THE JOURNAL OF FINANCE • VOL. XLV, NO. 4 • SEPTEMBER 1990

The Behavior of Eurocurrency Returns Across Different Holding Periods and Monetary Regimes

KAREN K. LEWIS*

ABSTRACT

Recent empirical studies of the risk premium across foreign exchange and other asset markets such as equity and longer term bonds have found conflicting evidence about the latent variable model restrictions of the consumption-based intertemporal capital asset pricing model. While studies using data for holding periods of one month or less generally reject the model, evidence using three-month holding periods indicates that the model cannot be rejected when including the returns on long relative to short deposit rates. This paper investigates the sources of differences in results using returns on foreign exchange and Eurocurrency deposits at three different maturities.

RECENT STUDIES TESTING THE restrictions implied by the intertemporal capital asset pricing model have found conflicting evidence. These restrictions, obtained from the first-order conditions of intertemporal utility maximization, imply that expected returns should move in constant proportion to each other, through a "single beta" latent variable.[1] This latent variable model has been estimated for various types of returns over different holding periods. Using a one-month forward contract horizon for a portfolio of foreign exchange excess returns, Hansen and Hodrick (1983) find that they cannot reject the restrictions of the model. However, using data beyond 1980, Hodrick and Srivastava (1984) reject the constraints and find the parameters sensitive to the sample period. Giovannini and Jorion (1987) use a one-week maturity horizon for foreign exchange and stock market returns to test and reject these restrictions. Thus, in general, tests of the latent variable restrictions have been rejected for returns with holding periods of one month or less.

By contrast, using a three-month holding period, Campbell and Clarida (1987) find that they cannot reject a constant single-beta model across excess returns on foreign exchange and on three-month relative to one-month Eurocurrency deposits.[2] Their study differs in other ways as well. For example, they include

* University of Pennsylvania, New York University, and National Bureau of Economic Research.
For useful comments, I am grateful to two anonymous referees, Bob Cumby, René Stulz, and seminar participants at the National Bureau of Economic Research Summer Institute, Columbia University, the Wharton School, University of Michigan, and New York University. I am also indebted to Alberto Giovannini and Philippe Jorion for providing the 7-day Eurocurrency deposit data, to Ross Levine and Steve Scott for assistance in using the DRI daily data tape, and to Naveen Seth for superb research assistance. Research support from the Glucksman Institute and the Salomon Brothers Center for Financial Studies is acknowledged with thanks. Any errors are mine alone.

[1] Hansen and Hodrick (1983) first tested the latent variable model restrictions that are implied by this model and are the focus of this paper. However, other studies such as Mark (1985) directly assume forms for the utility function to test restrictions implied by the same first-order conditions.

[2] Cumby (1988) also studies quarterly data using real excess returns to test a consumption-based asset pricing model, as will be discussed below.

the returns on three-month deposits in excess of one-month deposits together with the set of foreign exchange returns, while Cumby (1988) and Hodrick and Srivastava (1984) focus on foreign deposit returns alone, and Giovannini and Jorion (1987) study foreign deposit rates together with stock returns. The Campbell and Clarida study also differs by using data from 1976–1982, the end of the period of nonborrowed reserves targeting by the Federal Reserve.

Thus, the conflicting evidence leads naturally to the question: why are there differences in results? Do the results differ because of the maturity horizon, or because of the choice of the other market tested along with foreign exchange, or because of behavioral differences in excess returns across a change in "monetary regime"?

This paper investigates these questions using data for foreign exchange excess returns at three maturities: one week, one month, and three months. The analysis also incorporates data on the excess returns from holdings of long-term relative to short-term Eurocurrency deposits at two horizons: a one-month relative to a one-week deposit and a three-month relative to a one-month deposit. As a by-product, the returns from rolling over weekly deposits in excess of one-month deposits allow the first empirical investigation across foreign exchange and term structure returns at one-month maturities. The evidence below suggests that rejection of the model arises primarily from its sensitivity to the holding period. Further, even though the estimates of the model are unstable across the period of monetary targeting by the Federal Reserve, the tendency to reject the restrictions only over shorter holding periods remains after taking account of the shift during the period.

The plan of the paper is as follows. Section I studies the behavior of excess returns across one-week, one-month, and three-month horizons both of dollar-denominated assets relative to foreign-denominated assets and of longer-term deposits relative to shorter-term deposits. The analysis considers the sensitivity of the parameter stability to the period of nonborrowed reserves targeting by the Federal Reserve. Section II investigates the sensitivity of the results to the period of nonborrowed reserves targeting. Section III discusses different potential explanations for this tendency to reject over short holding periods. Concluding remarks follow.

I. The Behavior Across Holding Periods

A. The Basic Consumption-Based Intertemporal CAPM

The consumption-based asset pricing model uses the first-order conditions of the investor's intertemporal consumption decision to obtain a relationship between the conditional moments of asset returns and the intertemporal marginal rate of substitution in consumption. General equilibrium models within this framework include Stockman (1978), Breeden (1979), Stulz (1981), Grossman and Shiller (1982), and Lucas (1982). The restrictions follow from the Euler equation of intertemporal utility maximization of a representative investor with time-additive preferences. These first-order equations for any asset i and any k periods ahead are

$$U'(c_t) = \gamma^k E_t\{U'(c_{t+k})(1 + r^i_{t,k})\}, \tag{1}$$

where $U(c_t)$ is the period utility function of the representative investor, γ is the period discount rate in the utility function, E_t is the expectations operator with respect to information available at time t, and $r^i_{t,k}$ is the rate of return on any arbitrary asset i that matures in period $t + k$. Equation (1) must hold for any asset since the investor must be indifferent at the margin between consuming a unit of consumption today or saving the same amount in an asset with possible real payoffs $(1 + r^i_{t,k})$ at $t + k$.

Since this relationship holds for any asset i, the first-order condition can be rewritten in a convenient form by defining $r^c_{t,k}$ as the return on a benchmark asset that is perfectly conditionally correlated with the intertemporal marginal rate of substitution in consumption between t and $t + k$ and $r^{rf}_{t,k}$ as the risk-free rate. Using these definitions and equation (1) for both the benchmark asset and asset i, the first-order condition may be rewritten:[3]

$$E_t(r^i_{t,k} - r^{rf}_{t,k}) = \left(\frac{\operatorname{cov}_t(r^i_{t,k} - r^{rf}_{t,k},\, r^c_{t,k} - r^{rf}_{t,k})}{\operatorname{var}_t(r^c_{t,k} - r^{rf}_{t,k})}\right) E_t(r^c_{t,k} - r^{rf}_{t,k}). \tag{2}$$

For future reference, returns in excess of the conditionally risk-free rate will be defined as $er^i_{t,k} = r^i_{t,k} - r^{rf}_{t,k}$. Using this notation, it is clear that the expected return of any risky asset with the same holding period must satisfy the following conditional proportionality restriction:

$$E_t(er^i_{t,k}) = \beta^i_{t,k} E_t(er^c_{t,k}), \tag{3}$$

where

$$\beta^i_{t,k} = \operatorname{cov}_t(er^i_{t,k},\, er^c_{t,k})/\operatorname{var}_t(er^c_{t,k}).$$

That is, since $E_t(er^c_{t,k})$ is a common component to all asset returns, ex ante returns are conditionally proportional.

B. Testing the Restrictions

Although the expected return on the benchmark asset is unobserved, it depends upon a linear projection of the return on the current information set. Therefore, given the subset of variables in the agent's information set that is observed by the econometrician, x_t, the expected return may be written:

$$E_t(er^c_{t,k}) = \alpha'_k x_t + u_{t,k}, \tag{4}$$

where $u_{t,k}$ is the error at time t in measuring the expected return on the k-period benchmark portfolio. Combining equation (4) with equation (3) and using the ex post asset i returns together with rational expectations implies

$$er^i_{t,k} = \beta^i_{t,k}(\alpha'_k x_t + u_{t,k}) + \nu^i_{t+k} = \beta^i_{t,k}\alpha'_k x_t + \varepsilon^i_{t,k}, \tag{5}$$

[3] Hansen and Hodrick (1983) and Campbell (1987) discuss this relationship and its derivation.

where ν^i_{t+k} is the ex post forecast error for asset i and where $\varepsilon^i_{t,k}$ is the composite error of the measurement error and the forecast error; i.e., $\varepsilon^i_{t,k} = \beta^i_{t,k} u_{t,k} + \nu^i_{t+k}$.

The equations given in (5) imply restrictions in the movements across asset returns. In particular, these restrictions indicate that the coefficients in the regression of any set of n asset returns, r^i for $i = 1, \cdots, n$, on the information variables, x_t, should be proportional across equations.

Since the α_k and $\beta^i_{t,k}$ parameters are not separately identified in a linear regression, we must make an additional assumption concerning the behavior of the consumption betas to estimate the model. A typical assumption is that the ratio of consumption betas across returns is constant over time. To consider the usefulness of this assumption, notice that estimation of the system requires normalizing by an arbitrary reference asset since only $n - 1$ of the β^i are identified. Thus, choosing asset $i = 1$ as the arbitrary reference asset, the other $i = 2, \cdots, n$ assets may be written:

$$er^i_{t,k} = \left(\frac{\beta^i_{t,k}}{\beta^1_{t,k}}\right) \alpha'_k x_t + \varepsilon^i_{t,k}. \tag{6}$$

Then, under the assumption that the ratio of the betas is constant, the ex ante returns are proportional with the factor of proportionality equal to the ratio of consumption betas.

Therefore, testing these restrictions simply implies that, for the vector of k-period asset returns, $r_{t,k}$, projected onto the information variables, x_t, the components of the matrix of coefficients should be proportional across rows. Specifically, in the system of n equations, suppressing the subscript for a given holding period, k, we may rewrite (6) as

$$er^i_t = b'_i x_t + \varepsilon^i_t, \quad i = 1, \cdots, n, \tag{7}$$

such that $b_{ij} = (\beta^i/\beta^1)\alpha_j$ for each x component, $x_{j,t}$. The test described in Gibbons and Ferson (1985) implies the same restrictions since they substitute out the benchmark return in terms of a reference return. Arbitrarily choosing $er^1_{t,k}$ as the reference return and substituting this return in place of the benchmark return makes clear that the latent variable model based upon the benchmark return in equation (4) is observationally equivalent to one based upon the reference return where $E_t(er^1_{t,k}) = \alpha'_k x_t + u_{t,k}$.

C. Data Description and Construction

Since the first-order conditions hold for any asset return in excess of the risk-free rate, the restrictions in equation (7) should hold for any arbitrary set of risky returns held over the same period. To investigate the differences in the studies discussed above, the analysis below focuses on the returns of two types of risky strategies in excess of a risk-free rate for a given holding period: (a) the returns from holding foreign currency deposits and then converting the returns back into domestic currency at the prevailing spot rate at the time of maturity and (b) the returns from successively rolling over short deposits over the holding period. The first-order conditions in equation (1) imply that these risky returns with identical maturities should depend upon the same expected intertemporal rate of substitution in consumption.

The excess returns series were constructed following those found in other studies. The first type of excess returns, called "foreign exchange returns" below, are the returns on an open position of foreign currency deposits in excess of U.S. dollar deposits of the same maturity. Thus, the speculative returns on an open position in, for example, Deutschemark (DM) deposits have the basic form:

$$er_{k,t}^{DM} \equiv A_k((s_{t+k}^{DM} - s_t^{DM})/s_t^{DM}) + r_{k,t}^{DM} - r_{k,t}^{\$}, \tag{8}$$

where A_k is an annualization factor equal to 100 times the ratio of the number of days in the year to the number of days in the holding period, s_t^{DM} is the spot DM exchange rate at time t, and $r_{k,t}^{i}$ is the rate of return in currency i of a deposit for k periods. Hence, equation (8) is the ex post realized return on holding foreign exchange deposits in excess of a risk-free dollar deposit with the same maturity. To compare the different studies, we must calculate these returns for three different holding periods: that is, three months ($k = 3$), one month ($k = 1$), and one week ($k = W$).

The second type of excess return, called the "term structure return" below, is the realized profit on rolling over short deposits in excess of a longer maturity deposit. Campbell and Clarida (1987) consider the returns on rolling over one-month deposits in excess of three-month deposits. Since we will study returns with one-month holding periods below, we must also construct the returns from rolling over one-week deposits in excess of one-month deposits. Following the linearization in Campbell and Clarida (1987), these excess returns can be described as

$$er_{3,t}^{i} \equiv \left[(1/3) \sum_{j=0}^{2} r_{1,t+j}^{i}\right] - r_{3,t}^{i}, \quad \text{for } k = 3, \tag{9a}$$

$$er_{1,t}^{i} \equiv \left[(1/4) \sum_{j=0}^{3} r_{W,t+j}^{i}\right] - r_{1,t}^{i}, \quad \text{for } k = 1. \tag{9b}$$

Although these returns are not compounded, this simplification should have negligible effects upon the latent variable tests since the variances of the foreign exchange returns are several orders of magnitude greater than the variances of the returns on rolling over short-term deposits.

While empirical studies have constructed these returns using the same basic forms as given in equations (8) and (9), they differ in how they treat the timing of transactions. These differences compound the already difficult problem of properly aligning the available Eurocurrency deposit rate data, having fixed maturities of 7 days, 30 days, and 90 days, with the spot exchange rate data that carry a two-day settlement period. In particular, the empirical work below uses data for 7-day Eurocurrency rates from the *London Financial Times* available at the end of the week, provided by Alberto Giovannini and Philippe Jorion, and 30-day and 90-day Eurocurrency deposits and spot exchange rates from *Data Resources Incorporated* (DRI) available daily. The *Financial Times* data correspond to the close of the London market, or 12:00 noon (EST), while the DRI data correspond to 11:30 A.M. (EST) in New York. As a result of these differences, construction of the excess returns series requires a nontrivial tradeoff between following previous studies and accurately aligning the data within the constraints of data availability. In order to treat the series construction uniformly across holding periods, the excess returns were all based upon Friday observations of the raw series and were computed using the same method. The appendix describes this construction method as well as the estimation results using data series based on different construction methods.

Below we will investigate the behavior of these returns using as the instruments (i.e., x_t in equation (7)) the current interest differentials that match up with the left-hand-side variables in the set of ex post returns considered. That is, in the foreign exchange returns, the instruments are the differentials between the foreign deposit rate and the Eurodollar rate at that horizon. For example, a set of return equations including $er_{k,t}^{DM}$, the excess DM rate in equation (8), contain as instruments the interest differential between the Eurodollar and the EuroDM, i.e., $r_{k,t}^{DM} - r_{k,t}^{\$}$. By covered interest parity, these variables are approximately the forward premia. Both sets of variables have been used as instruments in a number of studies of the foreign exchange risk premium.[4]

For estimation involving the term structure returns, the instrument set includes the spread between long and short deposit rates. For example, estimation of the system in (7) with the returns on three-month Eurodollar rates in excess of rolling over one-month Eurodollar rates includes $r_{3,t}^{\$} - r_{1,t}^{\$}$, the spread between the three-month and the one-month Eurodollar rates, in the instrument list. Clearly, both sets of interest differentials are in the information sets of agents at time t and are therefore valid instruments.

D. *Econometric Issues*

Given the implications of the intertemporal capital asset pricing model in equation (7), we can test these restrictions for any arbitrary set of excess returns with the same holding period k by estimating the regressions of the returns r^i on the information variables x_t and then testing the restriction that the coefficients are proportional across equations. While the system of equations in (7) can be estimated by OLS, the variance-covariance matrix of estimates requires adjustment for two reasons. First, since the frequency of the data is weekly but some returns have holding periods longer than one week, i.e., $k = 1, 3$ in (8) and (9), the forecast errors in equation (7) are overlapping. Thus, for projections of returns with τ many weeks in the holding period onto current information, the residuals in equation (7) should contain a moving-average component of order $\tau - 1$ as Hansen and Hodrick (1980) have shown. A second estimation problem arises because these residuals appear to exhibit conditional heteroscedasticity as shown, for instance, by Cumby and Obstfeld (1984), Hodrick and Srivastava (1984), and Giovannini and Jorion (1987, 1989). Therefore, as reported below, the model was estimated by correcting the variance-covariance matrix for both heteroscedasticity and for the moving-average process. Using the overlapping

[4] See, for example, Cumby (1988), Hansen and Hodrick (1983), Hodrick and Srivastava (1984), and Giovannini and Jorion (1987).

forecast errors rule described above, the one-month returns were corrected for a MA(3) process while the three-month returns were corrected for a MA(12) process. Campbell and Clarida (1987) also correct the three-month returns for a MA(12) process.

The covariance matrix was estimated using the sample moments method described in Hansen (1982) when the estimated matrix was positive definite. When it was not positive definite, the covariance matrix was estimated with the method discussed in Newey and West (1987a). Following Campbell and Clarida (1987), the covariance matrix was estimated by averaging over 24 lags of the autocovariance matrices when using the Newey-West method. To check whether significant differences in the results might arise from differences in small sample properties of the covariance estimators, the latent variable model was also estimated using both estimators when the covariance matrix was positive definite. These estimates indicated that the latent variable model results are robust to using these two estimators (described in an appendix available upon request from the author).

To test the restrictions of the latent variable model, we could proceed by estimating the linear regressions given in equations (7) and then testing the proportionality of the coefficients across equations using a nonlinear Wald test. As discussed above, the latent variable model implies that the parameter coefficients should move in proportion to the ratios of each asset i's consumption beta to the reference asset's beta, i.e., that for a given holding period, $b_{ij} = (\beta^i/\beta^1)\alpha_j$. In general, however, we cannot directly test these conditions from the unconstrained model since the ratios of betas are not identified separately from the α.

On the other hand, we can obtain estimates of the ratio of betas by noting that the ratios of constants in the unconstrained projection equations to the constant in the reference return projection equation should equal the ratio of betas, i.e., $(b_{i0}/b_{10}) = (\beta^i_k/\beta^1_k)$. Using the ratios of constants in each equation to provide estimates of the ratio of betas, the latent variable model implies that $b_{ij} = (b_{i0}/b_{i1})b_{1j}$, for all $i \neq 1$, and $j \neq 0$. Using a nonlinear Wald test, this constraint can be tested by forming the gradient of the restrictions vector evaluated at the consistent parameter estimates as described in Harvey (1981). Clearly, the usefulness of this test for evaluating the latent variable restrictions depends upon how well the ratios of unconstrained constants in the projection equations measure the ratios of relative betas.

Alternatively, we could incorporate the nonlinear proportionality restrictions in estimating the system of equations (7) and directly obtain parameter estimates of the ratios of betas. Since these estimates of the beta ratios exploit proportionality information from all of the projection equation coefficients, they provide more efficient parameter estimates than the unrestricted constant coefficients. We may then directly test the constrained model's restrictions, as Hansen and Hodrick (1983) discuss. Specifically, a test of the latent variable model's restrictions is given by the number of sample observations times the criterion function of the restricted model evaluated at its minimized value. This statistic is distributed as chi-squared with degrees of freedom equal to the number of over-identifying restrictions. When the value of the criterion function is near its minimum, the over-identifying orthogonality conditions are close to zero and the restrictions of the model are not rejected. Below, we will use this statistic of the constrained model in addition to the Wald statistic on the unconstrained model to evaluate the latent variable model's restrictions.

E. Empirical Evidence

Since the data used here differ from data used in some of the other studies, preliminary investigation of the model began by comparing estimates of the model across the same time periods of other studies.[5] Overall, the results using the present data series appear consistent with those of other studies. Some description of the comparisons with other studies is provided in an appendix available from the author upon request.

To compare the results of the latent variable model across maturity horizons and also to test for the implications of including excess returns on long-relative-to-short-term deposits, the model was estimated both at two maturity horizons and for two sets of assets: foreign exchange returns alone, and foreign exchange together with term structure returns.[6] Tables I through III report the results of estimating the latent variable model in equation (7) for some of these assets over the period February 6, 1976 to May 19, 1986.

Tables I and II report unconstrained regression results for the asset returns and instrument sets that Campbell and Clarida (1987) studied at the three-month holding period. To investigate the sensitivity of the results to the holding periods, these regressions were estimated at the one-month holding period as well. Table I reports the results at two maturity horizons for a portfolio of excess returns on U.S. deposits relative to rolling over shorter term deposits and of the dollar relative to the Deutschemark and the British pound. As instrumental variables, the estimation used a constant, the forward premia for each currency, and the spread between the long and the short U.S. deposit rates. Table II gives the results of similar regressions for an expanded set of asset returns that includes the term structure returns for the EuroDM and Europound deposits, in addition to the returns used in Table I. Correspondingly, the instrument list includes the spread between the long relative to short rates for both the DM and pound deposits.

As the regression results indicate, the coefficients are surprisingly similar across holding periods. For example, in Table I, in the equations for the excess British pound foreign exchange returns over both holding periods, the coefficients on the forward premium of the dollar–DM rate and on the dollar–pound rate are about -3 and 2, respectively. Other similarities can be noted in the other equations as well. The evidence therefore suggests that returns and instruments follow a similar joint distribution at both holding periods.

Despite these similarities across holding periods in co-movements of returns as a function of the instruments, studies in the literature that test the latent variable model in equation (7) find the conflicting results discussed above. This

[5] For instance, Campbell and Clarida (1987) use Harris Bank data; Hansen and Hodrick (1983) and Hodrick and Srivastava (1984) use DRI data but with a different alignment in contract dates; Giovannini and Jorion (1987) use DRI data, with a similar data alignment as in this paper.

[6] These returns, as well as all of the models to be reported below, were estimated using the program described in Cumby and Huizinga (1988).

Table I

Unconstrained Regressions of Excess Returns from Rolling Over Eurodollar Deposits and from Open Foreign Exchange Positions upon Instrumental Variables

This table summarizes across holding periods of three months and one month the behavior of the ex ante predictable components of realized returns from rolling over short-term Eurodollar deposits in excess of longer-term Eurodollar deposits and from open positions of foreign Eurocurrency deposits in excess of Eurodollar deposits according to

$$r_{k,t} = b_0 + b_1 SUk_t + b_2 FGk_t + b_3 FBk_t + e_{k,t},$$

where the $r_{k,t}$ are either *RTUk*, the "term structure" returns of rolling over short Eurodollar deposits of k months maturity for $k = 1$, 3, or else *RFxk*, the "foreign exchange" returns on currency deposit x for currency $x = B$ (British pound), G (German DM), and for maturity $k = 1$, 3, and where *SUk* are the spreads between 3-month and 1-month Eurodollar rates for $k = 3$ and between 1-month and 1-week Eurodollar rates for $k = 1$, and where *Fxk* are the forward premia on a k-month contract on currency x. At each holding period, Test 1 is a nonlinear Wald test that $b_{ij} = (b_{i0}/b_{10})b_{1j}$, for $j \neq 0$, where b_{ij} is the coefficient b_j in row number i. Test 2 is the test of the latent variable model that the b_j's are proportional across equations for $er^i_{t,k} = \sum b_j x_{j,t}$, where the $x_{j,t}$'s are the instruments used in each panel. Test 3 is the Wald test that the b_j for 10/12/79 to 9/24/82 are the same as for the rest of the sample period. Estimation uses weekly frequency data sampled on Fridays over the period February 6, 1976 to May 19, 1986. The 7-day Eurocurrency deposit rates are from the *London Financial Times*, while all other data are from *Data Resources Incorporated.* Regressions are estimated by OLS correcting the variance-covariance matrix for overlapping forecast errors and conditional heteroscedasticity with the sample moments method described in Hansen (1982) or, if not positive-definite, with the method described in Newey and West (1987a).

A. Three-Month Holding Period ($k = 3$)

Excess Return ($er_{3,t}$)	b_0	b_1	b_2	b_3	R^2
1. *RTU*3	0.46	−0.68[a]	0.12	0.01	0.129
	(0.36)	(0.14)	(0.08)	(0.05)	
2. *RFG*3	−1.56	−0.78	−0.84	1.57	0.039
	(11.15)	(5.21)	(1.88)	(1.33)	
3. *RFB*3	−24.25[a]	3.25	−3.49[a]	2.47[b]	0.090
	(10.78)	(3.26)	(1.71)	(1.31)	
Test 1	6.30	Test 2	10.44	Test 3	89.16
$\chi^2(6)$	(.390)	$\chi^2(6)$	(.107)	$\chi^2(12)$	(<.000)

B. One-Month Holding Period ($k = 1$)

Excess Return ($er_{1,t}$)	b_0	b_1	b_2	b_3	R^2
1. *RTU*1	0.06	−0.79[a]	0.04	0.01	0.40
	(0.17)	(0.07)	(0.04)	(0.02)	
2. *RFG*1	1.83	3.88	0.37	1.44	0.02
	(11.68)	(6.00)	(2.42)	(1.31)	
3. *RFB*1	−22.80[a]	3.90	−2.74	2.30	0.03
	(9.33)	(5.04)	(1.91)	(1.17)	
Test 1	6.18	Test 2	18.74	Test 3	25.50
$\chi^2(6)$	(.403)	$\chi^2(6)$	(.005)	$\chi^2(12)$	(.013)

Standard errors are in parentheses under the parameter estimates. Marginal significance levels are in parentheses under Test 1, Test 2, and Test 3.

[a] Significantly different from zero at the 95% confidence level.

[b] Significantly different from zero at the 90% confidence level.

Table II

Unconstrained Regressions of Excess Returns from Rolling Over Eurocurrency Deposits and from Open Positions of Foreign Exchange upon Instrumental Variables

This table summarizes across holding periods of three months and one month the behavior of the ex ante predictable components of realized returns from rolling over short-term Eurocurrency deposits in excess of longer-term Eurocurrency deposits and from open positions of foreign Eurocurrency deposits in excess of Eurodollar deposits according to

$$er_{k,t} = b_0 + b_1SUk_t + b_2FGk_t + b_3FBk_t + b_4SGk_t + b_5SBk_t + e_{k,t},$$

where the $er_{k,t}$ are either $RTxk$, the "term structure" returns of rolling over short rates for holding periods of k months on deposits in currency $x = U$ (U.S.), B (British pound), G (German DM), and for maturity $k = 1, 3$, or else $RFxk$, the "foreign exchange" returns on currency deposit $x = B, G$ at maturity $k = 1, 3$, and where Sxk are the spreads between Eurocurrency rates of currency $x = U, B, G$, for 3-month and 1-month deposits for $k = 3$ and between 1-month and 1-week deposits for $k = 1$, and where Fxk are the forward premia for currency $x = B, G$, on a k-month contract, $k = 1, 3$. At each holding period, Test 1 is a nonlinear Wald test that $b_{ij} = (b_{i0}/b_{10})b_{1j}$, for $j \neq 0$, where b_{ij} is the coefficient b_j in row number i. Test 2 is the test of the latent variable model that the b_j's are proportional across equations for $er^i_{t,k} = \sum b_j x_{j,t}$, where the $x_{j,t}$'s are the instruments used in each panel. Test 3 is the Wald test that b_j for 10/12/79 to 9/24/82 are the same as the rest of the sample period. Estimation uses weekly frequency data sampled on Fridays over the period February 6, 1976 to May 19, 1986. The 7-day Eurocurrency deposit rates are from the *London Financial Times*, while all other data are from *Data Resources Incorporated*. Regressions are estimated by OLS correcting the variance-covariance matrix for overlapping forecast errors and conditional heteroscedasticity with the sample moments method described in Hansen (1982) or, if not positive-definite, with the method described in Newey and West (1987a).

Table II—*Continued*

A. Three-Month Holding Period ($k = 3$)

Excess Return ($er_{3,t}$)	b_0	b_1	b_2	b_3	b_4	b_5	R^2
1. *RTU*3	0.41	−0.64[a]	0.10	0.02	−0.56	0.07	0.135
	(0.38)	(0.19)	(0.09)	(0.04)	(0.44)	(0.23)	
2. *RFG*3	1.79	−4.27	0.52	1.72	17.16	11.46[a]	0.110
	(8.51)	(4.12)	(1.56)	(1.12)	(10.76)	(4.30)	
3. *RFB*3	−23.40[a]	2.05	−3.27[b]	2.76[a]	−1.98	6.86	0.110
	(9.69)	(3.33)	(1.73)	(1.11)	(9.41)	(4.76)	
4. *RTG*3	−0.48[a]	−0.17[a]	0.08[a]	0.04[a]	−0.20	−0.11	0.197
	(0.10)	(0.05)	(0.02)	(0.01)	(0.13)	(0.09)	
5. *RTB*3	0.33	−0.08	0.07	−0.04	0.02	−0.25	0.040
	(0.30)	(0.12)	(0.06)	(0.03)	(0.35)	(0.21)	
	Test 1	199.97	Test 2	17.36	Test 3	304.17	
	$\chi^2(20)$	(<.000)	$\chi^2(20)$	(.629)	$\chi^2(30)$	(<.000)	

B. One-Month Holding Period ($k = 1$)

Excess Return ($er_{1,t}$)	b_0	b_1	b_2	b_3	b_4	b_5	R^2
1. *RTU*1	0.04	−0.86[a]	0.04[b]	0.01	0.24[b]	0.02	0.417
	(0.10)	(0.07)	(0.02)	(0.01)	(0.13)	(0.04)	
2. *RFG*1	2.25	4.85	0.62	1.26	−3.56	3.95	0.026
	(9.07)	(5.92)	(1.90)	(1.18)	(5.07)	(2.75)	
3. *RFB*1	−22.17[a]	5.77	−2.65[a]	2.16	−6.11	0.64	0.035
	(9.47)	(5.10)	(1.77)	(1.32)	(5.26)	(2.95)	
4. *RTG*1	−0.35[a]	−0.10[a]	−0.05[a]	0.03[a]	−0.33[a]	0.02	0.243
	(0.10)	(0.04)	(0.02)	(0.01)	(0.08)	(0.03)	
5. *RTB*1	0.13	−0.00	0.04	−0.04[a]	−0.13	−0.40[a]	0.174
	(0.17)	(0.08)	(0.03)	(0.02)	(0.10)	(0.09)	
	Test 1	227.50	Test 2	37.37	Test 3	145.52	
	$\chi^2(20)$	(<.000)	$\chi^2(20)$	(.011)	$\chi^2(30)$	(<.000)	

Standard errors are in parentheses under the parameter estimates. Marginal significance levels are in parentheses under Test 1, Test, 2, and Test 3.

[a] Significantly different from zero at the 95% confidence level.

[b] Significantly different from zero at the 90% confidence level.

Table III

Unconstrained Regressions of Excess Returns from Open Foreign Exchange Positions upon Instrumental Variables

This table summarizes across holding periods of three months, one month, and one week (in Panels A, B, and C, respectively) the behavior of the ex ante predictable components of realized returns on open positions of foreign Eurocurrency deposits in excess of Eurodollar deposits according to

$$er_{k,t} = b_0 + b_1 FGk_t + b_2 FBk_t + b_3 FDk_t + b_4 FSk_t + e_{k,t},$$

where $er_{k,t}$ are $RFxk$, the excess foreign exchange returns for currencies $x = B$ (British pound), G (German DM), D (Dutch guilder), S (Swiss franc) over holding periods of $k = W$ (weekly), 1, 3, and where Fxk, is the forward premium for currency $x = B$, G, D, S at maturity $k = W$, 1, 3. At each holding period, Test 1 is a nonlinear Wald test that $b_{ij} = (b_{i0}/b_{10})b_{1j}$, for $j \neq 0$, where b_{ij} is the coefficient b_j in row number i. Test 2 is the test of the latent variable model that, at each holding period k, the b_j's are proportional across equations for $er^i_{t,k} = \sum b_j x_{j,t}$, where the $x_{j,t}$'s are the instruments used in each panel. Test 3 is the Wald test that b_j for 10/12/79 to 9/24/82 are the same as the rest of the sample period. Estimation uses weekly frequency data sampled on Fridays over the period February 6, 1976 to May 19, 1986. The 7-day Eurocurrency deposit rates are from the *London Financial Times*, while all other data are from *Data Resources Incorporated*. Regressions are estimated by OLS correcting the variance-covariance matrix for overlapping forecast errors and conditional heteroscedasticity with the sample moments method described in Hansen (1982) or, if not positive-definite, with the method described in Newey and West (1987a).

A. Three-Month Holding Period ($k = 3$)

Excess Return ($er_{3,t}$)	b_0	b_1	b_2	b_3	b_4	R^2
1. *RFG*3	8.41	−6.34[b]	0.21	3.21[a]	3.31	0.104
	(10.03)	(3.72)	(1.30)	(1.40)	(2.69)	
2. *RFB*3	−19.68[b]	−7.21[a]	2.72[a]	−0.03	3.11	0.119
	(11.85)	(2.71)	(1.48)	(2.08)	(2.12)	
3. *RFD*3	−3.72	−4.76	1.10	2.14	3.18	0.164
	(8.46)	(3.18)	(1.03)	(1.52)	(2.48)	
4. *RFS*3	−16.99[a]	−6.98[a]	3.54[a]	−0.10	4.14[b]	0.206
	(11.72)	(2.64)	(1.46)	(2.02)	(2.15)	
	Test 1	1739.7	Test 2	16.85	Test 3	109.51
	$\chi^2(12)$	(<.000)	$\chi^2(12)$	(.155)	$\chi^2(15)$	(<.000)

B. One-Month Holding Period ($k = 1$)

Excess Return ($er_{1,t}$)	b_0	b_1	b_2	b_3	b_4	R^2
1. *RFG*1	10.16	−3.54	0.17	2.55[b]	2.07	0.037
	(12.50)	(3.60)	(1.54)	(1.32)	(2.30)	
2. *RFB*1	−19.82[b]	−6.67[a]	3.04[a]	−1.22	3.37	0.050
	(11.88)	(3.00)	(1.45)	(1.97)	(2.22)	
3. *RFD*1	−6.65	−3.38	1.17	1.82	2.17	0.059
	(10.63)	(3.41)	(1.33)	(1.43)	(2.11)	
4. *RFS*1	−18.44	−7.01[a]	4.04[a]	−1.19	4.60[a]	0.095
	(11.83)	(3.07)	(1.45)	(1.94)	(2.31)	
	Test 1	33796.6	Test 2	21.82	Test 3	44.61
	$\chi^2(12)$	(<.000)	$\chi^2(12)$	(.040)	$\chi^2(20)$	(.001)

Table III—*Continued*

C. One-Week Holding Period (k = W)						
Excess Return ($er_{W,t}$)	b_0	b_1	b_2	b_3	b_4	R^2
1. *RFGW*	12.52	−0.53	−0.16	2.49[b]	0.58	0.011
	(10.21)	(2.70)	(1.30)	(1.47)	(0.92)	
2. *RFBW*	−19.64[a]	−4.06[b]	2.74[b]	−0.90	1.49	0.011
	(9.44)	(2.21)	(1.41)	(1.22)	(1.02)	
3. *RFDW*	0.56	−0.35	−0.34	3.52[a]	0.34	0.024
	(8.94)	(2.30)	(1.30)	(1.46)	(0.82)	
4. *RFSW*	−18.86[a]	4.05[b]	3.75[a]	−0.89	2.42[a]	0.028
	(9.35)	(2.18)	(1.40)	(1.21)	(1.01)	
	Test 1 $\chi^2(12)$	1568.4 (<.000)	Test 2 $\chi^2(12)$	55.39 (<.000)	Test 3 $\chi^2(20)$	26.44 (.152)

Standard errors are in parentheses under the parameter estimates. Marginal significance levels are in parentheses under Test 1, Test 2, and Test 3.

[a] Significantly different from zero at the 95% confidence level.

[b] Significantly different from zero at the 90% confidence level.

evidence naturally leads to the question: does the discrepancy in the results arise from including the term structure returns or does it arise from looking at a longer holding period? To address this question, the latent variable model restrictions were tested for the projection equations in (7) using both sets of returns for the one-month and three-month holding periods. With these tests, we can answer the question posed above by comparing the pattern in rejecting the latent variable model restrictions across holding periods and types of returns. If we do not reject the restrictions as the holding period returns become shorter than three months when we also include the term structure returns in estimation, then the discrepancy in results would seem to arise from the choice of returns. On the other hand, if we reject the restrictions over shorter holding periods even though we have included the term structure returns, then the conflicting evidence in the literature would appear to arise from differences in holding periods.

Tables I and II report the results of testing the latent variable model restrictions in the two ways described above. First, the Wald test that the ratios of unrestricted coefficients equal the ratio of constant coefficients is given as Test 1 in the tables, using the U.S. term structure returns as the reference return. For Table I, this test is not rejected at either holding period. However, for the larger set of returns in Table II, this restriction is rejected at all holding periods. Closer inspection of the coefficients suggests why these Wald tests differ. In Table I, only the British pound foreign exchange return contains a constant coefficient significantly different from zero. Thus, the restriction that the ratios of other coefficients are not significantly different from the ratios of these constants cannot be rejected. In Table II, on the other hand, the constant coefficient in the British term structure equation is also statistically significant. In this case, the other unrestricted projection coefficient estimates reject the hypothesis that they are statistically insignificant from the ratio of constant coefficients. Furthermore, this hypothesis is rejected at both holding periods. As discussed in Section I.*D* above, the usefulness of this Wald test as a test of the latent variable model depends strongly upon the ratio of unrestricted constant coefficients providing an accurate estimate of the relative consumption betas. Therefore, the restrictions were also tested by estimating the nonlinear latent variable model, thereby extracting the information about proportionality imbedded in the movement of the instrumental variables. Using this nonlinear model, we may form the test statistics of the over-identifying restrictions as described by Hansen and Hodrick (1983). These are reported as Test 2 in the tables.

These latent variable model tests of over-identifying restrictions (Test 2) in both Tables I and II follow the patterns noted elsewhere in the literature. At the three-month holding period, the results are consistent with those found by Campbell and Clarida (1987) for a shorter sample period. At standard significance levels, the chi-squared statistics do not reject the model at this horizon with marginal significance levels of 11% and 63% for Tables I and II, respectively. However, at the one-month holding period, the restrictions are strongly rejected with marginal significance levels of less than 2% in both cases. Furthermore, these restrictions were also tested for a number of different instrument lists that included (a) the spreads between long and short deposit rates on other currency deposits such as the Swiss franc, the Dutch guilder, and the French franc, (b) the squared forward premia, (c) lagged excess returns, and other combinations. In each case, the restrictions of the model tended to reject as the holding period shortened to one month relative to three months.

In summary, although the tests of the model's restrictions using the ratio of unrestricted constant coefficients as measures of the relative consumption betas appear sensitive to the instruments and returns, the tests of the over-identifying latent variable restrictions seem fairly robust to these same variables. The relative robustness of the latent variable model test seems reasonable since it incorporates the proportional movements latent in all of the information variables, thereby providing a more general test of the model. Using this general test, the restrictions of the model tend to be rejected as the holding period shortens.

The tests of over-identifying restrictions (Test 2) in both Tables I and II partially answer the question of why the literature has found conflicting evidence. Specifically, since the restrictions were rejected as the holding period shortened, they suggest that the discrepancy arises from differences in holding periods and not from including term structure returns in the return set. In order to more fully address this issue, these same restrictions will be tested below on a set of foreign exchange returns alone across holding periods.

F. Sensitivity to the Fed Operating Regime

A number of authors, such as Hodrick and Srivastava (1984) and Giovannini and Jorion (1987), have noted that ex ante returns in foreign exchange appear to behave differently before and after the change in Federal Reserve operating procedures in 1979.[7] Since the behavior of returns on longer term U.S. dollar

[7] Hodrick and Srivastava (1984) find evidence of a change in the behavior of the British pound and the French franc, and for the German mark if the break is assumed to occur later, in December 1980.

deposits is also likely to have been sensitive to this time period, the stability of the parameters in the unconstrained model was tested using a joint equation analogue to the test described in Hodrick and Srivastava (1984). Defining the stacked system version of the variance-covariance matrix given in equation (7) as Λ and the stacked parameter vector as **b**, these variables are estimated over the two subperiods, subscripted as 1 and 2 by period, to provide estimates of $\hat{\mathbf{b}}_1$, $\hat{\mathbf{b}}_2$, and $\hat{\Lambda}_1$, $\hat{\Lambda}_2$. Then the Wald statistic,

$$(\hat{\mathbf{b}}_1 - \hat{\mathbf{b}}_2)'\{(\hat{\Lambda}_1/T_1) + (\hat{\Lambda}_2/T_2)\}^{-1}(\hat{\mathbf{b}}_1 - \hat{\mathbf{b}}_2),$$

is asymptotically chi-squared with degrees of freedom equal to the number of parameters in **b**. These statistics are calculated for the subperiods between October 1979 and September 1982 and the rest of the sample.[8] They are reported as Test 3 in the tables. For the unconstrained parameter coefficients at the one-month and the three-month horizons and for both sets of assets, the hypothesis is strongly rejected with marginal significance levels less than 2%. Therefore, in Section II below, we investigate the behavior of these returns across the different monetary regime periods.

G. *Empirical Evidence Excluding Term Structure Returns*

The results reported above differ from those of other studies that have reported rejections of the latent variable model restrictions using foreign exchange returns without term structure returns. For example, Hodrick and Srivastava (1984) and Cumby (1988) use foreign exchange returns alone, while Giovannini and Jorion (1987) report the latent variable model results both for foreign exchange returns alone and for stock returns joint with foreign exchange returns. Thus, these results in conjunction with those found in Tables I and II suggest the following question: does the tendency to reject at short horizons arise only from including term structure returns? That is, if the tendency to reject at short horizons found in Tables I and II depends *only* upon including the returns on long relative to short deposit rates, then rejecting the restrictions cannot be related solely to the holding period. To answer this question, the latent variable model was estimated across three holding periods using foreign exchange returns alone.

Table III describes the results of testing the latent variable model with foreign exchange excess returns alone for three maturity horizons: three months, one month, and one week. The table reports the coefficient estimates of the unconstrained regressions for four currencies: the Deutschemark, the British pound, the Dutch guilder, and the Swiss franc. These regressions use as instruments the differential between the deposit rates for each of the currencies relative to the Eurodollar rate. Below each set of regressions, the table also reports the Wald

[8] Although the table reports these Wald tests, Newey and West (1987b) show that analogues of the Wald, likelihood ratio, Lagrange multiplier, and minimum chi-square test statistics for GMM estimation are mutually asymptotically equivalent. In addition to the Wald statistics reported in the text, a number of joint equation and single equation tests were also conducted to test the constancy of the parameters before and after this period. Joint equation estimates that included the long-term deposit excess returns generally could not reject constancy of the parameters before and after the interval of monetary targeting. However, the hypothesis was rejected for the foreign exchange excess returns against some of the individual currencies.

test statistic that the ratios of unconstrained coefficients equal the ratios of unconstrained constant coefficients (Test 1), the test of over-identifying restrictions of the latent variable model (Test 2), and of testing that the projection coefficients are stable over subperiods (Test 3).

A comparison of the three-month and one-month horizons in Table III shows that the basic relationships across the maturity horizon noted above continue to hold. With a few exceptions, the parameter coefficients are strikingly similar across maturity horizon. Again, as Test 1 shows, we very strongly reject the hypothesis that the coefficients equal the product of the counterpart in the first equation and the ratio of constant coefficients. Strikingly, for Test 2 the marginal significance level of the test of the constrained latent variable model declines as the maturity decreases, from 16% at the three-month horizon, to 4% at the one-month horizon, to less than 0.1% at the one-week horizon. Finally, for both the one-month and three-month maturities, the test of stability of the coefficients is very strongly rejected, although not for the one-week maturity returns.

To check the sensitivity of these results to the currencies being estimated, the latent variable model was also estimated including the French franc and the Japanese yen returns with the same returns as in Table III over the period of available data from October 1979 to May 1986 (not shown). In this case, the difference between the horizons becomes even more dramatic. Test 2 at the one-month horizon has a marginal significance level of 0.078 ($41.58 \sim \chi^2(30)$), while the statistic at the three-month horizon has a marginal significance level of 0.99 ($14.48 \sim \chi^2(30)$). However, at the one-week horizon, the model was soundly rejected at the 0.3% marginal significance level. Thus, the tendency to reject the single-beta latent variable model again arises from the difference in the holding period, not from excluding the term structure returns.

The results found in Table III differ somewhat from those found in Cumby (1988), who rejected the hypothesis that the foreign exchange returns over a three-month holding period follow a single-beta latent variable model. His results differ from the present investigation in two basic ways. First, he samples the returns monthly. Second, he uses an estimate of the variance-covariance matrix of parameters that requires estimating and then inverting a vector autoregression of the orthogonality conditions. An appendix available from the author shows that the difference in results between those obtained here and those found in Cumby (1988) arises from the method of calculating the variance-covariance matrix and not from using returns sampled monthly.

H. Summary of the Behavior Across the Holding Period

There are two basic findings in Tables I through III. First, using the full sample that includes the period of nonborrowed reserves targeting by the Federal Reserve, the latent variable model tends to be rejected only at the one-month, but not the three-month, holding period. This finding is not sensitive to including term structure returns in the set of asset returns investigated. The finding also appears to be robust to a number of different sets of instruments. By itself, such a finding suggests that, over relatively long holding periods, excess returns generally move together in constant proportion. Furthermore, it suggests that the restrictions implied by the intertemporal consumption-based asset pricing model are not rejected for longer holding periods, even though they are typically rejected over short holding periods. Therefore, such a finding is quite important to understanding the basic co-movements of asset returns in the economy.

However, the second basic finding based upon tests of parameter stability indicates that the joint behavior of returns and the variables used as instruments changed dramatically during the period of Federal Reserve operating procedures. Therefore, the next section investigates further the behavior of excess returns across this "monetary regime." Interestingly, even after taking account of the general shift in the distribution of returns, the latent variable model tends to reject only for short holding periods, as will be discussed below.

II. The Behavior of Returns Across Different Monetary Regimes

A striking feature of the results discussed in the previous section is the evidence of parameter instability around the periods of changes in operating procedures by the Federal Reserve. If the joint distribution of returns and other variables used as instruments shifted when the Federal Reserve changed its operating procedure, then the parameters in the projection equations given in equations (4) and (7) also shifted. However, if the change in the "monetary regime" was largely unanticipated, the latent variable model would continue to hold within each subperiod.[9] Therefore, below we estimate the latent variable model for each subperiod in order to account for this shift in monetary regime. Before proceeding, however, we should note that dividing the sample in this way substantially reduces the number of observations used in estimation. In particular, for the set of five asset return equations estimated in Table II, the relatively large number of orthogonality conditions is likely to over-parameterize estimation for the system over subperiods. Nevertheless, estimating the latent variable model over these periods may suggest how the relationships between returns changed with the shift in Fed policy.

Panel A of Table IV reports the latent variable parameter estimates and the test of over-identifying restrictions for the system of returns described in Table I, Panel A: i.e., the U.S. three-month interest rate relative to rolling over U.S. one-month interest rates, to an open Deutschemark position, and to an open British pound position. As before, the U.S. term structure excess return is arbitrarily chosen as the reference return with its beta normalized to unity. Thus, the coefficients reported in the other equations are the ratios of the betas for each currency over that of the U.S. term structure return, as described in equation (7). This normalization has the advantage of allowing comparisons between the signs of the covariances of each excess return with the benchmark portfolio. However, as the estimates of the relative betas show, this normalization has the

[9] On the other hand, Lewis (1991) investigates whether, during the period of nonborrowed reserves targeting by the Federal Reserve from 1979 to 1982, the market's belief that the Fed would return to interest rate targeting may have induced a "peso problem" in the excess returns on long relative to short rates. In related work, Hamilton (1988) finds the probability of a shift in regimes before and after the period was very small.

effect of making the estimates very large in absolute value since the variances of the foreign exchange returns are large relative to the variance of the term structure returns. Of course, choosing a different normalizing equation reduces substantially the relative beta estimates.

As the estimates in Panel A of Table IV indicate, the relationships between the relative betas differ over the time periods. Although the latent variable test indicates that the model cannot be rejected, the constrained parameters are somewhat imprecisely estimated. However, all of the constrained parameters are significantly different from zero for the three-month holding period during the nonborrowed reserves targeting period of the Federal Reserve.

The lower panel of Table IV reports the results for the latent variable model at the one-month horizon for these same assets and subperiods. As these results show, the restrictions of the model are rejected over each period except for the first subperiod.[10] Also, in contrast to the results for the three-month horizon, the ratios of betas are significantly different from zero during the third period. Thus, the overall results in Table IV confirm the results obtained in Section I above. As in earlier studies, the latent variable model appears to be rejected for one-month holding periods, but not when estimating the model only for the period before 1979. On the other hand, the restrictions are not rejected for any subsample at the three-month holding period.

Table V reports the constrained parameter results and the tests of the latent variable model at both holding periods for the set of returns in Table II. In general, the latent variable parameter estimates, α_i, tend to be more precisely estimated. At the three-month holding period, the restrictions of the latent variable model fail to be rejected as above. However, for this set of returns and instruments, the restrictions are also not rejected for the one-month returns. Since the estimation in Table V uses a larger subset of the returns in Table IV, this result appears somewhat unsatisfactory. If the U.S. term structure returns, Deutschemark foreign exchange returns, and British pound foreign exchange returns are not proportional over one-month holding periods as the evidence in Table IV indicates, then they cannot be jointly proportional when the term structure returns on the British pound and German DM Eurocurrency deposit rates are included as in Table V. Therefore, it seems likely that the lack of rejection arises from lack of power in estimation. The system may be overparameterized since the increased number of equations in Table V raises the number of orthogonality conditions from 12 to 30.

III. Interpreting the Behavior of Returns

As the results above indicate, returns on open positions in foreign exchange relative to Eurodollar deposits and on longer term deposits relative to rolling over short-term deposits tend to move together over longer holding periods, but not over short holding periods. This basic result appears to remain when allowing for the shift in the behavior of returns during the period of nonborrowed reserves targeting by the Federal Reserve from 1979 to 1982. Several potential explanations of this finding are discussed below.

[10] This result is consistent with the finding in Hansen and Hodrick (1983) using data from a similar time period.

Table IV

Parameter Estimates and Test of Restrictions Across Subsamples and Holding Periods of the Latent Variable Model for Excess Returns from Rolling Over Eurodollar Deposits and Open Foreign Exchange Positions

This table reports the parameter estimates of the restricted latent variable model for the three-equation set of excess returns from rolling over short-term Eurodollar deposits and open positions of EuroDM and Europound deposits in excess of Eurodollar deposits as defined below:

$$RTUk_t = \alpha_0 + \alpha_1 SUk_t + \alpha_2 FGk_t + \alpha_3 FBk_t + \varepsilon^1_{t,k},$$

$$RFGk_t = [\beta(RFGk)/\beta(RTUk)][\alpha_0 + \alpha_1 SUk_t + \alpha_2 FGk_t + \alpha_3 FBk_t] + \varepsilon^2_{t,k},$$

$$RFBk_t = [\beta(RFBk)/\beta(RTUk)][\alpha_0 + \alpha_1 SUk_t + \alpha_2 FGk_t + \alpha_3 FBk_t] + \varepsilon^3_{t,k},$$

where $RTUk$ are the "term structure" returns of rolling over short Eurodollar deposits for holding periods of k months maturity for $k = 1, 3$, and where $RFxk$ are the "foreign exchange" returns on currency deposit $x = B$ (British), G (German), at maturity $k = 1, 3$, and where SUk are the spreads between 3-month and 1-month Eurodollar rates for $k = 3$ and between 1-month and 1-week Eurodollar rates for $k = 1$, and where Fxk are the forward premia on k-month contracts, $k = 1, 3$, for currency $x = B, G$. For maturities of three months ($k = 3$) and one month ($k = 1$), Panels A and B, respectively, report the constrained parameter estimates on the reference asset, α_i, $i = 0, \cdots, 3$, and the ratio of the beta for each return over the reference beta, $\beta(er^i_k)/\beta(RTUk)$, using the return on rolling over Eurodollar rates as the reference asset. Each panel reports the estimates obtained over each subsample. The 7-day Eurocurrency deposit rates are from the *London Financial Times*, while all other data are from *Data Resources Incorporated.* Regressions are estimated by OLS correcting the variance-covariance matrix for overlapping forecast errors and conditional heteroscedasticity with the sample moments method described in Hansen (1982) or, if not positive-definite, with the method described in Newey and West (1987a). The last column reports the chi-squared test of the over-identifying restrictions described in Hansen (1982), a test of the latent variable model.

A. Three-Month Holding Period ($k = 3$)

Period	α_0	α_1	α_2	α_3	$\frac{\beta(RFG3)}{\beta(RTU3)}$	$\frac{\beta(RFB3)}{\beta(RTU3)}$	Latent Var. Test $\chi^2(6)$
Feb. 1976–Oct. 1979	0.14[a]	−0.04	−0.00	−0.01	119.36[a]	26.01	5.77
	(0.06)	(0.02)	(0.01)	(0.01)	(42.45)	(26.51)	(.449)
Oct. 1979–Sept. 1982	0.81[a]	−0.12[a]	0.12[a]	−0.05[a]	33.88[a]	−56.46[a]	5.38
	(0.21)	(0.04)	(0.03)	(0.02)	(9.03)	(13.06)	(.469)
Oct. 1982–May 1986	−0.00	0.04	0.00	0.01	685.60	686.02	6.28
	(0.02)	(0.07)	(0.01)	(0.02)	(1368.81)	(1318.29)	(.393)

B. One-Month Holding Period ($k = 1$)

Period	α_0	α_1	α_2	α_3	$\frac{\beta(RFG1)}{\beta(RTU1)}$	$\frac{\beta(RFB1)}{\beta(RTU1)}$	Latent Var. Test $\chi^2(6)$
Feb. 1976–Oct. 1979	−0.09	0.01	−0.01	−0.57[a]	13.99[a]	1.71	10.45
	(0.20)	(0.02)	(0.01)	(0.24)	(4.96)	(9.54)	(.107)
Oct. 1979–Sept. 1982	−0.53[a]	−0.04	0.07[b]	0.10	5.51	41.48[b]	12.23
	(0.25)	(0.04)	(0.04)	(0.14)	(26.04)	(24.01)	(.057)
Oct. 1982–May 1986	−0.17	0.00	0.04[a]	0.20[a]	113.93[b]	145.89[a]	15.62
	(0.17)	(0.04)	(0.02)	(0.09)	(58.71)	(56.72)	(.016)

Standard errors in parentheses under parameter estimates. Marginal significance levels in parentheses under "Latent Variable Test."

[a] Significantly different from zero at the 95% confidence level.

[b] Significantly different from zero at the 90% confidence level.

Table V

Parameter Estimates and Test of Restrictions Across Subsamples and Holding Periods of the Latent Variable Model for Excess Returns from Rolling Over Eurocurrency Deposits and Open Foreign Exchange Positions

This table reports the parameter estimates of the restricted latent variable model for the five-equation set of excess returns from rolling over short-term Eurocurrency deposits in U.S. dollars, German DM, and British pounds and from open positions of EuroDM and Europound deposits in excess of Eurodollar deposits as defined below:

$$RTUk_t = \alpha_0 + \alpha_1 SUk_t + \alpha_2 FGk_t + \alpha_3 FBk_t + \alpha_4 SGk_t + \alpha_5 SBk_t + \varepsilon^1_{t,k},$$

$$RFGk_t = [\beta(RFGk)/\beta(RTUk)][\alpha_0 + \alpha_1 SUk_t + \alpha_2 FGk_t + \alpha_3 FBk_t + \alpha_4 SGk_t + \alpha_5 SBk_t] + \varepsilon^2_{t,k},$$

$$RFBk_t = [\beta(RFBk)/\beta(RTUk)][\alpha_0 + \alpha_1 SUk_t + \alpha_2 FGk_t + \alpha_3 FBk_t + \alpha_4 SGk_t + \alpha_5 SBk_t] + \varepsilon^3_{t,k},$$

$$RTGk_t = [\beta(RTGk)/\beta(RTUk)][\alpha_0 + \alpha_1 SUk_t + \alpha_2 FGk_t + \alpha_3 FBk_t + \alpha_4 SGk_t + \alpha_5 SBk_t] + \varepsilon^4_{t,k},$$

$$RTBk_t = [\beta(RTBk)/\beta(RTUk)][\alpha_0 + \alpha_1 SUk_t + \alpha_2 FGk_t + \alpha_3 FBk_t + \alpha_4 SGk_t + \alpha_5 SBk_t] + \varepsilon^5_{t,k},$$

where the $RTxk$ are the "term structure" returns of rolling over short rates for holding periods of k months on deposits in currency $x = U$ (U.S.), B (British pound), G (German DM), and for maturity $k = 1, 3$, and where $RFxk$ are the "foreign exchange" returns on currency deposit $x = B, G$ at maturity $k = 1, 3$, and where Sxk are the spreads between Eurocurrency rates of currency $x = U, B, G$, for 3-month and 1-month deposits for $k = 3$ and between 1-month and 1-week deposits for $k = 1$, and where Fxk are the forward premia for currency $x = B, G$, on a k-month contract, $k = 1, 3$. For maturities of three months ($k = 3$) and one month ($k = 1$), Panels A and B, respectively, report the constrained parameter estimates on the references asset, α_i, $i = 0, \cdots, 3$, and the ratio of the beta for each return over the reference beta, $\beta(er^i_k)/\beta(RTUk)$, using the return on rolling over Eurodollar rates as the reference asset. Each panel reports the estimates obtained over each subsample. The 7-day Eurocurrency deposit rates are from the *London Financial Times*, while all other data are from *Data Resources Incorporated*. Regressions are estimated by OLS correcting the variance-covariance matrix for overlapping forecast errors and conditional heteroscedasticity with the sample moments method described in Hansen (1982) or, if not positive-definite, with the method described in Newey and West (1987a). The last column reports the chi-squared test of the over-identifying restrictions described in Hansen (1982), a test of the latent variable model.

Table V—*Continued*

A. Three-Month Holding Period ($k = 3$)

Period	α_0	α_1	α_2	α_3	α_4	α_5	$\beta(RFG3)/\beta(RTU3)$	$\beta(RFB3)/\beta(RTU3)$	$\beta(RTG3)/\beta(RTU3)$	$\beta(RTB3)/\beta(RTU3)$	Lat. Var. Test $\chi^2(12)$
Feb. 1976–Oct. 1979	−0.00	0.06[b]	0.01[b]	0.02[a]	0.02	0.02	410.9[b]	425.1[a]	0.08	0.54	7.71
	(0.01)	(0.00)	(0.00)	(0.01)	(0.02)	(0.01)	(218.5)	(211.9)	(0.30)	(0.47)	(.994)
Oct. 1979–Sept. 1982	0.47[a]	−0.02[b]	0.05[a]	−0.05[a]	−0.47[a]	−0.28[a]	−43.9[a]	−50.56[a]	−0.11[b]	0.22[a]	6.71
	(0.10)	(0.01)	(0.01)	(0.01)	(0.06)	(0.04)	(5.07)	(6.36)	(0.06)	(0.08)	(.998)
Oct. 1982–May 1986	0.17[a]	−0.03[a]	0.00	−0.01[a]	0.08[a]	−0.01[a]	109.1[a]	9.78	−0.34[a]	0.31	8.19
	(0.02)	(0.01)	(0.00)	(0.00)	(0.03)	(0.01)	(14.49)	(11.12)	(0.14)	(0.37)	(.991)

B. One-Month Holding Period ($k = 1$)

Period	α_0	α_1	α_2	α_3	α_4	α_5	$\beta(RFG1)/\beta(RTU1)$	$\beta(RFB1)/\beta(RTU1)$	$\beta(RTG1)/\beta(RTU1)$	$\beta(RTB1)/\beta(RTU1)$	Lat. Var. Test $\chi^2(12)$
Feb. 1976–Oct. 1979	−0.59[a]	−0.02[a]	0.02[a]	0.06	−0.18[a]	0.00	−30.33[a]	4.91	0.35[a]	0.79[a]	8.12
	(0.06)	(0.01)	(0.04)	(0.06)	(0.02)	(0.01)	(2.23)	(4.67)	(0.02)	(0.08)	(.991)
Oct. 1979–Sept. 1982	−0.14[b]	−0.01[b]	0.01[b]	0.10[b]	−0.10[b]	0.04	109.1	143.5	3.28[b]	1.53	6.62
	(0.08)	(0.01)	(0.00)	(0.06)	(0.06)	(0.03)	(67.8)	(88.0)	(1.94)	(0.95)	(.998)
Oct. 1982–May 1986	−0.06[a]	−0.00	0.02[a]	0.08[a]	0.04[a]	0.01	356.3[a]	423.1[a]	0.56[a]	−1.00[b]	7.92
	(0.03)	(0.00)	(0.01)	(0.03)	(0.01)	(0.01)	(144.0)	(16.17)	(0.15)	(0.53)	(.992)

Standard errors are in parentheses under the parameter estimates. Marginal significance levels are in parentheses under "Latent Variable Test."

[a] Significantly different from zero at the 95% confidence level.

[b] Significantly different from zero at the 90% confidence level.

In the context of the constant-ratios-of-betas form of the intertemporal capital asset pricing model described in Section I, the most immediate interpretation of these findings is that the restrictions of this model tend to hold for longer but not short holding periods. Within the framework of this model, there may be various explanations for the rejection over short holding periods. First, since empirical evidence suggests that the variances of different types of excess returns vary through time, conditional covariances between these returns and consumption and, therefore, "consumption betas," would seem to move over time as well.[11] If there are asset-specific movements in consumption betas, they would lead to rejection of the restrictions of the model implied by constant relative betas, as given in equation (7). Thus, the observed rejection of the latent variable model over short holding periods, but not over long ones, would be consistent with consumption betas that exhibit idiosyncratic variations over short holding periods that are dominated by common movements with consumption betas in other returns over longer holding periods.

Second, the evidence may be consistent with the intertemporal capital asset pricing model if the timing of information to the market looks more continuous over longer horizons. For example, suppose that information concerning the economy is only observed at relatively infrequent intervals, such as every quarter. Then, the forecast errors pertaining to quarterly returns will be white noise. However, for short holding periods such as a week or a month, expectations concerning announcements and other information that will be discovered beyond the holding period may induce persistent behavior in excess returns. In this case, the ex ante returns over short holding periods will depend in part upon serially correlated (nonoverlapping) forecast errors arising from expectations of a future discrete event.[12]

In addition to these potential explanations of the empirical evidence that depend upon the ICAPM, there are other possibilities as well. Using three-month holding periods significantly reduces the number of independent observations, relative to shorter holding periods. Although the sample covers a period of over ten years using weekly observations, the relatively small number of nonoverlapping observations for quarterly returns may not be enough to reject the model.[13] Perhaps with more years of data, the restrictions can be rejected.

Also, from a general perspective, the evidence simply implies that returns move together in proportion over longer holding periods. Returns could behave in this way if excess returns on different types of assets depend upon general uncertainty about the policy process. For example, Lewis (1989) shows how the tightening in the U.S. money market in the early 1980s induced systematic exchange rate

[11] Conditional heteroscedasticity has been found by Cumby and Obstfeld (1984), Giovannini and Jorion (1987), and Domowitz and Hakkio (1985) in foreign exchange returns and by Engle, Lilien, and Robins (1987) in excess long bond returns. Moreover, Cumby (1988) finds that he cannot reject the hypothesis that the covariance of three-month foreign exchange returns and consumption move in a constant proportion over time.

[12] This behavior of returns has been called the "peso problem." See Rogoff (1980) and Krasker (1980).

[13] See Stock and Richardson (1989) on the poor performance of mean reversion tests at multiyear horizons.

forecast errors if the agents in the market learned about the change only over time. This same learning process would also imply similar behavior in interest rate forecast errors. In this case, since the systematic nature in the forecast errors of interest rates and exchange rates is driven by learning about the same policy change, the predictable component of forecast errors that are observed ex post will be correlated across returns. Furthermore, due to the systematic nature of the forecast errors, this common component in returns is likely to be stronger over longer holding periods.

IV. Concluding Remarks

This paper has studied the co-movements in foreign exchange excess returns and interest rate term structure excess returns for some Eurocurrency deposits. One finding of the empirical investigation is that the maturity horizon has important implications for testing models such as the ICAPM. As the maturity horizon increases, the constant beta form of the ICAPM tends to be rejected less often. Thus, the conflicting evidence found in studies that tested this model across different asset markets and maturity horizons appears to arise from the difference in maturity rather than in the choice of asset returns.

The analysis in this paper also suggests that the behavior of excess returns differed markedly during the period of nonborrowed reserves targeting by the Federal Reserve in the early 1980s. This result is particularly evident in the behavior of excess returns on longer term relative to short-term U.S. dollar deposits, and less so for foreign exchange. Nevertheless, when estimated over subsamples, the tendency to reject only over short holding periods remained.

The empirical evidence from this paper therefore provides interesting results with potentially important implications. From the perspective of the intertemporal capital asset pricing model, the results suggest that the constant beta form of the latent variable model can be a fair approximation of the behavior of returns. Given recent evidence of conditional heteroscedasticity in returns, this result might seem surprising. However, rejection of the latent variable model can still be consistent with time-varying consumption betas if these betas themselves move in proportion as the holding period increases. Other explanations of the empirical evidence that are not mutually exclusive include the timing of the arrival of information to the market and the potential presence of uncertainty about discrete events such as policy regime changes. An interesting challenge for future research will be to determine the most likely explanations for the sensitivity of the latent variable model restrictions to the holding period.

Appendix

This appendix describes the construction of the excess returns series used in the text.

First, for the one-week holding period, $k = W$, the foreign exchange returns were calculated as in Giovannini and Jorion (1987). Current 7-day foreign Eurocurrency deposit rates, Eurodollar rates, and spot exchange rates for each

Friday were substituted for $r_{W,t}^{DM}$, $r_{W,t}^{\$}$, and s_t in (8), and the following Friday spot rate was used as the future spot exchange rate, s_{t+k}. The annualization factor was defined as $A_W = 100 \times (365/7)$. Given the settlements procedures in foreign exchange, collecting the spot exchange rates on the same day as the deposit rates may not accurately reflect the timing of transactions in the Eurocurrency market. Therefore, as will be discussed below, the series were also constructed using other methods without affecting the overall results.

For the one-month holding period, $k = 1$, construction of the returns required similar tradeoffs. For the foreign exchange returns, the Friday observations of the 30-day foreign Eurocurrency rate, Eurodollar rate, and spot exchange rate were again used for the current interest rate and spot rates, i.e., $r_{1,t}^{DM}$, $r_{1,t}^{\$}$, and s_t in (8). For the future spot exchange rate, the four-week or 28-day ahead spot rate was used and A_1 was set at $100 \times (365/28)$. Since the Eurocurrency rates are 30-day deposit rates while the changes in exchange rates are over 28 days, the excess returns are somewhat misaligned. To consider the importance of this misalignment, the series were also calculated using other methods, discussed below. However, in all cases, the basic results for the latent variable model were similar. They were also consistent with the findings in other studies such as Hansen and Hodrick (1983) and Hodrick and Srivastava (1984), studies that more carefully attend to the timing of transactions in constructing the foreign exchange returns series than does this paper.

Construction of the term structure returns at the one-month holding period follows the form of equation (9b), where the Friday 7-day and 30-day Eurocurrency deposit rates are used for $r_{W,t}^{i}$ and $r_{1,t}^{i}$, respectively. The returns from rolling over the 7-day rates are calculated by averaging over the deposit rates for the current Friday (t), one week ahead ($t + 1$), two weeks ahead ($t + 2$), and three weeks ahead ($t + 3$). Since the 30-day returns are not evenly divisible by the 7-day returns, the excess returns do not perfectly reflect the true returns. However, the overall results of the latent variable model remain when these returns are excluded.

Finally, for the three-month holding period, $k = 3$, construction of the foreign exchange and term structure returns follows the same basic approach as the one-week and one-month returns. For the foreign exchange returns, the Friday 90-day foreign Eurocurrency rate, Eurodollar rate, and spot exchange rate provided the current interest rate and spot rates, i.e., $r_{3,t}^{DM}$, $r_{3,t}^{\$}$, and s_t in (8). For the future spot exchange rate, the thirteen-week or 91-day ahead spot rate was used and A_1 was set at $100 \times (365/91)$. As with the other holding periods, this represents a slight difference between transactions timing of the 90-day Eurocurrency market and the 91-day ahead foreign exchange market. On the other hand, the 90-day Eurocurrency deposits are evenly divisible by the 30-day deposits so that the term structure returns should be essentially accurate as constructed in (9a), since $r_{1,t}^{i}$ are the 30-day deposits and $r_{3,t}^{i}$ are the 90-day deposits. Some misalignment may remain, however, since, for the 30-day Eurocurrency deposit rate, the 28-day ahead rate was used for $r_{1,t+1}^{i}$ and the 56-day ahead rate was used for $r_{1,t+2}^{i}$.

As this discussion makes clear, each of the series investigated contains some measurement error, either to be consistent with other studies or due to data availability and the need for a uniform calculation of returns across holding periods. Therefore, to consider whether these errors could have significant effects upon the empirical results, the data were also constructed in several different ways and then used to estimate the model. First, different values of A_i were used; specifically, $A_1 = (1200)$ and $A_3 = 400$. Second, instead of using Thursday as the default day when Friday was a holiday, the default day was set at the following Monday and also the previous Wednesday. Third, for the three-month holding period, the foreign exchange returns were also calculated with a 12-week, or 84-day, ahead spot exchange rate for s_{t+k} in (8). In all cases, there were slight changes in the unconstrained regression coefficients reported in Tables I, II, and III, but the basic implications for the latent variable model remain unchanged. Therefore, the errors from potential misalignments in the timing of transactions do not appear to have a significant effect upon the basic results concerning the latent variable model.

REFERENCES

Breeden, Douglas T., 1979, An intertemporal asset pricing model with stochastic consumption and investment opportunities, *Journal of Financial Economics* 7, 265–296.

Campbell, John Y., 1987, Stock returns and the term structure, *Journal of Financial Economics* 18, 373–399.

——— and Richard H. Clarida, 1987, The term structure of Euromarket interest rates: An empirical investigation, *Journal of Monetary Economics* 19, 25–44.

Cumby, Robert E., 1988, Is it risk? Explaining deviations from uncovered interest parity, *Journal of Monetary Economics* 22, 279–299.

——— and John Huizinga, 1988, Two-step two-stage least squares user's guide, New York University.

——— and Maurice Obstfeld, 1984, International interest-rate and price-level linkages under flexible exchange rates: A review of recent evidence, in J. F. O. Bilson and R. C. Marston, eds.: *Exchange Rates: Theory and Practice* (University of Chicago Press, Chicago, IL).

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.

Engle, Robert F., David M. Lilien, and Russell P. Robins, 1987, Estimating time-varying risk premia in the term structure: The ARCH-M model, *Econometrica* 55, 391–407.

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.

——— and Philippe Jorion, 1989, The time variation of risk and return in the foreign exchange and stock markets, *Journal of Finance* 44, 307–325.

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

Hamilton, James D., 1988, Rational expectations econometric analysis of changes in regime: An investigation of the structure of interest rates, *Journal of Economic Dynamics and Control* 12, 385–423.

Hansen, Lars P., 1982, Large sample properties of generalized method of moments estimators, *Econometrica* 10, 1029–1054.

——— and Robert J. Hodrick, 1980, Forward exchange rates as optimal predictors of the future spot rates: An econometric analysis, *Journal of Political Economy* 88, 829–853.

——— and Robert J. Hodrick, 1983, Risk averse speculation in the forward foreign exchange market: An econometric analysis of linear models, in Jacob A. Frenkel, ed.: *Exchange Rates and International Macroeconomics* (University of Chicago Press, Chicago, IL).

Harvey, Andrew C., 1981, *The Econometric Analysis of Time Series* (Philip Allan Publishers, Oxford).

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.

Krasker, William S., 1980, The "peso problem" in testing the efficiency of forward exchange markets, *Journal of Monetary Economics* 6, 269–276.

Lewis, Karen K., 1989, Changing belief about fundamentals and systematic rational forecast errors with evidence from foreign exchange, *American Economic Review* 79, 621–636.

———, 1991, Was there a "peso problem" in the U.S. interest rate term structure: 1979 to 1982?, *International Economic Review*, Forthcoming.

Lucas, Robert E., 1982, Interest rates and currency prices in a two-country world, *Journal of Monetary Economics* 10, 335–360.

Mark, Nelson, 1985, On time-varying risk premia in the foreign exchange markets: An econometric analysis, *Journal of Monetary Economics* 16, 3–18.

Newey, Whitney K. and Kenneth D. West, 1987a, A simple, positive semi-definite, heteroskedasticity and autocorrelation consistent covariance matrix, *Econometrica* 55, 703–708.

——— and Kenneth D. West, 1987b, Hypothesis testing with efficient method of moments estimation, *International Economic Review* 28, 777–787.

Rogoff, Kenneth S., 1980, Essays on expectations and exchange rate volatility, Unpublished Ph.D. dissertation, Massachusetts Institute of Technology, MA.

Stock, James H. and Matthews Richardson, 1989, Drawing inferences from statistics based on multi-year asset returns, Working Paper, Kennedy School of Government, Harvard University.

Stockman, Alan, 1978, Risk, information and forward exchange rates, in Jacob A. Frenkel and Harry G. Johnson, eds.: *The Economics of Exchange Rates: Selected Studies* (Addison Wesley, Reading, MA).

Stulz, René M., 1981, A model of international asset pricing, *Journal of Financial Economics* 9, 383–406.