currencies is tested below.

value of the criterion function, which is distributed as χ^2 with degrees freedom equal to the number of restrictions.

Hansen and Hodrick (1983) test the restrictions implied by a single-beta model of the foreign exchange risk premium. The test they carry out is equivalent to the Gibbons-Ferson test. Perhaps this is not surprising since both are tests of single-beta asset pricing models. The fact that the two tests are identical is obscured somewhat by differences in interpretation and motivation of the tests. Hansen and Hodrick assume that the betas are constant and treat the expected return on the benchmark portfolio as an unobserved latent variable assumed to be linearly related to some data X_t. Gibbons and Ferson, on the other hand substitute out the expected benchmark return by using an arbitrarily chosen reference asset and derive a set of proportionality restrictions that are identical to those obtained by Hansen and Hodrick.

Table 3 contains the results of the tests of the consumption-based models of the risk premium using the "real" X_t variables. Estimation of the full system of five equations each of which contains 11 regressors proved to be computationally infeasible. We therefore carry out the tests in reduced systems of three currencies each.[15] In each of these reduced systems there are 33 orthogonality conditions and 13 parameters to be estimated. There are thus 20 parameter restrictions in each system. The values of the criterion functions, which are χ^2 random variables with 20 degrees freedom, are 54.06 for the first set of currencies (Deutsche mark, Canadian dollar, Swiss franc), 83.38 for the second set

----

[15] Estimation requires the inversion of a matrix that is Nk x Nk, which is in this case 55 x 55. Attempts to compute this inversion proved unsuccessful. If the restrictions are rejected by the data, the use of the three smaller systems does not present any problems in interpreting the test results since the full system test would simply provide stronger rejections.

of currencies (Deutsche mark, U.K. pound, Swiss franc), and 88.01 for the third set of currencies (Deutsche mark, U.K. pound, French franc). The restrictions implied by the single-beta model are then rejected at any reasonable significance level in all three cases.[16]

IV. Modeling Conditional Covariances

The behavior of the conditional covariance of speculative returns and the rate of change of consumption plays a central role in consumption-based models of the risk premium. In this section we discuss modeling this conditional covariance with several goals in mind. First, if we are to explain ex-ante speculative profits as a risk premium using consumption-based models, the conditional covariance between consumption and real speculative returns must move over time.[17] In addition, the results discussed in section II suggest that ex-ante profits change sign over time. Since the expected excess return on $r^p_{t,k}$ cannot be negative, the conditional covariance must change sign over time as the risk premium changes sign. Second, it may be that the rejection of the restrictions implied by the consumption-beta model is due to time-varying relative consumption betas. If the conditional covariances can be modeled, the constancy of the relative consumption betas can be tested. Finally, if we find that the relative consumption betas change over time, we want to determine if the movement they exhibit can account

----

[16] Similar tests carried out using the DM and the U.K. pound as base currencies in place of the U.S. dollar as well as tests using the forward premia and squared forward premia as the relevant X_t. In all cases the proportionality restrictions are rejected at standard significance levels.

[17] Hodrick and Srivastava (1984) find that ex-ante profits cannot be explained solely by variation in the expected excess return on a benchmark portfolio.

for ex-ante speculative profits.

The estimation of the conditional covariance may be carried out by extending the results of Amemiya (1977) and Hasbrouck (1985).[18] We are interested in estimating the conditional covariance between the rate of change of consumption and the real return to forward speculation,

$$\sigma_{j,t} = \mathrm{cov}_t(r^j_{t,k}, \Delta c_{t+k}/c_t) = E_t\{[r^j_{t,k} - E_t(r^j_{t,k})][\Delta c_{t+k}/c_t - E_t(\Delta c_{t+k}/c_t)]\}$$

The econometrician, who is assumed to observe a set of variables, X_t, can use as an estimate of the conditional covariance the projection of $\sigma_{j,t}$ onto X_t,

$$\sigma_{j,t} = X_t\theta_j + \omega*^j_t.$$

It will prove convenient to rewrite the projection as,

$$\eta^j_{t,k}\eta^c_{t,k} = X_t\theta_j + \omega*^j_t + (\eta^j_{t,k}\eta^c_{t,k} - \sigma_{j,t}) = X_t\theta_j + \omega^j_t,$$

where $\eta^j_{t,k}$ and $\eta^c_{t,k}$ are the disturbances from projections of $r^j_{t,k}$ and $\Delta c_{t+k}/c_t$ onto X_t, respectively. Since $\eta^j_{t,k}$ and $\eta^c_{t,k}$ unobservable, we need to work with the residuals, $\hat{\eta}^j_{t,k} = \hat{\eta}^j_{t,k} - X_t(\hat{\alpha}_j - \alpha_j)$ and $\hat{\eta}^c_{t,k} = \eta^c_{t,k} - X_t(\hat{\alpha}_c - \alpha_c)$. The projection can then be rewritten in terms of observables,

$$(12)\quad \hat{\eta}^j_{t,k}\hat{\eta}^c_{t,k} = X_t\theta_j + \omega^j_t - \eta^c_{t,k}X_t(\hat{\alpha}_j - \alpha_j) - \eta^j_{t,k}X_t(\hat{\alpha}_c - \alpha_c) + X_t(\hat{\alpha}_c - \alpha_c)X_t(\hat{\alpha}_j - \alpha_j)$$

In the appendix we show that the OLS estimate of θ_j is consistent and asymptotically normal with a covariance matrix that can be consistently estimated using the techniques described in Hansen (1982) and Cumby,

----

[18] Hasbrouck (1985) extends the results in Amemiya (1977) in several important directions. Most importantly, he allows the regressors to be stochastic, does not require that the regression disturbance be normal, and allows the addition of a stochastic disturbance to the linear variance function.

[19] It should be pointed out that since we are examining the conditional covariances of the η and not the conditional covariances of the δ, the covariances we estimate are the sum of the covariances of the projection errors and the covariances real returns and real consumption. Therefore any inference about the movement of the conditional

Huizinga, and Obstfeld (1983).[19]

Once consistent estimates of θ_j and its asymptotic covariance matrix are obtained, we can test hypotheses about the conditional covariance of real consumption and the real return to forward speculation. The first of these hypotheses is the constancy of this conditional covariance, which implies that all elements of θ_j are zero except for the constant term. Next, if the hypothesis of a constant conditional covariance is rejected, we need to determine if the comovements of the conditional covariance and the returns to forward speculation are consistent with the consumption-based model of the risk premium. We can do this in three steps. First, we use the fitted values from the projections (12) to estimate the conditional covariances and to examine their movements over time. Next, we can test the assumption of constant relative consumption betas required for the Gibbons-Ferson test by using the projection equations (12). If relative consumption betas are constant over time, the conditional covariances must all move together over time. The hypothesis that the conditional covariances move together can be tested by determining if the coefficients in the projection equations (12) are proportional across currencies.

In estimating and testing hypotheses about the conditional covariances, the choice of the data to include in X_t must again be made. It seems natural to use the same information to estimate the behavior of conditional first moments and conditional second moments, so the "real"

----

covariance of real returns and real consumption over time based on the evidence presented here is conditional on assumptions we make concerning the movements of covariance of the projection errors. If the data X_t do a good job of describing the movements of the $r^j_{t,k}$ over time, we may reasonably assume that the covariance of projection errors is small. The estimates will then be dominated by movements in the conditional covariances of real returns and real consumption.

X_t variables are again assumed to make up the relevant information set. Prior to proceeding with estimation, a problem with consumption data should be confronted. Published data measure consumption over an interval rather than at a point in time. Using the results in Breeden, Gibbons, and Litzenberger (1986), it can be shown that if monthly data sampled quarterly are used and if the covariance between real returns and real consumption growth is constant, the estimate of the covariance obtained using interval consumption data will understate the true "spot" covariance by twenty percent. The dependent variables in the projections (12) are multiplied by 1.2 prior to estimation to correct for this bias. This will, of course, leave the test statistics unchanged but it will change the estimated magnitudes of the conditional covariances.

Table 4A contains the χ^2 statistics for testing the the constancy of conditional covariances. Recall that the the conditional covariances must vary over time if a time-varying risk premium is to be explained by the consumption-based models. In three cases the hypothesis of constant conditional covariances can be rejected at standard significance levels. Given that the covariances change over time, do they do so in a way that explains the behavior of ex-ante returns? First, we can examine plots of the conditional covariances over time, an example of which is Figure 4. The conditional covariance of real returns to a long forward position in DM and the rate of change of real consumption exhibits substantial fluctuations and frequently changes sign. As was the case with the ex-ante returns, the standard errors are generally large but periods of significantly positive and significantly negative conditional covariances can be discerned. Plots of the other four conditional covariances exhibit similarly large fluctuations and wide confidence intervals. Period in which the estimated conditional covariance is significantly different from zero can be discerned in all cases. The strong rejections of the constancy of the conditional covariances reported in Table 3, along with the fairly wide confidence intervals indicate that, while the data do not contain enough information to allow us to determine with great confidence the value of the conditional covariance at any point in time, they do contain enough information to allow us to determine that the conditional covariances are not constant over time. The relatively large standard errors suggest that part of the volatility exhibited by the point estimates of the conditional covariances is due to sampling variation.

Inspection of Figure 4 suggests that the comovements of the conditional covariance and the ex-ante speculative return are at least qualitatively consistent with the predictions of the consumption-based models of the risk premium. Importantly, the sharp drop in the ex-ante returns in all currencies relative to the dollar in 1981 coincides with a decrease in the conditional covariance in Figure 4. Similar results are found for each of the the other four currencies except the U.K. pound. While several other movements of the estimated expected returns coincide with similar movements of the estimated conditional covariance, not all significant sign changes of the estimated ex-ante return coincide with sign changes of the estimated conditional covariance.

Table 4B contains the results of the tests of proportionality of the conditional covariances for the three sets of currencies examined above. The tests are carried out using the GMM procedure used in carrying out the Gibbons-Ferson tests. Again there are 33 orthogonality conditions in the each system and 13 parameters to be estimated in each so that each system has 20 restrictions to be tested. The table reports the $\chi^2(20)$ statistics for the hypothesis that the conditional covariances are proportional. The proportionality constraints are not rejected at standard significance levels for any of the combinations. Thus a violation of the assumption of proportional conditional covariances cannot account for all of the strong rejections found when the Gibbons-Ferson and Hansen-Hodrick tests are carried out.

As a final check of the ability of the consumption-beta model to explain observed returns to forward speculation, a series of regressions (not reported) in run to determine if these returns are consistent with the equilibrium condition, (11), when the relative consumption betas are allowed to change over time. In each regression, realized returns to forward speculation in one currency relative to realized returns to forward speculation in a reference currency (DM) are regressed on a constant and the ratio of the fitted value for the conditional covariance for speculation in that currency relative to the conditional covariance for DM speculation. If relative returns depend linearly on relative consumption betas as the model predicts, we expect to find a slope coefficient of one. Instead, the estimated slope coefficients are close to zero in all four cases, and are in fact positive in only two cases. Even when conditional covariances are allowed to change over time, the consumption-based model of the risk premium does not appear to be able to explain observed real returns to forward speculation.

IV. Concluding remarks

This paper presents evidence of statistically significant ex-ante returns to forward speculation in five currencies relative to the dollar as well as for four currencies relative to the DM and the U.K. pound, and finds that these ex-ante returns exhibit considerable fluctuations over time and are positive in some periods and negative in others. The paper then goes on the determine whether these returns are consistent with models of risk premia that have been proposed in the literature. While the comovements of the estimated ex-ante returns to forward speculation and the estimated conditional covariances between these returns and consumption growth are broadly consistent with the predictions of the consumption-based model, on the whole the evidence suggests that the consumption-based model does not provide an adequate description of returns to forward speculation.

At least three possible explanation for the failure of the consumption-based model fully to explain observed returns to forward speculation apart from any weakness in the model come to mind. First, the failure may be due to data problems such as those encountered in measuring consumption or prices. A second explanation may lie in the possibility that agents may have rationally assigned finite probabilities to events such as policy changes that were not realized in the sample. If this is the case, in small samples we may find that the apparent ex-post bias in forward rates exceeds the true bias.[20] Finally, nonseparability over time of the utility function may account for the failure of the model to explain speculative returns.

----

[20] Lewis (1986) explores the implications of this problem in an explicit model of stochastic policy rules, and Stulz (1986) presents a model of learning behavior that can produce apparent ex-post forward rate bias.

References

Amemiya, Takeshi, 1977, A note on a heteroscedastic model, Journal of Econometrics, 6, 365-370.

Breeden, Douglas T., 1979, An intertemporal asset pricing model with stochastic consumption and investment opportunities, Journal of Financial Economics, 7, 265-196.

Breeden, Douglas T., Michael R. Gibbons, and Robert H. Litzenberger, 1986, Empirical tests of the consumption-oriented CAPM, working paper, Stanford University, 1986.

Cumby, Robert E., John Huizinga, and Maurice Obstfeld, 1983, Two-step two-stage least squares estimation in models with rational expectations, Journal of Econometrics, 11, 333-355.

Cumby Robert E. and Maurice Obstfeld, 1984, International interest rate and price level linkages under floating exchange rates: A review of recent evidence, in J. F. O. Bilson and R. Marston (eds), Exchange Rate Theory and Practice, Chicago: University of Chicago Press for the National Bureau of Economic Research.

Domowitz, Ian, and Craig S. Hakkio, 1985, Conditional variance and the risk premium in the foreign exchange market, Journal of International Economics, 19, 47-66.

Engel, Charles M., 1984, Testing for the absence of expected real profits from forward market speculation, Journal of International Economics, 17, 299-308.

Fama, Eugene F., 1984, Forward and spot exchange rates, Journal of Monetary Economics, 14, 319-338.

Frankel, Jeffrey A., 1982, In search of the exchange risk premium: A six-currency test assuming mean-variance optimization, Journal of International Money and Finance, 1, 255-274.

Frankel, Jeffrey A. and Charles M. Engel,1984, Do asset demand functions optimize over the mean and variance of real returns? A six currency test, Journal of International Economics, 17, 309-323.

Frenkel, Jacob A. and Assaf Razin, 1980, Stochastic prices and tests of efficiency of foreign exchange markets, Economics Letters, 6, 165-170.

Gibbons, Michael R. and Wayne Ferson, 1985, Testing asset pricing models with changing expectations of an unobservable market portfolio, Journal of Financial Economics, 14, 217-236.

Giovannini, Alberto and Philippe Jorion, 1987, Interest rates and risk premia in the stock market and in the foreign exchange market, Journal of International Money and Finance, 6, 107-123.

Grossman, Sanford J. and Robert J. Shiller, 1982, Consumption correlatedness and risk measurement in economies with non-traded assets and heterogeneous information, Journal of Financial Economics, 10, 195-210.

Hansen, Lars P., 1982, Large sample properties of generalized method of moments estimators, Econometrica, 50, 1029-1054.

Hansen, Lars P. and Robert J. Hodrick, 1983, Risk averse speculation in the forward foreign exchange market: An econometric analysis of linear models, in J.A. Frenkel (ed), Exchange Rates and International Macroeconomics, Chicago: University of Chicago Press for the National Bureau of Economic Research.

Hansen, Lars P. and Scott F. Richard, 1987, The role of conditioning information in deducing testable restrctions implied by dynamic asset pricing models, Econometrica, 55, 587-613.

Hasbrouck, Joel, 1985, Ex-ante uncertainty and ex-post variance in linear return models: An econometric analysis, working paper, New York University, Graduate School of Business Administration.

Hodrick, Robert J., 1987, The Empirical Evidence on the Efficiency of Forward and Futures Foreign Exchange Markets, forthcoming.

Hodrick, Robert J. and Sanjay Srivastava, 1984, An investigation of risk and return in forward foreign exchange, Journal of International Money and Finance, 3, 5-29.

Hodrick, Robert J. and Sanjay Srivastava, 1986, The covariation of risk premiums and expected future exchange rates, Journal of International Money and Finance, 5, S5-21.

Korajczyk, Robert A., 1985, The pricing of forward contracts for foreign exchange, Journal of Political Economy, 93, 346-368.

Lewis, Karen K., 1986, The implications of stochastic policy processes for the peso problem under flexible exchange rates, working paper, New York University Graduate School of Business Administration.

Mark, Nelson C., 1985, On time-varying risk premia in the foregn exchange market: An econometric analysis, Journal of Monetary Economics, 16, 3-18.

Meese, Richard A. and Kenneth J. Singleton, 1982, On unit roots and the empirical modeling of exchange rates, Journal of Finance, 37, 1029-1037.

Mishkin, F.S., 1981, The Real Interest Rate: An Empirical Investigation, in K. Brunner and A.H. Meltzer (eds), The Costs and Consequences of Inflation, Carnegie-Rochester Conference Series on Public Policy, Vol. 15. (Supplememt to the Journal of Monetary Economics).

Obstfeld, Maurice, 1985, Floating exchange rates: Experience and prospects, Brookings Papers on Economic Activity, 1985:2, 369-450.

Stockman, Alan, 1978, Risk, information, and forward exchange rates, in J. A. Frenkel and H. G. Johnson (eds), The Economics of Exchange Rates: Selected Studies, Reading MA: Addison Wesley, 159-178.

Stulz, Rene M., 1981, A model of international asset pricing, Journal of Financial Economics, 9, 383-406.

Stulz, Rene M., 1984, Pricing capital assets in an international setting: An introduction, Journal of International Business Studies, 15, 55-74.

Stulz, Rene M., 1986, Time varying risk premia, imperfect information, and the forward exchange rate, working paper, Ohio State University.

White, Halbert, (1984), Asymptotic Theory for Econometricians, New York: Academic Press. 59-178.

APPENDIX

This appendix shows the consistency and asymptotic normality of the least squares estimate of the parameters in the conditional covariance regression, (12). Recall that the model is,

$$\hat{\eta}^j_{t,k}\hat{\eta}^c_{t,k} = X_t\theta_j + \omega^j_t - \eta^c_{t,k}X_t(\hat{\alpha}_j-\alpha_j) - \eta^j_{t,k}X_t(\hat{\alpha}_c-\alpha_c) + X_t(\hat{\alpha}_c-\alpha_c)X_t(\hat{\alpha}_j-\alpha_j)$$

where $\hat{\eta}^j_{t,k}$ and $\hat{\eta}^c_{t,k}$ are the least squares residuals from

(8′) $r^j_{t,k} = X_t\alpha_j + u^j_{t,k} - \varepsilon^j_{t,k} = X_t\alpha_j + \eta^j_{t,k}$.

The following regularity conditions are assumed to hold.

A1) X_t, $\omega*^j_t$, and $\varepsilon^j_{t,k}$ are jointly stationary and ergodic.

A2) $E(\omega*^j_t|X_t) = 0$.

A3) $E(\varepsilon^j_t|X_t) = 0$.

A4) $E(u^j_t|X_t) = E(u^c_t|X_t) = E(u^j_t u^c_{t,k}|X_t) = 0$.

A5) $\text{plim}(X'X/T) = M$ exists and is of full rank, where X is a matrix with a typical row X_t.

A6) $\text{Lim}_{t\to\infty}(1/T)E(X'\omega\omega'X) = \Omega$ exists and is positive definite.

I. We now prove the consistency of the OLS estimate, θ_j.

$$(\hat{\theta}_j-\theta_j) = (X'X/T)^{-1}(1/T)\Sigma X_t'[\omega^j_t-\eta^c_{t,k}X_t(\hat{\alpha}_j-\alpha_j) - \eta^j_{t,k}X_t(\hat{\alpha}_c-\alpha_c) + X_t(\hat{\alpha}_c-\alpha_c)X_t(\hat{\alpha}_j-\alpha_j)]$$

Consider each element of this sum in turn.

(a) $(1/T)\Sigma X_t'\omega^j_t = (1/T)\Sigma X_t'(\omega*^j_t + (\eta^j_{t,k}\eta^c_{t,k} - \sigma_{j,t}))$

The first part of this, $(1/T)\Sigma X_t'\omega*^j_t$ converges in probability to zero by A2 and ergodicity. Next consider,

$$(1/T)\Sigma X_t'(\eta^j_{t,k}\eta^c_{t,k} - \sigma_{j,t}) = (1/T)\Sigma X_t'(\varepsilon^j_{t,k}\varepsilon^c_{t,k} + u^j_{t,k}u^c_{t,k} - \sigma_{j,t}).$$

Now, $\varepsilon^j_{t,k}\varepsilon^c_{t,k} - \sigma_{j,t}$ is just the deviation of a random variable from its conditional expectation and, since X_t is assumed to be in I_t, is orthogonal to X_t. The second part of this term then converges in probability to zero, leaving, $(1/T)\Sigma X_t'(u^j_{t,k}u^c_{t,k})$, which converges in probability to zero, by A4.[1]

(b) $(1/T)\Sigma X_t'\eta^j_{t,k}X_t(\hat{\alpha}_c-\alpha_c)$

First consider $(1/T)\Sigma X_t'\eta^j_{t,k}X_t = (1/T)\Sigma X_t'\varepsilon^j_{t,k}X_t - (1/T)\Sigma X_t'u^j_{t,k}X_t$. The first part of this expression converges in probability to zero by A3 and ergodicity. The second part of this expression converges in probability to zero By A4 and ergodicity. Finally, $(\hat{\alpha}_j-\alpha_j)$ converges in probability to zero since, by assumption, $\hat{\alpha}_j$ is a consistent estimate of α_j. Therefore the product of these three terms converges in probability to zero.

(c) $(1/T)\Sigma X_t'\eta^c_{t,k}X_t(\hat{\alpha}_j-\alpha_j)$

This third element of the sum can be shown to converge in probability to zero by an argument identical to that in (b).

(d) $(1/T)\Sigma X_t'X_t(\hat{\alpha}_c-\alpha_c)X_t(\hat{\alpha}_j-\alpha_j) = (1/T)\Sigma X_t'[\Sigma\Sigma X_{it}(\hat{\alpha}_{c,i}-\alpha_{c,i})X_{kt}(\hat{\alpha}_{j,k}-\alpha_{j,k})]$.

This can be rewritten as, $(1/T)\Sigma\Sigma\Sigma X_t'X_{it}(\hat{\alpha}_{c,i}-\alpha_{c,i})X_{kt}(\hat{\alpha}_{j,k}-\alpha_{j,k})$, a typical element of which is, $(1/T)(X_t'X_{it}X_{kt})(\hat{\alpha}_{c,i}-\alpha_{c,i})(\hat{\alpha}_{j,k}-\alpha_{j,k})$. Now, $(1/T)\Sigma\Sigma\Sigma X_t'X_{it}X_{kt}$ converges in probability to some vector and $(\hat{\alpha}_{c,i}-\alpha_{c,i})$ and $(\hat{\alpha}_{j,k}-\alpha_{j,k})$ each converge in probability to zero. Thus, the product converges in probability to zero.

Since each element of the sum converges in probability to zero and since $(X'X/T)^{-1}$ converges in probability to M^{-1} by A5, $(\hat{\theta}_j - \theta_j)$ converges in probability to zero.

II. Now we need to establish the asymptotic normality of the least

----

[1] We can see here that since we only measure expected returns and the expected rate of change of consumption with error, any of our inference about the comovements of the two are only valid if the projection error covariance is negligable. If, instead of A4, we were to assume that the conditional projection error covariance was constant but nonzero, the constant term in θ_j would be inconsistent but the slope coefficients would be unaffected.

squares estimator, θ_j. Consider

$$\sqrt{T}(\hat{\theta}_j-\theta_j) = (X'X/T)^{-1}(1/\sqrt{T})\Sigma X_t'[\omega_t^j - \eta_{t,k}^c X_t(\hat{\alpha}_j-\alpha_j) - \eta_{t,k}^j X_t(\hat{\alpha}_c-\alpha_c) + X_t(\hat{\alpha}_c-\alpha_c)X_t(\hat{\alpha}_j-\alpha_j)].$$

As above, we will consider each part of this expression in turn.

(a) $(1/\sqrt{T})\Sigma X_t\omega_t^j$

Under the assumptions set out above, this term is distributed asymptotically as $N(0,\Omega)$ where Ω is as defined in A6. See Hansen (1982).[2]

(b) $(1/\sqrt{T})\Sigma X_t'\eta_{t,k}^j X_t(\alpha_c-\alpha_c) = (1/T)[\Sigma X_t'\eta_{t,k}^j X_t]\sqrt{T}(\hat{\alpha}_c-\alpha_c)$

The arguments set out above can be used to show that the first part of this expression converges in probability to zero. The second converges in distribution to a normal random variable. The product thus converges in probability to zero.

(c) $(1/\sqrt{T})\Sigma X_t'\eta_{t,k}^c X_t(\hat{\alpha}_j-\alpha_j)$

This term can also be shown to converge in probability to zero using an argument identical to that in (b).

(d) $(1/\sqrt{T})\Sigma X_t'X_t(\hat{\alpha}_c-\alpha_c)X_t(\hat{\alpha}_j-\alpha_j)$.

A typical element of this can be written as, $\sqrt{T}(\hat{\alpha}_{j,i}-\alpha_{j,i})$ $(\hat{\alpha}_{c,k}-\alpha_{c,k})$ $(1/T)\Sigma X_t'X_{it}X_{kt}$. The first part of this product converges in distribution to a normal random variable. The second converges in probability to zero, while the third converges in probability to a vector. Thus the product converges in probability to zero.

We then have $\sqrt{T}(\hat{\theta}_j - \theta_j)$ is distributed asymptotically as $N(0,M^{-1}\Omega M^{-1})$.

----

2 An alternative to assuming joint stationarity is to allow X_t and ω_t^j to be drawn from different distributions over time and to assume that high order moments exist and use the mixing process theorems of White (1985).

DATA APPENDIX

c_t, real consumption is real consumption spending on nondurables and services per capita. Source: Survey of Current Business.

$F^j_{t,k}$, the k-period forward rate is calculated from three month eurocurrency deposit rates. $F^j_{t,3} = S^j_t(1 + r^{US}_{t,3})/(1 + r^j_{t,3})$.

IP_t, the industrial production index for the United States. In regressions for cross rates relative to the U.K. pound and the Deutsche mark, U.K. and West German data industrial production are used. Source: International Financial Statistics, June 1986 tape and various subsequent issues.

p_t, the U.S. consumer price index is the CPI-U measure of consumer prices and uses a rental equivalence measure for housing costs. Source: Survey of Current Business. In regressions for cross rates relative to the U.K. pound and the Deutsche mark, U.K. and West German CPI data are used. Source: International Financial Statistics, June 1986 tape and various subsequent issues.

$r^{US}_{t,3}$, $r^j_{t,3}$, the three-month eurocurrency deposit rate at the end of the month. Source: Morgan Guaranty, World Financial Markets.

S_t, the spot exchange rate at the end of the month. Source: International Financial Statistics, June 1986 tape and various subsequent issues.

TOT_t, the U.S. terms of trade is calculated as the ratio of the unit value index for U.S. exports to the unit value index for U.S. imports. Source: International Financial Statistics, June 1986 tape and various subsequent issues.

Table 1: Ex Ante Nominal Speculative Returns*

A. Autocorrelation of Residuals from U.S. Dollar Regressions

Currency	Lag 1	2	3	4	5	6	7	8	9	10	11	12
UK	0.61	0.30	-0.02	0.04	0.03	0.02	0.00	0.01	0.12	0.12	0.19	0.07
WG	0.59	0.32	-0.05	-0.03	-0.03	-0.02	0.00	-0.03	0.05	0.05	0.11	0.05
CA	0.50	0.07	-0.11	0.07	0.08	-0.10	-0.18	-0.09	0.12	0.23	0.19	-0.11
SW	0.64	0.33	-0.02	-0.03	-0.03	-0.04	-0.01	-0.05	0.04	0.04	0.10	0.07
FR	0.58	0.38	0.04	0.08	0.04	0.00	-0.05	-0.12	-0.05	-0.07	0.03	0.02

B. Chi-Square Tests

	$\chi^2(10)$ Base Currency		
Currency	U.S. dollar	U.K. pound	Deutsche mark
U.K. pound	49.63 (.30E-06)		
Deutsche mark	47.25 (.83E-06)	22.60 (.12E-01)	
Canadian dollar	48.59 (.48E-06)	35.78 (.92E-04)	34.83 (.13E-03)
Swiss franc	50.15 (.24E-06)	36.75 (.63E-04)	30.36 (.75E-03)
French franc	29.60 (.99E-03)	26.21 (.35E-02)	26.91 (.27E-02)

* The dependent variable is the percent nominal return to a long forward position in each of the five currencies relative to the base currency. The right-hand-side variables are the forward premia of each currency (relative to the base currency) and the squared forward premia of each currency. Marginal significance levels are in parentheses below the χ^2 statistics. The sample period is 1974:1 to 1986:12.

Table 2: Ex Ante Real Speculative Returns*

A. Chi-Square Tests with Forward Premia and Forward Premia Squared

Currency	$\chi^2(10)$ Base Currency U.S. dollar	U.K. pound	Deutsche mark
U.K. pound	49.60 (.30E-06)		
Deutsche mark	47.31 (.83E-06)	22.25 (.14E-01)	
Canadian dollar	48.48 (.54E-06)	35.93 (.86E-04)	34.54 (.15E-03)
Swiss franc	50.56 (.24E-06)	36.64 (.65E-04)	30.16 (.81E-02)
French franc	29.51 (.10E-02)	25.94 (.38E-02)	26.91 (.27E-02)

B. Chi-Square Tests with Forward Premia, Consumption Growth, Inflation, Terms of Trade, and Industrial Production Growth

Currency	$\chi^2(11)$ Base Currency U.S. dollar	U.K. pound	Deutsche mark
U.K. pound	55.81 (0.00)		
Deutsche mark	69.17 (0.00)	21.65 (.10E-01)	
Canadian dollar	37.14 (.54E-04)	64.36 (0.00)	41.73 (.37E-05)
Swiss franc	69.14 (0.00)	46.97 (.41E-06)	44.60 (.11E-05)
French franc	106.24 (0.00)	23.06 (.61E-02)	22.06 (.87E-02)

* The dependent variables are the percent real return to a long forward position in each of the five currencies relative to the base currency The right-hand-side variables in A are the same as in Table 1. In B, the squared forward premia are replaced with the real U.S. consumption growth, base country CPI inflation lagged three and twelve months, the base country terms of trade, and base country industrial production growth lagged three months. Marginal significance levels are in parentheses below the χ^2 statistics. The sample period is 1974:1 to 1986:12.

Table 3: Gibbons-Ferson Tests*

A. Real U.S. Dollar Returns

Currencies	$\chi^2(20)$
DM, CA, SW	54.06 (0.57E-04)
DM, UK, SW	83.38 (0.00)
DM, UK, FR	88.01 (0.00)

* The tests are χ^2 tests of the proportionality of the coefficients in the regressions in column 1 of Table 2B. Marginal significance levels are in parentheses below the χ^2 statistics. The sample period is 1974:1 to 1986:12.

Table 4: Conditional Covariances*

A. Tests of Constant Conditional Covariance

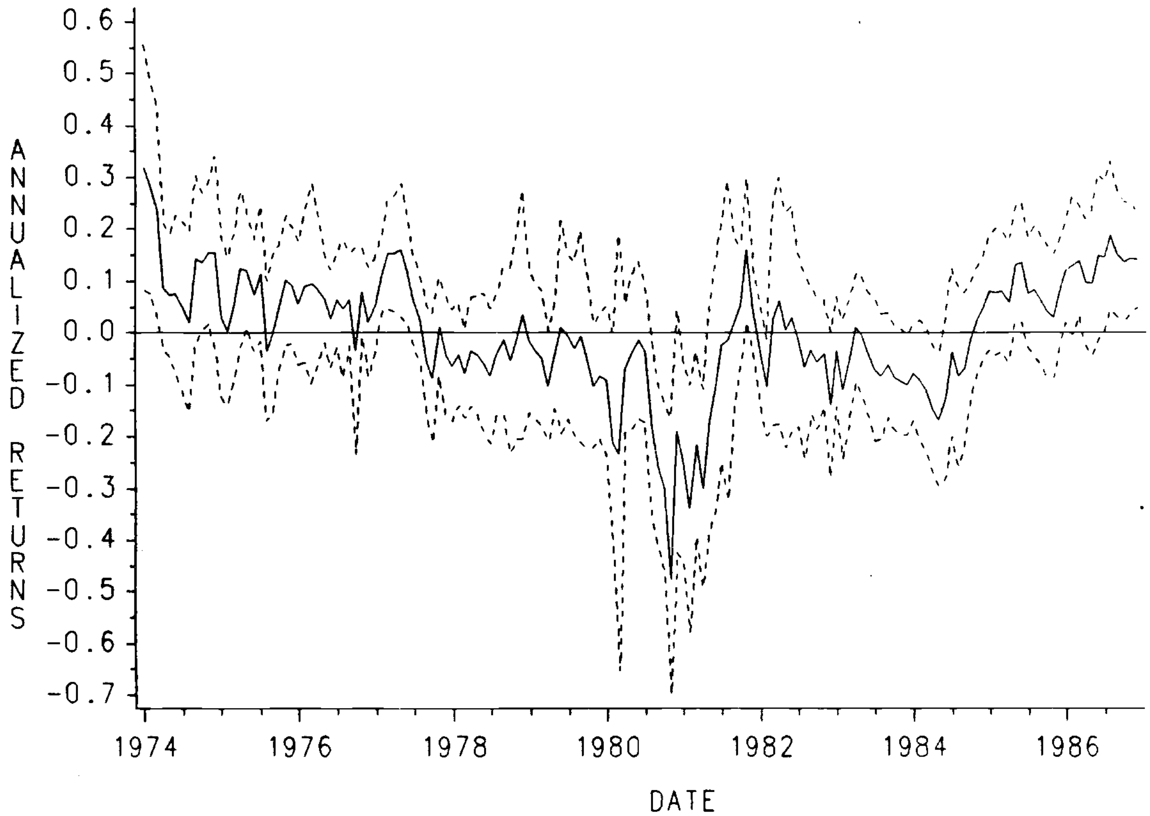
Currency	$\chi^2(10)$
U.K. pound	12.67 (0.24)
Deutsche mark	20.60 (.24E-01)
Canadian dollar	25.08 (.52E-02)
Swiss franc	14.04 (0.17)
French franc	47.32 (.83E-06)

B. Tests of the Proportionality of Conditional Covariances*

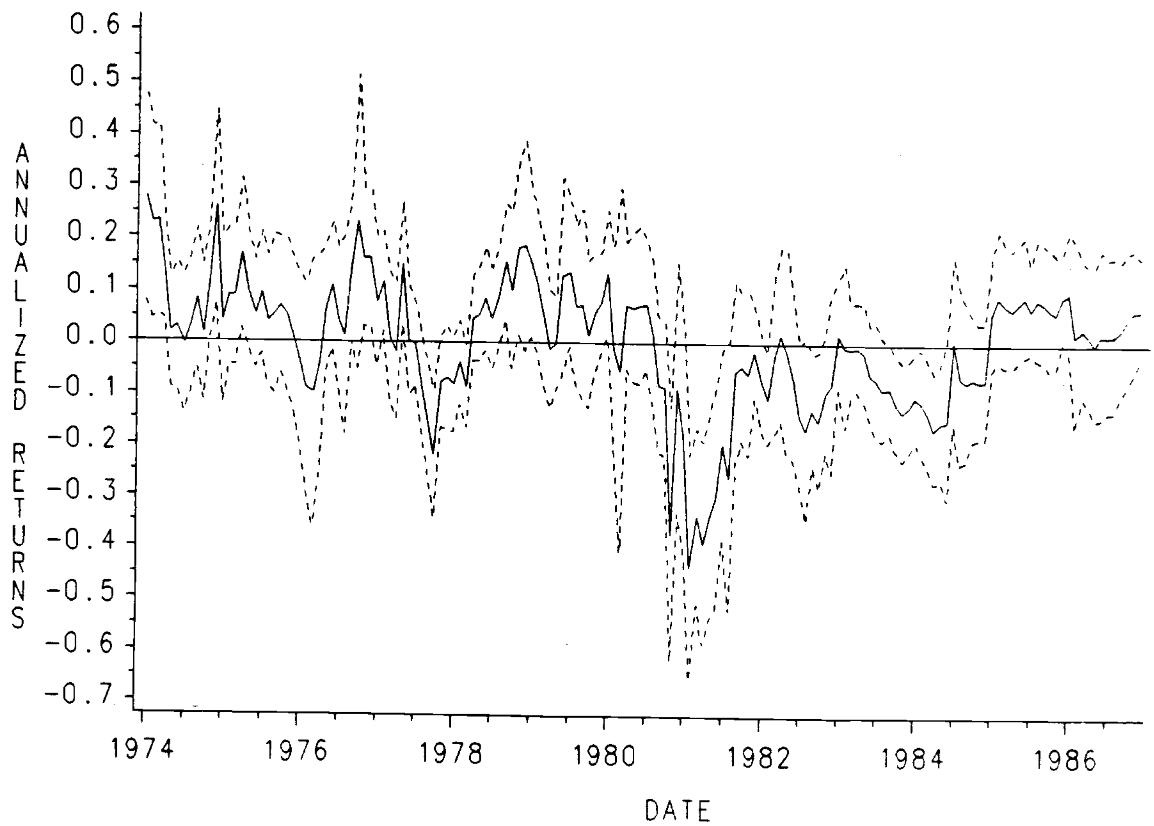
Currencies	$\chi^2(20)$
DM, CA, SW	21.36 (0.38)
DM, UK, SW	26.91 (0.14)
DM, UK, FR	21.83 (0.35)

* In A the depend variable is the product of the residuals from the regressions in column 1 of Table 2B and the residual of a regression of the percent growth in real per capita consumption on the same right-hand-side variables used in column 1 of Table 2B. The tests in A are tests of the hypothesis that all coefficients other than the constant term are zero. In B, the tests are χ^2 tests of the proportionality of the coefficients in the regressions in column 1 of Table 4A. Marginal significance levels are in parentheses below the χ^2 statistics. The sample period is 1974:1 to 1986:12.

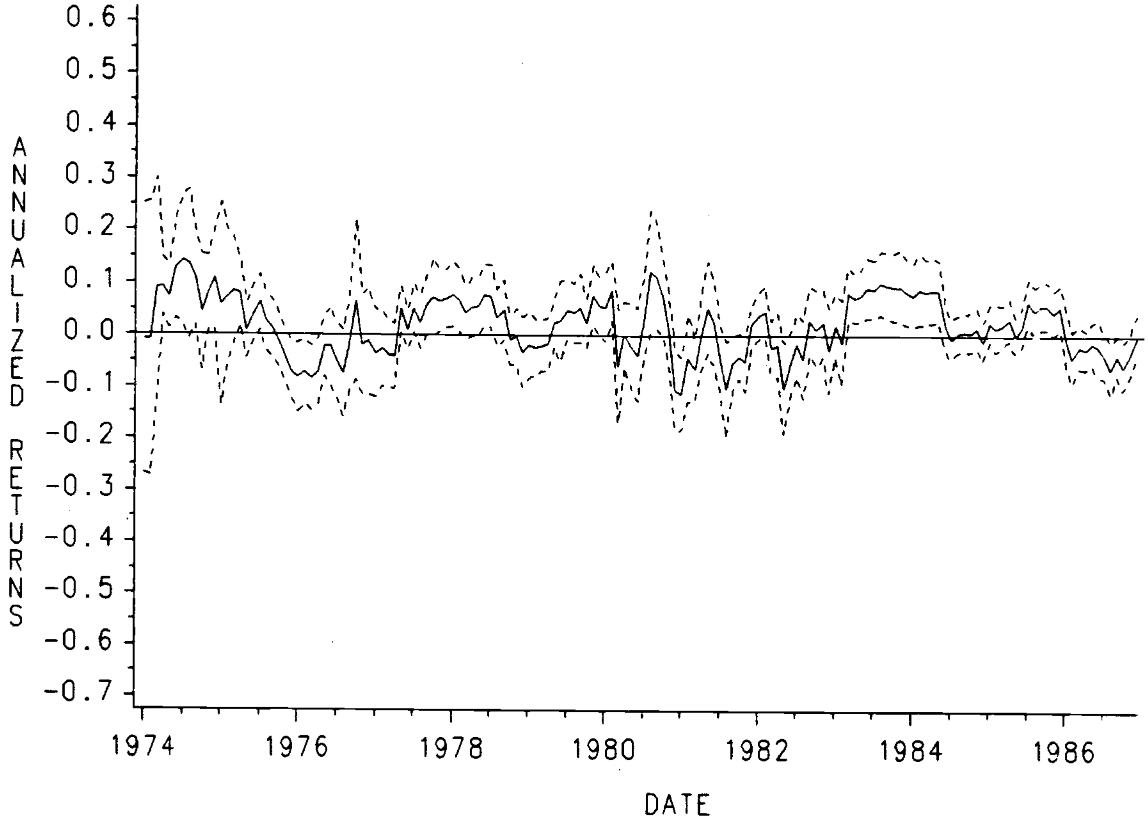
EX ANTE RETURNS: DEUTSCHE MARK



EX ANTE RETURNS: UK POUND



EX ANTE RETURNS: FRENCH FRANC/DM



CONDITIONAL COVARIANCE: DEUTSCHE MARK