NBER WORKING PAPERS SERIES

SHOULD THE HOLDING PERIOD MATTER FOR THE INTERTEMPORAL CONSUMPTION-BASED CAPM?

Karen K. Lewis

Working Paper No. 3583

NATIONAL BUREAU OF ECONOMIC RESEARCH
1050 Massachusetts Avenue
Cambridge, MA 02138
January 1991

I am grateful for useful comments by seminar participants at the World Congress of the Econometric Society, the University of Rochester, Harvard University, the National Bureau of Economic Research, Columbia University, the Wharton School, University of Michigan, the New York Federal Reserve, and New York University. I also benefitted from useful conversations with Dave Backus, Bob Cumby, Martin Evans, Ken Froot, and David Hsieh. I am indebted to Alberto Giovannini and Philippe Jorion for providing the 7-day Eurocurrency deposit data, to Ross Levine and Steve Scott for assistance with the DRI data tape, and to Naveen Seth for superb research assistance. Research support from the Glucksman Institute, the Salomon Brothers Center for Financial Studies, the Olin Foundation, and the National Science Foundation under grant #89-02794 is thankfully acknowledged. Any errors are solely my responsibility. This paper is part of NBER's research program in Financial Markets and Monetary Economics. Any opinions expressed are those of the author and not those of the National Bureau of Economic Research.

NBER Working Paper #3583
January 1991

SHOULD THE HOLDING PERIOD MATTER FOR THE INTERTEMPORAL CONSUMPTION-BASED CAPM?

ABSTRACT

Empirical studies of the restrictions implied by the intertemporal capital asset pricing model across different asset markets have found conflicting evidence. In general, restrictions from this model have been rejected over short holding periods, but not over longer holding periods such as a quarter. This paper asks whether an auxiliary assumption implicit in these tests could be responsible for the observed pattern of rejections. The auxiliary assumption requires that covariances of returns with consumption move in constant proportion over time. The paper first describes how this condition may break down within the context of a general equilibrium pricing relationship. Then, the condition is tested empirically using data on foreign exchange, bonds, and equity returns. Interestingly, the pattern of consumption covariances in foreign exchange and bonds indeed match the pattern of rejection in the intertemporal asset pricing relationship.

Karen K. Lewis
Department of Economics
McNeil Hall
University of
Pennsylvania
3718 Locust Walk
Philadelphia, PA 19104

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 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 this 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 eachother. Intuitively, since each risky return depends proportionally upon its own conditional covariance with the MRS, then all returns will move in a constant proportion to eachother if the relationship between covariances are constant over time.

Therefore, 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

---

[1]For example, these restrictions have been rejected for one-month holding periods by Hodrick and Srivastava (1984) for foreign exchange and 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 (1990) provides a survey.

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 pattern is exactly the rejection pattern found in the literature.

This paper asks whether the auxiliary assumption can explain the observed pattern of rejecting the intertemporal capital asset pricing restrictions. For this purpose, the conditional covariances are evaluated 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 pattern of rejecting the restrictions on returns, we must find that conditional covariances move in proportion over long but not short holding periods. The results reported below indeed find this pattern for most returns.

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. In order to explain the empirical regularity on returns, we should find that the covariances of some returns move idio-syncratically 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. That is, given that covariances move in proportion over long but not short holding periods, we can incorporate this information to provide a more powerful test. This last test therefore incorporates information across both the holding period and returns.

The paper proceeds as follows. Section 2 discusses how the typical latent variable model restrictions implied by the intertemporal CAPM depend upon the auxiliary hypothesis. This section briefly reviews the restrictions, the empirical regularity, and the effects of time-variation in conditional covariances. It then describes intuitively how time-varying covariances may arise within a simple general equilibrium model. 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 eachother arises from intertemporal utility maximization with the added assumption that all conditional covariances of returns with the MRS move in a constant proportion. To see this result, consider a representative agent that maximizes expected time-additive utility,

$$U = E_t \sum_{j=0}^{\infty} \gamma^j u(c_{t+j}) \tag{1}$$

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,

$$1 = E_t\left[\frac{(\gamma^k u'(c_{t+k})/p_{t+k})}{(u'(c_t)/p_t)} (1 + r^i_{t,k}) \right] \tag{2}$$

where p_t is the price of the consumption good at time t, and $r^i_{t,k}$ is the nominal return on an asset purchased at time t with payoffs *k* periods ahead. The key ingredient in (2) that influences the tests developed below is that 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:

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

Note also that $n_{t,k} = \prod_{i=0}^{k-1} n_{t+i,1}$, a relationship that will prove expositionally useful below.

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

$$E_t(r^i_{t,k} - r^r_{t,k}) = - \operatorname{Cov}_t(n_{t,k},\ r^i_{t,k})\ (1 + r^r_{t,k}),\ \forall\ i \tag{3}$$

where Cov_t is the covariance operator conditional upon current information. Equation (3) describes the risk premia on asset *i* relative to the risk free rate. The risk premium is positive if the intertemporal marginal rate of substitution covaries negatively with the return. In this case, states with low realized returns are correlated with high marginal utility. In equilibrium, agents will willingly hold these assets only if they have higher ***ex ante*** expected returns.

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

$$\begin{aligned} &E_t(r^i_{t,k} - r^r_{t,k}) \\ &\qquad = [\operatorname{Cov}_t(n_{t,k},\ r^i_{t,k})/ \operatorname{Cov}_t(n_{t,k},\ r^j_{t,k})]\ E_t(r^j_{t,k} - r^r_{t,k}) \qquad (4) \\ &\qquad\qquad \forall\ i,j,\ i \neq j. \end{aligned}$$

In other words, since all returns with the same holding period depend upon their conditional covariances with the intertemporal marginal rate of substitution over that same holding period, they move in proportion to eachother 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. In other words,

$$\left[\frac{\beta^i}{\beta^j}\right] = \frac{Cov_t(n_{t,k},\ r^i_{t,k})}{Cov_t(n_{t,k},\ r^j_{t,k})} \tag{5}$$

where the β's are constants. Early studies simply assumed that the conditional covariances were constant.[2] More recently, researchers have noted that condition (5) will generally hold as long as covariances move in proportion.[3]

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

$$r^i_{t,k} - r^r_{t,k} = x_t{}' \, b^i + \epsilon^i_{t+k} \tag{6}$$

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:

---

[2]See, for example, Hansen and Hodrick (1983).

[3]See, for example, Cumby (1988).

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

We may then test these latent variable model restrictions in (7).

Indeed, 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 3 months.[4] Interestingly, whether the restrictions in (7) are rejected appears to be robust to the types of returns included, but 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 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. To illustrate how this situation may arise, we next consider a simple example using a standard general equilibrium pricing model.

2.3 A General Equilibrium Example

Studies in the literature have tested the restrictions in (7) for different asset market returns including equity, bonds, and foreign exchange markets. We will look at each of these assets in turn.

---

[4]See the references in footnote 1.

First, if we define the nominal price of equity at t as q_t and the dividend stream as y_t, the rate of return for the stock held over one period is,[5]

$$(1 + r^q_{t,1}) = (q_{t+1} + y_{t+1}\, p_{t+1})/q_t \tag{8}$$

since equity pays a dividend equal to output in all future states. But the first-order condition for a share of stock yielding output in all future states implies that,

$$q_t = \gamma \sum_{j=0}^{\infty} \gamma^j \, E_t \, (n_{t,j+1} \, y_{t+j+1}) \tag{9}$$

which clearly depends upon the stochastic process of the marginal rate of substitution, $n_{t,k}$, and income, y_t.

Moreover, the marginal rate of substitution is an endogenous variable that depends upon the exogenous state variables in the economy. Defining s_t as the vector of state variables known at time t, the marginal rate of substitution between consumption at t and $t+k$ depends upon the history of the state up to $t+k$. Therefore, the MRS has the functional form:

$$n_{t,k} = n_{t,k}(s_{t+k}, \ldots, s_t, \ldots). \tag{10}$$

For a concrete example, suppose that utility is of the CRRA form, $u(c_t) = c^{1-\alpha}/(1-\alpha)$. Furthermore, consider an economy as in Lucas (1982) where income and money follow first-order Markov processes given by $y_{t+1} = Y_{t+1}\, y_t$, $m_{t+1} = M_{t+1}\, m_t$, respectively. Equilibrium implies that the price is given by $p_t = (y_t/m_t)$ and

[5]For more discussion of this equity return in general equilibrium contexts, see Lucas (1982), Mehra and Prescott (1985), Giovannini (1989), and Labadie (1989).

the nominal one-period ahead MRS is,[6]

$$n_{t,1} = (Y_{t+1}^{1-\alpha}/M_{t+1}). \tag{11}$$

In this simple case, the state vector is just $s_t = (Y_t, M_t)$ and the nominal MRS is a function of only the realized state: $n_{t,1} = n_{t,1}(s_{t+1})$. Furthermore, expected future prices in this Markovian world depend only upon the current state, $E_t n_{t,k} = E_t n_{t,k}(s_t)$.

This particular relationship holds only under the special assumptions described above. More generally, the marginal rate of substitution is determined by realizations of the state process through potentially complicated functional forms as captured by the relationship in (10). Furthermore, since the rate of return depends upon the MRS which is a function of the state vector process, these returns are themselves functions of the state. That is, substituting the MRS in (10) into (9) and the resulting expression into (8), the one period return on equity has the form:

$$r_{t,1}^q = r_{t,1}^q(\ n_{1,t}(s_{t+1},\ s_t,\ ),\ y_{t+1}) = r_{t,1}^q(s_{t+1},\ s_t,\ \ldots).$$

Now we can see that the auxiliary assumption in (5) regarding the conditional covariances between equity returns and the intertemporal marginal rate of substitution places conditions on,

$$\mathrm{Cov}_t(n_{t,1},\ r_{t,1}^q)$$

$$= \mathrm{Cov}(\ n_{t,1}(s_{t+1},\ s_t,\ \ldots),\ r_{t,1}^q(\ s_{t+1},\ s_t,\ \ldots) \mid s_t,\ s_{t-1},\ \ldots\ ) \tag{12}$$

[6]These equilibrium conditions assume that the asset market opens before the goods market every period as in Lucas (1982) so that the liquidity constraint is always binding. If the goods market opened before the asset market as in Svensson (1985), the pricing equations would be altered. See Giovannini (1989).

In other words, the equilibrium conditional covariance between the MRS and the return on equity depends upon the current and past states of the economy. Similar reasoning for holding equity for $k > 1$ periods implies that the k period equity return depends upon,

$$\mathrm{Cov}_t(n_{t,k},\ r_{t,k}^{q})$$
$$= \mathrm{Cov}(\ n_{t,k}(s_{t+k},\ s_{t+k-1},\ \ldots),\ r_{t,k}^{q}(\ s_{t+k},\ s_{t+k-1},\ \ldots) \mid s_t,\ s_{t-1},\ \ldots\ ) \quad (13)$$

Overall, the conditional covariances between excess returns on equity and the MRS for the same holding period depend upon the current state variables in the economy.

Term premia in bond markets are a second type of return studied empirically in the literature. Consider the returns in successively rolling over short-term bonds in excess of a risk-free bond of longer maturity.[7] For example, suppose an investor holds a one-period bond and then reinvests the proceeds in another one-period bond next period. Denoting this strategy as "b", the returns 2 periods hence are:

$$(1 + r_{t,2}^{b}) = (1 + r_{t,1}^{rf})\ (1 + r_{t+1,1}^{rf}) = 1/(E_t(n_{t,1})\ E_{t+1}(n_{t+1,1})). \quad (14)$$

That is, the returns are functions of the expected future marginal rates of substitution. But note from equation (10) that these marginal rates of substitution depend upon the state of the economy. We can therefore write the returns from rolling over one period bonds for two periods as $r_{t,2}^{b} = r_{t,2}^{b}(s_{t+1},\ s_t,\ \ldots)$, or more generally, the returns from rolling over these bonds for k periods as $r_{t,k}^{b} = r_{t,k}^{b}(s_{t+k-1},\ s_{t+k-2},\ \ldots)$.

---

[7]This type of term premium has been analyzed by Campbell and Clarida (1987) and Lewis (1990). Another type of term premium studied in this context is the excess return on holding longer term bonds for one period, as in Campbell (1987).

Therefore, the auxiliary condition in (5) for rolling over short bonds for k periods implies restrictions on the covariance given by,

$$\begin{aligned} &Cov_t(n_{t,k},\ r_{t,k}{}^b) \\ &= Cov_t(n_{t,k}(s_{t+k},\ s_{t+k-1},\ \ldots),\ r_{t,k}{}^b(s_{t+k-1},\ s_{t+k-2},\ \ldots) \mid s_t,\ s_{t-1},\ \ldots) \quad (15) \end{aligned}$$

As in the case of equity, these conditional covariances depend in turn upon the current and past states of the economy.

The third type of returns studied in the literature corresponds to the foreign exchange risk premium. In particular, this return arises from buying and holding a foreign currency denominated bond and then converting the proceeds back into domestic currency at the prevailing spot rate in the future. In the context of the general equilibrium framework above, the utility function must be modified to include two goods, a domestic and foreign endowment.[8] If each good may be purchased only with its own currency, then the price of the foreign good is determined by the counterpart to domestic price above, i.e, $p^* = (y^*/m^*)$, where $*$ refer to foreign variables. In this framework, Hodrick and Srivastava (1986) show that the return on holdings of foreign bonds also depends upon the ratios of the intertemporal marginal rate of substitutions of domestic and foreign currency. Specifically, denoting this risky strategy as "f", these returns are:

$$r_{t,k}{}^f - r_{t,k}{}^r = \left[\frac{E_t n_{t,k}}{E_t n_{t,k}{}^*}\right] \left[\frac{p_t}{p_t{}^*}\right] (u_{y,t}/u_{y*,t}) \qquad (16)$$

$$- E_t[(u_{y,t+k}/u_{y*,t+k})(p_t/p_t{}^*)].$$

where $n_{t,k}{}^*$ is the intertemporal marginal rate of substitution in the foreign good and money, and $(u_{y,t}/u_{y*,t})$ is the contemporaneous marginal rate of

[8]See Lucas (1982) and Svensson (1985b).

substitution between the domestic and foreign good. In equilibrium, the marginal rate of substitution between the two goods depends upon the state vector. Also, as noted before, the prices and the intertemporal rate of substitution depend upon the state vector. Therefore, as with the return on bonds and equity, the return on foreign exchange may be written as a function of the state vector:

$$
\begin{aligned}
& r_{t,k}{}^{f} \\
& = r_{t,k}{}^{f}(n_{t,k}(s_{t+k},\ s_{t+k-1},\ \ldots),\ n_{t,k}{}^{*}(s_{t+k},\ s_{t+k-1},\ \ldots),\ s_{t+k-1},\ s_{t+k-2},\ \ldots) \\
& \qquad = r_{t,k}{}^{f}(s_{t+k},\ s_{t+k-1},\ \ldots).
\end{aligned}
$$

For this reason, the conditional covariance between the foreign exchange return and the intertemporal rate of substitution has a form analogous to (13):

$$
\begin{aligned}
& \mathrm{Cov}_t(n_{t,k},\ r_{t,k}{}^{f}) \\
& = \mathrm{Cov}(\ n_{t,k}(s_{t+k},\ s_{t+k-1},\ \ldots),\ r_{t,k}{}^{f}(\ s_{t+k},\ s_{t+k-1},\ \ldots) \mid s_t,\ s_{t-1},\ \ldots\ ) \quad (17)
\end{aligned}
$$

In summary, the returns on all risky returns held over the same period depend upon the expected covariation of these returns with the intertemporal marginal rate of substitution in consumption. In turn, this expected covariation depends upon the current and past history of the economy.

2.3 Interpreting the Auxiliary Assumption of Conditional Covariances

As the example above illustrates, the conditional covariances of returns with the marginal rate of substitution depend upon the state vector of the economy. On the other hand, the auxiliary assumption in (5) requires that all covariances move in proportion over time. Given these relationships, we can now readily see how this auxiliary condition may break down.

To begin with, note that (5) will hold as long as the covariances are constant. Therefore, a violation of (5) requires conditional heteroscedasticity in the joint process of rates and the intertemporal MRS. Where does this heteroscedasticity come from? The covariances in (13), (15), and (17) make clear that 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. As an indication that the variance of the marginal rate of substitution may vary over time, Kandel and Stambaugh (1990) find that aggregate consumption growth displays heteroscedasticity, a result we will also find below. Second, since the returns can in general be complicated non-linear functions of the state variables, the functional form may induce heteroscedasticity in returns even if the state variables were homoscedastic.[9] This possibility has recently been discussed in studies of non-linearities in asset prices such as Scheinkman and LeBaron (1989) and Hsieh (1989).

As an empirical matter, variances that change with new information about the economic state has been found in many types of asset returns. For example, 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 (1989) for long bond returns.

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? For this purpose, recall that condition (5) requires the covariances of all returns to move in proportion over time. Therefore, the covariances of the equity premium in (13), the term premia in (15), and the foreign exchange premium in (17) must move together. On the other hand, if the underlying heteroscedasticity affects

[9]Hsieh (1990) provides an example when an asset price depends non-linearly upon state variables that are themselves i.i.d. The non-linearity of the functional form causes conditional heteroscedasticity in the asset price even though the underlying state process is homoscedastic.

the returns differently, conditional covariances will not move together. 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 non-linearity of returns. Therefore, we next ask whether a break-down in this auxiliary condition can explain the pattern of rejecting the latent variable model at shorter 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 holds 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. Table 1 in Part I defines these three types of returns all in excess of the risk-free rate. They are first, "foreign exchange returns" corresponding to equation (16), the returns from holding open positions in foreign currency deposits; second, the "term structure returns" from rolling over short rates as in equation (14) for two-periods; and third, the "equity returns" from holding equity, receiving dividends and capital gains, as in equation (8). Furthermore, 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 will begin by estimating equations (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^i, and, second, a set of some information variables, x_t, currently known by market traders. Table 1 under Part II defines the composition in the empirical estimation of both the portfolio of returns considered jointly and the set of information variables. In forming the portfolio and information sets, I attempted 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.[10] 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.[11] The information variables sets are also listed in Table 1 in part II. 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

---

[10]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 (1990).

[11]Foriegn 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 a portfolio set of one-week U.S. equity and foreign exchange returns, similar to the "Mixed Equity" set above.

Tests of the restrictions in (7) based upon estimating the projection equations (6) provide different results depending upon the holding period of the returns, k. Table 2 provides an example of this basic finding using the instrumental variable set A with weekly frequency data. 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.

A striking pattern emerges in these results: as the holding period shortens, the marginal significance levels of the latent variable hypothesis declines. As discussed and surveyed in Lewis (1990), this pattern is very robust to many factors including frequency of data, type of returns included in the portfolio, information variables set, and estimation method.

3.3 Ex Ante Returns and Conditional Covariances

As discussed in Section 2, the validity of the latent variable model as a test of intertemporal asset pricing relationships depends upon whether the consumption covariances with returns move in proportion over time. For this condition to be violated, these covariances must move over time. If so, the conditional second moments of returns and consumption depend upon current and lagged state variables, just as the conditional first moments do in the standard latent variable model. Therefore, to analyze the behavior of the conditional covariances, we will examine how they respond to changes in the current information set.

For this purpose, we will rewrite the equations (6), as:[12]

---

[12]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.

$$r^i_{t,k} - r^{rf}_{t,k} = x_t' \, b^i + \epsilon^i_{t+k} \qquad (6')$$

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

$$(\epsilon_{t,k}{}^i)^2 = (\sigma_{t,k}{}^i)^2 \exp[-(1/2)\mathrm{Var}(\nu^i) + \nu_{t+k}{}^i) \qquad (18)$$

where for non-overlapping forecast horizons, ν_t is an i.i.d. normally distributed random variable with variance $\mathrm{Var}(\nu)$.

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:

$$(\sigma_{t,k}{}^i)^2 = \delta_k{}^i \exp[\, z_t \, \theta_k{}^i - (1/2)\mathrm{Var}(w_k{}^i) + w_{t,k}{}^i) \qquad (19)$$

where $w_{t,k}$ is the error in measuring conditional variances by the econometrician and is normally distributed with variance, $\mathrm{Var}(w_k)$. Under these conditions and some standard regularity conditions, the conditional variance parameters, θ, can be estimated by OLS in the following regression:[13]

$$\log(\hat{\epsilon}_{t,k}{}^2) = -(1/2)[\mathrm{Var}(w_k) + \mathrm{Var}(\nu_k)] + \log(\delta_k) + z_t\theta_k + e_{t,k} \qquad (20)$$

where $e_{t,k} = w_{t,k} + \nu_{t+k}$ and where the superscript *i* has been suppressed for

---

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

notational simplicity. Note that the variances 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 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′). Based upon the regressions in equation (20), the sixth column reports the marginal significance levels for the Wald test statistic of the hypothesis that the θ coefficients are jointly zero. As the results indicate, asset returns appear to display considerable heteroscedasticity. These same test statistics using weekly frequency data provide even stronger evidence for 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:

$$(\Delta^k c_t/c_t) = x_t\, b_k{}^c + \epsilon_{t,k}{}^c \qquad (21)$$

where Δ^k is the forward difference operator k periods ahead, and $\epsilon_{t,k}{}^c$ is the residual to the consumption projection equation. Analogous to the equation (18), the cross products of the errors to returns and consumption depend upon the conditional covariance between consumption and asset returns according to,

$$\epsilon_{t,k}{}^i\, \epsilon_{t,k}{}^c = \sigma_{t,k}{}^{ic} \exp[-(1/2)\ \mathrm{Var}(\nu^{ic}) + \nu_{t+k}{}^{ic}] \qquad (22)$$

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,

$$\sigma_{t,k}^{i,c} = a_k^i \, (\sigma_{t,k}^i)^2 \tag{23}$$

where a_k^i is a constant. In other words, 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 equation (23) into equation (22) and taking the logarithm implies,

$$\log(\epsilon_{t,k}^i \, \epsilon_{t,k}^c) = \log(a_k^i) + \log(\delta_k^i) - (1/2)\mathrm{Var}(\nu^{ic}) - (1/2)\mathrm{Var}(w_k^i)$$

$$+ z_t \, \theta_k^i + e_{t,k}^{ic} \tag{24}$$

where $e_{t,k}^{ic} = w_{t,k}^i + \nu_{t+k}^{ic}$. In this form, we can directly evaluate the covariance restriction (23) by estimating equation (24) together with equation (20) 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 non-overlapping one-month returns.[14] As the table demonstrates, 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 case.

Therefore, 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 (23) holds so that $\sigma_t^{i,c} \propto (\sigma_{t,k}^i)^2$.

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 latent variable model if the conditional covariances of the MRS with returns 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.

---

[14]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).

3.4 Across Returns at Individual Holding Periods: Test 1

Using the process for variances together with the condition (23) tested above, we are now in a position to directly test the auxiliary assumption to latent variable tests given in (7). For this purpose, rewrite equation (19) given holding period k for assets i and j as:

$$(\sigma_{t,k}^{i})^2 = \delta_k^{i} \exp[\ z_t\ \theta_k^{i} - (1/2)\mathrm{Var}(w_k^{i}) + w_{t,k}^{i}) \qquad (25)$$
$$(\sigma_{t,k}^{j})^2 = \delta_k^{j} \exp[\ z_t\ \theta_k^{j} - (1/2)\mathrm{Var}(w_k^{j}) + w_{t,k}^{j}).$$

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 equations (19) across assets and testing the cross-equation restriction:

$$\text{Test 1: } \theta_k^{i} = \theta_k^{j},\ \forall\ i,j.$$

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 (19) 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 over-identifying 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 either 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 suggests that the larger equation system may be over-parameterized.

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 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.[15]

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

---

[15]Specifically, the conditional variance process in (18) requires that non-overlapping innovations ν_t be i.i.d. This hypothesis was tested using the "I-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 6 lags.

periods 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.[16] 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:

$$u_t = z_t \phi$$

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:

$$\left[\frac{(\partial\sigma^i_{k,t}/\sigma^i_{k,t})'}{\partial u_t}\right] = \eta_k^{\,i}$$

Rewriting the variance process in (19) in terms of the unit news variable over different holding periods such as $k = 1, 3$, yields:

---

[16]As described in the general equilibrium example in Section 2.3, this idiosyncratic variance behavior can occur since returns are different non-linear functions of the state process. Therefore, the variances of different types of returns would likely respond differently to changes in the state process regardless of whether the state process is itself heteroscedastic or the functional form of returns induces heteroscedasticity.

$$(\sigma_{t,1}^{i})^2 = \delta_1^{i} \exp[\eta_1^{i} u_t - (1/2)\mathrm{Var}(w_1^{i}) + w_{t,1}^{i}] \tag{26}$$
$$(\sigma_{t,3}^{i})^2 = \delta_3^{i} \exp[\eta_3^{i} u_t - (1/2)\mathrm{Var}(w_3^{i}) + w_{t,3}^{i}].$$

Now we may evaluate the cross-maturity elasticities for each individual asset return *i* directly by estimating equations (20) jointly for asset *i* at maturities $k = 1, 3$ and constraining the new information, u_t, to be the same across equations as in (26). 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. Therefore, we may test this hypothesis for 3-month relative to 1-month returns as:

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

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 $u_t = z_t \theta_1^{i}$. 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 over-identifying restrictions.[17]

---

[17]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, 2δ parameters and 1 conditional variance elasticity ratio), so that there remain N-2 over-

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 over-identifying 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 equation (19) jointly across one-week and one-month holding periods, and then across one-week and three-month holding periods. Denoting "w" as the one week horizon variance elasticity, we can test the hypotheses:

Test 2b: $(\eta_1^i/\eta_w^i) < 1$

Test 2c: $(\eta_3^i/\eta_w^i) < 1$.

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

---

identifying restrictions.

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 variances 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 3 month holding periods. That is,

$$\left[\frac{(\partial\sigma^i_{3,t}/\sigma^i_{3,t})}{\partial u_t}\right] = \eta_3^i = \eta_3, \qquad \forall i.$$

We can now 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 equations (26) across all returns, given (27) as a constraint. With these estimates, we can test whether,

Test 3: $(\eta_1^i/\eta_3) = (\eta_1^j/\eta_3)\ \forall\ i,\ j.$

Figures 1 through 3 depict the results of this estimation for the "Mixed Term Structure" and for the "Foreign Exchange" portfolio sets using information variables set A.[18] For the term structure/foreign exchange portfolio set, Figure 1 illustrates how the one-month conditional variances for each return reacts to news that induces a one percent 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 causes very different across the one-mont returns, ranging from 1.17 percent for the German foreign exchange return to -1.43 percent 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 .1% indicating a much stronger rejection than found in Table 4.[19]

The results of estimating these same relationships for the "Foreign Exchange" portfolio sets are depicted in Figures 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 Figure 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 .1%. The results of the estimation are given in Figure 3.

---

[18]This smaller IV set was used to avoid over-parameterizing 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.

[19]Furthermore, the over-identifying restrictions for this or any other portfolio were never rejected at the 95% confidence level.

In summary, using more efficient methods, evidence from Test 3 strengthen 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.

IV. Concluding Remarks

This paper has analyzed whether the relationship across holding periods of 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 that conditional covariances of the marginal rate of substitution in consumption and shorter-term returns move idio-syncratically. This type of behavior can arise from a general equilibrium relationship when either the state process or the function form of returns induce conditional heteroscedasticity. We must also find that these same conditional covariances move in proportion across returns over longer holding periods to explain the pattern.

This relationship was tested by looking both across returns such as equity, foreign exchange, and bonds, and across holding periods of one week, one month, and one quarter. Interestingly, although the equity process appeared to be misspecified, we found that the pattern of co-movements in consumption covariances matched the pattern in latent variable tests for foreign exchange and bonds. Therefore, this evidence suggests that the tendency to reject the intertemporal consumption-based asset pricing relationship at short horizons depends upon the inadequacy of an auxiliary assumption, not 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 deposti 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 Securuties 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).

References

Campbell, J. Y., 1987, Stock returns and the term structure, Journal of Financial Economics 18, 373-399.

Campbell, J. Y., and R. H. Clarida, 1987, The term structure of Euromarket interest rates: An empirical investigation, Journal of Monetary Economics 19, 25-44.

Christie, A., 1982, The stochastic behavior of common stock variances, Journal of Financial Economics 10, 407-432.

Cumby, R. E., 1990, Consumption risk and international equity returns: Some empirical evidence, Journal of International Money and Finance 9, 182-192.

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

Cumby, R. E., and J. Huizinga, 1990, Testing the autocorrelation structure of disturbances in ordinary least squares and instrumental variables regressions, National Bureau of Economic Research Technical Working Paper Series No. 92.

Cumby, R. E., and J. Huizinga, 1988, Two-step two-stage least squares user's guide, (New York University, New York City, New York).

Cumby, R. E., and M. 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, Illinois).

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

Evans, M. D. D., 1990, Interpreting the term structure using the intertemporal capital asset pricing model: An application of the non-linear ARCH-M model, Review of Financial Studies, forthcoming.

French, K. R., G. W. Schwert, and R. F. Stambaugh, 1987, Expected stock returns and volatility, Journal of Political Economy 19, 3-29.

Giovannini, A., 1989, Uncertainty and liquidity, Journal of Monetary Economics 23, 239-258.

Giovannini, A., and P. 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.

Hansen, L. P., 1982, Large sample properties of generalized method of moments estimators, Econometrica 10, 1029-1054.

Hansen, L. P., and R. J. Hodrick, 1983, Risk averse speculation in the forward foreign exchange market: An econometric analysis of linear models, in: J. A. Frenkel, ed., Exchange rates and international macroeconomics (University of Chicago Press, Chicago, Illinois).

Hodrick, R. J., and S. Srivastava, 1986, The covariation of risk premiums and expected future spot exchange rates, Journal of International Money and Finance 5, S5-S22.

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

Hsieh, D. A., 1990, A nonlinear stochastic rational expectations model of exchange rates, Working paper (Duke University, Durham, North Carolina).

Hsieh, D. A., 1989, Testing for nonlinear dependence in daily foreign exchange rates, Journal of Business 62, 339-368.

Kandel, S., and R. F. Stambaugh, 1990, Expectations and volatility of consumption and asset returns, The Review of Financial Studies 3, 207-232.

Labadie, P., 1989, Stochastic inflation and the equity premium, Journal of Monetary Economics 24, 277-298.

Lewis, K. K., 1990, The behavior of Eurocurrency returns across different holding periods and monetary regimes, Journal of Finance 45, 1211-1236.

Lucas, R. E., Jr., 1982, Interest rates and currency prices in a two-country world, Journal of Monetary Economics 10, 336-360.

Mehra, R. and E. Prescott, 1985, The equity premium: A puzzle, Journal of Monetary Economics 15, 145-162.

Poterba, J. M., and L. H. Summers, 1986, The persistence of volatility and stock market fluctuations, American Economic Review 76, 1142-1151.

Scheinkman, J., and B. LeBaron, 1989, Nonlinear dynamics and stock returns, Journal of Business 62, 311-337.

Schwert, G. W., and P. J. Seguin, 1989, Heteroskedasticity in stock returns, National Bureau of Economic Research Working Paper No. 2956.

Svensson, L. E. O., 1985a, Currency prices, terms of trade and interest rates: A general equilibrium asset-pricing cash-in-advance approach, Journal of International Economics 15, 17-41.

Svensson, L. E. O., 1985b, Money and asset prices in a cash-in-advance economy, Journal of Political Economy 93, 917-944.

Table 1

Summary of Variables and Portfolios Used in Reported Results

I. Definition of Returns for k Month Holding Period Returns

A. **"Foreign Exchange Returns"**: e.g., for German Deutschemark,

$$r_{k,t}^{DM} = (1200/k)(s_{t+k} - s_t) + \bar{r}_{k,t}^{DM} - \bar{r}_{k,t}^{\$}$$

B. **"Term Structure Returns"**[b]

$$r_{3,t}^{i} = (1/3) \sum_{j=0}^{3} \bar{r}_{1,t+j}^{i} - \bar{r}_{3,t}$$

$$r_{1,t}^{i} = (1/4) \sum_{j=0}^{3} \bar{r}_{.25,t+j} - \bar{r}_{1,t}$$

C. "Equity Returns" $r_{k,t} = A_k\,((P_{t+k} + D_{t,k})/P_t) - r_{k,t}^{\$}$

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"[c]

Holding Periods: 1 Week (7 days), 1 Month (28 days), 3 Months (91 days).

Returns: Foreign exchange returns for German Deutschemark, British Pound, French Franc, Japanese Yen.

Information Variables Set A: Current 1-month ahead forward premium for each of the four currencies against the dollar

Portfolio Set 2: "Mixed Term Structure/Foreign Exchange"

Holding Periods: 1 Month (28 days), 3 Months (91 days).

Returns: Foreign exchange returns for German Deutschemark, British Pound and interest rate term structure returns for U.S. dollar, German Deutschemark, and British Pound.

Information Variables Set A: Current 1-month ahead forward premium for German DM and the British pound against the dollar and current spread between 1-month and 1-week Eurocurrency deposits for the dollar, the DM, and the pound.

Portfolio Set 3: "Mixed Equity/Foreign Exchange"

Holding Periods: 1 Week (7 days), 1 Month (28 days), 3 Months (91 days).

Returns: Foreign exchange returns for German Deutschemark, British Pound and equity returns measured from the New York Stock Index.

Information Variables Set A: Current 1-month ahead forward premium for German DM and the British pound against the dollar and the previous monthly return on the New York Stock Index.

[a] s_t^{DM} is the spot price of 1 DM in terms of dollars at time t, $\bar{r}_{k,t}^{i}$ are the annualized k month Eurocurrency deposit rates in currency i, and A_k is an annualization factor: $A_t = 100x(365/N)$ where N is the number of days in the holding period. [b] $\bar{r}_{k,t}^{i}$ is the k-period Eurocurrency deposit denominated in currency i with maturity k. [c] Begins in October 1979

Table 2

Proportionality Test of the Single Beta Asset Pricing Model

Using Weekly Frequency Data

Portfolio Set 1: Foriegn Exchange

Holding Periods	*Jan. 76 - May 86*	*Oct. 79 - May 86*
	$\chi^2(12)$	$\chi^2(12)$
A. Three Month	NA	11.76 (.465)
B. One Month	NA	21.78 (.040)
C. One Week	NA	80.53 (<.000)

Portfolio Set 2: Mixed Term Structure/Foriegn Exchange

Holding Periods	*Jan 76-May 86*	*Oct 79-May 86*
	$\chi^2(20)$	$\chi^2(20)$
A. Three Months	15.76 (.731)	12.47 (.899)
B. One Month	39.27 (.006)	12.25 (.907)

Portfolio Set 3:Mixed Equity /Foriegn Exchange Returns

Holding Periods	*Jan 76 -May 86*	*Oct 79 -May 86*
	$\chi^2(6)$	$\chi^2(6)$
A. Three Months	4.62 (.593)	8.53 (.202)
B. One Month	6.31 (.389)	11.46 (.075)
C. One Week	20.47 (.002)	27.94 (<.000)

Notes: All variables, portfolio sets and information variables (I.V.) sets are defined in Table 1.

Table 3

Tests of Variance and Covariance Behavior
Using Monthly Frequency Data

Portfolio Set: Mixed Term Structure/Foreign Exchange
January 1976 - May 1986

Returns	*(1) Mean* $log(\hat{\varepsilon}^l \hat{\varepsilon}^c)$	*(2) S.D.* $log(\hat{\varepsilon}^l \hat{\varepsilon}^c)$	*(3) Ho: Const.* σ^{lc} *MSL*	*(4) Mean* $log(\hat{\varepsilon}^2)$	*(5) S.D.* $log(\hat{\varepsilon}^2)$	*(6) Ho: Const.* σ^l *MSL*	*(7) Ho: Const.* (σ^{lc}/σ^l) *MSL*
A. 3 Month Returns[a]							
Foreign Exchange:							
German DM	2.69	1.51	<.000	4.99	2.16	<.000	.892
UK pound	2.64	1.70	<.000	4.88	2.16	.049	.924
Term Structure:							
US Int. Rate	-1.44	1.71	<.000	-3.27	2.56	<.000	.869
German Int. Rate	-2.12	1.51	<.000	-4.64	2.04	<.000	.877
British Int. Rate	-1.42	1.59	<.000	-3.24	2.33	<.000	.990
B. 1 Month Returns							
Foreign Exchange:							
German DM	3.55	1.77	.185	5.60	2.29	.163	.495
UK pound	3.65	1.88	.211	5.80	2.47	.693	.373
Term Structure:							
US Int. Rate	-1.09	1.88	.030	-3.69	2.49	<.000	.104
German Int. Rate	-1.40	1.81	.001	-4.29	2.50	<.000	.112
British Int. Rate	-0.51	1.63	.697	-2.53	2.35	<.000	.073

Portfolio Set: Mixed Equity/Foreign Exchange
January 1976 - May 1986

Returns	*Mean* $log(\hat{\varepsilon}^l \hat{\varepsilon}^c)$	*S.D.* $log(\hat{\varepsilon}^l \hat{\varepsilon}^c)$	*Ho: Const.* σ^{lc} *M.S.L.*	*Mean* $log(\hat{\varepsilon}^2)$	*S.D.* $log(\hat{\varepsilon}^2)$	*Ho: Const.* σ^l *M.S.L.*	*Ho: Const.* (σ^{lc}/σ^l) *M.S.L.*
A. 3 Month Returns[a]							
Foreign Exchange:							
German DM	2.61	1.67	.147	4.95	2.32	.111	.225
UK pound	2.51	1.69	.062	4.75	2.40	.001	.251
US Equity	2.89	1.56	.217	5.51	2.15	.447	.255
B. 1 Month Returns							
Foreign Exchange:							
German DM	3.63	1.63	.007	5.68	2.22	.323	.899
UK pound	3.63	1.80	.013	5.68	2.61	.385	.999
US Equity	4.10	1.46	.856	6.63	2.04	.429	.293

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

Table 4

Test of Proportional Time Variation in Conditional Variances

Using Weekly Frequency Data

Portfolio Set 1: Foriegn Exchange

Holding Periods	*Jan. 76 - May 86*	*Oct. 79 - May 86*
	$\chi^2(24)$	$\chi^2(24)$
A. Three Month	NA	12.30 (.999)
B. One Month	NA	25.77 (.365)
C. One Week	NA	34.58 (.075)

Portfolio Set 2: Mixed Term Structure/Foriegn Exchange

Holding Periods	*Jan 76- May 86*	*Oct 79- May 86*
	$\chi^2(40)$	$\chi^2(40)$
A. Three Months	20.56 (.995)	13.36 (.999)
B. One Month	22.86 (.986)	13.39 (.999)
On Subset:		
	$\chi^2(20)$	$\chi^2(20)$
A. Three Months	33.12 (.033)	11.57 (.930)
B. One Month	33.84 (.027)	28.95 (.089)

Portfolio Set 3:Mixed Equity /Foriegn Exchange Returns

Holding Periods	*Jan 76 -May 86*	*Oct 79 -May 86*
	$\chi^2(12)$	$\chi^2(12)$
A. Three Months	23.46 (.024)	21.98 (.038)
B. One Month	9.60 (.651)	8.23 (.767)
C. One Week	29.56 (.002)	23.33 (.025)

Notes: All variables, portfolio sets and information variables (I.V.) sets are defined in Table 1.

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.)*	*Ho:* $\eta_{3m} < \eta_{1m}$ *t-stat.*	*Test of Restrict.*	*10/79 - 5/86* (η_{3m}/η_{1m}) *(S.E.)*	*Ho:* $\eta_{3m} < \eta_{1m}$ *t-stat.*	*Test of Restrict.*
1. Foreign Exchange Set						$\chi^2(7)$
German DM	NA	NA	NA	3.82 (2.87)	0.98	8.99 (.253)
UK pound	NA	NA	NA	2.90 (1.54)	1.23	12.00 (.101)
Japanese yen	NA	NA	NA	1.21 (0.18)	1.13	5.98 (.542)
French franc	NA	NA	NA	0.48 (0.13)	-3.90*	8.30 (.307)
2. Mixed Term/Foriegn Exchange						
Foreign Exchange:			$\chi^2(9)$			$\chi^2(9)$
German DM	0.69 (0.15)	-1.99*	9.18 (.421)	-0.30 (0.31)	-4.18*	7.83 (.551)
UK pound	-1.26 (0.55)	-4.12*	7.97 (.537)	-0.48 (0.37)	-4.00*	8.49 (.486)
Term Structure:						
U.S.	0.63 (0.10)	-3.82*	2.80 (.999)	0.99 (0.15)	-0.01	9.61 (.383)
German	0.74 (0.09)	-2.80*	23.21 (.006)	1.36 (0.25)	1.44	16.75 (.053)
UK	0.26 (0.13)	-5.85*	7.39 (.597)	0.81 (0.15)	-1.27	14.04 (.121)
3. Mixed Equity/Foriegn Exchange						
			$\chi^2(5)$			$\chi^2(5)$
German DM	1.68 (0.50)	1.36	5.64 (.343)	-2.60 (2.06)	-1.75*	4.64 (.475)
British pound	-2.51 (1.39)	-2.53*	9.41 (.094)	3.40 (2.74)	.88	6.35 (.274)
Equity	-1.44 (0.99)	-2.46*	3.29 (.655)	0.84 (0.27)	-.582	7.32 (.198)

Notes: *Test of Restrictions* is Hansen's J-Statistic of the over-identifying restrictions that one month and three month return variances move in the same proportion. * indicates significantly less than one at the 95% confidence level.

Table 6

Ratios of Conditional Variance Elasticities Over One Week Elasticities
Using Weekly Frequency Data

Portfolio Set 1: Foreign Exchange
Oct. 1979 - May 1986

Returns	(η_{1m}/η_{1w}) *(S.E.)*	*Ho:* $(\eta_{3m}/\eta_{1w})<1$ *t-stat.*	(η_{1m}/η_{1w}) *(S.E.)*	*Ho:* $(\eta_{1m}/\eta_{1w})<1$ *t-stat.*	*Test of Restrict. (M.S.L.)*
					$\chi^2(14)$
German	-2.91 (1.03)	-3.80[a]	-0.13 (0.34)	-3.32[a]	8.10 (.884)
UK	4.25 (2.49)	1.31	2.00 (0.91)	1.09	8.59 (.856)
Japan	2.11 (0.48)	2.31	1.79 (.47)	1.68	7.30 (.923)
France	-0.27 (0.26)	-4.88[a]	0.99 (0.28)	-0.04	10.06 (.758)

Portfolio Set 3: Mixed Equity/Foreign Exchange
January 1976 - May 1986

Returns	(η_{1m}/η_{1w}) *(S.E.)*	*Ho:* $(\eta_{3m}/\eta_{1w})<1$ *t-stat.*	(η_{1m}/η_{1w}) *(S.E.)*	*Ho:* $(\eta_{1m}/\eta_{1w})<1$ *t-stat.*	*Test of Restrict. (M.S.L.)*
					$\chi^2(10)$
German DM	.76 (.27)	-.88	.73 (.23)	-1.15	22.20 (.014)
UK pound	-.37 (.19)	-7.24[a]	.58 (.15)	-2.74[a]	18.60 (.046)
Equity	-.64 (.30)	-5.43[a]	.54 (.22)	-2.08[a]	6.98 (.727)

Portfolio Set 3: Mixed Equity/Foreign Exchange
October 1979 - May 1986

Returns	(η_{1m}/η_{1w}) *(S.E.)*	*Ho:* $(\eta_{3m}/\eta_{1w})<1$ *t-stat.*	(η_{1m}/η_{1w}) *(S.E.)*	*Ho:* $(\eta_{1m}/\eta_{1w})<1$ *t-stat.*	*Test of Restrict. (M.S.L.)*
					$\chi^2(10)$
German DM	-.70 (.30)	-5.62[a]	.41 (.24)	-5.87[a]	15.95 (.101)
UK pound	.74 (.29)	-.88	.03 (.25)	-3.90[a]	11.74 (.303)
Equity	.54 (.22)	-.48[a]	-.64 (.30)	-1.64[a]	6.98 (.727)

Notes: *Test of Restrictions* is Hansen's J-Statistic for the restrictions that one-week and longer period return variances move in proportion. [a] indicates significantly less than one at the 95% confidence level.

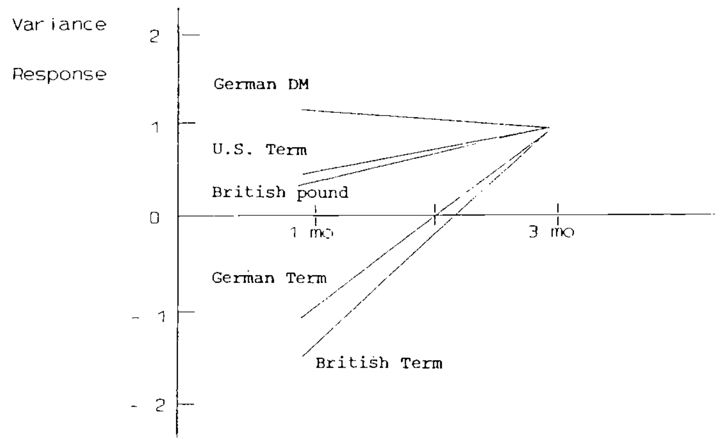

Fig 1--Joint Conditional Variance Response Across Term Structure and Foreign Exchange

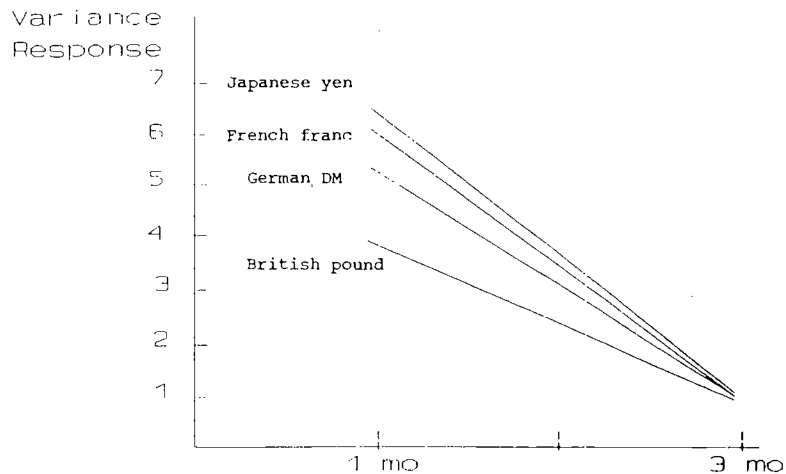

Fig. 2--Joint Variance Response Across One and Three Month Foreign Exchange

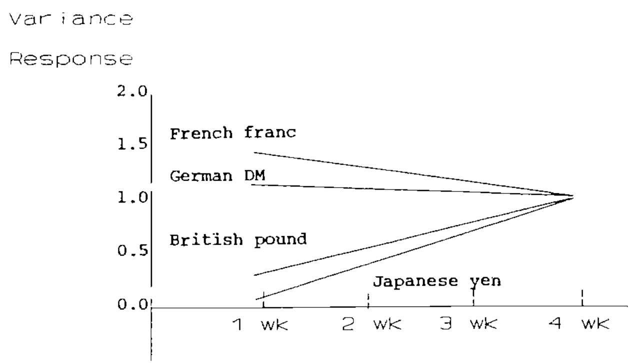

Fig 3.--Joint Conditional Variance Response Across One Week and One Month For. Exchange