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ABSTRACT

Departures from purchasing power parity imply that different countries have different prices for goods when a common numeraire is used. Stochastic changes in exchange rates are associated with changes in these prices and constitute additional sources of risk in asset pricing models. This article investigates whether exchange rate risks are priced in international asset markets using a conditional approach that allows for time variation in the rewards for exchange rate risk. The results for equities and currencies of the world's four largest equity markets support the existence of foreign exchange risk premia.

IN THE ASSET PRICING MODELS (APM) of Solnik (1974), Sercu (1980), Stulz (1981), and Adler and Dumas (1983), exchange rate risk is priced. Investors of different countries face different prices of goods at which they consume the income from their investments. In such a setting, the model contains risk premia that are based on the covariances of assets with exchange rates, in addition to the traditional premium based on the covariance with the market portfolio. These new premia are present because of deviations from purchasing-power parity (PPP).

We call "international" an APM that contains additional terms to reward exchange-rate risk, while we call "classic" an APM that does not contain such terms and in which there is only one risk premium based on the covariance of

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the asset return with the market portfolio.[1] In the presence of exact PPP (and no barriers to international investment), the classic APM would hold internationally. The main goal of the present article is to discriminate empirically between the two models and to test the null hypothesis that exchange-rate risk receives a zero price, against the international APM alternative.

Attempts have been made in the past at testing the international APM; see Solnik (1974) and Stehle (1977). They were based on the "unconditional" version of the international APM. These attempts, which will be replicated below, have been inconclusive.[2]

It is natural to test any APM in its "conditional" form. It is bad enough that the econometrician does not know what investors know and, therefore, *must* deal with a partially faulty APM. One would not want to carry the defect further by ignoring the conditioning information that is obviously available to investors, such as interest rates, equity prices, etc. These must appear in empirical tests in the form of instrumental variables.

Many methods have been proposed to test APMs in their conditional form. Of necessity, they rely on extraneous assumptions concerning the behavior of the moments of the probability distribution of returns in relation to the instrumental variables, and/or the behavior of the market prices of risk. Most methods constrain the behavior of second moments.[3]

An alternative is to constrain the behavior of the market prices of risk.[4] In the present article, we constrain these prices to be linear functions of the

[1] For a recent review and classification of international APMs, see Stulz (1992).

[2] Solnik (1977) expresses doubts that portfolio efficiency tests of the market portfolio of equities would discriminate between the classic and the international APM. Here we take a different approach and focus on the major difference between the two models, namely the relevance of foreign exchange risk premia in equilibrium pricing.

[3] In the latent variable approach, second moments are assumed to vary proportionally to each other, and market prices of risk (the latent variables in question) are related linearly to the instrumental variables (see Hansen and Hodrick (1983) and Hodrick and Srivastava (1984)). Generalized ARCH-M methods specify second moment behavior and assume that the market price of risk is constant (for an application of the GARCH-M methodology to international data, see Engel and Rodrigues (1989), Hodrick (1989), and Chan, Karolyi, and Stulz (1992). Chan, Karolyi, and Stulz conduct a conditional bivariate estimation of a GARCH-M model of two assets, a U.S. equity index and a foreign equity index.). In Ferson and Harvey (1991), as in Fama and MacBeth (1973), the second moments are obtained by a separate, first-pass estimation and then the model is tested. This last method relies on an assumption concerning the speed at which second moments vary. In Ferson and Harvey (1993), conditional betas are assumed to be linear in the information variables.

[4] Harvey (1989) proposes a method in which the prices of risk are assumed constant and then the second moments are free to move about in an unspecified way. Furthermore, for the special case in which the APM features only one risk premium, he proposed a substitution that allows both the market price of risk and the second moments to evolve in an unspecified way. In Harvey (1991), the classic APM is applied to international data using that technique. Unfortunately, the technique that allows both the second moments and the market prices to remain unspecified is not generalizable to more than one risk premium and has the undesirable implication that the APM is forced to be exactly verified when applied to the ex post return on the market portfolio. The market's pricing error—h_{mt} in Harvey's notation, as defined by his Equation (5) with $j = m$ — is *identically* equal to zero in all states of nature.

instrumental variables, in a manner similar to the latent-variable approach.[5] This assumption is parsimonious and buys a lot: it obviates the need to prespecify the behavior of first and second moments of returns over time.

For consistency, an APM that is conditional should also be intertemporal. If they anticipate information shifts, investors adjust their portfolio choice today in an attempt to hedge those shifts, thereby introducing in the APM more premia, which are based on the covariances of asset returns with changes in the instrumental variables. Hence, the theory leads to an equilibrium relation[6] that includes numerous risk premia linked to possible shifts in the information set. This is true equally in the classic and the international APMs. A test of a full-blown conditional, international, and intertemporal APM is beyond our computational capabilities at the present time. In most of the article we limit ourselves to a "static" (one-period) version of the APMs under consideration but, in the last section of this article, we attempt to throw some light on the choice one faces. Is it more important to take into account exchange risk premia or intertemporal hedging premia?

The article is organized as follows. Section I lays out the model and the econometric method to be used. Section II describes the data and some summary statistics concerning them. The purpose of Section III is to replicate earlier empirical attempts in testing for exchange risk premia in the unconditional APM. Section IV contains the main results of the article; we compare the international and the classic APMs and test for the price of exchange risk. We reject the hypothesis of zero price for exchange risk in the conditional, international APM. Section V presents some further analysis and interpretation of the time variation in prices of foreign exchange risk and world market risk. In Section VI, we perform numerous validation experiments, and in Section VII we test whether the world capital market—made up of the stock and the foreign exchange markets—is segmented; we do not reject the hypothesis that the world capital market is integrated. In Section VIII, we run a comparison between international and intertemporal APMs. We conclude in Section IX.

I. Model and Econometric Strategy

A. The Classic and International APMs with Time-Varying Moments

This study focuses on *conditional* asset pricing restrictions of a static nature. There are $L + 1$ countries, a set of $m = n + L + 1$ assets—other than the measurement-currency deposit—comprised of n equities or portfolios of equities, L nonmeasurement-currency deposits and the world portfolio of equities which is the mth and last asset. The nonmeasurement-currency deposits are

[5] See Assumption 2 below and the discussion of it.

[6] See Merton (1973). Stulz (1981) is an international, intertemporal, and conditional APM, but testing the full model requires a measurement of rates of consumption in the various countries. The static, international, and classic APMs can be derived as special cases of his model.

singled out by observing the above order in the list; i.e., they are the $(n + 1)$st to $(n + L)$th assets.

The international APM is equation (14) in Adler and Dumas (1983):[7]

$$E\left[r_{jt}|\Omega_{t-1}\right] = \sum_{i=1}^{L} \lambda_{i,t-1}\text{cov}\left[r_{jt}, r_{n+i,t}|\Omega_{t-1}\right] + \lambda_{m,t-1}\text{cov}\left[r_{jt}, r_{mt}|\Omega_{t-1}\right] \quad (1)$$

where r_{jt} is the nominal return on asset or portfolio $j, j = 1 \ldots m$, from time $t - 1$ to t, in excess of the rate of interest of the currency in which returns are measured, r_{mt} is the excess return on the world market portfolio, and Ω_{t-1} is the information set that investors use in choosing their portfolios. The time-varying coefficients $\lambda_{i,t-1}$, $i = 1 \ldots L$, are the *world prices of exchange rate risk*. The time-varying coefficient $\lambda_{m,t-1}$ is the world price of market risk.

Equation (1) is the result of an aggregation over the several categories of investors. In consuming their capital income, investors of different countries have access to goods at different prices.[8] Therefore, they view differently the returns from the same assets. The Japanese grant a premium on assets that protect their real purchasing power; this is a different premium from the one granted by the U.S. investors. Hence, the two premia show up as separate terms in the aggregated model. The covariance with exchange rates are present because of the optimal exchange-risk hedging behavior of the various investors. In truth, these should be covariances of rates of return on assets with changes in PPP deviations between countries. However, in this study we consider a world restricted to countries in which local currency inflation risk (to a one-month horizon) has been negligible. We, therefore, identify changes in PPP deviations with exchange rate changes.[9]

By contrast, the classic APM (CAPM) ignores investor diversity of that kind and assumes, in effect, that everyone in the world translates returns into consumption as do the residents of the reference currency country. Hence, no exchange-risk hedging premium appears. In the above notation, the restriction of the international APM to the classic APM is stated as:

$$\lambda_{i,t-1} = 0 \qquad i = 1 \ldots L, \forall t \quad (2)$$

B. The "Pricing-Kernel" Formulation

We now rewrite the international APM in the most parsimonious way that we can find. We introduce a new notation, M_t, referring to the marginal rate

[7] In Adler and Dumas (1983), the APM derivations rely on the assumption that first and second moments of returns are constant over time. Here, however, we use conditional moments, ignoring the implication that the movements in these moments would induce additional "intertemporal hedging" premia on the right-hand side of the APM. See Section VIII.

[8] The comparison is done after conversion into a common currency unit, of course.

[9] Although each investor looks at returns from his own currency's point of view, it is immaterial in which currency units the model is written. This is because the theoretical model assumes the absence of money illusion (i.e., homogeneity of degree zero of the investors' utility functions in nominal wealth and local prices).

of substitution between nominal returns at date t and at date $t - 1$. In addition, we call ρ_{t-1} the conditionally riskfree rate of interest in measurement currency for the period of time beginning at date $t - 1$. It is well known that the first-order condition of any portfolio choice problem may be written:

$$E[M_t(1 + \rho_{t-1})|\Omega_{t-1}] = 1, \tag{3a}$$

$$E\left[M_t r_{jt}|\Omega_{t-1}\right] = 0; \qquad j = 1, \dots m. \tag{3b}$$

Each APM specializes the form of M_t. The international APM equation (1) specializes M_t to be of the following form:

$$M_t = \left[1 - \lambda_{0,t-1} - \sum_{i=1}^{L} \lambda_{i,t-1} r_{n+i,t} - \lambda_{m,t-1} r_{mt}\right] \Big/ (1 + \rho_{t-1}) \tag{4}$$

That fact can be verified by direct substitution of equation (4) into equation (3b) (taking equation (3a) into account). The new time-varying term, $\lambda_{0,t-1}$, appears as a way of ensuring that equation (3a) is satisfied. It is not a market price of risk. It is the pure reflexion of the current level of the short rate of interest, ρ_{t-1}, compared to the current levels of the risk premia.

In the language of Hansen and Jagannathan (1991), M_t is the "pricing kernel" implied by the mean-variance theory of international asset pricing. In line with Hansen and Jagannathan (1991), equation (4) expresses the pricing kernel as a projection on a subspace of asset returns, namely the riskless return, the returns on the market portfolio, and the returns on currencies.[10] The innovation in this article, as compared to previous implementations, is that the projection has time-varying coefficients λ. Below (Assumption 2), we specify how those coefficients vary over time as a function of the instrumental (or state) variables.

C. Econometric Specification

In this subsection, we outline the empirical procedure that we are going to follow. We state two auxiliary assumptions that are needed.

ASSUMPTION 1: *The information Ω_{t-1} is generated by a vector of instrumental variables $\mathbf{Z}_{t-1}$.*

$\mathbf{Z}_{t-1}$ is a row vector of l predetermined instrumental variables that reflect everything that is known to the investor. We discuss in Section II below our choice of the $\mathbf{Z}$ variables; some of them will be variables endogenous to the financial market. The **Z**s in their limited number will not constitute a full description of the state of the real-world economy. Assumption 1 is a strong assumption that does not simply limit the information set of the econometrician; it limits the information set of the investors and, therefore, their

[10] Nonlinear pricing kernels are applied to international returns by Bansal, Hsieh, and Viswanathan (1993).

strategy space.[11] Testing an APM under information set Ω_{t-1} or under the information set generated by $\mathbf{Z}_{t-1}$ is not testing two different implications of the same theory of investment. The two tests refer to two different theories of investor behavior.

Next, we specify the way in which the market prices, λ, move over time. In a general equilibrium setting these would probably be nonlinear functions of the exogenous variables describing the state of the economy. Nonetheless, we now assume that the variables, Z, can serve as proxies for the state variables and that there exists an exact linear relationship between the λs and the Zs:[12]

ASSUMPTION 2:

$$\begin{aligned} \lambda_{0,t-1} &= -\mathbf{Z}_{t-1}\boldsymbol{\delta}, \\ \lambda_{i,t-1} &= \mathbf{Z}_{t-1}\boldsymbol{\phi}_i, \qquad i = 1, \ldots L \\ \lambda_{m,t-1} &= \mathbf{Z}_{t-1}\boldsymbol{\phi}_m. \end{aligned} \tag{5}$$

Here $\boldsymbol{\delta}$ and the $\boldsymbol{\phi}$s are time-invariant vectors of weights.

Defining the innovation, u_t, in the marginal rate of substitution:

$$u_t = 1 - M_t(1 + \rho_{t-1}), \tag{6}$$

and given Assumption 2 and the definition (4) of M_t, we have:

$$u_t = -\mathbf{Z}_{t-1}\boldsymbol{\delta} + \sum_{i=1}^{L} \mathbf{Z}_{t-1}\boldsymbol{\phi}_i r_{n+i,t} + \mathbf{Z}_{t-1}\boldsymbol{\phi}_m r_{mt}, \tag{7}$$

with u_t satisfying:

$$E[u_t|\Omega_{t-1}] = 0, \tag{8}$$

because of equation (3a).

Next, define:

$$h_{jt} = r_{jt} - r_{jt}u_t. \tag{9}$$

Equation (3b) implies that:

$$E[\mathbf{h}_t|\Omega_{t-1}] = 0. \tag{10}$$

We form the $1 + m$ vector of residuals $\boldsymbol{\epsilon}_t = (u_t, \mathbf{h}_t)$. Combining equations (8) and (10) under Assumption 1 yields: $E[\boldsymbol{\epsilon}_t|\mathbf{Z}_{t-1}] = 0$, which implies (but is by no means equivalent to) the unconditional condition:

$$E[\boldsymbol{\epsilon}_t\mathbf{Z}_{t-1}] = 0. \tag{11}$$

[11] By way of contrast, Hansen and Richard's (1987) theory of conditional versus unconditional portfolio efficiency deals with restrictions on the econometrician's (or the outside observer's) information set, keeping the investor's (or the fund manager's) information set the same.

[12] This specification has some similarities with the latent-variable approach of Hansen and Hodrick (1983) and Gibbons and Ferson (1985). Our λs are the conditional expected values of the "latent variables."

The sample version of this last restriction is the moment condition:

$$\mathbf{Z}'\boldsymbol{\epsilon} = 0, \tag{12}$$

where $\mathbf{Z}$ is a $T \times l$ matrix and $\boldsymbol{\epsilon}$ a $T \times (1 + m)$ matrix, T being the number of observations over time. This represents a total of $l \times (1 + m)$ moment conditions.

In order to run a test of the above specification, we choose a number of assets, m, which is large enough that the model imposes some overidentifying restrictions. We have l parameters $\boldsymbol{\delta}$, and $l \times (L + 1)$ parameters $\boldsymbol{\phi}$; we choose $L + 1 < m$ (more assets than countries) so that it is not possible to satisfy all the moment conditions exactly. We use Hansen's (1982) generalized method of moments (GMM) to minimize the average deviation from these moment conditions and thereby get the best estimates of the parameters $\boldsymbol{\delta}$ and $\boldsymbol{\phi}$. The deviations in the various moment conditions are weighted by a weighting matrix, $\mathbf{w}$, which is the inverse of a consistent estimate of the covariance matrix of sample moment conditions. Under the null hypothesis that equations (8) and (10) hold, the value of the resulting quadratic form is asymptotically distributed χ^2 with degrees of freedom equal to the number of orthogonality conditions minus the number of parameters. This being an asymptotic test, the number of observations plays no role in the determination of the degrees of freedom. Neither do distributional assumptions, provided that the conditions of the Central Limit Theorem are satisfied.

As is shown by Newey and West (1987a),[13] a similar statistic may be used to test the null hypothesis that some specific, additional restriction, or set of restrictions, binds the parameters. One computes the χ^2 statistic in the absence of the restriction and then again under the added constraints imposed on the model, but this time using the weighting matrix $\mathbf{w}$ of the unconstrained model. The difference in the two χ^2 statistics so obtained is itself distributed χ^2 with as many degrees of freedom as there are new restrictions.

II. Data and Preliminary Statistics

We consider the monthly excess return on equity and currency holdings measured in a common currency, the U.S. dollar. The excess return on an equity market is the return on that market (cum dividend) translated into dollars, minus the dollar one-month nominally risk-free rate. The return on a currency holding is the one-month interest rate[14] of that currency compounded by the exchange rate variation relative to the U.S. dollar, minus the dollar one-month risk-free rate.

In this study, we take four countries into account: Germany, the United Kingdom, Japan, and the United States. More precisely, we consider eight

[13] See also Gallant and Jorgenson (1979), Eichenbaum, Hansen, and Singleton (1988), Eichenbaum and Hansen (1990).

[14] These are Eurocurrency interest rates provided by Lombard Odier.

assets in addition to the U.S. dollar deposit: the equity index of each country,[15] a deutsche mark deposit, a pound sterling deposit, a yen deposit, and the world index of equities. In the implementation of the international APM, we include only three exchange risk premia, one for each of the currencies considered here. This creates a potential problem since in reality investors bear exchange risks from other currencies. For instance, French investors can hold the assets considered in this study and hence the exchange risks they bear will affect asset prices. However, it would be impossible for us to implement the model allowing for all exchange rate risks to affect asset prices. We discuss this problem in Section VI below.

The major shortcoming of any APM model is that the model does not provide any guidance as to the choice of instrumental variables. APMs are partial-equilibrium models reflecting the equilibrium in the financial market only. The link with underlying physical and monetary variables is left unspecified. General-equilibrium models would make that link explicit and would in this respect be preferable. However, most of the underlying physical variables of the economy are observed with measurement error and with, at best, quarterly frequency, whereas the financial variables that we wish to explain are observed with greater frequencies. Hence, even if the model being tested did specify the choice of underlying variables, these would have to be proxied by endogenous variables that are observed frequently, such as financial market variables. The search for proxies would be guided by the model and would, therefore, introduce a degree of arbitrariness smaller than that introduced by the arbitrary specification of instrumental variables in an APM. The methodological gain would be undeniable, but the end result may turn out to be very similar.

The number of instrumental variables is limited by the econometric methodology. Their choice is a sensitive question from both a conceptual and an empirical point of view. Conceptually, the APM theory fails to specify the list of instrumental variables; if it did, not one of them could be omitted in a proper test of the conditional APM. Empirically speaking, we verify below that the choice of instrumental variables does make a difference to the results.

Our specification borrows from the work of others which was done in unrelated contexts and on different data.[16] The dividend yield, the average bond yield, and the short-term interest rate provide some prediction of equity returns.[17] We also know that equity returns in January seem larger than during other months. Harvey (1991) shows that U.S. instruments have some power in predicting equity returns in foreign markets. The short-term inter-

[15] These are Morgan Stanley country indexes and the Morgan Stanley world index. See Harvey (1991) for an appraisal of these indexes.

[16] Hence, we are somewhat guilty of "data snooping" but not of "data mining."

[17] See Fama and French (1988), Breen, Glosten, and Jagannathan (1989), Ferson and Harvey (1991), Harvey (1991), and Solnik (1993), among others. Other variables, such as the default spread, have been shown to have some predictive power.

est-rate differential (equivalent to the forward premium) has been shown to help predict time variation in currency return.[18]

Our list includes six instrumental variables: a constant, the excess rate of return on the world index lagged one month, a January dummy,[19] the U.S. bond yield,[20] the dividend yield on the U.S. index, and the one-month rate of interest on a Eurodollar deposit. This choice of instruments allows for a direct comparison with published tests of the classic APM applied to international data (e.g., Harvey (1991)).

The prediction of non-U.S. returns, especially foreign currency returns, would have been aided by the inclusion of non-U.S. interest rates in the set of instrumental variables. Expanding the instrument set to include the three nondollar interest rates deteriorates the finite-sample properties of our estimates. Some of the proposed tests cannot be performed with as large a number of instruments. In Section VI, we return to that question.

Table I contains some preliminary statistics on the rates of return and the instrumental variables. Available index level data cover the period January 1970 to December 1991, which is a 264 data point series. However, we work with rates of return and we need to lag the rate of return on the world index by one month in the instrumental-variable set; that leaves 262 observations spanning March 1970 to December 1991.

Among instrumental variables, the U.S. bond yield, the U.S. dividend yield (both of them measured in excess of the Eurodollar rate), and the Eurodollar rate are fairly strongly correlated, and all three are somewhat correlated with the lagged world equity return.

There is a legitimate concern that some of the instrumental variables may not be stationary, thereby violating the assumption of the GMM. This may be the case for the dividend yield, the bond yield, and the short-term rate of interest. To alleviate this problem, we do two things. First, in the empirical implementation the dividend yield and the bond yield are replaced by their difference with the short-term rate.[21] Table I displays the autocorrelations of the instruments so constructed (except for the constant and the January dummy). Except for the lagged world rate of return, the degree of short-term serial dependence of these variables is large and positive. However, the positive serial dependence dies out at the two-year horizon and then becomes negative. As a second remedy, in Section VI below we compare the results obtained with the level of the short-term rate used as an instrument to the

[18] See Cumby and Obstfeld (1984), Kaminsky and Peruga (1990), Giovannini and Jorion (1989), Bekaert and Hodrick (1992), and others.

[19] In Japan, the fiscal year ends in March. There may be a case for introducing an April dummy variable.

[20] The index of bond yields is constructed by Lombard Odier; for a description, see Solnik (1993).

[21] This has the drawback of creating a strong multicollinearity between these instruments and the short rate. Multicollinearity does not affect the outcome of the tests that we perform below, only the values of the t-statistics posted in Tables I, IV, and V.

results obtained after first differencing the rate of interest. We verify that the results are not changed.

Table I also contains ordinary least squares (OLS) regressions of the various assets' returns on the instruments, in order to gauge the instruments' predictive ability. R^2s are small but a number of coefficients are significant. In all equations the Eurodollar rate plays a crucial role, although that is not always directly apparent in the table because the bond and dividend yield are used as spreads over the Eurodollar rate. Were the yields used by themselves, the Eurodollar rate would show up as significant and negative in all the equations. Admittedly, we are dealing with excess returns, which increases

Table I

Summary Statistics

Excess rates of return on assets are coded as 0.01 for a 1 percent rate of return per month. The instrumental variables that are yields or rates of return (except for the lagged world index rate of return) are coded as 1 for 1 percent per year. $r_m(-1)$ is the monthly rate of return on the world stock market lagged by one month. *USbony-E$* is the yield on an index of U.S. bond prices in excess of the Eurodollar deposit rate. *USDivy-E$* is the dividend yield on the U.S. stock index in excess of the Eurodollar deposit rate. *Euro$* is the one-month Eurodollar deposit rate. *JanD* is a dummy variable for the month of January.

Number of Observations = 262 (March 1970–December 1991)

Panel A: Excess Returns

	Mean/Month	Std. Dev./Month
German equity	0.0051	0.0623
U.K. equity	0.0066	0.0775
Japanese equity	0.0090	0.0659
U.S. equity	0.0025	0.0468
German deutsche mark	0.0017	0.0349
British pound	0.0017	0.0318
Japanese yen	0.0027	0.0332
World equity	0.0032	0.0436

Panel B: Instruments

	Mean	St. Dev.	Pairwise Correlations			
$r_m(-1)$	0.0361	0.5214	1.0	0.27	0.26	−0.25
USbony-E$	0.1545	2.1343		1.0	0.78	−0.74
USDivy-E$	−4.7019	2.6823			1.0	−0.97
Euro$	8.9063	3.2765				1.0

Panel C: Instrument Auto-Correlations

	rho1	rho2	rho3	rho4	rho8	rho12	rho24	rho36
$r_m(-1)$	0.11	−0.03	0.03	−0.01	−0.03	0.05	0.00	0.03
USbony-E$	0.91	0.82	0.76	0.70	0.60	0.47	0.06	−0.23
USDivy-E$	0.93	0.85	0.79	0.73	0.62	0.48	0.05	−0.24
Euro$	0.95	0.89	0.85	0.80	0.71	0.59	0.21	0.05

Table I—*Continued*

Panel D: Excess Returns Regressed on the Instruments*				
	German Equity	U.K. Equity	Japanese Equity	U.S. Equity
Constant	−0.0005 (−0.03)	−0.0241 (−1.25)	0.0236 (1.23)	−0.0222 (−1.65)
$r_m(-1)$	0.0032 (0.35)	0.0015 (0.14)	0.0105 (1.02)	−4E-05 (−0.01)
JanD	−0.0093 (−0.73)	0.0457 (1.68)	0.0052 (0.48)	0.0191 (1.53)
USbon-E$	0.0025 (0.92)	0.0004 (0.11)	0.0019 (0.68)	0.0014 (0.62)
USDiv-E$	0.0076 (1.41)	0.0216 (2.90)	0.0031 (0.55)	0.0135 (2.70)
Euro$	0.0047 (1.12)	0.0144 (2.80)	−0.0001 (−0.02)	0.0097 (2.65)
R^2	0.037	0.080	0.050	0.075
Residual auto-correlation	−0.018	0.033	−0.029	0.015

Panel E: Excess Returns Regressed on the Instruments				
	Deutsche Mark	British Pound	Japanese Yen	World Equity
Constant	0.0276 (2.77)	0.0167 (2.12)	0.0179 (1.93)	−0.0128 (−1.00)
$r_m(-1)$	−0.0112 (−2.75)	−0.0059 (−1.50)	−0.0027 (−0.64)	0.0030 (0.52)
JanD	−0.0149 (−2.00)	−0.0017 (−0.28)	−0.0113 (−1.81)	0.0174 (1.66)
USbon-E$	−0.0012 (−0.75)	−0.0015 (−1.04)	0.0009 (0.59)	0.0015 (0.71)
USDiv-E$	0.0015 (0.48)	0.0028 (0.98)	0.0012 (0.35)	0.0116 (2.86)
Euro$	−0.0019 (−0.79)	−0.0001 (−0.06)	−0.0010 (−0.37)	0.0077 (2.50)
R^2	0.085	0.036	0.062	0.087
Residual auto-correlation	0.048	0.090	0.043	0.005

* Ordinary Least Squares with heteroskedasticity consistent standard errors. First row of each cell: value of the coefficient. Second row: t-statistic in parentheses.

the size of that coefficient by one. Even without that, these would remain negative and significant. The U.S. dividend yield plays a significant role in predicting U.S. stock returns.

The January dummy and the U.S. bond yield play a significant role in predicting deutsche mark returns. The January dummy plays a similar role in the case of the Japanese yen. Whereas other researchers have amply documented day-of-the-week effects in foreign exchange rates, we are not aware of independent findings of month-of-the-year effects in currencies.

In order to show that our instruments are not "poor instruments,"[22] for each asset we run a Newey-West χ^2 test of the hypothesis that the coefficients of asset returns linearly regressed on the time-varying instruments are equal to zero. The first column of Table II shows the results of these tests. Except for German equities and the British pound, the tests indicate that the variables we have chosen are relevant instruments for the first moments of returns. Similarly, in the second column of Table II, we determine whether they are relevant instruments for second moments. Here, we use an auxiliary assumption that the first moments are linearly related to the instruments; i.e., the second moments are computed from residuals of a first-pass regression of returns. The second moments we consider are those that will be relevant in our test of the international and classic APMs, namely the covariances of asset returns with the three currencies and the world market. For each asset, we test the joint null hypothesis that these four moments are constant, against the alternative hypothesis that they are linearly related to the instruments. For all assets except the British pound we reject the null.

In comparing the R^2s of the regressions of the various assets on the common U.S. instrument set, we may appraise the extent to which it is legitimate to use these U.S. instruments rather than own-country instruments. The point to notice in Table I is that the R^2s of the regressions for non-U.S. assets exhibit no tendency of being lower than the R^2 of U.S. equity. Foster and Smith (1992) point out that the set of instruments of one country may appear to help predict another country's rate of return even though it is the wrong set of predictive variables, simply as a result of the correlation between the two countries' asset rates of returns. However, in calibrated simulations they indicate that, if such were the case, the spurious R^2 would be about 50 to 60 percent lower than that of the own-country prediction regression.[23] We see no such pattern in Table I. The reason why U.S. variables are legitimate instruments for non-U.S. returns is presumably that asset returns are related to business cycles and that the U.S. cycle has led or coincided with other cycles during the period under study.

III. Estimation and Tests under the Unconditional Versions of the APMs

In this section, we estimate the unconditional forms of the two contending APMs by setting $\mathbf{Z} = 1$. The results of these two separate estimations are

[22] See Nelson and Startz (1990).

[23] See their Table 5.

Table II

Chi-Square Tests of Linear Dependence between the First and Second Moments and the Instruments

The null hypothesis is that the moments under consideration (first moment in the first column, four relevant second cross-moments in the second column) are constant. The alternative is that they are linearly related to the instruments. Realized excess returns regressed on the instruments (as in Table I) produce a χ^2 equal to zero. When the coefficients of that regression (other than the constant) are constrained to be equal to zero, the χ^2 (built on the Hansen weighting 5×5 matrix of the unconstrained regression) rises to, e.g., 10.11 in the case of German equity. The residuals of these regressions are multiplied with each other (e.g., the German residual is multiplied successively with the deutsche mark, British pound, Japanese yen, and World equity residuals). These products are then regressed on instruments. When the coefficients of those regressions (other than the constants) are constrained to be equal to zero, the joint χ^2 (built on the overall Hansen weighting 20×20 matrix of the unconstrained regression) rises to, e.g., 50.60 in the case of German equity. Values of χ^2 are shown in the table, with corresponding p-values in parentheses.

	First Moment 5 Degrees of Freedom	Four Second Moments 20 Degrees of Freedom
German equity	10.11	50.60
	(0.072)	(0.0002)
U.K. equity	17.28	49.60
	(0.004)	(0.0003)
Japanese equity	15.28	52.57
	(0.009)	(0.0001)
U.S. equity	15.95	44.73
	(0.007)	(0.0012)
Deutsche mark	20.99	48.39
	(0.0008)	(0.0004)
British pound	8.16	22.26
	(0.148)	(0.327)
Japanese yen	15.25	51.32
	(0.009)	(0.0001)
World market	19.20	57.60
	(0.0018)	(0.00002)

displayed in Table III. The classic and the international APMs imply estimated constant relative risk aversions equal to 1.703 and 1.133 respectively. According to the theory in Adler and Dumas (1983), the market prices of foreign-exchange risk should be negative when the risk aversion of each investor subpopulation is greater than 1. Here, however, none of the exchange risk premia have significant t-statistics. The p-values found are 0.161 for the classic, unconditional APM and 0.049 for the international, unconditional APM.

Statement 1: In their unconditional versions, the international APM is marginally rejected while the classic APM is not rejected.

Table III

Estimation of the Unconditional Asset Pricing Models

Generalized method of moments tests of the moment conditions (11) (for the international asset pricing model (APM)) with the instruments reduced to: $\mathbf{Z} = 1$. λ_0 is an intercept term to make sure that the riskless rate is on the APM market line. λ_m is interpreted as market risk aversion. λ_{DEM}, λ_{GBP}, and λ_{JPY} are rewards per unit of risk for exchange rate risk. For the classic APM, similar moment conditions are constructed with λ_{DEM}, λ_{GBP}, and λ_{JPY} set equal to zero. Excess rates of return on assets are coded as 0.01 for a 1 percent rate of return per month. The instrumental variables that are yields or rates of return (except for the lagged world index rate of return) are coded as 1 for 1 percent per year. $r_m(-1)$ is the monthly rate of return on the world stock market lagged by one month. *USbony-E$* is the yield on an index of U.S. bond prices in excess of the Eurodollar deposit rate. *USDivy-E$* is the dividend yield on the U.S. stock index in excess of the Eurodollar deposit rate. *Euro$* is the one-month Eurodollar deposit rate. *JanD* is a dummy variable for the month of January. Number of observations = 262. Estimated coefficient: first row of each cell. t-statistic: second row in parentheses.

Panel A: Classic APM (Number of Factors = 1)

λ_0	λ_m
0.0055	1.7032
(0.59)	(1.16)

Chi-square: 10.53
Right tail p-value: 0.161
Degrees of freedom: 7

Panel B: International APM (Number of Factors = 4)

λ_0	λ_{DEM}	λ_{GBP}	λ_{JPY}	λ_m
0.0099	−0.3607	0.3417	2.2238	1.1334
(0.82)	(−0.14)	(0.13)	(0.91)	(0.74)

Chi-square: 9.54
Right-tail p-value: 0.049
Degrees of freedom: 4

Since the classic and the international APMs are nested, we can use the classic APM as the null hypothesis and the international APM as the alternative, to test the significance of exchange risk pricing. We do this test in the context of the unconditional APMs in order to replicate earlier empirical attempts. We fit the classic, unconditional APM using the covariance matrix of moment conditions obtained from the international, unconditional APM estimation. The result of this hypothesis test appears in Table VI. The p-value for the hypothesis of zero risk pricing is 0.74.

Statement 2: The hypothesis that exchange rate risk is not priced (H_0: $\lambda_i = 0$; $i = 1 \ldots L$) in the unconditional version of the international, static APM is not rejected.

This is the sense in which earlier empirical attempts based on the unconditional APMs are inconclusive.

IV. The Significance of Risk Premia in the Conditional, International APM

In this section, we reach the main results of this article concerning the relative desirability of the classic versus international, conditional APMs and we show the significance of exchange risk pricing.

The first order of business is to determine whether the international APM which we have formulated is viable. Table IV contains the result of estimating the conditional, international APM on the data and the instrumental variables that we have described. The table displays the estimates of the δ and the ϕ coefficients, which represent the way the world prices of risk vary over time. Several have significant t-statistics. The χ^2 test of the overidentifying restrictions created by the APM has a p-value equal to 0.228. Hence:

Table IV

Estimation of the Conditional, International Asset Pricing Model

Generalized method of moments tests of the moment conditions equation (11) with the full set of instruments. The vectors, $\boldsymbol{\delta}$, $\boldsymbol{\phi}_{DEM}$, $\boldsymbol{\phi}_{GBP}$, $\boldsymbol{\phi}_{JPY}$, and $\boldsymbol{\phi}_m$ contain the coefficients of the linear relationships between λ_0, λ_{DEM}, λ_{GBP}, λ_{JPY}, λ_m, and the instrument, $\mathbf{Z}$. λ_0 $(= -\mathbf{Z}\boldsymbol{\delta})$ is an intercept term to make sure that the riskless rate is on the asset pricing model market line. λ_m $(= \mathbf{Z}\boldsymbol{\phi}_m)$ is interpreted as market risk aversion. λ_{DEM} $(= \mathbf{Z}\boldsymbol{\phi}_{DEM})$, λ_{GBP} $(= \mathbf{Z}\boldsymbol{\phi}_{GBP})$, and λ_{JPY} $(= \mathbf{Z}\boldsymbol{\phi}_{JPY})$ are rewards per unit of risk for exchange rate risk. The time series of estimated values of the λs is displayed in Figure 1. Excess rates of return on assets are coded as 0.01 for a 1 percent rate of return per month. The instrumental variables that are yields or rates of return (except for the lagged world index rate of return) are coded as 1 for 1 percent per year. $r_m(-1)$ is the monthly rate of return on the world stock market lagged by one month. *USbony-E\$* is the yield on an index of U.S. bond prices in excess of the Eurodollar deposit rate. *USDivy-E\$* is the dividend yield on the U.S. stock index in excess of the Eurodollar deposit rate. *Euro\$* is the one-month Eurodollar deposit rate. *JanD* is a dummy variable for the month of January. Number of observations = 262. Estimated coefficient: first row of each cell. t-Statistic: second row in parentheses.*

	$\boldsymbol{\delta}$	$\boldsymbol{\delta}_{DEM}$	$\boldsymbol{\phi}_{GBP}$	$\boldsymbol{\phi}_{JPY}$	$\boldsymbol{\phi}_m$
Constant	0.1003	22.4735	19.3737	18.2531	−18.5363
	(0.38)	(1.53)	(0.83)	(1.01)	(−2.42)
$r_m(-1)$	−0.0991	−371.31	180.101	100.625	31.6038
	(−0.07)	(−3.27)	(2.00)	(1.30)	(1.00)
JanD	1.3406	−60.460	40.6945	0.8343	22.8588
	(2.47)	(−3.20)	(1.83)	(0.03)	(2.66)
USbony-E\$	−0.0358	−1.7018	−7.9977	3.4880	1.9627
	(−0.93)	(−0.44)	(−2.00)	(1.28)	(1.68)
USDivy-E\$	−0.0197	0.4544	5.6074	−3.7700	10.1277
	(−0.23)	(0.09)	(1.17)	(−1.04)	(4.53)
Euro\$	0.0195	−1.6611	0.5133	−3.6828	7.6852
	(0.30)	(−0.50)	(0.12)	(−1.21)	(4.52)

* Chi-square: 28.8039; right-tail p-value: 0.2276; degrees of freedom: 24.

Statement 3: The conditional version of the international APM is not rejected by the data.

Next, we estimate the conditional, classic APM, i.e., not including the risk premia on exchange-rate covariance risk. Table V contains the results of that estimation. The p-value is 0.005. Hence:

Statement 4: The conditional version of the classic APM applied to international data is rejected.

This estimation resembles Harvey's (1991) estimation except for the addition of currencies in the list of assets.[24]

In this article, we are interested in determining whether exchange risk is priced. Considering how sharp the rejection of the classic APM is and in view of the nonrejection of the international APM, we should have no doubt that exchange risk is priced in the international financial market. Nevertheless, a proper test is needed.

Since the classic and the international APMs are nested, we can again treat the classic as the null hypothesis and the international as the alternative hypothesis, in testing for the significance of exchange rate pricing. The null is that every element of the $\boldsymbol{\phi}_i$ vectors is zero (see restriction (2), and Assumption 2, equation). Following the Newey-West (1987a) prescription, we estimate the classic, conditional APM, holding the weighting matrix at its value estimated under the international, conditional APM. Table VI show the result. The χ^2 constrained by the null is 83.84, which is much higher than the unrestricted χ^2 that had been found for the international APM (28.80); the null hypothesis is rejected at all conventional levels of significance.

Statement 5: The hypothesis of zero price on exchange rate risk ($\boldsymbol{\phi}_i = 0$; $i = 1 \ldots L$) in the conditional version of the international APM is rejected.

By way of comparison, recall (Statement 1) that, in the absence of conditioning information, we reject the international APM, while we do not reject the classic APM and that exchange risk premia are not significant (Statement 2). Conditioning information plays a crucial role in discriminating between the two models.

It is useful to determine which asset(s) of our list cause the classic APM to be rejected, while the international one is not. We seek an answer to this question by decomposing the χ^2 statistic into components, each one of which represents the contribution of one asset. A perfect decomposition of this statistic is not possible because the covariances in the moment conditions applied to two different assets are not generally equal to zero. A partial response to the question is obtained by approximating the covariance matrix

[24] Harvey (1992) uses the default spread on bonds as an additional instrumental variable and does not include the dollar short interest rate as a separate instrument. He writes a projection equation for the first moment of each asset return, not just for the marginal rate of substitution, as we do here (equation (7)) with parsimony in mind.

Table V

Estimation of the Conditional Classic Asset Pricing Model

Generalized method of moments tests of the moment conditions (11) with λ_{DEM}, λ_{GBP}, and λ_{JPY} set equal to zero and with the full set of instruments. The vectors, δ and ϕ_m contain the coefficients of the linear relationships between λ_0, λ_m, and the instruments, $\mathbf{Z}$. λ_0 $(= -\mathbf{Z}\delta)$ is an intercept term to make sure that the riskless rate is on the asset pricing model market line. λ_m $(= \mathbf{Z}\phi_m)$ is interpreted as market risk aversion. Excess rates of return on assets are coded as 0.01 for a 1 percent rate of return per month. The instrumental variables that are yields or rates of return (except for the lagged world index rate of return) are coded as 1 for 1 percent per year. $r_m(-1)$ is the month rate of return on the world stock market lagged by one month. *USbony-E$* is the yield on an index of U.S. bond prices in excess of the Eurodollar deposit rate. *USDivy-E$* is the dividend yield on the U.S. stock index in excess of the Eurodollar deposit rate. *Euro$* is the one-month Eurodollar deposit rate. *JanD* is a dummy variable for the month of January. Number of observations = 262. Estimated coefficient: first row of each cell. t-statistic: second row in parentheses.*

	δ	ϕ_m
Constant	−0.0637	−9.6176
	(−0.61)	(−1.64)
$r_m(-1)$	0.5866	24.6887
	(1.08)	(0.99)
JanD	0.1264	0.0802
	(1.03)	(0.03)
USbony-E$	−0.0425	0.0442
	(−2.36)	(0.05)
USDivy-E$	0.1090	11.1430
	(2.03)	(6.12)
Euro$	0.0792	7.5439
	(2.06)	(5.45)

* Chi-square: 69.1189; right-tail p-value: 0.0053; degrees of freedom: 42.

of moment conditions by the elements that belong to its diagonal blocks, each block corresponding to one asset.[25]

Table VII shows the decomposition of various χ^2 statistics based on neglecting the elements outside the diagonal blocks. The last two lines of each column serve as a check that the χ^2 so reconstructed is very close to the statistic measured originally. The first column of Table VII contains the decomposition of the χ^2 statistic that served to test the overidentifying restrictions of the classic APM (supporting Statement 4); the second column contains the decomposition of the χ^2 of the international APM (Statement 3), and the third column pertains to the χ^2 difference of the test of significance of exchange risk premia (Statement 5). It is apparent that the rejection of the

[25] The procedure we use here is a summary procedure. After identifying the asset which causes rejection, we run a formal test (see below).

Table VI

Hypotheses Tests

This table assembles the test results that support Statements 2 and 5 to 10 of the text. The first column refers to the number of the statement under consideration. The second and third columns make explicit the null hypothesis being tested and the alternative. The fourth, fifth, and sixth columns spell out the test results. The vectors ϕ_{DEM}, ϕ_{GBP}, ϕ_{JPY}, and ϕ_m contain the coefficients of the linear relationships between λ_{DEM}, λ_{GBP}, λ_{JPY}, λ_m, and the instruments, $\mathbf{Z}$. λ_m $(= \mathbf{Z}\phi_m)$ is interpreted as market risk aversion. λ_{DEM} $(= \mathbf{Z}\phi_{DEM})$, λ_{GBP} $(= \mathbf{Z}\phi_{GBP})$, and λ_{JPY} $(= \mathbf{Z}\phi_{JPY})$ are rewards per unit of risk for exchange rate risk. To save space in the table below, we write ϕ_i or λ_i $(i = 1, 2, 3)$ in lieu of ϕ_{DEM}, ϕ_{GBP}, and ϕ_{JPY} or λ_{DEM}, λ_{GBP}, and λ_{JPY}, respectively. Concerning Statement 7, the γ_is are coefficients of a hypothesized proportionality between the vectors ϕ_i and ϕ_m. Concerning Statement 8, the superscripts "e" and "FX" refer to estimates measured on the equity and the foreign exchange market respectively.

Statement Number	Null Hypothesis	Alternative (Unrestricted Model)	χ^2 Difference [Newey-West (1987a) D-Statistic]	Degrees of Freedom	p-Value
2	$\lambda_i = 0$; $i = 1, 2, 3$	International unconditional APM	10.80 -9.54 1.26	3	0.740
5	$\phi_i = 0$; $i = 1, 2, 3$	International conditional APM	83.84 -28.80 55.04	18	< 0.001
6	λ_m time invariant	International conditional APM	66.90 -28.80 38.10	5	< 0.001
	λ_i time invariant $i = 1, 2, 3$	International conditional APM	80.93 -28.80 52.12	15	< 0.001
7	$\phi_i = \gamma_i \phi_m$ $i = 1, 2, 3$	International conditional APM	78.11 -28.80 49.31	15	< 0.001
8	$\phi_i^e = \phi_i^{FX}$ $i = 1, 2, 3$ Integration	International conditional APM allowing different ϕs	24.80 $-$ 8.69 16.11	18	0.584
8	$\phi_i^e = \phi_i^{FX}$ $i = 1, 2, 3$ Integration	International conditional APM allowing different ϕs	21.19 $-$ 7.67 13.51	18	0.760
10	ϕs of prices of intertemporal risk are zero	Classic, intertemporal APM	62.88 -26.16 36.72	18	0.006

classic APM as well as the rejection of the classic APM null are due primarily to the deutsche mark deposit, and secondarily to the Japanese yen deposit.

As a confirmation of this role of the deustche mark, we also run a formal test in the manner of Newey and West (1987a), Eichenbaum, Hansen, and

Table VII

Analysis of Chi Squares

The table shows the contribution of subsets of moment conditions to some of the χ^2s used in this study. The decompositions of χ^2s are based on neglecting the elements outside the diagonal blocks of the variance-covariance matrix of moments. The last two lines of each column serve as a check that the χ^2 so reconstructed is very close to the statistic measured originally. The first column contains the decomposition of the χ^2 statistic that served to test the overidentifying restrictions of the classic APM (supporting Statement 4); the second column contains the decomposition of the χ^2 of the international asset pricing model (APM) (statement 3) and the third column pertains to the χ^2 difference of the test of significance of exchange risk premia (Statement 5).

Moment Conditions Concerning	χ^2 of Classic APM	χ^2 of International APM	χ^2 of Significance Test
Marginal rate of substitution	0.95	2.65	2.91
German equity	2.29	5.59	5.86
U.K. equity	3.87	6.75	4.39
Japanese equity	5.45	2.02	2.75
U.S. equity	1.63	0.94	0.69
deutsche mark	25.76	0.79	20.52
British pound	6.57	1.56	7.82
Japanese yen	13.84	0.82	9.58
World equity	1.70	1.81	1.00
Sum (reconstructed χ^2)	62.06	22.92	55.52
Actual χ^2	69.12	28.80	55.04

Singleton (1988, Appendix C) and Gallant (1987). To do that, we remove the deutsche mark from the list of assets and reestimate the parameters holding the weighting matrix as its value estimated under the full set of assets, removing from the inverse weight matrix the rows and columns corresponding to the deutsche mark moment conditions. The drop in the objective function value is distributed χ^2 with as many degrees of freedom as there are instrumental variables. The reduced χ^2 is found to be equal to 48.37, while the χ^2 under the full list of assets for the classic APM was 69.12. The p-value of the null hypothesis that the deutsche mark fits the classic APM is equal to 0.002.

In view of the result of Hansen and Richard (1987)—that unconditionally efficient portfolios are also conditionally efficient—there is an apparent contradiction between our Statement 1 and our Statement 4, as far as the classic APM is concerned. Indeed, Statement 1 says that the market portfolio may be considered unconditionally efficient, whereas Statement 4 says that it is not conditionally efficient. Did we wrongly reject the classic APM in Statement 4?

The classic APM is a first-order condition applied to the market portfolio. The appropriate form of a first-order condition depends on the space of strategies being considered. No statement of efficiency of a portfolio should be made without specifying the space of strategies within which that portfolio is efficient.

Hansen and Richard's result must be understood *ceteris paribus*. Their result means that a portfolio which is unconditionally efficient within a space of strategies is also conditionally efficient within that *same space* of strategies. The restriction that we have tested under the name "conditional (classic) APM" is a moment condition that is implied by the first-order condition of conditional efficiency, in the vast space of strategies that depend, in an unspecified fashion,[26] on the lagged information variables, $\mathbf{Z}$. The moment condition that we have tested under the name "unconditional (classic) APM" is the first-order condition of unconditional efficiency in the space of strategies with constant portfolio weights ($\mathbf{Z} = 1$). Hansen and Richard's result does not apply to this comparison.[27]

In effect, the special case $\mathbf{Z} = 1$, under which Statements 1 and 2 were obtained, not only restricts the conditioning information of the econometrician; it also restricts the investors' information and, therefore, their space of strategies.

V. The Evolution of Risk Premia and the Goodness-of-Fit

Much interest in recent years has been focused on the question of knowing what moves risk premia in financial markets. Is the movement in risk premia mostly the result of the time-variation of risk or of the time-variation of the market reward per unit of risk? In the latent variables methodology, the time-variation in expected asset returns is driven by variation in the prices of risk. The work of Ferson and Harvey (1991, 1993) is specifically geared at separating the effects of the two variations. In the present model, we allow both quantities to vary; the risk varies in a fashion that is unspecified; world prices of risk (λ_i's, λ_m) vary according to Assumption 2. We are therefore in a position to throw some light on this debate.[28]

We test the hypotheses that λ_m, and then the λ_i's, are time-invariant. The results are displayed in Table VI. The p-value of the nulls of time invariance are extremely small.

[26] The fact that we constrain the marginal rate of substitution to have the form given in equation (7) with respect to $\mathbf{Z}$ restricts the space of strategies in some ways. However, it does not specifically restrict it to the space of linear strategies because the MRS is a nonlinear function of consumption and consumption may be a nonlinear function of returns.

[27] To illustrate the role played by the strategy space in deriving first-order conditions, consider the problem of writing the unconditional efficiency of a portfolio policy among the strategies with portfolio weights that are linear in the information variables, $\mathbf{Z}$. In the notation of Section I above, the first-order conditions of that problem would be:

$$E[h_j Z_k] = 0, \qquad \text{for all } j \text{ and } k.$$

These conditions are, in fact, identical to the particular implication of conditional efficiency that we have selected to express the conditional APM (equation (12)). They do not resemble the condition that we have called the "unconditional APM."

[28] But we cannot quantify the relative importance of the two types of variations since we have left the second moments unspecified.

Statement 6: The hypothesis that the world price of market risk is time-invariant in the conditional version of the international APM is rejected. The hypothesis that the world prices of foreign exchange risks are time-invariant is also rejected.

It is possible to visualize the time variations of the various world prices of risk. Panels a to d of Figure 1 display the evolution of the three world prices of foreign exchange risk and of world market risk, the latter being also compared between the international and the classic APM. To facilitate the comparisons, we display their deviations from their unconditional values based on the unconditional means and the seasonal effect of the January dummy. As can be seen in Panel d of Figure 1, the prices of market risk estimated in the classic and international APM vary in an almost identical fashion. The addition of foreign exchange risk premia does not affect the estimation of the price of market risk. The prices of foreign-exchange risk tend to be more volatile than the price of market risk. This is especially the case for the deutsche mark premium. The time series correlations across factors is quite small (ranging from -0.48 to 0.51) although they are projected on the same six instruments. The price of British pound risk may have decreased since 1982 while the price of the Japanese yen risk may have increased over the same period. The actual values of λ_m, which should be the market-average risk aversion, in fact fluctuate widely between positive and negative values. We have already mentioned that, provided risk aversions are greater than 1, the prices of exchange rate risk should be negative. Here, however, they also fluctuate widely between positive and negative values. As a result, the estimated marginal rate of substitution (not shown) is excessively volatile by any measure of excess volatility.

While we have been able to conclude that exchange risk premia are statistically significant, the technique used here by itself does not allow an estimate of the relative sizes of the various risk premia. Risk premia are equal to market prices of risk times the ex ante measures of risk (covariances with exchange rates and with the market portfolio). Unfortunately, ex ante measures of risk are not available in our estimation. As has been mentioned, the econometric method used here (see Section I) has the property that it does not require the specification of the behavior of second moments. The flip side of that coin is that we cannot indicate the size of risk premia only on the basis of the assumptions made so far.

Suppose, however, that for the purpose of interpretation only, we now make the additional assumptions that second moments do not vary over time and that first moments are linearly related to instrumental variables. In that case, we can display and compare month after month: (i) the expected returns of the eight assets based on a freely estimated linear statistical model, (ii) the expected returns under the restriction of the international APM with four factors, and (iii) the expected returns under the classic APM restriction with one factor. This is done in the eight panels of Figure 2, which provide a visual evaluation of the goodness of fit of the two models. The actual goodness of fit

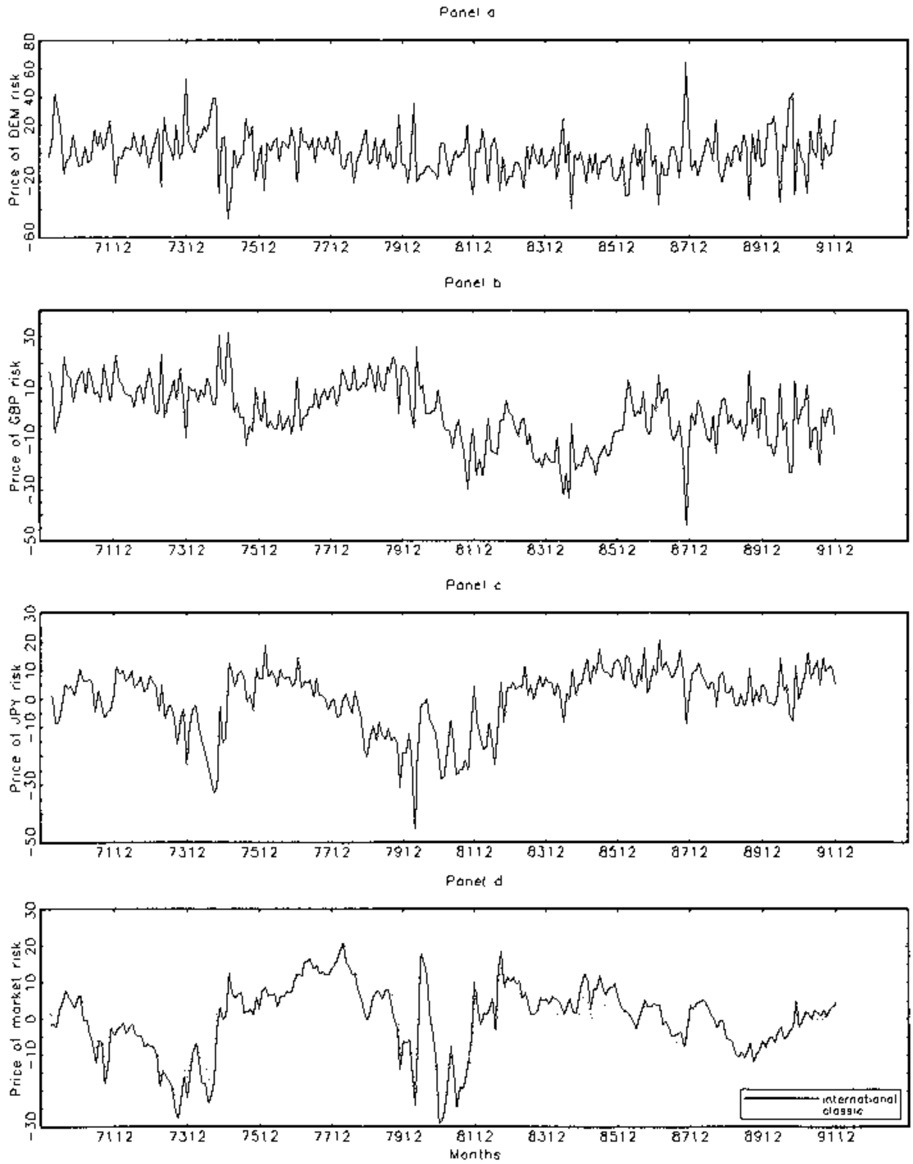

Figure 1. Time variation of world market prices of risk. Panels a to d show the time series (months are on the abscissa) of the estimated prices of deutsche mark (DEM), British pound (GBP), Japanese yen (JPY) and world market risks respectively. These are estimated as linear functions of the instrumental variables (see Assumption 2; e.g., $\lambda_{DEM} = \mathbf{Z}.\boldsymbol{\phi}_{DEM}$). The figure shows the deviations of the market prices of risks from their unconditional values that are based on the overall mean and the January seasonal variable. Notice that the price of world market risk is approximately the same in the international and the classic asset pricing models; the addition of foreign exchange premia does not materially affect the estimation of the risk premium linked to the market. There are 262 monthly observations. Month 1 is March 1970; month 262 is December 1991.

is, of course, better than is shown in this figure because the estimated model is not dependent on the two additional assumptions that are made for the purpose of drawing the diagram.

Observe that both models' expected returns track statistical expected returns fairly well. This implies that the model restrictions to four and one factor are not excessively far from the truth. It also implies that the actual estimation of the two APMs provides an estimate that is close to the one that would have been obtained, had we assumed second moments to be constant, even though we have allowed them to move in an unspecified manner. The movement of the market prices of risk, λ, is close to the movement in expected returns of the assets. In fact, one could verify that the market price of DEM risk is approximately proportional to the conditionally expected return on the DEM, and similarly for each risk premium.

The international model tracks statistical expected returns somewhat more closely on a month-to-month basis. The difference between the two models is more pronounced for the deutsche mark and the Japanese yen, which confirms our statistical analysis of Section IV.

VI. Discussion of Validity

In this section,[29] we evaluate the robustness of our results.[30] We are particularly concerned that some of the assumptions of the GMM method may not be satisfied. Chief among those is the finite size of our sample while the χ^2 tests are asymptotic (Section VI.*A*). Other possible violations of GMM assumptions may arise from nonstationarity of the variables (rates of return and instrumental variables) and from serial dependence in the sample moments. We look at these two aspects in Sections VI.*B* and VI.*C*. In addition, we want to check whether our results depend on the currency unit in which returns have been measured (Section VI.*D*), on missing exchange-risk premia (Section VI.*E*), on the inclusion of foreign interest rates in the set of instruments (Section VI.*F*), or on an incorrect definition of the market portfolio (Section VI.*G*).

A. Finite Sample Size

Given the relative parsimony of our econometric specification (Section I.*B*), the finite-sample problem may not be as acute in the present application as it may be in others. With 8 assets, 6 instrumental variables, 30 parameters, and 262 months of observations, the total number of numerical items to be estimated, equal to the number of parameters plus the elements of the weight

[29] Detailed outputs for all the validity tests presented in this section are not reported but are available from the authors.

[30] We already know that they are sensitive to the choice of conditioning information.

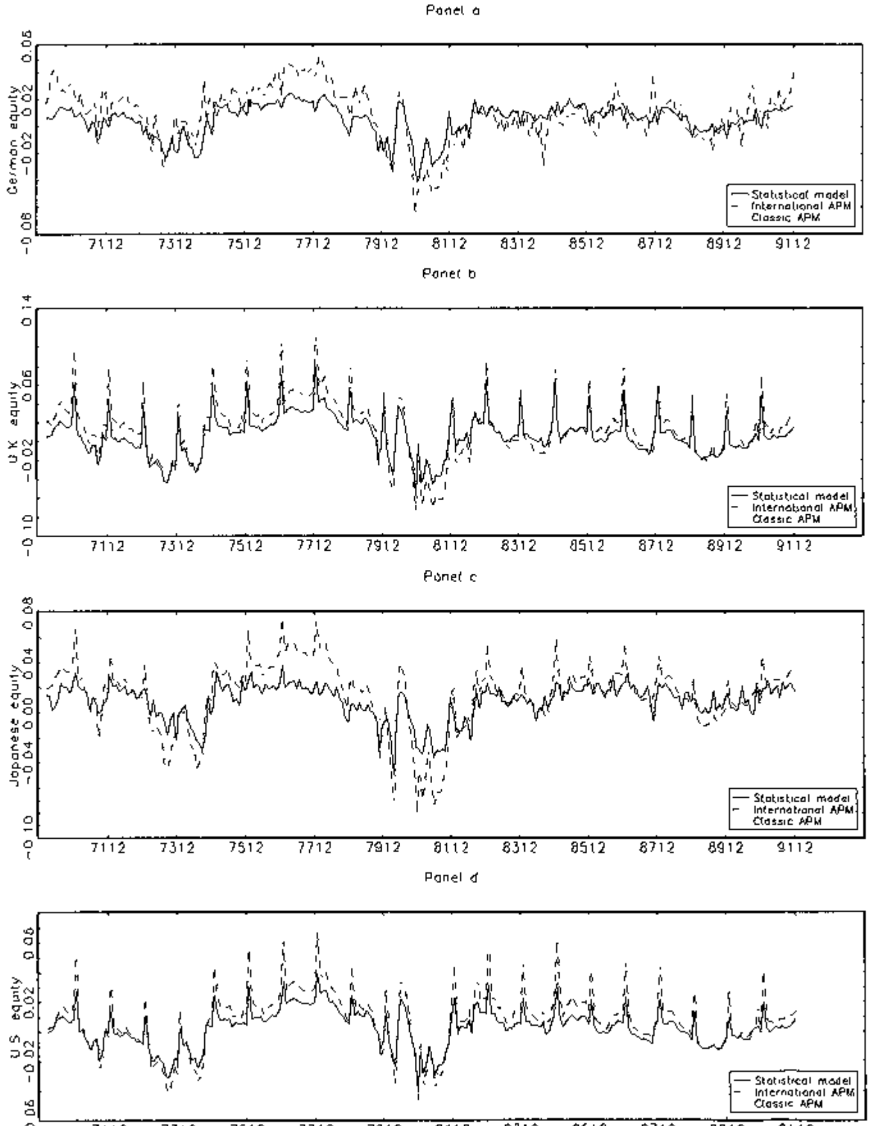

Figure 2. Time variation of conditional expected returns on assets. Panels a to h show respectively for each asset in our sample (German equity, U.K. equity, Japanese equity, U.S. equity, deutsche mark, British pound, Japanese yen, World equity portfolio), the time series of three estimates of expected returns. The first estimate labeled "statistical model" is given by a multiple regression of ex post rates of return on information variables. The second and third estimates are given by the market prices of risk of the classic and international APMs premultiplying a set of second moments, which are artificially assumed constant. The figure affords a visual evaluation of the time-series goodness-of-fit of the two models under the auxiliary assumption of time invariant second moments. The actual fit of the models tested in the article is better than this representation indicates, because the tests do not require the auxiliary assumption. There are 262 monthly observations. Month 1 is March 1970; month 262 is December 1991.

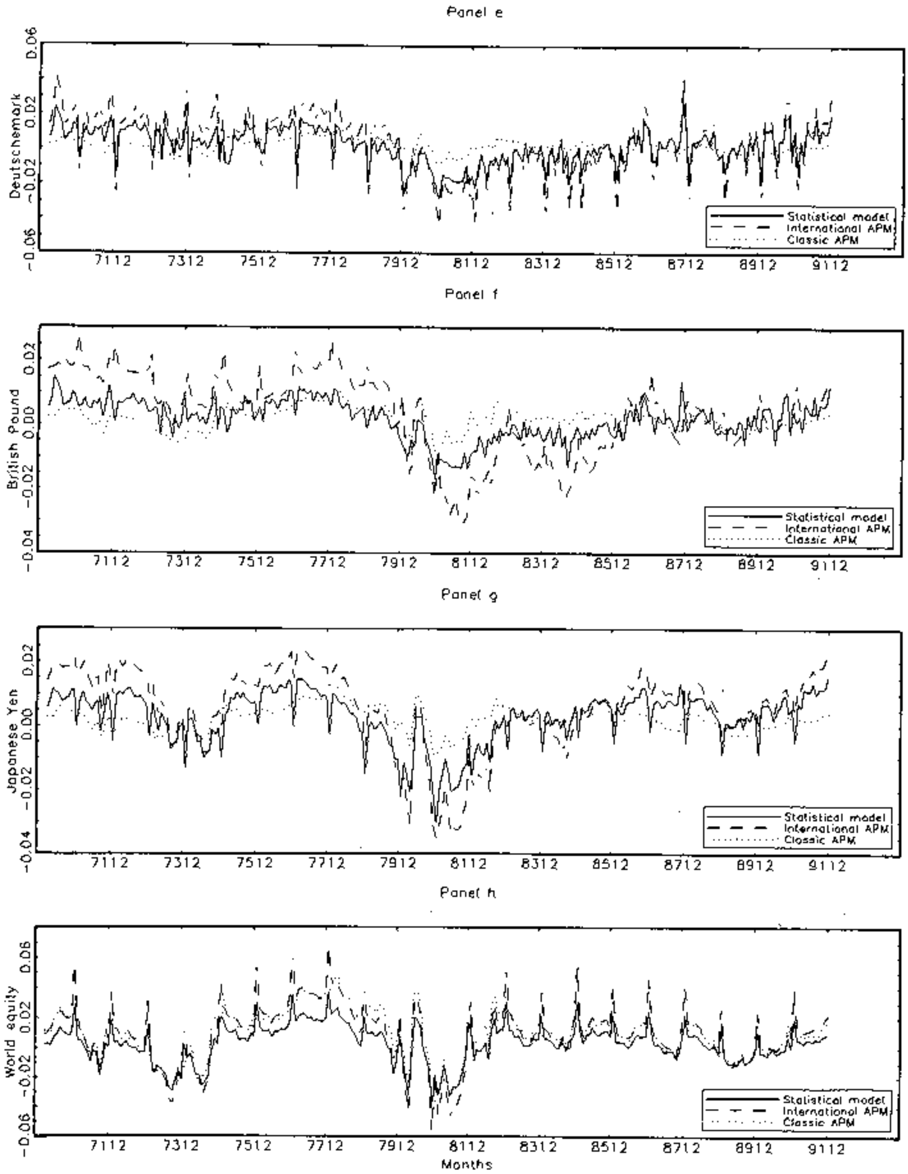

Figure 2.-*Continued.*

matrix, is equal to 1515 ($= 30 + 54 \times 55/2$), while the total number of datapoints concerning asset returns and instruments is equal to 3668 ($= 262 \times (8 + 6)$). The adjustment for finite sample size suggested by Ferson and Foerster (1994) implies that estimated variances of estimates should be multiplied by $1.7 (= 3668/(3668-1515))$, or that the standard errors should be increased (and the t-statistics should be reduced) by 30 percent only.

Monte Carlo experiments are the traditional answer to the question of finite sample size.[31] They require that one simulate the stochastic processes for all the variables of the model, including the instruments that are autocorrelated. Their stochastic processes must, therefore, either be known from theory—which they are not in our case—or be estimated prior to obtaining the sampling distributions on the coefficients of the APM. Hence, Monte Carlo experiments themselves would be questionable in this case.

Instead of simulating the variables, data may be obtained from out-of-sample observations. Indeed, the ultimate test of the validity of a model is its ability to help in forecasting. In sample, raw statistical estimates from an OLS regression provide the best possible fit; the overidentifying restrictions of a model only deteriorate the fit. Out of sample, however, a theoretical model can beat OLS. Unfortunately, splitting our sample of 262 observations into two subsamples proved perilous because each subsample was small.[32] χ^2 statistics computed out of sample were excessively large. It does not seem possible at this point to validate the use of asymptotic statistics by such out-of-sample tests.

B. Stationarity

The only potential problem of nonstationarity concerns the Eurodollar rate of interest. Other variables, such as the dividend yield and the bond yield, have been introduced as spreads relative to the Eurodollar rate. They are unlikely to be nonstationary.

In order to allay possible concerns with this issue, we recalculated the critical estimation and test using as an instrumental variable the first difference of the Eurodollar rate rather than its level. The χ^2 for the overidentifying restrictions of the international APM is equal to 22.73 which produces a p-value equal to 0.536. The χ^2 for the overidentifying restrictions of the classic APM is equal to 59.23 and produces a p-value equal to 0.04. The χ^2 for the test of significance of the foreign-exchange risk premia is equal to 66.61 and leads to an extremely small p-value. Hence, the results contained in Statements 3, 4, and 5 are not overturned.

C. Serial Dependence in the Vector of Sample Moments

Autocorrelation in the vector of sample moments may affect the consistency of the variance-covariance matrix of that vector. Newey and West (1987b) propose a weighting scheme for covariance smoothing. However, the number of terms to be included in the scheme is not easy to optimize. We have conducted experiments to determine the impact of covariance smoothing, depending on the number of months included in the scheme, on the test of overidentifying restrictions of the international APM. The p-values of that

[31] See Ferson and Foerster (1994).

[32] For instance, subsamples of equal sizes contain $3668/2 = 1834$ datapoints. This number is to be compared to 1515, the number of numerical elements to be estimated.

test increase with the number of moving average terms. The test result is, therefore, not overturned.

D. Measurement Currency

The theoretical APMs themselves are invariant to the choice of the measurement currency. To see this, consider first the classic APM expressed in nominal dollars. Translating rates of return into a different currency only changes the intercept from the dollar interest rate to the new currency interest rate. Hence, translating returns into the new currency and measuring excess returns relative to the new currency interest rate leaves the intercept equal to zero. Consider now the international APM expressed in nominal dollar excess returns. Translating dollar excess returns into, e.g., deutsche mark excess returns removes the deutsche mark risk premium (which is absorbed into the deutsche mark interest rate) and replaces it with a dollar risk premium. Exchange risk premia only permute with each other.[33]

Unfortunately, auxiliary Assumption 2 is not invariant to measurement currency. To ascertain the potential damage, we again perform the estimation of the two APMs and the tests of significance, using the deutsche mark, instead of the U.S. dollar, as the unit in terms of which nominal rates of return are measured. The χ^2 for the overidentifying restrictions of the international APM, equal to 28.69, produces a p-value equal to: 0.232. The χ^2 for the overidentifying restrictions of the classic APM, equal to 71.11, produces a p-value equal to 0.003. The χ^2 for the test of significance of the foreign-exchange risk premia, equal to 68.31, leads to an extremely small p-value. The results contained in Statements 3, 4, and 5 stand.

E. Missing Exchange Risk Premia

We have already pointed out that the international APM should in principle contain as many risk premia as there are national investor subpopulations in the world. Adler and Dumas (1983) explain that the sizes of the market prices of risk are related to the wealth of the various investor groups. It is to be assumed that the U.S., the German, the British, and the Japanese investor groups capture a large share of world wealth. Furthermore, the addition of other foreign exchange risk premia is only expected to render more acceptable the international APM, which already has a p-value in excess of 20 percent. And, of course, the rejection of the classic APM does not hinge on the number of exchange risk premia that we include in the alternative international APM.

The addition of foreign exchange risk premia is limited by the number of degrees of freedom and the finite-sample properties. In an attempt to mea-

[33] In the international APM, the return on the world market portfolio in the market risk premium can equivalently be expressed as a weighted average of country equity rates of return all expressed in dollars, or expressed in own currency. The difference between the two formulations would be compensated by the coefficient of the exchange risk premia.

sure the impact of the specification of exchange risk premia, we perform two comparisons. First, in line with Ferson and Harvey (1993) and Bansal, Hsieh, and Viswanathan (1993), we examine the model with a single GDP-weighted exchange risk premium. The premium is a weighted average of deutsche mark, pound, and yen premia, with 1981 GDP weights. This leads to a drastic reduction in the p-value of the international APM ($p = 0.077$). Because the international APM and the single-exchange risk premium models are nested, we can run a formal test of one against the other. The χ^2 of the restricted APM with a unique premium is found to be equal to 70.95. The gap relative to the unrestricted χ^2 of 28.8, with 12 degrees of freedom, produces an extremely small p-value for the null hypothesis of GDP weights. This illustrates that constant GDP weights do not capture the relative importance of the separate risk premia (see also Statement 6).

Second, we remove one exchange risk premium. We exclude the currency from the smaller economy, namely the British pound. In this case, the international APM is still accepted although marginally so (p-value = 0.1). The estimation of the coefficients of the price of market risk, ϕ_m, is robust to the exclusion of the pound premium. The correlation between the two estimated time series of λ_m is equal to 0.91.

F. Foreign Interest Rates

Foreign interest rates, in addition to the Eurodollar rate, no doubt have some ability to predict rates of return, particularly non-U.S. rates of return and returns on foreign currencies. It would seem relevant to include them as instruments. We have performed an experiment in which interest rate differentials between the deutsche mark, the British pound, and the yen on the one hand and the dollar on the other were added as instruments. The p-value on the international APM improved somewhat from its value obtained under the benchmark set of instruments: 0.334 as opposed to 0.228. Expanding the instrument set may not be advisable as it may deteriorate the finite-sample properties of the estimators.

G. Nominal Bonds in the Market Portfolio

The significance of foreign exchange risk premia in the international APM could be an artifact linked to the large outstanding amount of government bonds which have nominal denominations in various currencies.[34] These bonds are in investors' portfolios and should, perhaps, have been included in the market portfolio as has been suggested by Frankel (1982).[35] It is at least conceivable that the foreign-exchange risk premia, which we measured em-

[34] We are grateful to Ingrid Werner for this observation.

[35] Under the principle of Ricardian equivalence they should not be, because government bonds, considered as an asset in investors' portfolios, are offset by future taxes that investors will have to pay to allow the government to pay off the bonds. But it is not clear to what extent this principle holds in practice.

pirically, are just proxies for the missing bond returns. We now test this possibility. For lack of a reliable theory indicating what share government bonds should receive in the market portfolio, we make the auxiliary assumption that the shares of those bonds in the market portfolio are constant. This joint hypothesis implies that there exist constants γ_i such that:

$$\lambda_{i,t-1} = \gamma_i \lambda_{m,t-1}, \qquad \text{for all } t, \tag{13}$$

or, equivalently:

$$\boldsymbol{\phi}_i = \gamma_i \boldsymbol{\phi}_m. \tag{14}$$

γ_i is the share of bond i in the market portfolio divided by one minus the share of all bonds (i.e., the share of equities). The overidentifying restrictions of the international APM, with restriction equation (14) added, were tested and rejected with a $\chi^2 = 65.22$ and 39 degrees of freedom leading to a p-value equal to 0.053. The null hypothesis equation (14) was tested against the unrestricted international APM. The p-value is extremely small (Table VI).

Statement 7: The foreign-exchange risk premia that we have identified cannot be interpreted as proxies for missing bonds in the market portfolio.

VII. Testing for Segmentation

We have already accepted jointly all the overidentifying restrictions of the international conditional APM. It may be superfluous to test one of them in isolation. Nevertheless, we observe that the conditional, classic APM is rejected, in large part as a result of the introduction of currencies in the list of assets (see Section IV). One may therefore raise the question of whether currencies are priced differently from equity securities, i.e., whether the foreign exchange market is segmented from the stock market.

In testing for segmentation, we cannot allow each asset to have a different $\boldsymbol{\phi}$ matrix, as the model would then be underidentified. Instead, we group assets into categories. There are many possible groupings. In line with the above observation, we recognize two groups or markets: equities and currencies. We estimate a conditional version of the international APM in which we relax the APM restriction that the $\boldsymbol{\phi}$s are the same for both asset categories. Then we test the hypothesis of equality of the $\boldsymbol{\phi}$s between the two categories.[36] We only allow the prices of exchange rate risk to differ.

The definition (7) of the residual of the marginal rate of substitution forecast, u_t, involves the parameters $\boldsymbol{\phi}$. There are two versions of the test that differ according to which set of parameters $\boldsymbol{\phi}$ is used in the formulation of the constrained alternative: those of the foreign-exchange market or those of the equity market. The results are not sensitive to the choice of specification.

[36] See Harvey (1991) for a test in the same spirit.

In the three panels of Figure 3, we compare the market prices of exchange risk, λ_is, measured in the equity market with those measured in the foreign exchange market. The two track each other fairly closely, indicating an absence of segmentation between the two markets. The correlation of the time-varying prices of foreign-exchange risk in the equity and foreign-exchange markets ranges from 0.68 for the deutsche mark factor to 0.84 for the yen factor.

A formal comparative test is obtained by estimating the standard form of the international, conditional APM, with ϕs restricted to being equal in the two financial markets, but using the weighting matrix of the relaxed model. The calculations for the corresponding Newey-West test are in Table VI. The p-value is 0.584 or 0.760 depending on the version of the test. Hence:

Statement 8: In the international, conditional APM, the hypothesis of integration of the international stock and foreign exchange markets is not rejected.

When we compare this result to the existing literature, the ability of the international APM simultaneously to price equities and foreign exchange should perhaps be regarded as progress in the right direction. Giovannini and Jorion (1987, 1988), within a latent-variable framework, and Giovannini and Jorion (1989), within a GARCH framework, reject the classic APM on a dataset including equities and currencies. Mark (1988) and McCurdy and Morgan (1991), in a GARCH framework, test the classic APM on the excess returns of foreign currencies. They find that the systematic component of returns related to the world market portfolio represents a significant risk premium, but they do not compare the premium to the one applying to stocks. Korajczyck and Viallet (1992), in a APT framework that puts some restriction on the way in which betas can vary over time, find that excess returns on foreign currencies cannot be explained by their covariations with equity markets, a fact that they interpret as evidence of segmentation between equity and foreign exchange markets.

VIII. International Versus Intertemporal APMs

We pointed out in the introduction that a conditional, static APM is internally inconsistent. If it is conditional, it should be intertemporal since investors anticipate the future variations of the instrumental variables and hedge them over their lifetime. Indeed, recent research in international financial markets has made it a priority to use intertemporal models. These come in two classes. The first class contains so-called "consumption APMs" as in Breeden (1979), Lucas (1978, 1982), and Stulz (1981); these are Euler conditions based on one investor's consumption stream; they are valid equally in the domestic or international context; their empirical test requires the

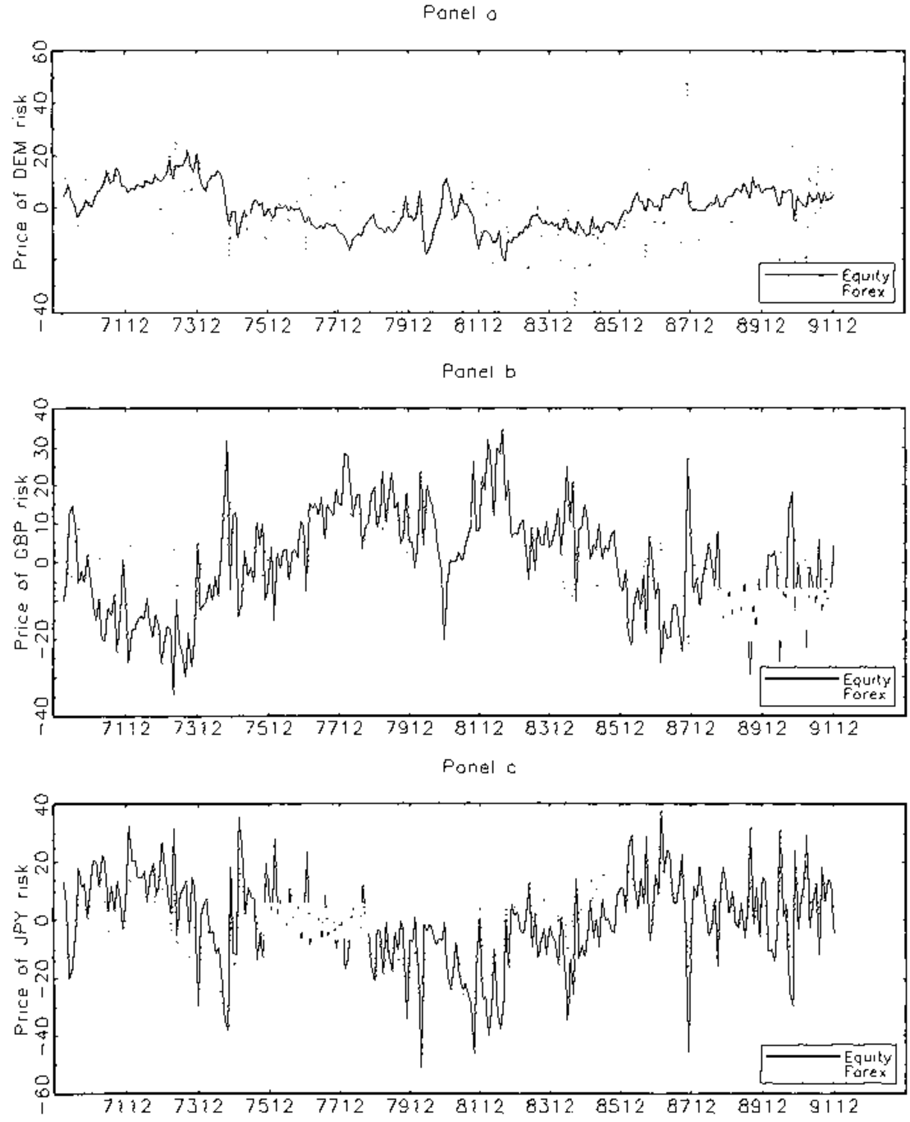

Figure 3. Integration between equity and foreign exchange markets. Panels a to c show respectively the world prices of deutsche mark, pound, and yen risks in the international APM, estimated on two data sets. The first data set includes the rates of return on equity securities (German equity, U.K. equity, Japanese equity, U.S. equity, and world equity). The corresponding curves in the three panels are labeled "equity." The second data set includes the rates of return on currencies, (deutsche mark, British pound, and Japanese yen). The curves are labeled "Forex." Under integration, the two subsets of securities should be priced on the basis of equal prices of foreign exchange risk. The figure illustrates that the two sets of prices are indeed quite close. The formal test described in the text does not reject the hypothesis of equality of the prices of exchange risk. Estimation and testing are done under the assumption that the world price of market risk is the same for both subsets of securities. There are 262 monthly observations. Month 1 is March 1970; month 262 is December 1991.

measurement of consumption.[37] The second class is based on the world market portfolio as in Merton (1973).[38] Since APMs of this second class are obtained by aggregation, the composition of the world population of investors here again makes a difference.

The test of a full conditional, international, intertemporal Merton-APM would require an extremely large number of assets as otherwise the statistical model is not overidentified. We are not in a position to run a test with enough assets. But we can run a "horse race" between the two types of models in order to determine which should be our fist priority: the introduction of hedging risk premia as in the intertemporal, classic APM or the introduction of exchange risk premia as in the static, international APM.

We applied to the international data an intertemporal model in which all the independent, stochastic instrumental variables,[39] with a one-period lead, are also priced risk factors. The p-value was found equal to 0.345.

Statement 9: The intertemporal, classic APM is not rejected by our international data.

The "horse race" is not conclusive. The conditional, international APM and the intertemporal APM are both accepted by the data.

Finally, we perform a test to verify the significance of intertemporal hedging premia in the intertemporal APM (i.e., we test the hypothesis of zero prices, λ, falling on the intertemporal factors). This involves re-estimating the conditional, static APM, using the weighting matrix of the intertemporal model. The resulting χ^2 is equal to 26.16. Calculations for this test appear in Table VI. The p-value is 0.006.

Statement 10: The hypothesis of zero prices on intertemporal hedging risks in the intertemporal APM is rejected by our international data.

We have no evidence to conclude in favor of either model.

Researchers may never be able to distinguish APMs with multiple risk premia once the investment opportunity set is allowed to vary over time. Imagine, for instance, that the intertemporal APM holds with as many risk premia as there are state variables. In general equilibrium, asset prices are functions of the state variables so that ex post rates of return on assets are functions of the state variables at the beginning and at the end of the investment period. If these are globally invertible functions, a subset of asset prices, e.g., exchange rates, can serve as proxies for state variables when constructing risk premia. In that sense, exchange rate risk premia may be equivalent to intertemporal risk premia.

[37] In the latent-variable method (Hansen and Hodrick (1983), Gibbons and Ferson (1985), and Hodrick and Srivastava (1984)), the assumption that each security's measure of risk is constant obviates the need to observe consumption. See, however, Wheatley's (1989) critique.

[38] See Ferson and Harvey (1991, 1993).

[39] This includes all of them except the constant and the January dummy, which are not stochastic, and the rate of return on the world market index which is colinear with the market factor. That leaves the U.S. bond yield, the U.S. dividend yield, and the Eurodollar interest rate.

IX. Conclusion

Based on a sample of securities that includes equities and currencies, our results show that foreign-exchange risks premia are a significant component of securities rates of return in the international financial market, and that the international APM dominates the classic APM.

The parsimonious econometric specification that we use here has undoubtedly helped in reaching this unequivocal conclusion. The parsimony is achieved in two ways. First, the estimation technique used makes it possible to test the various APMs without specifying the behavior of the second moments of rates of return. Second, by relying on the concept of kernel estimation, we also avoid having need to estimate the behavior of the first moments of rates of return. Only the behavior of the investors' marginal rate of substitution has to be estimated.

The parsimony in the specification increases the power of the tests by reducing the number of parameters that have to be estimated. But, it makes it impossible to answer two interesting questions. Because the second moments are not specified, it is not possible to determine how large the foreign exchange risk premia are relative to the classic reward for market covariance risk. However, after adding an assumption of time-invariance of second moments, we find that the exchange-risk premia yielded a visibly improved explanation of conditionally expected returns.

We are able to determine which assets cause the classic APM to be rejected in our sample. We find that foreign exchange risk premia are needed chiefly in order to explain rates of return on currencies. Once these risk premia are included in the APM, no evidence of segmentation between currency markets and stock markets is found.

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