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Should the holding period matter for the intertemporal consumption-based CAPM?

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Empirical studies of the restrictions implied by the intertemporal capital asset pricing model across different asset markets have found conflicting evidence. This paper asks whether an auxiliary assumption implicit in these tests could be responsible for the pattern of rejections. This auxiliary assumption requires that covariances of returns with consumption move in constant proportion over time. The paper tests this condition empirically using data on foreign exchange, bonds, and equity returns. Interestingly, the evidence suggests that the tendency to reject the intertemporal consumption-based asset pricing relationship depends upon the inadequacy of the auxiliary assumption, not necessarily the relationship itself.

1. Introduction

Recent empirical studies have focused upon restrictions implied by the first-order conditions of intertemporal utility maximization for different asset markets and over different holding periods. These restrictions imply that the expected return on any risky investment strategy must depend upon the conditional covariance between this return and the intertemporal marginal rate of substitution in consumption (hereafter, the MRS). Interestingly, whether these restrictions are rejected in the data appears to depend upon

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the holding period of assets. That is, studies using returns for holding periods of one month or less have summarily rejected these restrictions, while studies with longer three-month holding periods have not rejected the same restrictions.[1] This evidence clearly raises the question: why should the holding period affect how closely returns conform to implications of the consumption-based asset pricing model?

This paper considers the question by focusing upon an auxiliary condition implicit in these tests. This condition requires that the covariance of the MRS with any return move in proportion to the covariance of the MRS with any other return. Given this assumption, the first-order conditions imply that all risky returns held over a particular period must also move in proportion to each other.

The analysis below tests whether the conditional covariances between returns and the MRS in fact move in proportion over time. If this condition is invalid, we would expect to reject proportionality of returns, even if the first-order conditions held. Hence, the holding period should matter for testing the proportionality restrictions of returns if the auxiliary condition itself depends upon the holding period. For example, suppose that conditional covariances of the MRS with long holding period returns move in proportion, but the covariances of the MRS with shorter holding period returns vary idiosyncratically according to the type of return. Then, if the first-order conditions of intertemporal utility maximization hold, we will not reject proportionality of returns over the longer holding period but we will reject over short holding periods.

This paper asks whether the auxiliary assumption can explain the observed rejection pattern by evaluating the conditional covariances from three different perspectives. *First*, the paper analyzes the behavior across returns at each holding period. Specifically, at holding periods of one week, one month, and three months, the study tests whether covariances of returns move in proportion. To explain the rejections, we must find that conditional covariances move in proportion over long but not short holding periods.

Second, the paper examines the behavior across maturities of each individual return. In particular, for each return, the analysis tests whether new information causes the conditional covariances to react more strongly over short holding periods relative to longer holding periods. To explain the empirical regularity on returns, we should find that the covariances of some

[1] For example, these restrictions have been rejected for one-month holding periods by Hodrick and Srivastava (1984) for foreign exchange and by Campbell (1987) for bond and stock returns, and for one-week holding periods by Giovannini and Jorion (1987) for foreign exchange and stock returns. But, at the three-month holding period, Campbell and Clarida (1987) do not reject these restrictions using foreign exchange and bond returns, and Cumby (1989) does not reject the restrictions using equity returns across countries. Lewis (1990a) provides a survey. A notable exception to this pattern is Ferson (1990), who rejects the restrictions using bond and equity returns for quarterly data from 1947 to 1985.

returns move idiosyncratically at short horizons, but revert to moving proportionally with other covariances over longer horizons.

Third, the paper evaluates the joint behavior across returns and over maturities. Specifically, given that covariances move in proportion over long but not short holding periods, we can incorporate this information to provide a more powerful test.

Using these three approaches, the empirical results indicate that covariances indeed tend to move in proportion as the holding period lengthens. Therefore, this evidence suggests that rejections in the intertemporal consumption-based asset pricing relationship at short horizons arises from an inadequate auxiliary assumption, not necessarily from the relationship itself.

The paper proceeds as follows. Section 2 discusses how the latent variable model restrictions implied by the intertemporal CAPM depend upon the auxiliary hypothesis. This section reviews the restrictions, the empirical regularity, and the effects of time variation in conditional covariances. Section 3 analyzes the pattern of conditional covariance behavior across holding periods using the three tests discussed above. Concluding remarks follow.

2. The latent variable model and time-varying consumption covariances

2.1. *Intertemporal utility maximization and the latent variable model*

The restriction that returns move in proportion to each other arises from intertemporal utility maximization with the added assumption that all conditional covariances of returns with the MRS move in a constant proportion. Consider a representative agent that maximizes expected time-additive utility,

$$U_t = \mathrm{E}_t \sum_{j=0}^{\infty} \gamma^j u(c_{t+j}), \tag{1}$$

where E_t denotes the expectation operator conditional on information known at time t, $u(\cdot)$ is the period utility function, c_t is consumption at date t, and $\gamma < 1$ is the discount factor. Then, any asset with nominal payoffs k periods ahead must satisfy the first-order conditions

$$1 = \mathrm{E}_t\left[\frac{\left(\gamma^k u'(c_{t+k})/p_{t+k}\right)}{\left(u'(c_t)/p_t\right)}\left(1 + r_{t,k}^i\right)\right], \tag{2}$$

where p_t is the price of the consumption good at time t and $r_{t,k}^i$ is the nominal return on an asset purchased at time t with payoffs k periods ahead. From (2), any asset maturing at $t+k$ depends upon the nominal intertemporal rate of substitution in consumption. For notational simplicity, we will define this variable as

$$n_{t,k} \equiv \frac{\left(\gamma^k u'(c_{t+k})/p_{t+k}\right)}{\left(u'(c_t)/p_t\right)}.$$

Since eq. (2) holds for all assets, it also holds for the risk-free rate over holding period k, implying that this return is $(1+r^r_{t,k}) = (1/\mathrm{E}_t n_{t,k})$. Using this result, eq. (2) can be rewritten as

$$\mathrm{E}_t\left(r^i_{t,k} - r^r_{t,k}\right) = -\mathrm{cov}_t\left(n_{t,k}, r^i_{t,k}\right)\left(1 + r^r_{t,k}\right), \quad \forall i, \tag{3}$$

where cov_t is the covariance operator conditional upon current information. Eq. (3) describes the risk premia on asset i relative to the risk-free rate.

Since (3) holds for all assets i, we may substitute for any other asset j to obtain

$$\mathrm{E}_t\left(r^i_{t,k} - r^r_{t,k}\right) = \left[\mathrm{cov}_t\left(n_{t,k}, r^i_{t,k}\right)/\mathrm{cov}_t\left(n_{t,k}, r^j_{t,k}\right)\right]\mathrm{E}_t\left(r^j_{t,k} - r^r_{t,k}\right),$$
$$\forall i, j, \quad i \neq j. \tag{4}$$

In other words, since all returns with the same holding period depend upon their conditional covariances with the MRS over that same holding period, they move in proportion to each other according to the ratios of these conditional covariances.

2.2. *The latent variable model*

Studies of these restrictions have proceeded under the auxiliary assumption that the ratios of consumption covariances are constant over time so that

$$\left(\frac{\beta^i}{\beta^j}\right) = \frac{\mathrm{cov}_t\left(n_{t,k}, r^i_{t,k}\right)}{\mathrm{cov}_t\left(n_{t,k}, r^j_{t,k}\right)}, \tag{5}$$

where the β's are constants.[2]

Given that (5) holds, the first-order conditions in (4) imply restrictions on the projections of excess returns on information variables known at time t. To see this, consider the projections of excess returns from any asset i upon a

[2]Early studies simply assume that the conditional covariances were constant. More recently, researchers have noted that condition (5) will hold as long as covariances move in proportion. See Hansen and Hodrick (1983), Cumby (1988), and Wheatley (1989).

subset of the information set, $x_t = (x_{1t}, x_{2t}, \ldots, x_{Nt})'$,

$$r^i_{t,k} - r^r_{t,k} = x'_t b^i + \varepsilon^i_{t+k}, \tag{6}$$

where $b^i = (b^i_1, b^i_2, \ldots, b^i_N)'$ is a parameter vector and where ε^i_{t+k} is a composite error, the sum of an error in measuring expected returns and a k-step-ahead forecast error. Then the first-order conditions, (4), together with the maintained auxiliary condition (5) imply the restrictions

$$[b^i_1, b^i_2, \ldots, b^i_N] = (\beta^i/\beta^j)[b^j_1, b^j_2, \ldots, b^j_N], \qquad \forall i,j, \quad i \neq j. \tag{7}$$

These restrictions have been tested for a number of different types of returns and for different holding periods. The types of excess returns studied include open positions on foreign exchange, stock market returns, and bond returns of various maturities, while the holding periods have ranged from one week to three months. Interestingly, whether the restrictions in (7) are rejected appears to depend strongly upon the length of the holding period. In particular, the restrictions in (7) are rejected over holding periods of one month or less, but are generally not rejected for quarterly holding periods.

This pattern would be perfectly consistent with the intertemporal Euler equations, however, if the auxiliary condition in (5) depended upon the holding period. Specifically, if covariances move in proportion over longer holding periods such as a quarter, but not over shorter holding periods, then condition (5) in turn would hold for longer holding periods and not short ones. As a result, we would reject the restrictions over these shorter holding periods simply because the auxiliary assumption was violated – not because the model was wrong.

2.3. Interpreting the auxiliary assumption of conditional covariances

We now consider how this auxiliary condition may break down. First note that (5) will automatically hold if the covariances are constant. Therefore, any violation of (5) requires conditional heteroscedasticity in the joint process of rates and the intertemporal MRS. As an empirical matter, variances that change with new information about the economic state have been found in many types of asset returns.[3]

[3]For asset returns, this heteroscedasticity has been found by Cumby and Obstfeld (1984), Giovannini and Jorion (1987), and Domowitz and Hakkio (1985) in foreign exchange returns; by Christie (1982), Poterba and Summers (1986), Schwert and Seguin (1989), and French, Schwert, and Stambaugh (1987) for stock returns; and by Evans (1990) for long bond returns. Kandel and Stambaugh (1990) find that aggregate consumption growth, an ingredient in the MRS, displays heteroscedasticity.

Where does this heteroscedasticity come from? In general equilibrium, both returns and the MRS are determined by the state process of the economy.[4] Hence, there are at least two potential sources. First, the state variables themselves may be conditionally heteroscedastic. In this case, changing variances of the state process will make the variances of the returns and MRS processes change over time as well. Second, since the returns can in general be complicated nonlinear functions of the state variables, the functional form may induce heteroscedasticity in returns even if the state variables are homoscedastic.[5]

Given that heteroscedasticity exists and depends upon the state of the economy, we may next ask: what pattern in conditional consumption covariances would give the observed pattern of rejecting the latent variable model? If the underlying heteroscedasticity affects the returns differently, conditional covariances will not move together, violating the auxiliary assumption (5). Since each of the returns functions are distinct functions of the state process, idiosyncratic movements in the covariances seem likely whether the cause is the primitive process of the states or the nonlinearity of returns. On the other hand, time aggregation may mitigate the importance of these nonlinearities and other reasons for idiosyncratic behavior in variances. If so, then this auxiliary assumption may be valid over longer, but not short, holding periods.

3. Does the holding period matter for consumption covariances?

As shown above, the latent variable restrictions would be rejected even when the intertemporal asset pricing relations hold if the covariances of consumption and returns do not move in proportion over time. This section begins by briefly summarizing the findings in the literature concerning the latent variable restrictions. Then, consumption covariances are examined over holding periods and types of returns to evaluate whether their behavior can account for the pattern of rejecting the latent variable restrictions.

3.1. Data definitions and definition of variables

To investigate the three types of returns discussed in section 2, the data series were constructed for equity premia, bond term premia, and foreign exchange risk premia. Part I in table 1 defines these three types of returns all

[4]Lewis (1990b) provides some general equilibrium examples in the context of models similar to Lucas (1982) and Mehra and Prescott (1985).

[5]This possibility has recently been discussed in studies of nonlinearities in asset prices such as Scheinkman and LeBaron (1989) and Hsieh (1989). Hsieh (1990) shows that, when an asset price depends nonlinearly upon its state variables, the asset price can exhibit conditional heteroscedasticity even if the underlying state process is conditionally homoscedastic.

Table 1

Summary of variables and portfolios.

I. Definition of Returns for k-Month Holding Period Returns

A. Foreign exchange returns[a]

$$r^i_{k,t} \equiv (1200/k)(s^i_{t+k} - s^i_t) + \hat{r}^i_{k,t} - \hat{r}^s_{k,t} \text{ for currency } i$$

B. Term structure returns[b]

$$r^i_{3,t} \equiv \frac{1}{3}\sum_{j=0}^{2} \hat{r}^i_{1,t+j} - \hat{r}^i_{3,t}$$

$$r^i_{1,t} \equiv \frac{1}{4}\sum_{j=0}^{3} \hat{r}^i_{w,t+j} - \hat{r}^i_{1,t}$$

C. Equity returns[c]

$$r_{k,t} \equiv A_k((P_{t+k} + D_{t,k})/P_t) - r^s_{k,t}$$

II. Composition of Portfolio and Information Variables Sets

Structure of Information Variables Sets

Set A – See under each portfolio set below.
Set B – Set A plus variables in set A squared.
Set C – Set A plus quarterly growth rates of consumption, inflation lagged three and twelve months, industrial production, and the U.S. terms of trade.

Portfolio set 1: 'Foreign Exchange'

Holding periods: 1 week, 1 month, 3 months
Returns: Foreign exchange returns for Deutsche mark, British pound, French franc, Japanese yen.
Information variables set A: Current one-month-ahead forward premium for included currencies.

Portfolio set 2: 'Mixed Term Structure / Foreign Exchange'

Holding periods: 1 month, 3 months
Returns: Foreign exchange returns for Deutsche mark, British pound, and term structure returns for U.S. dollar, Deutsche mark, and British pound.
Information variables set A: Current one-month-ahead forward premium and current spread between one-month and one-week Eurocurrency deposits for included currencies.

Portfolio set 3: 'Mixed Equity / Foreign Exchange'

Holding periods: 1 week, 1 month, 3 months
Returns: Foreign exchange returns for Deutsche mark, British pound, and New York Stock Index returns.
Information variables set A: Current one-month-ahead forward premium for included currencies and the previous monthly return on the NYSE.

[a] s^i_t is the spot price of one unit of currency i in terms of dollars at time t, $\hat{r}^i_{k,t}$ are the annualized k-month Eurocurrency deposit rates in currency i, and $A_t \equiv 100 \times (365/N)$ where N is the number of days in the holding period.

[b] The terms $r_{3,t}$, $r_{1,t}$, and $r_{w,t}$ correspond to the three-month, one-month, and one-week deposits, respectively.

[c] P_t is time t stock price and $D_{t,k}$ is dividend payment between t and $t+k$.

in excess of the risk-free rate. They are, first, 'foreign exchange returns', the returns from holding open positions in foreign currency deposits; second, the 'term structure returns' from rolling over short rates for longer periods; and third, the 'equity returns' from holding equity, receiving dividends and capital gains. In order to analyze the behavior across holding periods, these return series were calculated for one-week, one-month, and three-month holding periods.

We begin by estimating eq. (6) and testing the latent variable model restrictions in (7) across holding periods. Testing the restrictions in (7) requires, first, a set of returns as left-hand-side variables, r', and, second, a set of some information variables, x_t, currently known by market traders. Part II in table 1 defines the composition in the empirical estimation of both the portfolio of returns considered jointly and the set of information variables.

The portfolio and information sets were formed to match different groups of studies in the literature. In the first portfolio, 'Foreign Exchange', returns on open positions in German mark, British pound, Japanese yen, and French franc bonds against the dollar bonds are examined jointly.[6] The second, 'Mixed Term Structure/Foreign Exchange', portfolio set consists of five returns: three returns on longer-term Eurocurrency deposits relative to rolling over short-term deposits for three currencies, the German mark, the British pound, and the U.S. dollar, and two foreign exchange returns for the German mark and the British pound. The third, 'Mixed Equity/Foreign Exchange', portfolio set is the excess return on U.S. equity plus the two foreign exchange returns for the German mark and the British pound.[7] The information variables sets are also listed in part II of table 1. Set A includes standard variables that appear to be correlated with the left-hand-side variables. Set B includes the squares of these same variables. Finally, instead of squared variables, set C substitutes some real variables that are likely correlated with current consumption. The data appendix describes these sets in more detail as well as the sources of all the data series.

3.2. The latent variable model

Tests of the restrictions in (7) based upon estimating the projection equations, (6), provide different results depending upon the holding period of

[6] The Japanese yen and the French franc data do not begin until October 1979. However, using estimation periods that start earlier with other currencies do not alter the basic conclusions below. See Lewis (1990a).

[7] Foreign exchange groups of returns have been analyzed in Hansen and Hodrick (1983), Hodrick and Srivastava (1984), and Cumby (1988), among others. Campbell and Clarida (1987) examine the same five 'Mixed Term Structure/ Foreign Exchange' return set as in the text for a three-month holding period. Giovannini and Jorion (1987) test the restrictions for the portfolio set of one-week U.S. equity and foreign exchange returns, similar to the 'Mixed Equity' set above.

Table 2

Proportionality test of the single beta asset pricing model using weekly frequency data.[a]

Holding periods	January 76–May 86	October 79–May 86
	Portfolio set 1: 'Foreign Exchange'	
	$\chi^2(12)$	$\chi^2(12)$
A. Three months	NA	11.76 (0.465)
B. One month	NA	21.78 (0.040)
C. One week	NA	80.53 (< 0.000)
	Portfolio set 2: 'Mixed Term Structure / Foreign Exchange'	
	$\chi^2(20)$	$\chi^2(20)$
A. Three months	15.76 (0.731)	12.47 (0.899)
B. One month	39.27 (0.006)	12.25 (0.907)
	Portfolio set 3: 'Mixed Equity / Foreign Exchange Returns'	
	$\chi^2(6)$	$\chi^2(6)$
A. Three months	4.62 (0.593)	8.53 (0.202)
B. One month	6.31 (0.389)	11.46 (0.075)
C. One week	20.47 (0.002)	27.94 (< 0.000)

[a]All variables, portfolio sets, and information variables sets are defined in table 1.

the returns, k. Table 2 provides an example of this basic finding using the instrumental variable set A with weekly frequency data.[8] For the 'Foreign Exchange' and the 'Mixed Equity/Foreign Exchange' portfolio sets, the restrictions are rejected at marginal significance levels of less than 1%. For the 'Mixed Term Structure/Foreign Exchange' set estimated over the full sample, the restrictions are also strongly rejected at the one-month holding period, but not for the three-month holding period.

3.3. *Ex ante returns and conditional covariances*

Since the validity of the latent variable model as a test of intertemporal asset pricing relationships depends upon the behavior of consumption covari-

[8]These test statistics were estimated based upon GMM estimators that are asymptotically efficient. See Gibbons and Ferson (1985) and Hansen and Hodrick (1983). Ferson and Foerster (1990) have recently shown that these estimators may behave poorly in finite samples and that an iterated GMM approach may have better properties in systems with many asset equations.

ances, we will examine how the covariances respond to the current information set.

For this purpose, we first rewrite eq. (6) as[9]

$$r^{i}_{t,k} - r^{r}_{t,k} = x'_t b^i + \varepsilon^{i}_{t+k}, \tag{6'}$$

where $\varepsilon^{i}_{t+k}|x_t \sim$ i.i.d. $(0,(\sigma^{i}_{t,k})^2)$. In particular, the *ex post* realized squared residual depends upon the market's true conditional variance forecast and a disturbance term, ν. The relationship between the conditional variance and the squared residuals is given by

$$\left(\varepsilon^{i}_{t,k}\right)^2 = \left(\sigma^{i}_{t,k}\right)^2 \exp\left[-\tfrac{1}{2}\operatorname{var}(\nu^i) + \nu^{i}_{t+k}\right], \tag{8}$$

where for nonoverlapping forecast horizons ν_t is an i.i.d. normally distributed random variable with variance var(ν).

Although the market's conditional variance of returns is unobserved by the econometrician, we can use the same logic here as in the latent variable model. That is, given that the econometrician observes a subset of the current information set, z_t, he observes the true conditional variance with error according to

$$\left(\sigma^{i}_{t,k}\right)^2 = \delta^{i}_{k} \exp\left[z_t\theta^{i}_{k} - \tfrac{1}{2}\operatorname{var}\left(w^{i}_{k}\right) + w^{i}_{t,k}\right], \tag{9}$$

where $w_{t,k}$ is the error in measuring conditional variances by the econometrician and is normally distributed with variance var(w_k). Under these conditions and some standard regularity conditions, the conditional variance parameters, θ, can be estimated by OLS in the following regression:[10]

$$\log\left(\hat{\varepsilon}^{2}_{t,k}\right) = -\tfrac{1}{2}\left[\operatorname{var}(w_k) + \operatorname{var}(\nu_k)\right] + \log(\delta_k) + z_t\theta_k + e_{t,k}, \tag{10}$$

where $e_{t,k} \equiv w_{t,k} + \nu_{t+k}$ and where the superscript i has been suppressed for notational simplicity. Note that the variance of the measurement error in conditional variances, var(w), the variance of the disturbance to conditional variances, var(ν), and the scale factor in conditional variances, δ_k, are not independently observable. Therefore, the θ parameters will be relatively inefficiently estimated. Although in principle the variables that help explain

[9]In this equation, the residuals to the projection equations are assumed to be the true innovations. However, the measurement error in expectations biases the results away from the hypothesis considered below. This result is demonstrated in an appendix available upon request from the author.

[10]These conditions are described in an appendix available upon request from the author.

the conditional variances, z_t, need not be the same as those that explain the conditional means, x_t, they are assumed the same below.

Table 3 provides summary statistics and tests of the hypothesis of constant variances for the returns in each portfolio set of monthly frequency data. Columns 4 and 5 give the means and standard deviations, respectively, of the logarithms of the squared residuals from the projection equations, (6′). The sixth column reports the marginal significance levels for the Wald test statistic of the hypothesis that the θ coefficients are jointly zero based upon the regressions in eq. (10). These results corroborate other studies finding that asset returns appear to display considerable heteroscedasticity.

3.4. Do conditional covariances move in proportion to return variances?

To construct measures of the covariances of these returns with consumption, we also require a projection equation for consumption as in the following:

$$\left(\Delta^k c_t / c_t\right) = x_t b_k^c + \varepsilon_{t,k}^c, \tag{11}$$

where Δ^k is the forward difference operator k periods ahead and $\varepsilon_{t,k}^c$ is the residual to the consumption projection equation. Analogous to eq. (8), the cross-products of the errors to returns and consumption depend upon the conditional covariance between consumption and asset returns according to

$$\varepsilon_{t,k}^i \varepsilon_{t,k}^c = \sigma_{t,k}^{ic} \exp\left[-\tfrac{1}{2}\operatorname{var}(\nu^{ic}) + \nu_{t+k}^{ic}\right], \tag{12}$$

where $\sigma_{t,k}^{ic}$ is the covariance conditional upon time t information between consumption and asset i returns over the next k periods and where ν^{ic} is a normally distributed disturbance. Thus, ν^{ic} is the *ex post* innovation in the conditional covariance, $\sigma_{t,k}^{ic}$.

Since the standard deviations of the residuals in the returns projections are much larger than the standard deviation of residuals in the consumption projections, much of the variation in the conditional covariances of returns and consumption may arise from movements in the variances of returns. If so, then we may exploit this information to provide more precise measures of the behavior of covariances.

For this purpose, note that if all of the movement in this covariance arises from movement in the variance of returns, then these variables will obey the restriction

$$\sigma_{t,k}^{ic} = a_k^i \left(\sigma_{t,k}^i\right)^2, \tag{13}$$

Table 3

Tests of variance and covariance behavior using monthly frequency data; January 1976–May 1986.[a]

Returns	(1) Mean $\log(\hat{\varepsilon}^{\prime}\hat{\varepsilon}^{c})$	(2) S.D. $\log(\hat{\varepsilon}^{\prime}\hat{\varepsilon}^{c})$	(3) H_0: Const. $\sigma^{\prime c}$ MSL	(4) Mean $\log(\hat{\varepsilon}^2)$	(5) S.D. $\log(\hat{\varepsilon}^2)$	(6) H_0: Const. $\sigma^{\prime}$ MSL	(7) H_0: Const $(\sigma^{\prime c}/\sigma^{\prime})$ MSL
Portfolio set: 'Mixed Term Structure / Foreign Exchange'							
A. Three-month returns[a]							
Foreign exchange:							
German DM	2.69	1.51	< 0.000	4.99	2.16	< 0 000	0.892
U.K. pound	2.64	1.70	< 0.000	4.88	2.16	0.049	0.924
Term structure:							
U.S. int. rate	−1.44	1.71	< 0.000	−3.27	2.56	< 0 000	0.869
German int. rate	−2.12	1.51	< 0.000	−4.64	2 04	< 0.000	0.877
British int. rate	−1.42	1.59	< 0.000	−3.24	2.33	< 0.000	0.990
B. One-month returns							
Foreign exchange:							
German DM	3.55	1.77	0.185	5 60	2.29	0.163	0.495
U.K. pound	3.65	1.88	0.211	5.80	2.47	0.693	0 373
Term structure:							
U.S. int. rate	−1.09	1.88	0.030	−3.69	2.49	< 0.000	0.104
German int. rate	−1.40	1.81	0.001	−4.29	2.50	< 0.000	0.112
British int. rate	−0.51	1.63	0.697	−2.53	2.35	< 0.000	0.073
Portfolio set: 'Mixed Equity / Foreign Exchange'							
A. Three-month returns[a]							
Foreign exchange.							
German DM	2.61	1.67	0.147	4.95	2.32	0 111	0.225
U.K. pound	2.51	1.69	0.062	4.75	2.40	0.001	0 251
U.S. equity	2.89	1.56	0.217	5.51	2.15	0.447	0.255
B. One-month returns							
Foreign exchange:							
German DM	3.63	1.63	0.007	5.68	2.22	0.323	0.899
U.K. pound	3.63	1.80	0.013	5.68	2.61	0.385	0.999
U.S. equity	4.10	1.46	0.856	6.63	2.04	0.429	0.293

[a]Standard errors are adjusted for a first-order moving average process using the sample moments method described in Hansen (1982).

where a_k^i is a constant. This restriction says that the conditional covariances vary over time in constant proportion with the conditional variances in returns.

We can test this restriction using the relationship between *ex post* residuals and conditional variances. Substituting eq. (13) into eq. (12) and taking the logarithm implies

$$\log(\varepsilon_{t,k}^i \varepsilon_{t,k}^c) = \log(a_k^i) + \log(\delta_k^i) - \tfrac{1}{2}\operatorname{var}(\nu^{ic}) - \tfrac{1}{2}\operatorname{var}(w_k^i) + z_t \theta_k^i + e_{t,k}^{ic}, \tag{14}$$

where $e_{t,k}^{ic} \equiv w_{t,k}^i + \nu_{t+k}^{ic}$. In this form, we can directly evaluate the covariance restriction (13) by estimating eq. (14) together with eq. (10) and testing the restrictions that all of the components in the θ vector are equal; i.e., $\theta_k^i = \theta_k^j$, for all i, j.

The Wald tests of this restriction are given in the last column of table 3. As the results indicate, this restriction is not rejected at the 95% confidence level for any of the nonoverlapping one-month returns.[11] Furthermore, the hypothesis that the time-varying coefficients are equal across the returns variances and consumption covariances is not rejected at the 80% confidence level for any of the three-month returns.

Thus, we cannot reject the hypothesis that the changes in the covariance in consumption and returns depend only upon changes in the variance of returns, a result that should be important in future research. This result implies that we may focus upon the behavior of return variances alone in order to understand the behavior of consumption and return covariances. Hence, for the rest of the analysis below, we will assume that (13) holds so that $\sigma_t^{ic} \propto (\sigma_{t,k}^i)^2$.

We can now directly address the question of whether the holding pattern should matter. Recall that we would expect to see the observed pattern of rejection in the latent variable model if the conditional covariances move in constant proportion over long holding periods but not short holding periods. We will next test this relationship using three different tests. Test 1 asks: for a given holding period k, do consumption covariances across different returns move proportionally over time? Test 2 asks: for individual returns i, do consumption covariances with each return across different holding periods tend to react more strongly to new information as the holding period shortens? Test 3 uses information both across assets and holding periods to obtain a more powerful test of both questions jointly.

[11]For the three-month returns, the residuals are likely to be autocorrelated due to the shock to the cross-products of *ex post* projection errors, i.e., ν_{t+k}. Since evidence of serial correlation was found, the reported results are corrected for a moving average process using the sample moments method described in Hansen (1982). The degree of serial correlation in the residuals was tested with the 'l-test' using Cumby and Huizinga (1988, 1990).

3.5. Across returns at individual holding periods: Test 1

Using the variance process together with condition (13), we may now directly test the auxiliary assumption to latent variable tests (7). For this purpose rewrite eq. (9), given holding period k for assets i and j, as

$$\begin{aligned}\left(\sigma_{t,k}^{i}\right)^{2} &= \delta_{k}^{i}\exp\left[z_{t}\theta_{k}^{i}-\tfrac{1}{2}\operatorname{var}\left(w_{k}^{i}\right)+w_{t,k}^{i}\right],\\ \left(\sigma_{t,k}^{j}\right)^{2} &= \delta_{k}^{j}\exp\left[z_{t}\theta_{k}^{j}-\tfrac{1}{2}\operatorname{var}\left(w_{k}^{j}\right)+w_{t,k}^{j}\right].\end{aligned} \tag{15}$$

If conditional variances move in proportion over holding period k, the coefficients on the time-varying z_t processes must be the same across all returns. Hence, we may test the proportionality of variances by estimating eq. (10) across assets and testing the cross-equation restriction:

$$\text{Test 1:}\quad \theta_{k}^{i}=\theta_{k}^{j},\quad \forall\, i,j.$$

If the holding period matters for violations of this assumption, then the holding period will also matter for testing the latent variable restrictions. In particular, we should find that Test 1 is rejected over short periods, but not over longer holding periods of three months.

Table 4 reports the results of these tests across holding periods for each portfolio. We evaluate Test 1 by first estimating (10) jointly for all of the assets in each portfolio set with Hansen's (1982) GMM, constraining $\theta^{i}=\theta^{j}$ for all i and j. The table reports the chi-squared statistics of the overidentifying restrictions along with the marginal significance levels in parentheses. Strikingly, the test statistics on the 'Foreign Exchange' portfolio set mirror the relationship across holding periods found in the latent variable model estimates. The marginal significance levels of the proportionality conditions increase with the holding period.

These restrictions are not rejected at the one-month or the three-month horizon for the 'Mixed Term Structure/Foreign Exchange' portfolio set for either the full or subsample periods. To check whether this result arises from the relatively large number of parameters, the restrictions were also tested for the same information variables but using a smaller portfolio set with only the Eurodollar term structure returns and the British pound and German mark returns. If the restrictions do not hold for these three equations, they should also be rejected for the larger system of five equations. However, as table 4 indicates, these restrictions are strongly rejected for the full sample period for both the one-month and the three-month holding periods. The lower marginal significance levels when estimating fewer equations suggest that the larger equation system may be overparameterized.

Lastly, the equity portfolio set displays an odd pattern. The return variances appear relatively constant over the one-month period and the restrictions are not rejected over this horizon. However, they are rejected at the

Table 4

Test of proportional time variation in conditional variances using weekly frequency data.[a]

Holding periods	January 76 –May 86	October 79 –May 86
	Portfolio set 1: 'Foreign Exchange'	
	$\chi^2(24)$	$\chi^2(24)$
A. Three months	NA	12.30 (0.999)
B. One month	NA	25.77 (0.365)
C. One week	NA	34.58 (0.075)
	Portfolio set 2: 'Mixed Term Structure / Foreign Exchange'	
	$\chi^2(40)$	$\chi^2(40)$
A. Three months	20.56 (0.995)	13.36 (0.999)
B. One month	22.86 (0.986)	13.39 (0.999)
	On subset	
	$\chi^2(20)$	$\chi^2(20)$
A. Three months	33.12 (0.033)	11.57 (0.930)
B. One month	33.84 (0.027)	28.95 (0.089)
	Portfolio set 3: 'Mixed Equity / Foreign Exchange Returns'	
	$\chi^2(12)$	$\chi^2(12)$
A. Three months	23.46 (0.024)	21.98 (0.038)
B. One month	9.60 (0.651)	8.23 (0.767)
C. One week	29.56 (0.002)	23.33 (0.025)

[a]All variables, portfolio sets, and information variables sets are defined in table 1.

one-week and three-month horizons. Further inspection of the conditional variances of equity returns indicated that the regularity conditions necessary for estimation did not hold.[12]

In summary, direct tests of the condition that variances move together suggest a pattern consistent with the pattern found in rejecting the intertem-

[12]Specifically, the conditional variance process in (8) requires that nonoverlapping innovations ν_t be i.i.d. This hypothesis was tested using the 'l-statistic' described in Cumby and Huizinga (1990). Although this hypothesis could not be rejected for most returns, the variance on equity displayed significant evidence of serial correlation up to six lags.

poral CAPM latent variable model. The basic pattern can be found in foreign exchange and term premia, but equity appears to be misspecified by the conditional variances model. Note that Test 1 above tests the behavior across returns at given holding periods, k, but does not incorporate behavior across holding periods. Therefore, more information about the pattern may be gleaned by investigating variances across holding periods directly.

3.5. Across holding periods for individual returns: Test 2

The evidence above suggests that, as the holding period shortens, conditional variances tend to move idiosyncratically as a function of the state process of the economy. One explanation for this behavior is that, upon viewing new information, investors change their beliefs about the variances of short holding returns more strongly than the variances of long holding returns. Hence, investors' beliefs about the longer-term returns variances are relatively unchanged. If investors assess the returns process in this way, we should find this behavior empirically across holding periods for individual returns.

To analyze the reaction of the variances to current information, we will define a unit 'news' information variable, u_t. This variable is a linear combination of variables in the current information set,

$$u_t \equiv z_t \phi,$$

where ϕ is vector of parameters. We may then measure the relative variance response for returns at different holding periods by estimating how the variances react to the same set of new information. For this purpose, we will estimate the elasticity of variances with respect to news, defined for each asset i, as

$$\left[\frac{\left(\partial\sigma_{k,t}^i/\sigma_{k,t}^i\right)}{\partial u_t}\right] \equiv \eta_k^i .$$

Rewriting the variance process in (9) in terms of the unit news variable, over different holding periods such as $k = 1,3$, yields

$$\begin{aligned}\left(\sigma_{t,1}^i\right)^2 &= \delta_1^i \exp\left[\eta_1^i u_t - \tfrac{1}{2}\operatorname{var}\left(w_1^i\right) + w_{t,1}^i\right],\\ \left(\sigma_{t,3}^i\right)^2 &= \delta_3^i \exp\left[\eta_3^i u_t - \tfrac{1}{2}\operatorname{var}\left(w_3^i\right) + w_{t,3}^i\right].\end{aligned} \tag{16}$$

We may evaluate the cross-maturity elasticities for each individual asset return i directly by estimating eq. (10) jointly for asset i at maturities $k = 1,3$

and constraining the new information, u_t, to be the same across equations as in (16). Using these estimates provides a measure of the variance reaction to 'news' through the conditional variances ratio, (η_3^i/η_1^i). When this ratio is equal to one, then new information causes investors to change their forecasts of the one-month and the three-month returns variances in the same proportion. In this case, idiosyncratic changes in the conditional variances over short-term horizons persist to longer-term horizons. On the other hand, when the conditional variance ratio is less than one, investors faced with new information revise their forecasts of the one-month variance more strongly than the three-month variance. Thus, we may test this hypothesis for three-month relative to one-month returns as:

$$\text{Test 2a:} \quad \left(\eta_3^i/\eta_1^i\right) < 1.$$

Table 5 reports the results of estimating these conditional variance elasticity ratios for the three portfolio sets and for the full sample and post-1979 subsample. The projection vector in the one-month equation was normalized as the 'news' variable so that $u_t \equiv z_t\theta_1^i$. The first column in the table reports the conditional variance elasticity ratio together with its standard error. The second column reports the t-statistic for the hypothesis that the ratio is less than one. The third column gives the chi-squared statistic of the test of overidentifying restrictions.[13]

As the table demonstrates, the full sample estimation provides fairly precise results. For the 'Mixed Term Structure/Foreign Exchange' portfolio set, all of the point estimates of the elasticity ratios are significantly less than one, as the first and second columns show. Also, the overidentifying restrictions are not rejected except for the German term structure case. For the full sample estimates of the 'Mixed Equity/Foreign Exchange' portfolio, the conditional variance elasticity ratios are also significantly less than one except for the German foreign exchange returns. For the period since October 1979, the elasticity ratios are generally less precisely estimated and, perhaps as a consequence, the results appear more mixed. Overall, however, most of these ratios are significantly less than one for the full sample.

Similarly, table 6 reports these same elasticity variance estimates for the one-month relative to the one-week returns and the three-month relative to the one-week returns. As above, these elasticities come from estimating eq. (10) jointly across one-week and one-month holding periods, and then across

[13]If N is the number of z_t variables, there are $2N$ orthogonality conditions but only $N+2$ parameters to estimate ($N-1$ many θ parameters, 2 δ parameters, and 1 conditional variance elasticity ratio), so that there remain $N-2$ overidentifying restrictions.

Table 5

Ratios of conditional variance elasticities over one-month elasticities using weekly frequency data.

Returns	1/76–5/86 (η_{3m}/η_{1m}) (S.E.)	H_0: $\eta_{3m} < \eta_{1m}$ t-stat.	Test[a] of restrictions	10/79–5/86 (η_{3m}/η_{1m}) (S.E.)	H_0: $\eta_{3m} < \eta_{1m}$ t-stat.	Test[a] of restrictions
1. 'Foreign Exchange'						
						$\chi^2(7)$
German DM	NA	NA	NA	3.82 (2.87)	0.98	8.99 (0.253)
U.K. pound	NA	NA	NA	2.90 (1.54)	1.23	12.00 (0.101)
Japanese yen	NA	NA	NA	1.21 (0.18)	1.13	5.98 (0.542)
French franc	NA	NA	NA	0.48 (0.13)	-3.90^{b}	8.30 (0.307)
2. 'Mixed Term Structure / Foreign Exchange'						
Foreign exchange						
			$\chi^2(9)$			$\chi^2(9)$
German DM	0.69 (0.15)	-1.99^{b}	9.18 (0.421)	−0.30 (0.31)	-4.18^{b}	7.83 (0.551)

U.K. pound	−1.26 (0.55)	−4.12[b]	7.97 (0.537)	−0.48 (0.37)	−4.00[b]	8.49 (0.486)
Term structure						
U.S.	0.63 (0.10)	−3.82[b]	2.80 (0.999)	0.99 (0.15)	−0.01	9.61 (0.383)
German	0.74 (0.09)	−2.80[b]	23.21 (0.006)	1.36 (0.25)	1.44	16.75 (0.053)
U.K.	0.26 (0.13)	−5.85[b]	7.39 (0.597)	0.81 (0.15)	−1.27	14.04 (0.121)
			3. 'Mixed Equity / Foreign Exchange'			
			$\chi^2(5)$			$\chi^2(5)$
German DM	1.68 (0.50)	1.36	5.64 (0.343)	−2.60 (2.06)	−1.75[b]	4.64 (0.475)
British pound	−2.51 (1.39)	−2.53[b]	9.41 (0.094)	3.40 (2.74)	0.88	6.35 (0.274)
Equity	−1.44 (0.99)	−2.46[b]	3.29 (0.655)	0.84 (0.27)	−0.582	7.32 (0.198)

[a]Hansen's J-statistic of the overidentifying restrictions that one-month and three-month return variances move in the same proportion.
[b]Significantly less than one at the 95% confidence level.

Table 6

Ratios of conditional variance elasticities over one-week elasticities using weekly frequency data.

Returns	(η_{1m}/η_{1w}) (S.E.)	H$_0$: $(\eta_{3m}/\eta_{1w}) < 1$ t-stat.	(η_{1m}/η_{1w}) (S.E.)	H$_0$: $(\eta_{1m}/\eta_{1w}) < 1$ t-stat.	Test of[a] restrictions (M.S.L.)
Portfolio set 1: 'Foreign Exchange', October 1979–May 1986					
					$\chi^2(14)$
German	−2.91	−3.80[b]	−0.13	−3.32[b]	8.10
	(1.03)		(0.34)		(0.884)
U.K.	4.25	1.31	2.00	1.09	8.59
	(2.49)		(0.91)		(0.856)
Japan	2.11	2.31	1.79	1.68	7.30
	(0.48)		(0.47)		(0.923)
France	−0.27	−4.88[b]	0.99	−0.04	10.06
	(0.26)		(0.28)		(0.758)
Portfolio set 3: 'Mixed Equity / Foreign Exchange', January 1976–May 1986					
					$\chi^2(10)$
German DM	0.76	−0.88	0.73	−1.15	22.20
	(0.27)		(0.23)		(0.014)
U.K. pound	−0.37	−7.24[b]	0.58	−2.74[b]	18.60
	(0.19)		(0.15)		(0.046)
Equity	−0.64	−5.43[b]	0.54	−2.08[b]	6.98
	(0.30)		(0.22)		(0.727)
Portfolio set 3: 'Mixed Equity / Foreign Exchange', October 1979–May 1986					
					$\chi^2(10)$
German DM	−0.70	−5.62[b]	0.41	−5.87[b]	15.95
	(0.30)		(0.24)		(0.101)
U.K. pound	0.74	−0.88	0.03	−3.90[b]	11.74
	(0.29)		(0.25)		(0.303)
Equity	0.54	−0.48[b]	−0.64	−1.64[b]	6.98
	(0.22)		(0.30)		(0.727)

[a]Hansen's J-statistic for the restrictions that one-week and longer-period return variances move in proportion.

[b]Significantly less than one at the 95% confidence level.

one-week and three-month holding periods. Denoting 'w' as the one-week horizon variance elasticity, we can test the hypotheses:

$$\text{Test 2b:} \quad (\eta_1^t/\eta_w^t) < 1,$$

$$\text{Test 2c:} \quad (\eta_3^t/\eta_w^t) < 1.$$

These ratios, as reported in columns 1 and 3, are estimated for the two portfolio sets with weekly returns, i.e., the 'Mixed Equity/Foreign Exchange' and the 'Foreign Exchange' sets. Interestingly, for the 'Mixed' set, conditional variance ratios relative to a one-week holding period appear to mirror the results relative to the one-month holding period. For example, as in table 5 for this portfolio, the variance elasticity ratios are significantly less than one for the British foreign exchange returns and the equity returns, but not the German foreign exchange returns. For the 'Foreign Exchange' set, weekly holding period data provide more efficient estimates for the French franc and German mark elasticities. For instance, these estimates indicate that the weekly German mark returns react significantly more to new information relative to both three-month and one-month return variances. Taking the three-month and the one-month horizon variance results in table 6 together, the conditional variance response for the weekly returns exceeded at least one of the longer-period variance returns for all cases except for the Japanese yen.

In summary, testing the conditional variance behavior of individual assets across holding periods suggests that investors facing new information would revise their short-term variance forecasts by more than in the longer term. This behavior is not sufficient to argue that conditional variances move idiosyncratically over short but not long holding periods, since the variances on all returns could potentially react in the same direction. Therefore, we will next consider the behavior of conditional variances *both* across returns and holding periods. In so doing, we will incorporate information from Test 1 and Test 2 to provide a more powerful test of the underlying hypothesis.

3.6. Across returns and holding periods: Test 3

Evidence from Test 1 suggested that three-month conditional variances move together so that $\theta_3^i = \theta_3^j$ for all i, j. If so, then the elasticity of the variances with respect to new information as defined above for Test 2 must be the same for all returns with three-month holding periods,

$$\left[\frac{\left(\partial\sigma_{3,t}^i/\sigma_{3,t}^i\right)}{\partial u_t}\right] \equiv \eta_3^i = \eta_3, \quad \forall\, i. \tag{17}$$

We can estimate the conditional variance elasticity ratios of all one-month returns individually relative to the joint three-month variance elasticity. That is, we can construct a joint equation version of Test 2 by estimating eq. (16) across all returns, given (17) as a constraint, and then test whether

$$\text{Test 3:} \quad \left(\eta_1^i/\eta_3\right) = \left(\eta_1^j/\eta_3\right), \quad \forall\, i, j.$$

Figs. 1 through 3 depict the results of this estimation for the 'Mixed Term Structure' and for the 'Foreign Exchange' portfolio sets using information

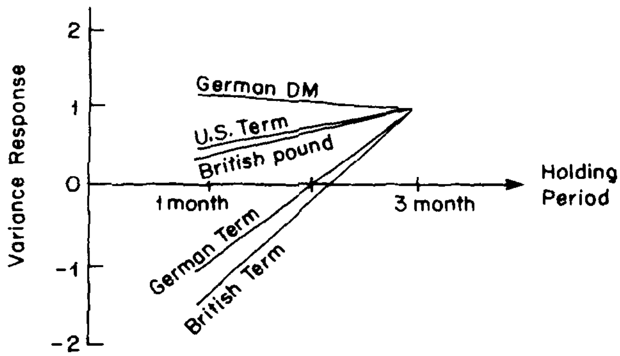

Fig. 1. Joint conditional variance response across term structure and foreign exchange.

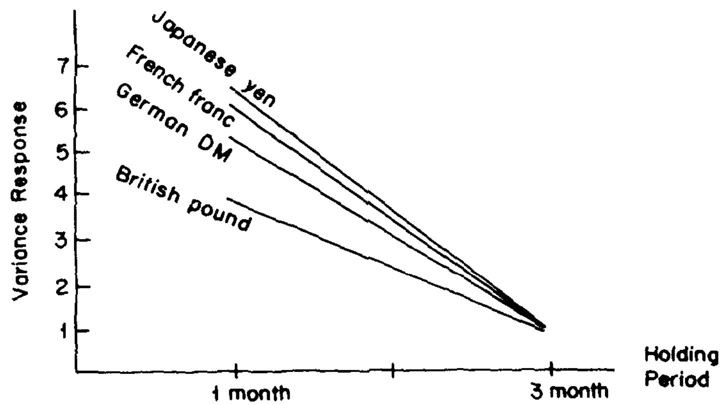

Fig. 2. Joint conditional variance response across one- and three-month foreign exchange.

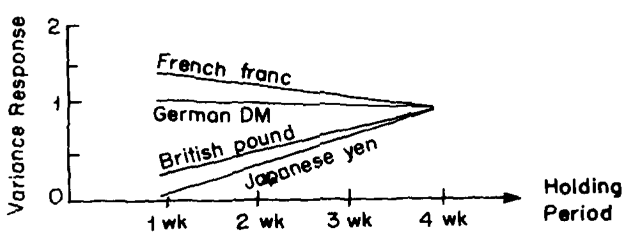

Fig. 3. Joint conditional variance response across one-week and one-month foreign exchange.

variables set A.[14] For the 'Term Structure/Foreign Exchange' portfolio set, fig. 1 illustrates how the one-month conditional variances for each return reacts to news that induces a 1% change across the conditional variances in all three-month returns. Clearly, even though this news affects the three-month variances in the same way, it induces very different reactions among the one-month returns, ranging from 1.17% for the German foreign exchange return to -1.43% for the British interest rate term structure return. For the Test 3 hypothesis that all five one-month elasticities are equal, the Wald test statistic had a marginal significance level below 0.1% indicating a much stronger rejection than found in table 4.[15]

The results of estimating these same relationships for the 'Foreign Exchange' portfolio sets are depicted in figs. 2 and 3 for the one- to three-month response and the one-week to one-month response, respectively. Although the point estimates of the conditional variance responses in fig. 2 are quite a bit higher than for three months and are generally far apart, they are also measured imprecisely. A Wald statistic for Test 3 only had a marginal significance level of 31%. However, this same relationship tested at the one-month to one-week horizons gave marginal significance levels less than 0.1%. The results of the estimation are given in fig. 3.

In summary, using more efficient methods, evidence from Test 3 strengthens the relationships found in Tests 1 and 2 above. Conditional variances appear to move together over longer holding periods in response to changes in current information about the state of the economy. But conditional variances react idiosyncratically over shorter holding periods in response to this same information.

4. Concluding remarks

This paper has analyzed whether the relationship among conditional covariances between asset returns and consumption can explain the tendency to reject latent variable models of the consumption-based asset pricing model. To explain this pattern, we must find these conditional covariances move idiosyncratically over short, but not longer holding periods. This relationship was tested by looking across returns such as equity, foreign exchange, and bonds, and across holding periods of one week, one month, and one quarter. Although the equity process appeared to be misspecified, we found that the pattern of co-movements in consumption covariances matched the pattern in most latent variable tests for foreign exchange and

[14]This smaller IV set was used to avoid overparameterizing the system since the system of equations is now larger. The variances were normalized by the three-month German mark foreign exchange returns in all cases.

[15]Furthermore, the overidentifying restrictions for this or any other portfolio were never rejected at the 95% confidence level.

bonds. Interestingly, this evidence suggests that rejections in the intertemporal consumption-based asset pricing relationship at short horizons depend upon the inadequacy of an auxiliary assumption, not necessarily upon the relationship itself.

Data appendix

Deposit rates for one-week, one-month, and three-month holding periods from the Eurocurrency market comprise the interest rate series. The spot exchange rates and the one-month and three-month deposit are from *Data Resources Incorporated*, while the one-week Eurocurrency deposit rates are from the *London Financial Times* and were provided by Philippe Jorion. The weekly returns on equity were calculated from the daily New York Stock Index at the University of Chicago *Center for Research in Securities and Prices* and were also provided by Philippe Jorion.

The information variables sets are described in table 1. Information variable set A only includes with a constant the forward premia plus the spread between current long and short rates for the term structure returns and the lagged equity returns for the equity portfolio set. In addition to these, information variable set B also includes the squares of these same variables in set A. Hodrick and Srivastava (1984), Giovannini and Jorion (1987), and Cumby (1988) find that the squares of the forward premia help explain the *ex ante* returns on foreign exchange. Instead of squared variables, information variable set C substitutes several real variables used in Cumby (1988). They are the monthly growth rates of consumption, the consumer price level, and industrial production, all lagged three months, plus the consumer price level lagged twelve months, and the current U.S. Terms of Trade. I am grateful to Bob Cumby for providing me with his data series, described more fully in Cumby (1988).

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