Journal of Financial Economics 32 (1992) 137-167. North-Holland

Global financial markets and the risk premium on U.S. equity*

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Received August 1991, final version received February 1992

There is a significant foreign influence on the risk premium for U.S. assets. Using a bivariate GARCH-in-mean process, we find that the conditional expected excess return on U.S. stocks is positively related to the conditional covariance of the return of these stocks with the return on a foreign index but is not related to its own conditional variance. Further, we are unable to reject the international version of the CAPM. We present evidence for different model specifications, multiple-day returns, and alternative proxies for foreign stock returns.

1. Introduction

What drives the risk premium on U.S. equity? If the U.S. capital markets are segmented from foreign markets or if U.S. assets constitute most of the wealth traded internationally, the risk premium on U.S. assets should be determined solely in the U.S. Until the mid-1970s, these conditions seemed to prevail. Since

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*The authors thank participants in seminars at Concordia University, Queen's University, Vanderbilt University, the University of Western Ontario, Indiana University, the University of Chicago, Massachusetts Institute of Technology, the University of Minnesota, Ohio State University, Notre Dame University, the NBER Conference on International Asset Pricing in Philadelphia, the Western Economics Association meetings in Seattle, the European Finance Association meetings in Rotterdam, the Johnson Symposium at the University of Wisconsin, and the Pacific-Basin Capital Markets Research Conference in Seoul for useful comments. We are grateful to Warren Bailey, Tim Bollerslev, Stephen Brown, George Constantinides, Robert Cumby, Wayne Ferson, Stephen Foerster, Robert Hodrick, John Huizinga, Ravi Jagannathan, Patric Hendershott, Robert Korajczyk, Craig Lewis, Andrew Lo, Thomas McCurdy, Richard Roll, Paul Seguin, Lemma Senbet, Henri Servaes, Robert Stambaugh, Simon Wheatley, and especially the referee, Campbell Harvey, and editor, William Schwert. James Tompkins and Darrell Lee provided helpful research assistance. K.C. Chan and Andrew Karolyi thank the Dice Center for Financial Economics at Ohio State University for partial funding.

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then, capital markets have become increasingly integrated internationally and the market value of U.S. assets has become a smaller fraction of world wealth. These developments suggest that the risk premium on U.S. assets may now be determined primarily on world capital markets. We investigate this issue and find empirical evidence of a significant foreign influence on the risk premium of U.S. assets.

The capital asset pricing model (CAPM) implies that the risk of the market portfolio is measured by the variance of its returns, so that the risk premium for the market portfolio increases with the variance of its returns. Further, with risk-averse investors, the addition of a mean-preserving spread to the distribution of the market portfolio return increases the risk premium, so that even if the CAPM does not hold, one would typically expect a positive relation between the risk premium and the variance of the market portfolio. Merton (1980) estimates the relation between the risk premium and volatility using contemporaneous variance estimates. Using a generalized autoregressive conditional heteroskedasticity (GARCH) representation developed by Engle (1982) and Bollerslev (1986), French, Schwert, and Stambaugh (1987) provide evidence that the conditional expected excess returns on the market portfolio and the conditional variance of its returns are positively related. They use daily data from January 1928 to December 1984 for the excess returns on the market with Standard and Poor's (S&P) 500 index representing the market portfolio.

For most of the sample period studied by French, Schwert, and Stambaugh, foreign capital markets were substantially segmented from the U.S. capital markets and their capitalization value was much smaller. There is now substantial evidence that stock markets are reasonably well integrated. For instance, Cho, Eun, and Senbet (1986), Wheatley (1988), Korajczyk and Viallet (1989), Gultekin, Gultekin, and Penati (1989), Cumby (1990), Harvey (1991), Bekaert and Hodrick (1992), and Campbell and Hamao (1992) provide evidence of integration, especially during the 1980s, using a variety of asset pricing models with monthly data. Further, as fig. 1 indicates, the capitalization of U.S. stocks has become a smaller part of the capitalization of the world market portfolio of common stocks. These two developments suggest that the S&P 500 is unlikely to have been an adequate proxy for the world market portfolio over the last decade.

If we use the CAPM, the evidence that stock markets are fairly well integrated suggests that the risk premium on the market portfolio of U.S. assets depends on the covariance of its returns with the returns on the world market portfolio. This covariance is a weighted average of the variance of the market portfolio of U.S. assets and the covariance of the returns on the market portfolio of U.S. assets with the market portfolio of non-U.S. assets, where the weights are the proportions of U.S. and foreign stocks in the world market portfolio. Even if the CAPM does not hold, this covariance may be important in determining the risk premium on the U.S. market portfolio. For instance in Merton's (1973) intertemporal CAPM, the risk premium on a portfolio still increases with the covariance of the returns of that portfolio with the market portfolio.

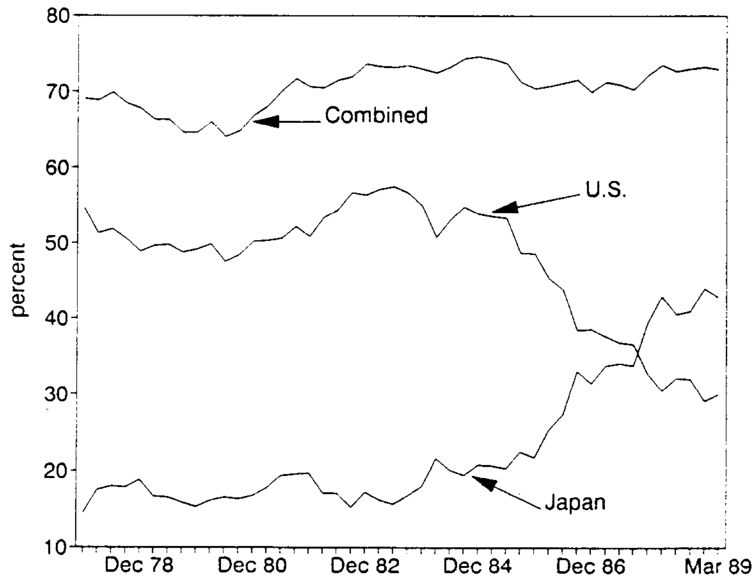

Fig. 1. Market capitalization of the U.S. and Japanese equity markets from 1978 to 1989 as proportion of total world equity market capitalization. The market values for the U.S. and Japanese stocks are drawn from various quarterly issues of Morgan Stanley's *Capital International Perspectives*.

In our empirical work, we model the daily excess returns of the S&P 500 and of a portfolio of non-U.S. assets jointly, using a bivariate GARCH-in-mean representation. We find that the conditional expected excess returns on the U.S. market portfolio are significantly related to the conditional covariance of the S&P 500 with Japan's Nikkei 225 index, but not significantly to the conditional variance of the S&P 500. We obtain a similar though weaker result if the market portfolio of foreign assets is approximated by the Morgan Stanley EAFE index (which is a value-weighted index comprising stocks from Europe, Australia, and the Far East) or the Morgan Stanley Japan index. We also show that this result is robust to a number of alternative econometric specifications as well as to different measurement intervals for the returns data. This result suggests that the impact of foreign stock markets on the risk premium of the U.S. market portfolio cannot be neglected.

When we impose the restrictions of the international version of the CAPM, we find that we cannot reject the model at the 5 percent significance level. Given the small number of assets in our tests, however, one might argue that this significance level is too low. At the 10 percent level, the model is rejected when we use the Nikkei indices, but not otherwise. Further, at that level of significance, we cannot reject a two-factor model that implies that Japanese and U.S.

market risks are priced identically across countries when the Nikkei or the EAFE indices are used. We view this evidence as supporting the notion of international integration among equity markets.

The paper proceeds as follows. The estimated model is introduced in section 2. Section 3 presents the main empirical results. Section 4 investigates the robustness of the results to alternative econometric specifications and returns measurement intervals, and section 5 presents concluding remarks.

2. The model

We assume that:

(A.1) Markets are internationally integrated.

(A.2) Investors are period-by-period mean-variance optimizers in a common numeraire, which we take to be the U.S. dollar.

(A.3) The aggregate relative risk aversion, $\lambda = \sum_{i=1}^{N} \lambda_{it} W_{it} / \sum_{i=1}^{N} W_{it}$, is constant, where λ_{it} is the ith individual's relative risk aversion, W_{it} is his wealth at time t, and N is the number of investors.

Assumption (A.1) implies that assets with the same risk have the same expected excess returns irrespective of where they are traded. Assumption (A.2) ensures that the CAPM holds. The assumption of a common numeraire implies that there are no preferred currency habitats for individual investors.[1] Assumption (A.3) implies that the risk premium changes only when the volatility of the market portfolio changes.

An alternative to assumption (A.2) used in some of our tests is:

(A.2′) Investors use the currency of their home country as the numeraire.

Assumption (A.2′) corresponds to a world in which the exchange rate reflects changes in the relative price of goods across countries and investors consume different baskets of goods in different countries. With this assumption, each stock position in the world market portfolio is financed in the country in which the stock is traded. Hence, the excess returns on a stock equal the returns of that stock in its own currency net of the returns of the risk-free asset in that currency. The interpretation of these own-currency returns is that they are equal to dollar returns hedged against currency risk with a hedge ratio of one: that is, one goes short one unit of foreign currency for each purchase of one unit of foreign

[1] See Alder and Dumas (1983) and Stulz (1983) for reviews of the literature on international asset pricing and discussions of the role of preferred currency habitats.

currency stock. This approach, pioneered by Solnik (1974), leads to a linear pricing relation between expected own-currency excess returns and their conditional covariance with the world market portfolio of own-currency excess returns.[2]

With these assumptions, we have

$$E(R_{mt+1} - R_{0t+1} | \Omega_t) = \lambda \text{var}(R_{mt+1} - R_{0t+1} | \Omega_t), \tag{1}$$

where E(·) is the conditional expectation operator, Ω_t is the investors' information set at t, R_{mt+1} is the return on the world market portfolio from t to $t+1$, and R_{0t+1} is the risk-free rate for the same period of time. Var $(R_{mt+1} - R_{0t+1} | \Omega_t)$ is the conditional variance of the excess return on the world market portfolio for the period t to $t+1$ given the information set Ω_t. With assumption (A.2), all excess returns are computed in dollars, whereas with the preferred currency habitats assumption (A.2′), excess returns are own-currency returns, so that the excess return on the market portfolio is a weighted average of own-currency returns.

Using the CAPM and (1), we have

$$E(r_{dt+1} | \Omega_t) = \lambda \text{cov}(r_{dt+1}, r_{mt+1} | \Omega_t) \tag{2}$$

where lower-case r denotes excess returns and $\text{cov}(r_{dt+1}, r_{mt+1} | \Omega_t)$ is the conditional covariance of the excess returns on the U.S. domestic market portfolio with the excess return on the world market portfolio.

The return on the world market portfolio can be written as

$$r_{mt+1} = \omega_{dt} r_{dt+1} + (1 - \omega_{dt}) r_{ft+1}, \tag{3}$$

where ω_{dt} is the capitalization of the U.S. market portfolio as a fraction of world wealth at time t, r_{dt+1} is the excess return on the U.S. domestic market portfolio, and r_{ft+1} is the excess return on the foreign market portfolio.

Using (3), we can rewrite (2) as

$$E(r_{dt+1} | \Omega_t) = \lambda[\omega_{dt} \text{var}(r_{dt+1} | \Omega_t) + (1 - \omega_{dt}) \text{cov}(r_{dt+1}, r_{ft+1} | \Omega_t)]. \tag{4}$$

In this paper, we focus on the relation in (4). If the U.S. market portfolio equals the world market portfolio, the weights become $\omega_{dt} = 1$ for all t and (4) reduces to the relation examined by French, Schwert, and Stambaugh (1987).

[2]There is no theoretical reason for the hedge ratio of one to be the minimum-variance hedge. When it is not, Sercu (1980) shows that the appropriate returns are those of self-financing portfolios of stocks hedged against exchange rate risk using the minimum-variance hedge. Bailey, Ng, and Stulz (1992) provide evidence, however, that a hedge ratio of one cannot be improved on for the Nikkei 225 by using conditional minimum-variance hedge ratios.

Even if the returns of the U.S. market portfolio are not related to the returns of the market portfolio of non-U.S. assets, (4) still differs from that of French, Schwert, and Stambaugh (1987), since the conditional variance of the U.S. market portfolio is multiplied by the weight of the U.S. portfolio in the world portfolio.

With daily data from a single country, (4) has a straightforward interpretation because it corresponds to a portfolio strategy that can be easily implemented: buy the market portfolio at the close of trading and hold it until the next day's close. With daily data from several countries, the trading hours generally do not overlap in calendar time. Using close-to-close returns for each country, it is not possible to implement directly a portfolio strategy of buying the world market portfolio at the close of U.S. trading at time t and selling it at the close of U.S. trading at time $t + 1$, because the foreign markets typically close before the U.S. market. To reduce the problem created by nonsynchronous trading hours, one could use returns measured over longer periods, so that the nonoverlapping period becomes a smaller fraction of the measurement interval. We choose instead to use daily data and address the problem of nonsynchronous trading hours directly. We use daily data for two reasons. First, the period for which it is reasonable to investigate the effect of foreign stock markets on the U.S. risk premium is reasonably short, so our tests would lack power with infrequently sampled data. Since existing tests of international asset pricing models that focus on stock returns use monthly data, the use of daily data is a contribution to the international asset pricing literature.[3] Second, the results of French, Schwert, and Stambaugh (1987) that most strongly support the positive relation between conditional expected excess returns and conditional variances are obtained using daily returns.

In using daily returns, we consider an investor who makes his portfolio decisions at the close of trading in the U.S. markets. When the investor forms his portfolio, he uses expected returns and the variance–covariance matrix of returns conditional on information available at that time. If we had foreign returns computed from the close of the U.S. market to the next close of the U.S. market, we would use them as part of the information set. Since foreign returns measured over the same calendar time as U.S. returns are not available, we have to model the joint dynamics of the U.S. and foreign returns to reflect the nonsynchronism of trading hours. Specifically, at the close of the U.S. market the investor knows the return on the U.S. market and the return on the foreign market at its close earlier in the day. Since the U.S. market closes after most foreign markets, part of the next day's return of the foreign market has already accrued when the investor makes his portfolio decision. As long as contemporaneous returns on

[3] Engel and Rodriguez (1989) and Giovannini and Jorion (1989) use weekly data but test asset pricing models using returns on foreign currencies rather than on foreign stocks. Chang, Pinegar, and Ravichandran (1991) use daily returns in a study of the integration of European equity markets where the problem of nonsynchronous trading is not significant.

the domestic and foreign markets are correlated, the investor can use today's U.S. return to improve his forecast of tomorrow's foreign return. Our approach accounts explicitly for this correlation.

The information in the U.S. close-to-close return today about the next day's foreign return gives the appearance that foreign market returns can be predicted from the previous day's U.S. return. This predictability is spurious, however, since it cannot be exploited through a portfolio strategy if foreign markets are efficient. The foreign return does not contain similar information about the next-calendar-day U.S. return, since on the previous calendar day the foreign markets closed before the U.S. market and hence the information in the foreign returns is incorporated in the previous-day U.S. return.[4]

Since the lack of synchronism in trading hours affects our modeling strategy, it is important to make sure it does not bias our results. We therefore set up two further experiments. First, we know that as the measurement interval is lengthened, the lack of synchronism becomes less important. Our robustness tests indeed show that our finding of a significant foreign effect on the U.S. risk premium holds if we use longer measurement intervals for the returns. Second, we demonstrate that it holds if we use open-to-open U.S. returns and hence allow the investor to form his expectations at the open of U.S. markets.

In the following, we assume that the information set of investors, Ω_t, is approximated by a set of instruments that includes only past returns of the U.S. and the foreign portfolios for the sample period.[5] We therefore propose the parameterization

$$r_{dt+1} = a_d + \beta_{dv}\omega_{dt}h_{dt+1} + \beta_{dc}(1-\omega_{dt})h_{ct+1} + \theta_{d1}\varepsilon_{dt} + \theta_{d2}\varepsilon_{dt-1} + \varepsilon_{dt+1}, \tag{5a}$$

$$r_{ft+1} = a_f + \beta_{fv}\omega_{ft}h_{ft+1} + \beta_{fc}(1-\omega_{ft})h_{ct+1} + \theta_{f1}\varepsilon_{ft} + \theta_{f2}\varepsilon_{ft-1} + \phi_f\varepsilon_{dt} + \varepsilon_{ft+1}, \tag{5b}$$

where the conditional variances, h_{ft+1} and h_{dt+1}, and covariances, h_{ct+1}, depend only on past returns. The specification for each country allows the returns to depend on two lagged disturbances to incorporate the effects of infrequent

[4]If we had used the close of trading in foreign markets as the time when the investor makes his portfolio decision instead, today's foreign return would be useful to forecast the same-calendar-day U.S. return, but the U.S. return would not be useful to forecast tomorrow's foreign return. Hence, in this case, the returns dynamics would have to be modeled differently.

[5]This is in contrast of the instrumental variables approach of Harvey (1991), which conditions the market risk premium on a set of macroeconomic variables.

trading on the dynamics of the index returns[6]. We also allow the foreign returns to depend on one lagged disturbance of the domestic returns through ϕ_f to take into account the lack of synchronism in trading hours. As explained above, there should be no ϕ_d in the domestic returns equation.[7]

An important step in implementing the model empirically is specifying the dynamics of the conditional variances and covariances. We use a general specification drawn from the ARCH process originally proposed by Engle (1982), which allows the conditional covariance matrix to be related to its own past values and past squared disturbances. We posit the following general process:

$$\varepsilon_{t+1} \sim \mathrm{N}(0, \mathbf{H}_{t+1}) \quad \text{and} \quad \mathbf{H}_{t+1} = \begin{bmatrix} h_{dt+1} & h_{ct+1} \\ h_{ct+1} & h_{ft+1} \end{bmatrix}, \tag{6a}$$

$$\mathbf{H}_{t+1} = \mathbf{P}'\mathbf{P} + \mathbf{F}'\mathbf{H}_t\mathbf{F} + \mathbf{G}'\varepsilon_t\varepsilon_t'\mathbf{G}\,. \tag{6b}$$

where $\mathbf{H}_{t+1}$ denotes the 2×2 variance–covariance matrix conditional on information at time t and ε_{t+1} denotes the vector of disturbances from eqs. (5a)–(5b). **P** is an upper triangular matrix of coefficients, whereas **F** and **G** are free matrices of coefficients. This model was originally proposed by Baba, Engle, Kraft, and Kroner (1989) (BEKK). The important feature of this specification is that it builds in sufficient generality, allowing the conditional variances and covariances of the two stock markets to influence each other, and, at the same time, does not require us to estimate a large number of parameters (eleven for the bivariate system used here). Even more importantly, perhaps, the BEKK process guarantees that the covariance matrices in the system are positive definite. To see the level of generality we gain by implementing this specification, we can contrast it with the multivariate ARCH model adopted by Bollerslev, Engle, and Wooldridge (1988) (BEW), in which the conditional variances depend only on past squared residuals and covariances on past products of disturbances. The important cross-market effects, highlighted by Hamao, Masulis, and Ng (1990) for the national stock markets of the U.S., U.K., and Japan, are ignored in the BEW specification. A more general BEW process could be specified to capture these effects, but positive semi-definiteness of the

[6]Stoll and Whaley (1990) show how the effects of infrequent trading and bid–ask spreads can cause stock index returns to follow an ARMA-type process. See also Muthuswamy (1990).

[7]Our method of addressing this problem with nonsynchronous trading hours could be improved in future research, since we omit the covariance attributable to the unanticipated return on the foreign market between the close of that market and the close of the U.S. market because that return is unobservable. Instead, we implicitly assume that the true covariance is a constant proportion of the estimated covariance. If this assumption is incorrect, we would expect our results to be sensitive to alternative measurement intervals, but we show in section 4 that they are not.

conditional covariance matrix in that system is no longer assured. We discuss estimation results with alternative processes in section 4.

3. Description and interpretation of empirical results

3.1. Data

We use three indices for non-U.S. assets. The first is the Nikkei 225 Stock Average. The data for the Nikkei 225 index are collected from the *Asian Wall Street Journal* as are the exchange rates used to obtain dollar-denominated returns. The Nikkei 225 is a price-weighted index of stocks traded in the first section of the Tokyo stock market. For this index, we have data from January 3, 1978 to December 31, 1989. We also use two value-weighted indices published in Morgan Stanley's *Capital International Perspectives* that are available from January 3, 1980 to December 31, 1989. These are the Morgan Stanley Japan index in yen and the Morgan Stanley EAFE index in dollars. We use the indices only on calendar days when they are available in both countries and make no distinction between single- and multiple-day returns. The Morgan Stanley Japan index is the Japanese component of EAFE and hence is a value-weighted index. The EAFE index does not include all stocks for each market, since it generally ignores small-capitalization stocks. It is widely used by practitioners as a benchmark for the performance of non-U.S. stocks, however, and internationally-diversified index funds generally try to replicate the performance of that portfolio. It includes no adjustment for cross-holdings in Japan and other countries, so it may give more weight to Japan than is warranted, as argued by McDonald (1989) and French and Poterba (1991).

Table 1 provides summary statistics for the data, including the cross-correlations among the various portfolio excess returns. The U.S. index and dollar-denominated foreign equity index excess returns are obtained using the three-month U.S. Treasury bill yields; those for the yen-denominated Nikkei 225 index are computed with the three-month Gensaki interest rate. Panel A indicates that the excess returns series show significant negative skewness for the S&P 500 and the Nikkei yen excess returns and generally positive excess kurtosis. These results suggest that the returns distribution deviates from normality. The Kolmogorov D statistic and tests based on Bera and Jarque (1982) confirm this formally.

The autocorrelation coefficients are shown in panel B for the raw excess returns and squared returns series. The significant positive and declining autocorrelations for the squared series indicate some second-order dependence that the GARCH models in this study seek to capture.

Finally, the cross-correlations between the S&P 500 returns and those of the various foreign index returns are shown in panel C. The significant leading

Table 1

Summary statistics for daily U. S. and foreign equity market excess returns (in percent) from January 1978 to December 1989.

The U.S. equity index is Standard and Poor's 500 stock index and the foreign market indices are the U.S. dollar- and yen-denominated Nikkei 225 index, the Morgan Stanley Japan index (yen-denominated), and the Morgan Stanley EAFE (dollar-denominated) index. All dollar-denominated index returns are computed net of the three-month U.S. Treasury bill yield and the yen-denominated index returns are computed net of the three-month Gensaki interest rate. The returns associated with October 16, 19, 20, and 21, 1987 are omitted. The Kolmogorov–Smirnoff D statistic tests the null hypothesis of normality with critical values of 0.0256, 0.0281, and 0.0271 at the 5 percent significance level and 0.0307, 0.0337, and 0.0325 at the 1 percent significance level for 2819, 2338, and 2522 degrees of freedom, respectively. The Bera–Jarque B test statistic for normality is based on the excess skewness and kurtosis coefficients and is asymptotically distributed χ^2 with two degrees of freedom with critical values at the 5 percent significance of 5.99 and at the 1 percent significance level of 9.21. The tests for deviations from normality for the skewness and kurtosis statistics are based on D'Agostino, Belanger, and D'Agostino (1990). Cross-correlations are given between the S&P 500 index daily excess returns, r_t (and squared returns, r_t^2), and those of the index shown in the table.

Panel A: Distributional statistics

Statistic	S&P 500	Nikkei ($)	Nikkei (yen)	MS Japan	MS EAFE
Nobs.	2819	2819	2819	2338	2522
Mean	0.0254	0.0693	0.0619	0.0688	0.0495
Std. dev.	0.9760	1.0648	0.7658	0.9414	0.8917
Skewness	− 0.4409[a]	0.0535	− 0.2912[a]	0.1323	0.0327
Kurtosis	5.6594[a]	2.2764[a]	5.8208[a]	8.8262[a]	3.0575[a]
Kolmogorov D	0.0525[a]	0.0469[a]	0.0759[a]	0.0858[a]	0.0357[a]
Bera–Jarque B	72.576[a]	11.489[a]	75.706[a]	157.09[a]	19.570[a]

Panel B: Autocorrelations of daily excess returns

Statistic	S&P 500	Nikkei ($)	Nikkei (yen)	MS Japan	MS EAFE
			Series: r_t		
ρ_1	0.0552[a]	0.0975[a]	0.0934[a]	0.1229[a]	0.1059[a]
ρ_2	0.0077	0.0027	− 0.0604[a]	− 0.0646[a]	0.0258
ρ_3	− 0.0234	0.0246	− 0.0024	0.0051	0.0191
ρ_4	− 0.0042	0.0345	0.0134	0.0167	0.0300
ρ_5	− 0.0086	− 0.0295	− 0.0507[a]	− 0.0695[a]	0.0079
ρ_6	− 0.0047	− 0.0004	− 0.0379	− 0.0296	− 0.0100
			Series: r_t^2		
ρ_1	0.0675[a]	0.1492[a]	0.2617[a]	0.1156[a]	0.1493[a]
ρ_2	0.0792[a]	0.1025[a]	0.1826[a]	0.2121[a]	0.0619[a]
ρ_3	0.1641[a]	0.0600[a]	0.1853[a]	0.2134[a]	0.1009[a]
ρ_4	0.1227[a]	0.0451[a]	0.1003[a]	0.0466	0.1123[a]
ρ_5	0.0824[a]	0.0749[a]	0.1211[a]	0.1413[a]	0.0866[a]
ρ_6	0.1463[a]	0.0206	0.0966[a]	0.0711[a]	0.0219

Table 1 (continued)

	Panel C: Cross-correlations of daily S&P 500 and foreign market excess returns							
	Nikkei ($)		Nikkei (yen)		MS Japan		MS EAFE	
Lag	r_t	r_t^2	r_t	r_t^2	r	r_t^2	r	r_t^2
−6	−0.0167	0.0377	−0.0394[a]	0.0689[a]	−0.0263	0.0363	0.0034	0.0404[a]
−5	0.0188	0.0567[a]	0.0177	0.0786[a]	0.0255	0.0418[a]	0.0061	0.0630[a]
−4	−0.0051	0.0487[a]	−0.0153	0.0929[a]	−0.0319	0.0783[a]	−0.0022	0.1539[a]
−3	0.0029	0.0609[a]	−0.0144	0.1268[a]	−0.0171	0.0782[a]	0.0390	0.0746[a]
−2	0.0571[a]	0.0770[a]	0.0689[a]	0.1451[a]	0.1004[a]	0.1061[a]	0.0456[a]	0.0545[a]
−1	0.1916[a]	0.1516[a]	0.2707[a]	0.2139[a]	0.2531[a]	0.1642[a]	0.3305[a]	0.1854[a]
0	0.0838[a]	0.1522[a]	0.1201[a]	0.2003[a]	0.1095[a]	0.1269[a]	0.1629[a]	0.2190[a]
1	0.0137	0.1021[a]	−0.0033	0.2194[a]	−0.0041	0.1999[a]	0.0062	0.0954[a]
2	0.0015	0.0754[a]	−0.0164	0.1906[a]	−0.0342	0.1806[a]	0.0099	0.0504[a]
3	0.0536[a]	0.0356	0.0411[a]	0.0747[a]	0.0422[a]	0.1556[a]	0.0195	0.0749[a]
4	−0.0356	0.0675[a]	−0.0471[a]	−0.1129[a]	−0.0281	0.0889[a]	−0.0353	0.0555[a]
5	−0.0206	0.0444[a]	−0.0284	0.1274[a]	−0.0483[a]	0.1036[a]	−0.0288	0.0202
6	−0.0044	0.0441[a]	−0.0198	0.0647[a]	−0.0044	0.0336	−0.0001	0.0464[a]

[a]Significant at the 1 percent level.

correlations from U.S. to foreign stocks in the raw returns reflect the nonsynchronous trading hours for the respective markets. Eqs. (5a) and (5b) of the model proposed in the previous section attempt to control for this effect. More interestingly, the cross-correlations computed for the squared returns series suggest that what happens in one market affects the volatility of returns in another. Again, the dynamics of the conditional covariances applying the BEKK process of (6) can capture such cross-market dependence in the volatility of the returns.

Eqs. (5a) and (5b) require computation of market weights. We obtain these weights from Morgan Stanley's *Capital International Perspectives*. Since the weights are published quarterly, we use the beginning-of-quarter weights reported by Morgan Stanley and compute the weights within each quarter by adjusting the beginning-of-quarter weights dynamically, using the realized returns on the indices. For each quarter, we extend the interpolation to the beginning of the next quarter and compare our weights with those reported by Morgan Stanley. In this comparison, the reported weights are always very close to the predicted weights.

3.2. Empirical results

Panel A of table 2 reproduces estimates of (5a)–(5b) that use, respectively, the dollar-denominated excess returns on the Nikkei 225 index, the yen-denominated excess returns on the Nikkei 225 index and the Morgan Stanley Japan index, and the dollar-denominated excess returns on the EAFE index as the returns on the foreign market portfolio. The estimates are obtained using dummy variables in (5a)–(5b) for October 16, 19, 20, and 21, 1987, when it appears the distribution of returns differs dramatically from the distribution on the other days in our sample. The point estimates obtained without the dummy variables for the parameters of interest are quite similar to those reported here, but statistical inference becomes more complicated because of the serious departures from the normality assumption.

The estimates of the equation for U.S. returns with the dollar returns of the Nikkei 225 show that the conditional covariance of the U.S. returns with the Nikkei return has a significant positive effect on U.S. conditional expected returns, but that the conditional variance of the U.S. returns has no effect. This result holds whether we use standard t-statistics or the alternative t-statistics robust to departures from normality suggested by Bollerslev and Wooldridge (1990).[8] Hence, this result supports the hypothesis that the U.S. risk premium is determined on global markets during the sample period. The lack of significance of the coefficient on the variance is surprising. Since the conditional variance and covariance have a significant correlation (0.57 if we use the dollar-denominated Nikkei returns), one might be tempted to attribute the lack of significance of the conditional variance to multicollinearity. If we estimate a univariate model for the U.S. returns for our sample period, however, the coefficient on the variance is not significant either. A more plausible explanation is that, if the returns are measured with error, the error decreases the precision of our variance estimate more than the precision of our covariance estimate, since returns squared involve taking the square of the error, whereas products of domestic and foreign returns involve multiplying two errors that are imperfectly correlated. The estimates obtained with the yen-denominated excess returns of the Nikkei 225 are qualitatively similar to those obtained with dollar-denominated returns.

The estimates using the Morgan Stanley Japan index and EAFE as foreign market portfolios are consistent with those obtained using the Nikkei 225. In both equations, the covariance has a positive coefficient that is significantly different from zero using standard t-statistics. Using robust t-statistics, the covariance coefficient is significant at the 10 percent level for the Morgan Stanley Japan index, but is insignificant at the same level for the EAFE index. The point estimates of the covariance coefficients are almost the same for the equations using the Nikkei 225 and EAFE. The variance of the U.S. returns has a positive but statistically insignificant coefficient in both equations.

Table 2 also reports results for foreign excess returns. The results are not completely symmetric to those for the U.S. When we use the dollar excess returns on the Nikkei, the covariance with the U.S. returns has a positive but

[8]Because of the large sample size of this analysis, the appropriate criteria for statistical significance for sample statistics and estimated coefficients are unclear. Throughout the text and tables, we highlight critical values at conventional significance levels but caution the reader that a more conservative cutoff may be appropriate.

insignificant effect on the expected excess returns of the index, but the variance of the Nikkei has a significant positive coefficient. The same result holds for the equation estimated with the yen-denominated excess returns on the Nikkei, except that in this case the covariance has a significant positive coefficient with the standard t-statistic and an insignificant coefficient with the robust t-statistic. With the Morgan Stanley Japan and EAFE indices, the variance of the foreign index has an insignificant positive coefficient; the covariance has a negative insignificant coefficient for the EAFE but a positive significant coefficient for the Morgan Stanley Japan index at the 5 percent level using the standard t-statistic and at the 10 percent level using the robust t-statistic.

For all indices, we test the hypothesis that the expected returns are determined in the home country – i.e., that Japanese expected returns do not depend on the covariance of Japanese returns with U.S. returns and that U.S. expected returns do not depend on the covariance of U.S. returns with Japanese returns. Though the results are not reported, this hypothesis is easily rejected for all indices. Consequently, irrespective of the index used as a proxy for foreign stocks, the covariance of foreign stock returns with domestic stock returns has an important effect on the expected returns on domestic stocks.

Table 2 also measures the fraction of the total variation of returns captured by our explanatory variables. Whereas the fraction of the total variation explained for the U.S. returns, denoted R_1^2, is small but respectable, given the difficulty of forecasting daily returns, it would be virtually negligible if we ignored the international effect presented in this paper. For instance, if we set the coefficient on the conditional covariance equal to zero in the U.S. returns equation when the dollar return on the Nikkei 225 is used as the foreign return, the R_1^2 falls from 1.16 to 0.29 percent; in contrast, it falls to only 0.94 percent if we set the coefficient on the conditional variance equal to zero and to 0.25 percent if both coefficients on conditional variance and covariance equal zero. If the coefficients for the conditional variance and covariance are set at zero, the R^2 captures the proportion of the total variation that is due to the lagged moving average terms that are posited to capture the effects of nonsynchronous trading of the component stocks within each of the domestic and foreign market indexes. Further, the R^2 for the U.S. returns equation in this case is substantially larger than that for a simple univariate model (0.28 percent). Although the fraction of the total variation explained for the foreign returns, R_2^2, is much larger than the fraction explained for the U.S. return, it becomes comparable to the fraction explained for the U.S. return if ϕ_f is set equal to zero, as one would expect.

Figs. 2 and 3 plot the conditional expected excess returns and conditional variances and covariances for the U.S. when the foreign index uses dollar-denominated returns on the Nikkei 225; all the series exhibit substantial variation over time. For the U.S. returns, the conditional daily expected excess returns have a standard deviation of 0.09 percent and a mean of 0.04 percent, whereas the unconditional daily excess returns have a mean of 0.025 percent and

Table 2

Estimates from the bivariate model of daily expected excess returns for U.S. and foreign equity markets.

U.S. equity returns, r_{dt+1}, are given by Standard and Poor's 500 stock index and the foreign equity market returns, r_{ft+1}, by the U.S. dollar- and yen-denominated Nikkei 225 index (model 1 and 4), the Morgan Stanley Japan yen-denominated index (MSJP) (model 2), or the Morgan Stanley EAFE index (model 3). All dollar-denominated indices are computed net of the three-month U.S. Treasury bill yield; yen-denominated indices are computed net of the three-month Gensaki interest rate. The market-value weights, ω_{jt}, for each stock market j on day t, with $j = d$ for U.S. equity and $j = f$ for foreign equity markets, are daily interpolations from quarterly estimates published by Morgan Stanley's *Capital International Perspectives*. The market weights sum to unity in each regression. The number of observations for each model corresponds to that shown in table 1. R_1^2 and R_2^2 denote the ratio of the explained to total variation in the excess returns associated with the S&P 500 and foreign index markets. Standard t-statistics are presented in parentheses and robust t-statistics computed with quasi-maximum-likelihood methods are in brackets. The model parameters are given by the following system of equations:

$$r_{dt+1} = \alpha_d + \beta_{dv} h_{dt+1} \omega_{dt} + \beta_{dc} h_{ct+1}(1 - \omega_{dt}) + \theta_{d1}\varepsilon_{dt} + \theta_{d2}\varepsilon_{dt-1} + \varepsilon_{dt+1}$$

$$r_{ft+1} = \alpha_f + \beta_{fc} h_{ct+1}(1 - \omega_{ft}) + \beta_{fv} h_{ft+1}\omega_{ft} + \theta_{f1}\varepsilon_{ft} + \theta_{f2}\varepsilon_{ft-1} + \phi_f \varepsilon_{dt} + \varepsilon_{ft+1}$$

$$\begin{bmatrix} \varepsilon_{dt+1} \\ \varepsilon_{ft+1} \end{bmatrix} \sim \mathrm{N}(0, \mathbf{H}_{t+1}) \quad \text{where} \quad \mathbf{H}_{t+1} = \begin{bmatrix} h_{dt+1} & h_{ct+1} \\ h_{ct+1} & h_{ft+1} \end{bmatrix} \quad \text{and} \quad \mathbf{H}_{t+1} = \mathbf{P'P} + \mathbf{F'H}_t\mathbf{F} + \mathbf{G'}\varepsilon_t\varepsilon_t'\mathbf{G}$$

Panel A: Conditional expected excess returns

	U.S. excess returns					Foreign excess returns						
Model	α_d	θ_{d1}	θ_{d2}	β_{dv}	β_{dc}	α_f	θ_{f1}	θ_{f2}	ϕ_f	β_{fc}	β_{fv}	Log-likelihood
1. Nikkei$	0.0362	0.0516	−0.0062	−3.555	87.38	−0.0205	0.0769	0.0104	0.2028	27.57	5.730	18198.926
1979–89	(0.76)	(2.62)[a]	(−0.31)	(−0.41)	(3.07)[a]	(−0.49)	(3.82)[a]	(0.53)	(10.92)[a]	(2.38)[a]	(0.25)	$R_1^2 = 1.16\%$
	[1.16]	[2.55][a]	[−0.32]	[−0.58]	[2.34][a]	[−0.47]	[3.42][a]	[0.49]	[8.95][a]	[2.36][a]	[0.28]	$R_2^2 = 4.52\%$
2. MSJP	−0.0035	0.0387	−0.0057	2.791	56.01	0.0335	0.1453	0.0232	0.1692	9.289	53.44	15559.932
1980–89	(−0.07)	(1.85)[b]	(−0.27)	(0.32)	(2.40)[a]	(1.28)	(6.29)[a]	(1.11)	(10.98)[a]	(1.08)	(2.32)[a]	$R_1^2 = 0.95\%$
	[−0.13]	[1.81][b]	[−0.29]	[0.49]	[1.76][b]	[1.37]	[5.61][a]	[0.88]	[9.68][a]	[0.77]	[1.79][b]	$R_2^2 = 7.27\%$
3. EAFE	0.0230	0.0526	0.0018	1.722	57.30	0.0329	0.0735	0.0252	0.2871	18.95	−10.91	16855.220
1980–89	(0.47)	(2.63)[a]	(0.09)	(0.13)	(2.57)[a]	(0.73)	(3.23)[a]	(1.08)	(18.59)[a]	(1.52)	(−0.31)	$R_1^2 = 0.67\%$
	[0.57]	[2.74][a]	[0.21]	[0.16]	[1.60]	[0.76]	[3.32][a]	[1.24]	[15.32][a]	[1.55]	[−0.30]	$R_2^2 = 11.1\%$
4. Nikkei yen	−0.0060	0.0508	−0.0047	2.044	72.93	0.0196	0.0801	0.0026	0.1694	26.16	35.33	19397.461
1979–89	(−0.12)	(2.65)[a]	(−0.24)	(0.24)	(3.20)[a]	(1.12)	(4.03)[a]	(0.14)	(14.58)[a]	(2.55)[a]	(1.86)[b]	$R_1^2 = 0.56\%$
	[−1.65][b]	[2.56][a]	[−0.24]	[0.59]	[1.91][a]	[1.05]	[3.39][a]	[0.11]	[12.43][a]	[2.12][a]	[1.44]	$R_2^2 = 6.13\%$

Panel B: Conditional covariance dynamics

Model	**P**			**G**				**F**			
	P_{11}	P_{12}	P_{22}	G_{11}	G_{21}	G_{12}	G_{22}	F_{11}	F_{21}	F_{12}	F_{22}
1. Nikkei$	0.1359	−0.0629	0.4334	−0.1842	0.0136	−0.1402	−0.3344	0.9707	0.0217	−0.0204	0.8332
	(10.6)[a]	(−0.87)	(14.4)[a]	(−15.9)[a]	(0.82)	(−7.49)[a]	(−16.2)[a]	(291.0)[a]	(1.96)[a]	(−1.97)[a]	(41.2)[a]
	[5.08][a]	[−0.86]	[9.17][a]	[−6.57][a]	[0.90]	[−3.87][a]	[−10.6][a]	[138.0][a]	[3.29][a]	[−1.41]	[27.1][a]
2. MSJP	0.1184	−0.1182	0.1609	−0.1594	−0.0342	0.0664	−0.3869	0.9791	−0.0009	0.0219	0.8966
	(7.53)[a]	(−3.09)[a]	(6.19)[a]	(−16.6)[a]	(−2.36)[a]	(6.21)[a]	(−23.7)[a]	(292.0)[a]	(−0.142)	(4.37)[a]	(107.0)[a]
	[4.57][a]	[−1.64][b]	[2.29][a]	[−5.67][a]	[−3.25][a]	[1.63]	[−11.8][a]	[169.0][a]	[−16.5][a]	[1.61]	[59.4][a]
3. EAFE	0.1043	0.0080	0.2451	−0.1575	0.0134	−0.0496	−0.3006	0.9792	0.0177	−0.0095	0.9060
	(7.91)[a]	(0.178)	(10.5)[a]	(−17.4)[a]	(0.774)	(−4.02)[a]	(−15.6)[a]	(413.0)[a]	(2.15)[a]	(−1.83)[b]	(64.8)[a]
	[3.57][a]	[0.104]	[8.45][a]	[−4.23][a]	[0.545]	[−1.42]	[−8.76][a]	[132.0][a]	[1.25]	[−0.763]	[49.8][a]
4. Nikkei yen	0.1143	−0.0885	0.1619	−0.1627	−0.0671	0.0445	−0.4050	0.9786	−0.0139	0.0170	0.8812
	(7.29)[a]	(−2.57)[a]	(8.59)[a]	(−15.5)[a]	(−3.92)[a]	(4.22)[a]	(−27.1)[a]	(281.0)[a]	(−1.57)	(3.30)[a]	(119.0)[a]
	[4.09][a]	[−1.46]	[4.35][a]	[−6.86][a]	[−4.44][a]	[1.97][b]	[−12.7][a]	[163.0][a]	[−4.63][a]	[1.95][b]	[49.5][a]

[a]Significant at the 5 percent level

[b]Significant at the 10 percent level.

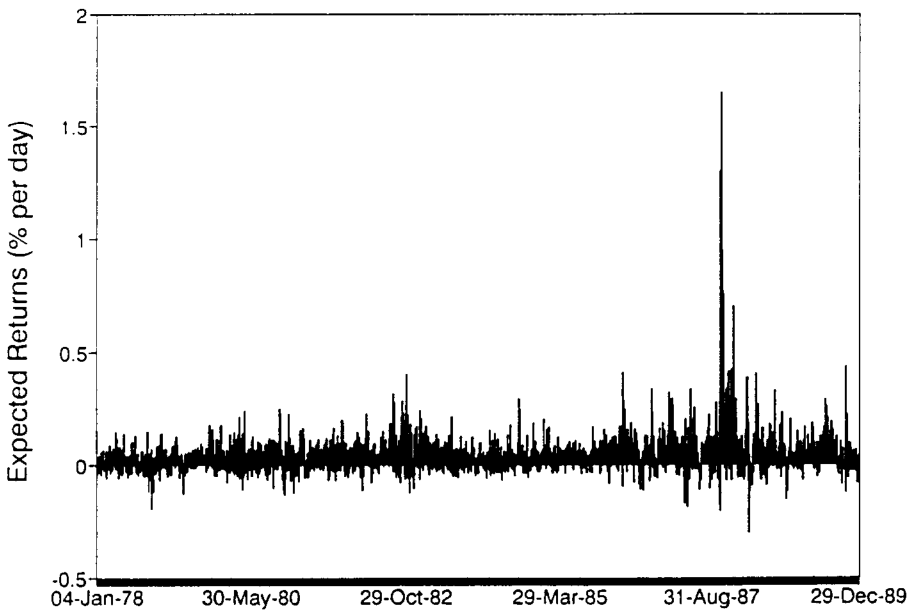

Fig. 2. Conditional expected risk premium on U.S. equity from 1978 to 1989. The fitted values for the expected risk premium of S&P 500 stocks are computed from estimates of the bivariate GARCH model using the dollar-denominated Nikkei 225 index returns as the foreign market (table 2, panel A). The figure excludes the four days around the October 1987 market crash (October 16, 19, 20, and 21).

a standard deviation of 0.97 percent. Further, the sample mean and standard deviation of the daily conditional variance are, respectively, 0.94 and 0.54, whereas for the covariance they are 0.09 and 0.22. As a measure of the economic significance of the international effect on the U.S. risk premium, our results suggest that a doubling of the conditional covariance from its mean value of 0.09 to 0.18 doubles the conditional expected excess return on the U.S. portfolio from 0.04 percent to 0.075 percent. A doubling of the conditional covariance is an increase of slightly less than half the standard deviation of the conditional covariance, indicating that such an increase is not unusual. Similarly, a doubling of the conditional risk premium is also an increase of slightly less than half its standard deviation.

Panel B of table 2 presents the estimates of the coefficients of the dynamics of the covariances from (6). The individual coefficient estimates are difficult to interpret, but the significance of the diagonal coefficients of the F and G matrices suggests that the GARCH effects are pervasive and strong. Moreover, most of the off-diagonal coefficients are also significant. This result confirms the evidence of Hamao, Masulis, and Ng (1990) that modeling the cross-market

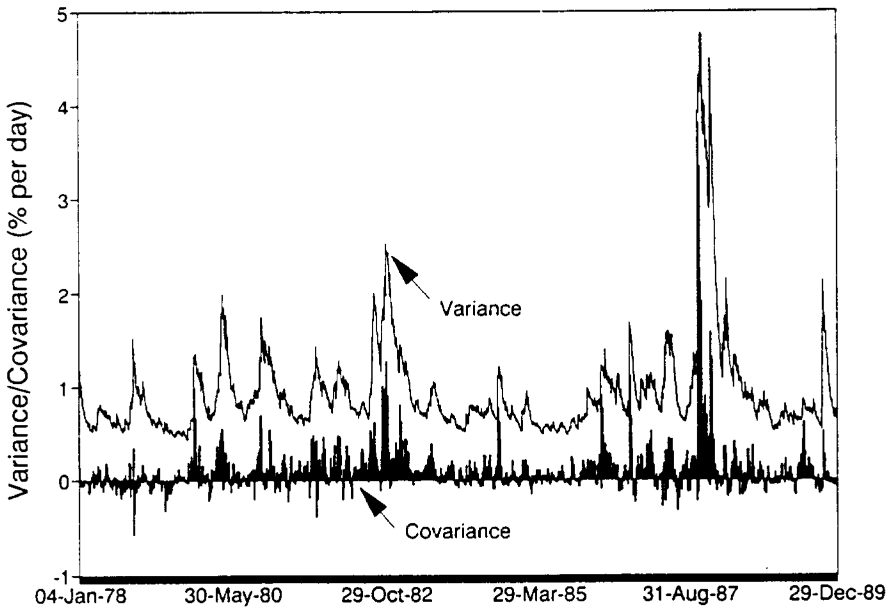

Fig. 3. Conditional variance and covariance of the risk premium on U.S. equity from 1978 to 1989. The fitted values for the conditional variance (*line*) and covariance (*solid bar*) of the risk premium of S&P 500 stocks are computed from estimates of the bivariate GARCH model using the dollar-denominated Nikkei 225 index returns as the foreign market (table 2, panel B). The figure excludes the four days around the October 1987 market crash (October 16, 19, 20, and 21).

dependence in the conditional volatility of U.S. and foreign stock returns is important.

Table 3 provides evidence on whether the coefficients on conditional variance and covariances of table 2 conform to the predictions of the international CAPM (ICAPM). The ICAPM implies that all the β coefficients of table 2 should be equal.[9] Table 3 presents estimates of the models of table 2 with the β coefficients constrained to be equal. Irrespective of the foreign portfolio used, the world price of risk is positive and significantly different from zero. With the ICAPM, the estimate for the price of risk corresponds to the representative investor's coefficient of relative risk aversion. The estimates for this coefficient vary from 9.6 to 18.1, depending on the foreign portfolio used. Although the estimate is high, it is not altogether unreasonable in light of the literature. For instance, although larger than the estimates obtained by Friend and Blume (1975) and Hansen and Singleton (1983), it is substantially smaller than those

[9]In footnote 7, we argue that our approach may yield measures that understate the true covariance. If it does, our test of the ICAPM is more conservative, because our estimate of the covariance effect, β_{dc}, is biased upward and hence away from the other β_{ij} coefficients.

Table 3

Tests of the international CAPM and alternative specifications using the bivariate model of daily expected excess returns for U.S. and foreign equity markets.

U.S. equity returns are given by Standard and Poor's 500 stock index and the foreign equity market returns by the U.S. dollar- and yen-denominated Nikkei 225 index (model 1 and 4), the Morgan Stanley Japan index (MSJP) (model 2), and the Morgan Stanley EAFE index (model 3). The following restrictions are tested:

ICAPM model: $\beta_{dv} = \beta_{dc} = \beta_{fv} = \beta_{fc}$

Alternative 1: $\beta_{dv} = \beta_{dc}$, $\beta_{fv} = \beta_{fc}$

Alternative 2: $\beta_{dv} = \beta_{fc}$, $\beta_{fv} = \beta_{dc}$

Alternative 1 assumes that the prices of risk in the U.S. and foreign market differ due to market segmentation; alternative 2 assumes a two-factor model with U.S. and foreign market risk as the respective loadings. Only the coefficient estimates for the conditional variance and covariance terms in the S&P 500 and foreign market excess returns equations are reported, with associated robust t-statistics computed with quasi-maximum-likelihood methods in brackets and χ^2 values of the likelihood ratio test of the restrictions.

Model	Specification	U.S. returns		Foreign returns		χ^2
		β_{dv}	β_{dc}	β_{fv}	β_{fc}	(p-value)
1. Nikkei $ 1979–89	Unrestricted	−3.555 [−0.58]	87.38 [2.34][a]	27.57 [2.36][a]	5.730 [0.28]	
	ICAPM	9.647 [1.90][a]	9.647	9.647	9.647	7.428 (0.0594)
	Alternative 1	7.912 [2.23][a]	7.912	12.54 [1.29]	12.54	7.202 (0.0273)
	Alternative 2	2.492 [0.06]	28.35 [2.50][a]	28.35	2.492	3.752 (0.1532)
2. MSJP 1980–89	Unrestricted	2.791 [0.49]	56.01 [1.76][b]	9.289 [0.77]	53.44 [1.79][b]	
	ICAPM	11.96 [2.01][a]	11.96	11.96	11.96	5.714 (0.1264)
	Alternative 1	8.853 [1.15]	8.853	14.33 [1.21]	14.33	5.294 (0.0709)
	Alternative 2	8.164 [0.86]	15.91 [1.53]	15.91	8.164	5.436 (0.0660)
3. EAFE 1980–89	Unrestricted	1.722 [0.16]	57.30 [1.60]	18.95 [1.55]	−10.91 [−0.30]	
	ICAPM	12.86 [1.75][b]	12.86	12.86	12.86	4.994 (0.1722)
	Alternative 1	17.87 [1.73][b]	17.87	6.820 [0.68]	6.820	3.904 (0.1420)
	Alternative 2	5.585 [0.45]	20.56 [1.73][b]	20.56	5.585	4.324 (0.1151)

Table 3 (continued)

Model	Specification	U.S. returns		Foreign returns		χ^2
		β_{dv}	β_{dc}	β_{fv}	β_{fc}	(p-value)
4. Nikkei yen 1979–89	Unrestricted	2.044 [0.59]	72.93 [1.91][a]	26.16 [2.12][a]	35.33 [1.44]	
	ICAPM	18.06 [2.91][a]	18.06	18.06	18.06	6.675 (0.0830)
	Alternative 1	12.09 [1.58]	12.09	23.02 [2.87][a]	23.02	4.984 (0.0827)
	Alternative 2	7.582 [1.41]	32.53 [2.92][a]	32.53	7.582	3.844 (0.1463)

[a]Significant at the 5% level.
[b]Significant at the 10% level.

presented by Wheatley (1988) and comparable to those of Harvey (1991) in tests of international asset pricing models.

The last column of table 3 presents χ^2 statistics that compare the unrestricted and restricted versions of the model. The restricted version cannot be rejected at the 5 percent level. One might argue that a more conservative approach is warranted, however, since we have only two assets. If we use the 10 percent significance level as our benchmark, we cannot reject the ICAPM when the foreign portfolio is EAFE or Morgan Stanley Japan; we can, however, reject the ICAPM if the foreign portfolio is the Nikkei, whether denominated in dollars or in yen. This limited success with the Nikkei may reflect the fact that the Nikkei is not a value-weighted index and hence gives more weight to small stocks. Hence, although our results support the ICAPM, the lack of rejection may also reflect the limited power of our tests.

4. Further tests

In this section, we first relate the results to earlier work. We then examine the robustness of the results and conclude with more tests of the international capital asset pricing model.

4.1. Relation to previous work

French, Schwert, and Stambaugh (1987) use a univariate GARCH-in-mean process to relate conditional expected returns on the S&P 500 index to the conditional volatility of the index and find a positive relation. Their test is not nested in the test discussed in section 3, since they estimate the process followed by the S&P 500 using U.S. data only. We replicated the French, Schwert, and Stambaugh study over our sample periods for the U.S. and Japan, but do not present the estimates here.[10] Their result does not hold for U.S. returns, as the conditional expected returns on the U.S. market portfolio are not significantly related to the conditional variance of the returns of that portfolio.[11] In contrast, however, there is evidence that the conditional expected returns of the foreign indices are related to their conditional volatility.[12] These results are consistent with the results obtained in table 2, since in that table the coefficient on the conditional volatility of the U.S. index in the U.S. returns equation was insignificant, whereas the coefficient on the conditional volatility of the foreign index in the foreign returns equation was significant for the Nikkei 225 whether denominated in yen or in dollars.

Bollerslev, Engle, and Wooldridge (1988) investigate a trivariate conditional CAPM that includes three portfolios: a portfolio of stocks, a portfolio of long-term bonds, and a portfolio of short-term bonds. Their approach is similar to ours in that they model the joint distribution of returns using a GARCH-in-mean process, but they posit dynamics for the variance–covariance matrix of returns that allow variances and covariances to depend only on their past values, in contrast to the dynamics we posit, which allow for additional cross-market dependence. They find a positive price of covariance risk, which is analogous to our finding of a positive price of risk in the equations estimated in table 3.

Existing research on international asset pricing has focused on unconditional returns. The most notable exception is Harvey (1991), who allows the price of risk to change over time in a pooled cross-sectional time-series test using monthly returns for a variety of countries. He models asset returns dynamics with instrumental variables and uses a generalized method-of-moments estimator to test the restrictions of the ICAPM. He finds evidence that the world price of covariance risk is not constant and rejects the hypothesis that Japanese stocks are priced according to the ICAPM.

4.2. Robustness and stability of the international effect

4.2.1. Nonsynchronous trading hours

As discussed in section 2, the use of daily data raises the issue of lack of synchronism in the returns measurement intervals across countries. There is no way to eliminate this lack of synchronism, but using longer measurement intervals should reduce its importance, since as the interval increases, the

[10] French, Schwert, and Stambaugh estimate a GARCH (1, 2) process, whereas the estimates reproduced in this paper are obtained for a GARCH (1, 1) process. None of the conclusions of this paper are affected if this alternative specification is introduced.

[11] Glosten, Jagannathan, and Runkle (1989) demonstrate the instability of the relation estimated by French, Schwert, and Stambaugh for different specifications of the conditional variance.

[12] See Chan and Karolyi (1991) for similar findings on the Nikkei 225 Stock Average index.

fraction of the interval for which returns are observed on both markets also increases. So, one way to evaluate the robustness of our results and the appropriateness of our approach for dealing with the nonsynchronous trading hours is to reestimate our equations on longer observation intervals. Results with the dollar-denominated Nikkei excess returns are reproduced in table 4 for two-day, three-day, and five-day returns. It is clear that our conclusion that the conditional U.S. expected excess returns are positively related to the conditional covariance of the U.S. excess returns with foreign-market returns still holds when we lengthen the observation interval. Importantly, as the returns are computed over longer periods, the apparent predictability of foreign returns from U.S. returns due to the overlapping of the measurement intervals disappears, as expected.[13]

The lack of synchronism in trading times leads to lower estimates of conditional covariances, since the same information can be incorporated in different indices on different calendar days. Although we try to estimate conditional covariances approximating those that would prevail if trading were synchronous across countries, our approach is imperfect. When returns are measured over longer horizons, the nonoverlapping part of the returns becomes a smaller fraction of the measurement interval, so that the conditional covariances should increase.[14] This difference between long-run and short-run conditional covariance does not affect our conclusion that there is an international effect on the U.S. risk premium, since that effect is robust to increases in the measurement interval. Although the coefficient β_{dc} falls when returns are measured over longer intervals, the decrease is not significant.

Another way to reduce the effects of the lack of synchronism in trading hours between the two countries is to use open-to-open U.S. returns with close-to-close Japanese returns. Transactions data of intraday quotes on the S&P 500 from the Chicago Mercantile Exchange were available to us for the 1984–89 subperiod. In table 4, we provide estimates of our model where U.S. returns are computed from prices one hour after the market opening. We use returns one hour after the opening because opening returns have different properties from other intraday returns.[15] These estimates fully support the conclusions drawn from the estimates that use the close-to-close returns. Interestingly, these estimates

[13]As the measurement interval is extended, however, one would expect the estimates to become less precise because the number of independent observations falls. Further, our regression framework becomes less appropriate, since longer-horizon returns may have different time-series properties than shorter-horizon returns. Nevertheless, in a regression not reproduced here, we extended the measurement interval to ten days and found that the β_{dc} coefficient estimate is also statistically significant with a value of 142.8.

[14]See Bailey and Stulz (1990) for evidence on how covariance estimates between U.S. and Japanese returns depend on the measurement interval because of the nonsynchronous trading hours.

[15]See Lin, Engle, and Ito (1991) for evidence of the importance of this distinction for the measurement of spillover effects across markets.

Table 4

Tests for robustness of estimates of the bivariate model of daily expected excess returns for U.S. and foreign equity markets.

U.S. equity returns are given by Standard and Poor's 500 stock index and the foreign market returns by the U.S. dollar-denominated Nikkei 225 index. The estimates for the covariance dynamics are not reported. Standard t-statistics are presented in parentheses and robust t-statistics computed with quasi-maximum-likelihood methods are in brackets. The model parameters are those for the system of equations displayed in table 2 and in the text.

	U.S. return equation					Foreign return equation						
Model	α_d	θ_{d1}	θ_{d2}	β_{dv}	β_{dc}	α_f	θ_{f1}	θ_{f2}	ϕ_f	β_{fv}	β_{fc}	Log-likelihood
Two-day returns	−0.0765	0.0129	−0.0271	2.464	80.59	−0.1309	−0.0156	0.0460	0.1768	26.24	20.03	7886.333
(1411 observations)	(−0.71)	(0.42)	(−1.02)	(0.29)	(3.04)[a]	(−1.21)	(−0.52)	(1.52)	(5.87)[a]	(2.68)[a]	(0.99)	
	[−1.31]	[0.51]	[−0.95]	[0.45]	[2.22][a]	[−0.87]	[−0.42]	[1.41]	[5.51][a]	[2.14][a]	[0.85]	
Three-day returns	−0.3307	−0.0358	−0.0165	11.87	70.27	0.2231	0.0044	0.0176	0.1259	6.450	−11.88	4820.689
(941 observations)	(−1.69)[b]	(−0.96)	(−0.52)	(1.39)	(1.99)[a]	(0.88)	(0.12)	(0.47)	(3.12)[a]	(0.55)	(−0.58)	
	[−3.23][a]	[−1.11]	[−0.45]	[1.93][b]	[2.14][a]	[0.98]	[0.11]	[0.52]	[3.13][a]	[0.65]	[−0.52]	
Five-day returns	−0.3023	−0.0347	−0.0496	10.26	50.72	0.3786	0.0567	0.0796	0.0819	2.436	4.438	2671.088
(564 observations)	(−1.55)	(−0.81)	(−1.18)	(1.46)	(1.86)[b]	(1.42)	(1.17)	(1.71)[b]	(1.74)[b]	(0.21)	(0.30)	
	[−1.13]	[−0.90]	[−1.18]	[1.16]	[1.85][b]	[1.42]	[1.32]	[1.96][a]	[1.62]	[0.23]	[0.28]	
Open open U.S. returns	0.0793	−0.0212	0.0103	−12.61	87.49	0.0198	0.1013	0.0555	0.1339	16.22	49.78	8115.645
(1984/08 1989/12)	(1.29)	(−0.76)	(0.35)	(−1.03)	(2.13)[a]	(0.27)	(3.34)[a]	(1.95)[b]	(4.88)[a]	(1.04)	(1.52)	
(1264 observations)	[0.77]	[−0.41]	[0.09]	[−0.45]	[1.64][b]	[0.19]	[2.96][a]	[1.67][b]	[3.84][a]	[1.17]	[1.34]	
Holiday/weekend dummy	0.0267	0.0708	−0.0058	3.014	58.51	−0.0535	0.0498	−0.0071	0.2214	−2.875	38.17	18150.384
	(0.56)	(3.59)[a]	(−0.31)	(0.36)	(2.58)[a]	(−1.71)	(2.32)[a]	(−0.31)	(12.6)[a]	(−0.17)	(4.54)[a]	
	[1.00]	[3.73][a]	[−0.29]	[0.54]	[1.59]	[−1.18]	[2.09][a]	[−0.24]	[9.86][a]	[−0.12]	[3.13][a]	
No MA coefficients	0.0129			0.556	91.61	−0.0222				28.46	−1.694	18037.485
	(0.29)			(0.07)	(4.29)[a]	(−0.63)				(3.49)[a]	(−0.08)	
	[0.32]			[0.12]	[3.04][a]	[−0.41]				[2.04][a]	[−0.10]	
No market weights	0.0181	0.0512	−0.0075	−0.908	66.99	−0.0351	0.0781	0.0111	0.2036	20.84	19.94	18197.939
	(0.38)	(2.59)[a]	(−0.37)	(−0.08)	(2.67)[a]	(−0.51)	(3.85)[a]	(0.56)	(11.0)[a]	(1.47)	(0.65)	
	[0.72]	[2.52][a]	[−0.38]	[−0.13]	[2.22][a]	[−0.53]	[3.48][a]	[0.53]	[9.01]a	[1.52]	[0.59]	

[a]Significant at the 5 percent level.
[b]Significant at the 10 percent level.

correspond to a subperiod that approximates the second half of our sample of Nikkei data. Yet for this second half, the coefficient for the international effect is essentially the same as for the whole sample when we use close-to-close returns.

4.2.2. Conditioning information

The model estimated in table 2 includes lagged disturbances to account for the effects of nonsynchronous trading of component stocks in the index on the measurement of stock index returns. The significant positive effect of the first lagged disturbance is consistent with our infrequent trading motivation for the inclusion of these lagged disturbance terms.[16] Ignoring this lag structure may spuriously increase the importance of the international effect. This is shown in table 4, where we provide estimates of a version of our model that excludes the lagged disturbances. We also estimated the model with ϕ_f set equal to zero. These estimates, though not reproduced here, show a stronger international effect when ϕ_f is not included. Another alternative we examined but do not report was to roll back the investors' conditioning information by one day, so that ε_{t-1} replaced ε_t in (5) and (6); this approach still produced similar results. Rolling back the conditioning information set has three purposes. First, it is an attempt to see how the results change with changes in the information set.[17] Second, it shows the effect of removing the spurious predictability of foreign returns due to the lack of synchronism in trading hours. Finally, it shows that the covariance effect found for the U.S. expected excess returns is unlikely to be due to asynchronous trading in the S&P 500 stock index. When we roll back the conditioning information by one calendar day, the conditional covariance estimate does not incorporate information from the previous day and is not likely to be affected by information that may not have been fully incorporated into stock prices because of asynchronous trading. In summary, our results on the existence of an international effect are not sensitive to the lag structure included in the estimated models.

4.2.3. Day-of-the-week effects

The estimates in table 2 use trading time as the measurement unit for returns; i.e., all returns are computed from the end of a trading period to the end of the next trading period. In table 4, we provided model estimates where we allow different intercepts in the expected excess returns and the variance–covariance

[16]One would expect the foreign indices to have more evidence of infrequent trading, so the first lagged disturbance should have a greater effect on the foreign returns equations. It is puzzling, however, that the first lagged disturbance should have a greater effect in the foreign equation using the value-weighted Morgan Stanley index than on the equations using the price-weighted Nikkei 225.

[17]Our econometric models attempt to approximate (1). These models are approximations because we condition on an observed information set that is necessarily coarser than Ω_t, the true information set.

equations for multiple calendar days because of weekends or holidays. The results are similar, although the robust t-statistic on the international effect falls to 1.59. In this case, however, the coefficient on the first lagged disturbance becomes strongly significant for both countries and the coefficient in the Nikkei equation for the lagged U.S. return increases in value.

4.2.4. *Alternative specifications*

Multivariate GARCH is a full information maximum likelihood (FIML) estimation method. Although it is more efficient than conventional instrumental variables methods, its parameter estimates can be unstable if the model is misspecified. To investigate whether our inference about the foreign effect is sensitive to these estimation methods, we also use a two-pass procedure. We first estimate the model assuming constant conditional means to generate the conditional variance and covariance series and then regress the excess returns on the conditional variances and covariances using a seemingly unrelated regression (SUR) model. The conclusions are similar to those with the GARCH-in-mean approach. In particular, for the model with dollar-denominated Nikkei returns, using this two-pass approach, the coefficient β_{dc} is 54.02 with a heteroskedasticity-consistent t-statistic of 3.10.

Since the expression in (6) for the dynamics of the variance–covariance matrix is a projection equation without theoretical foundation, we also estimated our model using alternative specifications. One version uses (6) without imposing the BEKK constraints on the coefficient matrices. This unrestricted version, however, does not ensure that the variance–covariance matrix is positive semi-definite. With this version, the point estimates for the variance and covariance effects on the mean returns were similar to those from the BEKK process but less precise. The lack of precision was probably due to the larger number of parameters required for implementation. We also estimated the model using the constant correlation model of Schwert and Seguin (1990) and the Bollerslev, Engle, and Wooldridge (1988) process. The coefficients are again similar to those reported in table 2 except for the coefficient on the conditional covariance in the U.S. returns equation, which is 174 [with a standard (robust) t-statistic of 1.81 (1.40)] for the Bollerslev, Engle, and Wooldridge process and 284 [with a standard (robust) t-statistic of 1.44 (1.40)] for the Schwert and Seguin model. Though these additional formulations impose strong restrictions on the covariance dynamics – constant correlation in Schwert–Seguin and no cross-market spillover effects in BEW – the covariance effect is still important.

The market weights used in the estimation are constructed dynamically from quarterly observations provided by Morgan Stanley's *Capital International Perspectives*. We investigated the sensitivity of our results to the estimates for the market weights by reestimating the model assuming constant weights of 0.5.

These estimates are reproduced in table 3 and show that the international effect holds strongly with constant market weights.

4.2.5. Residual diagnostics

Table 5 reports some residual diagnostic tests for the bivariate and univariate models in this paper. Panel A presents the cross-sectional and time-series statistics for the scaled residuals of the bivariate models for the U.S. and foreign excess returns. The residuals are scaled by the square root of the conditional variance, $\hat{\varepsilon}_{it}/\sqrt{h_{it}}$. Tests indicate that the average standardized residuals are small compared with their standard deviation (using conventional t-statistics, they would not be significantly different from zero), suggesting that we should not be overly concerned that overfitting might lead to a bias in our estimates of the conditional means. The significant positive excess kurtosis and negative skewness observed for the raw returns in table 1 are reduced considerably, but the Kolmogorov D and Bera–Jarque test statistics for normality still indicate significant deviations from normality, suggesting that our focus on t-statistics robust to deviations from normality is warranted.[18] Panel C demonstrates that the cross-correlation patterns noted in the raw returns of table 1 are mostly absorbed by the bivariate model. In general, only contemporaneous correlations between the foreign and U.S. returns appear to be significantly different from zero.

4.3. Additional asset pricing tests

In table 3, we are unable to reject the ICAPM at the 5 percent level, but can do so at the 10 percent level for the Nikkei indices. A better way to understand our results is to compare them with tests of simple two-factor models. One can think of two possible such models in the context of this paper. First, there could be barriers to international investment that limit investors' abilities to benefit from international diversification. We refer to this version as the segmented model. Second, investors might not use a mean-variance framework, but rather hold securities also for the purpose of hedging against unanticipated changes in various state variables; we call this the hedging model. In the latter case, we posit the domestic and foreign market excess returns as factor-mimicking portfolios.

With segmented markets, the price of risk can differ across countries. In table 3, we report estimates of a model that allows for this variation. In this model, the coefficients on the weighted conditional variances and covariances are constrained to be the same within countries, but not across countries. This

[18] Nelson (1992) and Nelson and Foster (1991a,b) provide some conditions under which ARCH models may perform reasonably in estimating and even in forecasting conditional covariances even when some evidence of model misspecification (e.g., nonnormality in standardized residuals) remains.

Table 5

Residual diagnostics for bivariate GARCH models for daily U.S. and foreign equity market excess returns from January 1978 to December 1989. The corresponding model estimates are from table 2.

The U.S. equity index is Standard and Poor's 500 stock index and the foreign market is the U.S. dollar- and yen-denominated Nikkei 225 index, the Morgan Stanley Japan index, or the Morgan Stanley EAFE index. The Kolmogorov–Smirnoff D statistic tests the null hypothesis of normality with critical values of 0.0256, 0.0281, and 0.0271 at the 5 percent significance level and 0.0307, 0.0337, and 0.0325 at the 1 percent significance level for 2819, 2338, and 2522 degrees of freedom, respectively. The Bera–Jarque B test statistic for normality is based on the excess skewness and kurtosis coefficients and is asymptotically distributed χ^2 with two degrees of freedom with critical values at the 5 percent significance level of 5.99 and at the 1 percent significance level of 9.21. The tests for deviations from normality for the skewness and kurtosis statistics are based on D'Agostino, Belanger, and D'Agostino (1990). ρ_j denotes the autocorrelation coefficients of order j for both the raw and squared standardized residuals. Cross-correlations are given between the residuals from the S&P 500 index daily excess returns, $\hat{\varepsilon}_t$ (and squared returns, $\hat{\varepsilon}_t^2$), and those of the index shown in the table.

Panel A: Bivariate model standardized residuals

	Model 1		Model 2		Model 3		Model 4	
Statistic	S&P 500	Nikkei ($)	S&P 500	MS Japan	S&P 500	MS EAFE	S&P 500	Nikkei (yen)
Nobs.	2819	2819	2338	2338	2522	2522	2819	2819
Mean	−0.0148	−0.0124	−0.0082	−0.0327	−0.0154	−0.0233	−0.0044	−0.0335
Std. dev.	0.9915	0.9936	0.9916	0.9905	0.9956	0.9948	0.9886	0.9904
Skewness	−0.4119[a]	0.1256	−0.4523[a]	0.0479	−0.4529[a]	0.0608	−0.4274[a]	−0.3898[a]
Kurtosis	4.1108[a]	2.0331[a]	4.2673[a]	3.8949[a]	4.1796[a]	1.8164[a]	4.2162[a]	2.5917[a]
Kolmogorov D	0.0378[a]	0.0379[a]	0.0411[a]	0.0547[a]	0.0368[a]	0.0393[a]	0.0408[a]	0.0461[a]
Bera–Jarque B	38.886[a]	9.2839	38.336[a]	30.582[a]	38.270[a]	6.9354	40.940[a]	16.204[a]
				Series: $\hat{\varepsilon}_t$				
ρ_1	0.0327	0.0091	0.0326	0.0051	0.0199	0.0086	0.0292	0.0003
ρ_2	0.0153	0.0109	0.0179	−0.0002	0.0191	0.0099	0.0136	0.0056
ρ_3	−0.0094	0.0209	−0.0257	0.0071	−0.0145	0.0084	−0.0112	0.0088
ρ_4	−0.0085	0.0309	−0.0109	0.0019	−0.0268	0.0351	−0.0091	0.0053
ρ_5	0.0082	−0.0042	0.0021	−0.0242	−0.0123	0.0252	0.0066	−0.0129
ρ_6	−0.0003	0.0146	0.0022	−0.0009	−0.0013	0.0108	0.0018	−0.0077

Series: $\hat{\varepsilon}_t^2$

ρ_1	0.0015	0.0201	0.0111	− 0.0029	0.0136	0.0079	0.0157	0.0308
ρ_2	0.0125	0.0020	0.0259	0.0191	0.0237	0.0028	0.0198	0.0252
ρ_3	0.0015	− 0.0039	0.0201	0.0509[a]	0.0131	0.0195	0.0132	0.0145
ρ_4	0.0099	− 0.0215	0.0258	− 0.0413[a]	0.0326	− 0.0094	0.0201	− 0.0434[a]
ρ_5	− 0.0134	0.0064	− 0.0041	− 0.0185	0.0084	0.0065	− 0.0086	− 0.0269
ρ_6	0.0282	− 0.0342	0.0336	− 0.0175	0.0019	− 0.0292	0.0307	− 0.0442[a]

Panel B: Cross-correlations of standardized residuals of bivariate models of S&P 500 and foreign market returns

	Model 1 Nikkei ($)		Model 2 MS Japan		Model 3 MS EAFE		Model 4 Nikkei (yen)	
Lag	$\hat{\varepsilon}_t$	$\hat{\varepsilon}_t^2$	$\hat{\varepsilon}_t$	$\hat{\varepsilon}_t^2$	$\hat{\varepsilon}_t$	$\hat{\varepsilon}_t^2$	$\hat{\varepsilon}_t$	$\hat{\varepsilon}_t^2$
− 6	− 0.0009	0.0018	− 0.0018	− 0.0085	0.0209	− 0.0147	− 0.0123	0.0005
− 5	0.0254	− 0.0091	0.0421[a]	0.0088	0.0169	0.0053	0.0227	− 0.0084
− 4	− 0.0083	− 0.0139	− 0.0144	− 0.0116	0.0200	0.0186	− 0.0073	− 0.0089
− 3	− 0.0013	0.0005	− 0.0102	0.0067	0.0223	0.0287	− 0.0078	0.0257
− 2	0.0457[a]	0.0136	0.0879[a]	0.0370	0.0356	0.0199	0.0571[a]	0.0449[a]
− 1	− 0.0003	0.0166	0.0366	0.0478[a]	0.0103	0.0523[a]	0.0263	0.0410[a]
0	0.0579[a]	0.0959[a]	0.0762[a]	0.0379	0.1349[a]	0.1315[a]	0.0747[a]	0.0686[a]
1	0.0072	0.0559[a]	0.0099	0.1350[a]	− 0.0047	0.0346	0.0035	0.0687[a]
2	0.0086	0.0183	0.0060	0.0266	0.0285	0.0223	0.0157	0.0298
3	0.0549[a]	− 0.0054	0.0371	− 0.0160	0.0318	0.0281	0.0420[a]	− 0.0195
4	− 0.0174	0.0323	0.0052	0.0125	− 0.0278	− 0.0065	− 0.0155	0.0158
5	− 0.0143	0.0274	− 0.0261	0.0611[a]	− 0.0008	0.0078	− 0.0205	0.0255
6	− 0.0109	− 0.0082	− 0.0093	0.0151	0.0151	0.0172	− 0.0264	0.0116

[a]Significant at the 1 percent level.

model does not appear to perform better than the ICAPM, so the weak performance of the ICAPM is apparently not due to barriers to international investment.

The hedging model does perform better than the ICAPM. With this model, the coefficients on the weighted domestic conditional variance (covariance) and the foreign weighted conditional covariance (variance) are constrained to be the same. At the 10 percent level, the hedging model is rejected only when we use the Morgan Stanley Japan index. These results suggest that more work on a multi-factor hedging model could lead to a better understanding of how risky assets are priced in an international setting. Evidence by Cumby (1990) and Wheatley (1988) on the Stulz (1981) international consumption asset pricing model supports this conjecture. Additional evidence is provided by Campbell and Hamao (1992) and Bekaert and Hodrick (1992) in their work showing that changes in expected returns can be captured by models with more than one latent variable.

So far, we have interpreted our tests of the ICAPM in terms of alternatives positing that an asset's risk is not measured simply by the conditional beta of its returns in relation to the world market portfolio. Since all our tests assume that the world relative risk aversion is constant, an alternative to the models in this section is one in which the world relative risk aversion changes over time.[19] Harvey (1991) pursues this approach using instrumental variables in a generalized method-of-moments framework. With daily data, an approach that models the dynamics for the price of risk explicitly will be difficult to implement because the instruments used in Harvey and related analyses are typically not observed daily. Nevertheless, in the spirit of Harvey we investigated the hypothesis that the price of risk changes during our sample period by allowing it to differ in the two halves of our sample when the index uses dollar-denominated returns on the Nikkei 225. We find that it is indeed significantly higher in the second half of the sample, suggesting that further research should model the time variation of the price of risk explicitly.

5. Concluding remarks

We find that there is a significant foreign influence on the time-varying risk premium for U.S. stocks. We investigate the relation between the conditional expected excess returns on the S&P 500 portfolio and its conditional risk as measured by its conditional volatility and its conditional covariance with the returns on some index of foreign stocks. We find that the conditional covariance of the S&P 500 returns with the returns of foreign stocks is significantly positively related to the conditional expected excess returns of the S&P 500 over

[19]McCurdy and Morgan (1991) employ a multivariate GARCH framework that allows for a CAPM-type specification with time variation in the conditional betas. They examine the risk premiums in deviations from uncovered interest rate parity in weekly spot currency prices.

the 1978–89 period, when the returns to foreign stocks are approximated by the returns of the Nikkei 225 index, Japan's Morgan Stanley index, or, to a lesser extent, by the EAFE index. This result holds whether the model is tested using dollar-denominated excess returns or the yen-denomianted excess returns for the Nikkei 225. In contrast, there is no significant relation between the conditional expected excess return on the S&P 500 and its conditional variance. We show that the international effect is economically significant, since an increase in the daily conditional covariance between U.S. and foreign returns that corresponds to half its sample standard deviation increases the daily U.S. risk premium by about 0.035 percent, or slightly less than one-half of its standard deviation.

We still find an international effect on the U.S. risk premium when we use alternative measurement intervals of the returns and alternative econometric specifications. The effect holds even if we use two-, three-, and five-day returns rather than daily returns and if we use open-to-open returns for the S&P 500 rather than close-to-close returns. Further, our results hold if the elements of the conditional variance–covariance matrix obtained from the GARCH approach are used as explanatory variables in a two-pass approach using seemingly unrelated regression models.

We are unable to reject the international CAPM at the 5 percent level of significance, but can reject it at the 10 percent level when the Nikkei 225 index is used as the foreign index. In general, a two-factor model in which the domestic and foreign indices stand for unspecified factors performs better than the ICAPM. In particular, this two-factor model cannot be rejected at the 10 percent level for the Nikkei indices. Our tests therefore support the hypothesis that markets are internationally integrated over the sample period we consider.

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