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International Asset Pricing and Portfolio Diversification with Time-Varying Risk

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ABSTRACT

We test the conditional capital asset pricing model (CAPM) for the world's eight largest equity markets using a parsimonious generalized autoregressive conditional heteroskedasticity (GARCH) parameterization. Our methodology can be applied simultaneously to many assets and, at the same time, accommodate general dynamics of the conditional moments. The evidence supports most of the pricing restrictions of the model, but some of the variation in risk-adjusted excess returns remains predictable during periods of high interest rates. Our estimates indicate that, although severe market declines are contagious, the expected gains from international diversification for a U.S. investor average 2.11 percent per year and have not significantly declined over the last two decades.

IN THIS ARTICLE, WE TEST a conditional version of the capital asset pricing model (CAPM) in an international setting, and analyze its implications for international portfolio diversification. First, we investigate both the cross-sectional and time-series restrictions of the model using data from the eight largest equity markets in the world. Then, we take the perspective of a U.S. investor and use the estimated model to examine how the ex ante benefits of international diversification have changed over the last two and one-half decades in response to changing conditions in international security markets. In particular, we focus our attention on two issues. First, we analyze the effects of the increasing level of integration among financial markets on the expected gains from international diversification. Second, we study whether, over short horizons, an internationally diversified portfolio provides a good hedge against large declines in the U.S. market.

The evidence shows that the world price of covariance risk is equal across countries and changes over time in a predictable way, whereas the price of

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country-specific risk is not different from zero. This is consistent with the CAPM and supports the hypothesis of international market integration. However, we find that when the price of market risk is restricted to be positive, as the theory implies, some of the variation in residual returns remains predictable. Interestingly, the predictability disappears when we relax the nonnegativity restriction on the price of risk. This might be due to the inability of the model to accommodate negative risk premia during periods of high short-term interest rates. We conclude that although the conditional version of the traditional CAPM provides useful information on the dynamics of market premia, a more adequate model of international asset pricing should probably include additional factors. Along these lines, recent work by Dumas and Solnik (1995) and De Santis and Gérard (1997) suggests that, in international markets, currency risk is priced in addition to market risk.

Our results have interesting implications for investors who want to diversify their portfolios internationally. We find that severe U.S. market declines are contagious at the international level, and often imply a significant reduction in the gains from holding an internationally diversified portfolio. Our estimates indicate that, on average, the expected gains from international diversification are equal to 2.11 percent on an annual basis and have not been significantly affected by the increasing level of integration of international markets.

To implement the tests, we extend the multivariate generalized autoregressive conditional heteroskedasticity (GARCH) parameterization recently proposed by Ding and Engle (1994) to accommodate GARCH-in-Mean effects, which are essential to most asset pricing models. Since this specification is considerably more parsimonious than those used in previous studies, we can test the pricing restrictions of the model and analyze international comovements for a relatively large number of markets. This is an important methodological contribution of our article. Finally, we develop an alternative to the test of the conditional CAPM proposed by Bollerslev, Engle, and Wooldridge (1988). Our specification can be used to test the cross-sectional restrictions of the model on any subset of the assets, even when observations of asset supplies at each point in time are unavailable to the econometrician. Implementation of the test only requires returns on a market-wide index. This specification can easily be extended to accommodate multiple risk factors.

The rest of the article is organized as follows. Section I presents the asset pricing model and the methodology used to test it. Section II describes the data. Section III contains the estimation and test results. Section IV discusses the implications of our findings for international portfolio diversification. Section V concludes the article.

I. Model and Empirical Methods

A. A Model of International Asset Pricing

One of the most widely used models in finance is the CAPM originally derived by Sharpe (1964) and Lintner (1965). In a two-period framework, the model predicts that the expected return on any traded asset, in excess of a risk-free return, is proportional to the systematic risk of the asset, as measured by its covariance with a market-wide portfolio return.

Merton (1973) shows that, in an intertemporal model, economic agents need to hedge against changes in the investment opportunity set. This implies that the expected return on any asset is a function of the covariances between its return and the return on a number of hedging portfolios. A simple specification where the market portfolio is the only pricing factor can be obtained by introducing additional assumptions, for example, by assuming that investors have logarithmic preferences.

We use the conditional CAPM with one factor as our benchmark model and then test alternative specifications. Formally, the asset pricing equations can be written

$$\mathrm{E}(R_{it}|\Im_{t-1}) - R_{ft} = \delta_{t-1}\mathrm{cov}(R_{it}, R_{mt}|\Im_{t-1}) \quad \forall i \tag{1}$$

where R_{it} is the return on asset i between time $t-1$ and t; R_{ft} is the return on a (conditionally) risk-free asset; R_{mt} is the return on the market portfolio and $\Im_{t-1}$ is the set of market-wide information available at the end of time $t-1$. Equation (1) implies that δ_{t-1} can be interpreted as the price of covariance risk. Because the same equation has to hold for the market portfolio, δ_{t-1} is also referred to as the price of market risk.

The same model is often used in an international framework (see, among others, Giovannini and Jorion (1989), Harvey (1991) and Chan, Karolyi, and Stulz (1992)). In most applications, all returns in equation (1) are measured in a common currency, usually the U.S. dollar. This approach assumes that investors do not cover their exposure to exchange rate risk or, equivalently, that the price of exchange risk is equal to zero.[1]

B. Empirical Methods

Equation (1) appears to be the natural relation to use in empirical tests of the CAPM because it takes into account investors' use of newly acquired information in making portfolio decisions. The model requires equation (1) to hold for every asset, including the market portfolio. Therefore, in an economy with N risky assets, the following system of pricing restrictions has to be satisfied, at each point in time

$$\begin{aligned}
\mathrm{E}(R_{1t}|\Im_{t-1}) - R_{ft} &= \delta_{t-1}\mathrm{cov}(R_{1t}, R_{mt}|\Im_{t-1}) \\
&\vdots \\
\mathrm{E}(R_{N-1t}|\Im_{t-1}) - R_{ft} &= \delta_{t-1}\mathrm{cov}(R_{N-1t}, R_{mt}|\Im_{t-1}) \\
\mathrm{E}(R_{mt}|\Im_{t-1}) - R_{ft} &= \delta_{t-1}\mathrm{var}(R_{mt}|\Im_{t-1}).
\end{aligned} \tag{2}$$

[1] A more general framework with optimal hedging is discussed in Solnik (1974), Sercu (1980), Stulz (1981, 1995), and Adler and Dumas (1983).

The system includes only $(N - 1)$ risky securities plus the market portfolio to avoid redundancies. If all the risky assets were included in the system, the last equation would just be a linear combination of the first N equations in each period. In empirical work, any subset of the assets can be used if N is too large. However, the cost of reducing the size of the system is that information on cross-correlations is lost and tests of the asset pricing restrictions imposed by the model have less power.

Formally, let R_t denote the $(N \times 1)$ vector which includes $(N - 1)$ risky assets and the market portfolio. Then, the following system of equations can be used as a benchmark model to test the conditional CAPM:

$$R_t - R_{ft}\iota = \delta_{t-1} h_{Nt} + \epsilon_t \qquad \epsilon_t | \Im_{t-1} \sim N(0, H_t) \tag{3}$$

where ι is an N-dimensional vector of ones, H_t is the $(N \times N)$ conditional covariance matrix of asset returns, h_{Nt} is the Nth column of H_t and contains the conditional covariance of each asset with the market.

Equation (3) follows directly from the conditional CAPM. However, the model does not impose any restrictions on the dynamics of the conditional second moments. GARCH processes can be used to fill this gap and obtain a testable version of the model.

A popular GARCH (1, 1) parameterization for H_t is

$$H_t = C'C + A'\epsilon_{t-1}\epsilon'_{t-1}A + B'H_{t-1}B \tag{4}$$

where C is an $(N \times N)$ symmetric matrix and A and B are $(N \times N)$ matrices of constant coefficients.[2]

The specification in equation (4) is very appealing, given the success of univariate GARCH processes in fitting financial time series. It is also very difficult to estimate due to the large number of unknown parameters. For this reason, most studies that use multivariate GARCH processes limit the analysis to a small number of assets and/or impose several restrictions on the process generating H_t. Often, either the correlations are restricted to be constant (Bollerslev (1990) and Ng (1991)) or both A and B are restricted to be diagonal matrices (see, for example, Bollerslev, Engle, and Wooldridge (1988)). The latter restriction implies that the variances in H_t depend only on past squared residuals and an autoregressive component, while the covariances depend upon past cross-products of residuals and an autoregressive component. We choose the diagonal representation because the constant correlation model appears to be too restrictive for the issues we want to address. In particular, some authors (see, for example, Longin and Solnik (1995) and Karolyi and Stulz (1996)) have suggested that correlations among asset returns change with market conditions, and this feature cannot be accommodated by a constant correlation model. One could argue that the diagonal parameterization is also quite restrictive in light of the cross-market dependences in conditional volatility documented in recent studies (see, for example,

[2] See Engle and Kroner (1995) for a detailed discussion of this parameterization.

Hamao, Masulis, and Ng (1990) and Chan, Karolyi, and Stulz (1992)). However, most of those studies use high frequency data. Because we use returns measured at monthly frequencies, we believe that the spillover in volatility may not be very strong. In the data section we provide some evidence in support of this conjecture.

In this case, equation (4) can be written in a simpler form

$$H_t = C'C + aa' * \epsilon_{t-1}\epsilon'_{t-1} + bb' * H_{t-1} \tag{5}$$

where a and b are $(N \times 1)$ vectors which include the diagonal elements of A and B respectively, and $*$ denotes the Hadamard matrix product (element by element).

Unfortunately, if the researcher is interested in analyzing a relatively large number of assets, even the diagonal model is still difficult to estimate, unless more restrictions are imposed on the process driving H_t. To further reduce the number of parameters, we adopt a specification recently proposed by Ding and Engle (1994) which assumes the process to be covariance stationary. The next subsection describes the details of this parameterization.

C. Parsimonious Multivariate GARCH-in-Mean Parameterization

Consider the following system of equations, in which the conditional expectation of the vector of returns R_t is not explicitly parameterized

$$R_t = E_{t-1}(R_t) + \epsilon_t \qquad \epsilon_t | \Im_{t-1} \sim N(0, H_t)$$

and the dynamics for H_t are as described in equation (5). If the ϵ_t process is covariance stationary, its unconditional variance-covariance matrix is equal to

$$H_0 = C'C * (\iota\iota' - aa' - bb')^{-1}.$$

Ding and Engle suggest to replace $C'C$ in equation (5) with $H_0 * (\iota\iota' - aa' - \mathrm{bb}')$ during estimation. If a consistent estimator of H_0 is available, optimization needs to be performed only with respect to the parameters in a and b, but not with respect to the parameters in $C'C$. In a diagonal system with N assets, this implies that the number of unknown parameters in the conditional variance equation is reduced from $2N + N(N + 1)/2$ to $2N$.[3]

The essential ingredient to implement this methodology is H_0. If there is no GARCH-in-Mean component in the model, estimation can be performed in two stages. First, a series of estimated residuals is obtained from a model without a GARCH correction for the second moments. Then, the residuals from the first stage are used to estimate the multivariate GARCH process.

In most asset pricing applications, however, the pricing restrictions postulate a relation between conditional expected returns and conditional second moments. In this case, the information matrix of the parameters is no longer

[3] In our implementation, with nine assets, the total number of parameters to estimate for the GARCH process is reduced from 63 to 18 by adopting the Ding and Engle specification.

block-diagonal, because of the GARCH-in-Mean feature of the model, and the two-stage estimation procedure suggested by Ding and Engle must be modified.

If investors have rational expectations, the error term that measures the difference between actual returns and their conditional expectations must be orthogonal to the conditioning information. Therefore, the following relation exists between the unconditional covariance matrix of the returns and the unconditional covariance matrix of the residuals

$$\operatorname{cov}(R_t) = \operatorname{cov}[E_{t-1}(R_t)] + H_0.$$

Obviously, H_0 is not directly observable. However, an iterative procedure can be implemented as follows. In the first iteration H_0 is set equal to the sample covariance matrix of the returns. Thereafter it is updated using the covariance matrix of the estimated residuals at the end of each iteration.[4]

In either case, equation (5) is replaced by

$$H_t = H_0 * (\iota\iota' - aa' - bb') + aa' * \epsilon_{t-1}\epsilon_{t-1}' + bb' * H_{t-1}. \qquad (6)$$

The advantages of this multivariate GARCH-in-Mean parameterization are obvious to researchers in asset pricing. First, the reduction in the number of parameters to estimate makes the model applicable to relatively large cross-sections of assets, while preserving flexibility in the dynamics of the conditional second moments. Second, it restricts the unconditional second moments implied by the estimated model to closely match the corresponding sample moments.

Inevitably, parsimony comes at a cost. If the assumption of covariance stationarity does not hold over the sampling period, the inferences are incorrect. Further, some generalizations of standard GARCH processes, like the asymmetric response of conditional second moments to past innovations, can be hard to implement.[5]

An alternative approach to study large systems of assets is the Factor-GARCH model of Engle, Ng, and Rothschild (1990) and King, Sentana and Wadhwani (1994). In this case, the dynamics of the conditional covariance matrix are driven by a limited number of factors that follow GARCH processes. Such a parameterization is a parsimonious alternative to the one proposed here. However, also in this case, parsimony can only be obtained by sacrificing the flexibility of the process. For example, if only one factor is used to model the entire covariance matrix, and the parameters of the model are assumed to be constant, then the chances to detect changes in covariances and differences in volatility across markets are limited (see De Santis (1993)).

[4] The procedure is similar to iterated GMM estimation and is likely to improve the small sample properties of the estimators, even when the two-stage procedure delivers consistent results (see Ferson (1995)).

[5] If the asymmetric response is modeled using information variables not observed at time zero, this parameterization requires the use of truncated multivariate distributions.

We use equations (3) and (6) as our benchmark model. Under the assumption of conditional normality, the log-likelihood function can be written as follows:

$$\ln L(\theta) = -\frac{TN}{2} \ln 2\pi - \frac{1}{2} \sum_{t=1}^{T} \ln|H_t(\theta)| - \frac{1}{2} \sum_{t=1}^{T} \epsilon_t(\theta)' H_t(\theta)^{-1} \epsilon_t(\theta) \qquad (7)$$

where θ is the vector of unknown parameters in the model. Since the normality assumption is often violated in financial time series, we estimate the model and compute all our tests using the quasi-maximum likelihood (QML) approach proposed by Bollerslev and Wooldridge (1992). Under standard regularity conditions, the QML estimator is consistent and asymptotically normal and statistical inferences can be carried out by computing robust LM or Wald statistics. Optimization is performed using the BHHH (Berndt, Hall, Hall, and Hausman (1974)) algorithm.

D. Related Work

Multivariate GARCH models have been used widely to test the pricing restrictions of the conditional CAPM. An incomplete list includes Bollerslev, Engle, and Wooldridge (1988), Giovannini and Jorion (1989), Ng (1991), and Chan, Karolyi, and Stulz (1992). The parameterization used in these studies can be formalized as follows[6]

$$R_t - R_{ft}\iota = \delta H_t \omega_{t-1} + \epsilon_t \qquad \epsilon_t | \Im_{t-1} \sim N(0, H_t) \qquad (8)$$

where ι and H_t are defined earlier, and ω_{t-1} is an $(N \times 1)$ vector of market weights for the risky assets, measured at the end of time $t - 1$.

As discussed earlier, the parsimonious GARCH parameterization that we propose allows us to test the restrictions of the model simultaneously on a large number of assets while preserving the flexibility of the most general GARCH processes. In addition, our specification of the mean equation (equation (3)) differs from equation (8) in that we replace the last element of vector R_t with the return on the market portfolio. This simple variation has two advantages.

First, once the dynamics of H_t are specified, the model in equation (8) can be estimated and tested as long as the market weights are observed by the econometrician at each time t. Often, when market capitalizations are not available, the net supply ω_t of each asset is estimated from multiple sources, thus introducing measurement errors. The parameterization that we propose

[6] This notation does not necessarily reflect all the features of each of the studies cited above. However, it provides a good framework to characterize the common aspects among them.

requires a market index, but not the individual asset weights, to be available to the econometrician at each t.[7]

Second, in our approach, equation (3) can be extended to the case where asset returns depend on multiple risk factors, by augmenting R_t with the returns on factor portfolios and adding the risk premium for each factor to the right-hand side of the equation.[8]

A different approach to testing the conditional version of the international CAPM is proposed by Harvey (1991) who combines the hypothesis of conditional mean-variance efficiency of the world portfolio with the auxiliary assumption that investors use a linear filter to predict asset returns. His test is based on Hansen's (1982) generalized method of moments (GMM) and can be implemented without prespecifying the dynamics of the conditional second moments.

Our approach differs from that of Harvey in two respects. First, in our auxiliary assumption, we parameterize the dynamics of the conditional second moments instead of the first moments. Because a large body of research in finance shows that models that predict second moments have been more successful than models that predict first moments, our approach is likely to have more power.[9] Moreover, using Harvey's method, many variables of interest that are functions of conditional second moments cannot be recovered. Obviously, these variables could be estimated by imposing additional moment restrictions. Unfortunately, it is well known that the properties of GMM-based tests deteriorate quickly as the dimension of the system increases.[10]

Second, we use maximum likelihood to test the model. As long as the model is correctly specified, our approach is more efficient. Of course, the method may yield incorrect inferences if the assumptions underlying the likelihood function are violated. However, this problem can be alleviated by using QML estimates for the standard errors, which are robust to violations of the normality assumption.

In a recent article, Dumas and Solnik (1995) use an instrumental variable approach similar to Harvey's to test whether currency risk is priced in international financial markets, when investors use optimal hedging strategies. In their study, they parameterize only the dynamics of the intertemporal marginal rate of substitution, while leaving conditional first and second moments

[7] In a recent study, Bekaert and Harvey (1995) use a similar approach. However, due to the complex parameterization of their model, their implementation is limited to the world index and one country at a time.

[8] With multiple factors equation (3) would become:

$$R_t - R_{ft}\iota = \delta_{1,t-1}h_{(N-K+1)t} + \delta_{2,t-1}h_{(N-K+2)t} + \cdots + \delta_{K,t-1}h_{Nt} + \epsilon_t, \quad \epsilon_t|\Im_{t-1} \sim N(0, H_t)$$

where R_t denotes the $(N \times 1)$ vector of returns which includes $(N - K)$ risky assets and the K factor mimicking portfolios, $h_{N-K+k,t}$ is the $(N - K + k)$th column of H_t and contains the conditional covariances of each asset with factor mimicking portfolio k.

[9] See Engel, Frankel, Froot, and Rodrigues (1995) for a discussion of this issue.

[10] See, for example, Ferson (1995). For this reason, in most cases Harvey (1991) tests the pricing restrictions for one country at a time against the world portfolio.

unspecified. Also in this case, the methodology is very general and parsimonious; however, the lack of parameterization of the conditional second moments limits the number of questions that can be addressed. For example, as the authors point out, they cannot compare the dimension of the exchange risk premia relative to the reward for market risk. Our methodology can be extended to test their model and answer some of the questions that cannot be addressed by their method. However, since this extension raises nontrivial issues that go beyond the scope of this study, we leave it to future work.[11]

II. Data

We use monthly dollar-denominated returns on stock indices for the G7 countries (Canada, France, Germany, Italy, Japan, United Kingdom, and U.S.A.) and Switzerland, the largest European market not included in the G7. To approximate the market portfolio, we use a value-weighted world index. All the data are from Morgan Stanley Capital International (MSCI) and cover the period from January 1970 through December 1994.[12]

Table I contains summary statistics for the U.S. dollar returns on the eight national indices and on the value-weighted world index. All values are computed in excess of the return on the U.S. T-Bill closest to 30 days to maturity, as reported in the CRSP risk-free files. Panel A in the table contains means, standard deviations, skewness, kurtosis, Bera-Jarque (1982) statistics, and the sample correlation matrix. The index of kurtosis shows that the unconditional distribution of excess returns has heavier tails than the normal for all the countries in our sample. The Bera-Jarque statistic also indicates that the hypothesis of normality is rejected in all instances.

Most correlations in the table are below 0.5 and the average, excluding the world index, is equal to 0.435. These numbers tend to be lower than the correlations between portfolios of U.S. assets. For example, Elton and Gruber (1992) document that, during the period 1980–1988, the correlation between a value-weighted index of the largest 1000 stocks traded in the U.S. and a value-weighted index of the next 2000 largest stocks is equal to 0.92.

Panel B reports autocorrelations for the excess returns and excess returns squared. The lack of statistically significant autocorrelations in the return series reveals that we do not need to include an AR correction in the mean equations. On the other hand, autocorrelation is detected, at short lags, in the squared returns, suggesting that a GARCH parameterization for the second moments might be appropriate.

Finally, the table contains the cross-correlations of squared returns, at different leads and lags, between the world and the other countries. With few exceptions, only the contemporaneous correlations are statistically significant. The same holds for cross-correlations between other pairs of countries. Of the

[11] See De Santis and Gérard (1997).

[12] The MSCI data set has been widely used in previous work. The MSCI methodological notes provide a detailed description of the data.

Table I

Summary Statistics of Excess Returns

Monthly dollar-denominated returns on the equity indices of eight countries, and the value-weighted world index are from MSCI (Morgan Stanley Capital International). Excess returns are obtained by subtracting the return on the United States T-Bill that is closest to 30 days to maturity. R^e_{it} denotes excess returns in percent per month. The sample covers the period January 1970 through December 1994 (300 observations). B–J is the Bera–Jarque test for normality; Q is the Ljung–Box statistic of order 12.

Panel A: Summary Statistics

	Canada	Japan	France	Germany	Italy	Switzerland	United Kingdom	United States	World
Weights									
Jan 70	0.035	0.039	0.023	0.028	0.014	0.011	0.090	0.691	1.00
Dec 82	0.038	0.182	0.013	0.034	0.009	0.018	0.079	0.555	1.00
Dec 94	0.021	0.289	0.033	0.041	0.012	0.030	0.101	0.361	1.00
Mean	0.33	1.00	0.64	0.57	0.23	0.63	0.72	0.38	0.45
Std. Dv.	5.50	6.70	6.97	6.06	7.76	5.60	7.46	4.48	4.22
Skewness	−0.34*	0.18	0.00	−0.09	0.31*	0.01	1.31**	−0.20	−0.35*
Kurtosis[a]	2.07**	0.77**	1.40**	0.88**	0.87**	1.72**	9.90**	2.44**	1.57**
B–J	56.46**	8.37*	22.95**	9.28*	13.47**	23.68**	1264.67**	72.95**	35.23**
Q_{12}	11.90	16.71	12.40	18.93	21.64*	6.73	14.50	9.29	12.86

Unconditional Correlations of R^e_{it}

	Canada	Japan	France	Germany	Italy	Switzerland	United Kingdom	United States	World
Canada	1	0.273	0.423	0.293	0.288	0.464	0.513	0.701	0.702
Japan		1	0.396	0.387	0.381	0.441	0.364	0.269	0.687
France			1	0.599	0.431	0.610	0.540	0.437	0.620
Germany				1	0.379	0.701	0.426	0.348	0.556
Italy					1	0.373	0.337	0.215	0.427
Switzerland						1	0.571	0.504	0.687
United Kingdom							1	0.508	0.684
United States								1	0.828

144 nonsimultaneous correlations (of order −2, −1, 1, 2) between all pairs of indices, only eight are statistically significant, which can be imputed to chance. This evidence suggests that, at least in our sample, cross-market dependences in volatility are not strong and, therefore, the diagonal GARCH parameterization that we adopt is not too restrictive.

Table II contains summary statistics for the conditioning variables. We select a set of instruments that have been widely used in the international asset pricing literature (see, among others, Bekaert and Hodrick (1992), Ferson and Harvey (1993), and Bekaert and Harvey (1995)). In particular, in addition to a January dummy, our instruments include the lagged dividend-price ratio for the world index in excess of the risk-free rate; the month-to-month change in the U.S. term premium, measured by the yield on the ten-year U.S. Treasury note in excess of the one-month T-Bill rate; the U.S.

Table I—*Continued*

Panel B: Autocorrelations and Cross-Correlations

Autocorrelations of R^e_{it}

Lag	Canada	Japan	France	Germany	Italy	Switzerland	United Kingdom	United States	World
1	0.002	0.088	0.080	0.008	0.099	0.062	0.091	0.015	0.095
2	−0.086	−0.008	−0.008	−0.019	−0.038	−0.053	−0.099	−0.032	−0.042
3	0.065	0.086	0.128*	0.089	0.100	0.060	0.062	0.007	0.045
4	−0.036	0.044	0.032	0.073	0.072	0.005	0.014	−0.008	0.004
5	0.094	0.021	0.014	−0.109	0.009	0.027	−0.130*	0.098	0.092
6	0.044	−0.021	0.005	0.011	0.136*	−0.049	−0.043	−0.096	−0.083

Autocorrelations of $(R^e_{it})^2$

Lag	Canada	Japan	France	Germany	Italy	Switzerland	United Kingdom	United States	World
1	0.070	0.125*	0.061	0.176**	0.188**	0.084	0.167*	0.117*	0.066
2	0.228**	0.061	−0.002	−0.033	−0.010	−0.108	0.087	0.050	0.010
3	0.016	−0.000	0.046	0.076	0.022	0.031	0.052	0.129*	0.035
4	0.079	0.024	0.167*	0.058	0.023	0.036	0.029	0.035	0.047
5	0.043	0.053	0.018	0.060	0.088	0.041	0.112	0.002	0.094
6	−0.006	0.053	0.035	0.038	0.035	−0.001	−0.002	0.028	0.033

Cross-correlations of $(R^e_{it})^2$—World and Country j

Lag	Canada	Japan	France	Germany	Italy	Switzerland	United Kingdom	United States
−6	−0.010	−0.084	−0.051	0.025	0.007	−0.013	−0.032	−0.074
−5	0.047	0.069	0.069	0.049	0.078	0.079	−0.093	0.108
−4	−0.035	0.046	−0.007	−0.051	0.025	−0.060	−0.014	−0.001
−3	0.035	0.071	0.116*	0.042	0.066	0.058	0.093	0.011
−2	−0.075	−0.003	−0.030	−0.043	0.048	−0.070	−0.049	−0.048
−1	0.120*	0.143*	0.036	0.061	0.076	0.032	0.054	0.035
0	0.702**	0.687**	0.620**	0.556**	0.427**	0.687**	0.684**	0.828**
1	0.019	0.073	0.083	0.011	0.138*	0.099	0.065	0.099
2	−0.038	−0.015	−0.011	0.006	−0.045	−0.054	−0.040	−0.009
3	0.044	−0.003	0.070	0.016	0.057	0.024	0.030	0.055
4	0.005	0.007	0.043	0.009	0.041	0.014	0.003	−0.031
5	0.007	0.038	0.034	0.075	0.001	0.085	0.053	0.086
6	−0.061	−0.033	−0.081	−0.031	−0.074	−0.085	−0.068	−0.083

Number of significant cross-correlations of order (−2, −1, 1, 2): 8 out of 144.

[a] Equal to zero for the normal distribution.

* and ** denote statistical significance at the 5 percent and 1 percent levels, respectively.

default premium, measured by the yield difference between Moody's BAA and AAA rated bonds; and the month-to-month change in the U.S. interest rate. The table shows that the instruments carry nonredundant information, since their correlations are quite low.

Table II

Summary Statistics of Information Variables

The information set includes the world dividend yield in excess of the return on the United States (U.S.) T-Bill with maturity closest to 30 days (XDPR), the change in U.S. term premium (ΔUSTP), the U.S. default premium (USDP), and the change in the 30-day U.S. T-Bill return (ΔUSRF). The world dividend yield is the dollar-denominated dividend yield on the MSCI world index. The U.S. term premium is equal to the yield on 10-year U.S. T-Notes in excess of the yield of the 3-month U.S. T-Bill. The U.S. default premium is the yield on Moody's BAA rated bonds in excess of the yield on Moody's AAA rated bonds. The sample covers the period January 1970 through December 1994 (300 observations).

					Correlations		
	Mean	Std. Dev.	Min.	Max.	ΔUSTP	USDP	ΔUSRF
XDPR	−0.249	0.189	−0.930	0.029	−0.018	−0.356	0.162
ΔUSTP	0.010	0.476	−2.190	3.330	1	0.135	−0.427
USDP	1.198	0.440	0.580	2.690		1	−0.148
ΔUSRF	−0.001	0.074	−0.467	0.276			1

	Autocorrelations						
Lag	1	2	3	4	5	6	12
XDPR	0.922	0.854	0.792	0.744	0.715	0.694	0.501
ΔUSTP	0.227	−0.150	−0.039	−0.077	−0.087	−0.138	−0.155
USDP	0.952	0.889	0.842	0.806	0.769	0.714	0.479
ΔUSRF	−0.063	−0.051	−0.110	−0.093	−0.032	0.026	0.058

III. Empirical Evidence

As discussed in Section I, the traditional CAPM, applied to international financial markets, postulates that the conditional expected return on any asset is linearly related to the conditional covariance between that asset and the return on a world-wide portfolio. The model, which we refer to as the benchmark model, can be rewritten as follows for convenience

$$R_{it} - R_{ft} = \delta_{t-1}\text{cov}(R_{it}, R_{mt}|\Im_{t-1}) + \epsilon_{it} \qquad \forall i. \tag{9}$$

If international markets are fully integrated and world market risk is the only relevant factor, the price of covariance-risk δ_{t-1} should be positive and equal across all markets.

We compare the performance of the benchmark model to two alternative specifications which can provide useful insights, should the standard CAPM fail to hold. The first alternative allows for some level of market segmentation. In particular, we introduce two changes. First, we assume that country-specific risk, as measured by the conditional variance, is priced in each market in addition to world-wide risk. Second, we introduce a constant in each pricing equation to accommodate other forms of segmentation, like differences in tax treatment or other institutional arrangements that cannot be captured by the benchmark model. Formally, the first alternative to equation (9), which we refer to as the partial segmentation model, is specified as follows:

$$R_{it} - R_{ft} = \alpha_i + \delta_{t-1}\mathrm{cov}(R_{it}, R_{mt}|\Im_{t-1}) + \gamma_i \mathrm{var}(R_{it}|\Im_{t-1}) + \epsilon_{it} \qquad \forall i. \qquad (10)$$

Under the null hypothesis of full market integration both α_i and γ_i must be equal to zero for all markets.

The second alternative examines whether the variation in conditional expected returns is fully explained by world market risk. Let z_{t-1} be a $(k \times 1)$ vector of variables included in $\Im_{t-1}$. We estimate the following system of equations:

$$R_{it} - R_{ft} = \alpha_i + \lambda' z_{t-1} + \delta_{t-1}\mathrm{cov}(R_{it}, R_{mt}|\Im_{t-1}) + \epsilon_{it} \qquad \forall i. \qquad (11)$$

If the traditional CAPM is misspecified, some of the variables in z_{t-1} could have statistical power in explaining the dynamics of the expected returns.

A. The Conditional CAPM with Constant Price of Risk

We start our investigation by introducing the additional assumption that the price of market risk δ_{t-1} is constant. This implies that, although both the conditional risk-free rate and the conditional mean-standard deviation frontier can change in each period, the slope of the capital market line is fixed. This restriction has been imposed in many studies of the conditional CAPM (e.g., Giovannini and Jorion (1989) and Chan, Karolyi, and Stulz (1992)) and, in this sense, it represents an interesting starting point.

Panel A in Table III contains QML estimates of the parameters for the benchmark model. The point estimate for the price of market risk is equal to 2.37,[13] which is reasonable both in size and sign. However, the 1.61 value for the robust standard error is relatively large and implies that the coefficient is statistically significant only at the 15 percent level.

Next, consider the estimated parameters for the GARCH process. With the exception of Germany, all the elements in the vectors a and b are statistically significant at any conventional level. In addition, the estimates satisfy the stationarity conditions for all the variance and covariance processes.[14] As it is typical in most studies that use GARCH models, all processes display high persistence and the estimates of the b_i coefficients (which link second moments to their lagged value) are considerably larger than the corresponding estimates of the a_is (which link second moments to their past innovations).

Panel B includes a variety of diagnostic statistics to evaluate this specification of the model. The average pricing errors are a relatively small fraction of the corresponding average excess returns. However, for each index, the fraction of total variation in the excess returns explained by the model, which we

[13] The number reported in the table is 0.0237 because the estimation uses percentage returns.

[14] Theorem 1 in Bollerslev (1986) implies that for each process in H_t to be covariance stationary, the condition $a_i a_j + b_i b_j < 1 \ \forall i, j$ has to be satisfied.

Table III

Quasi-Maximum Likelihood (QML) Estimates of the Conditional International CAPM with Constant Price of Risk

Estimates are based on monthly dollar-denominated returns from January 1970 through December 1994. Data for the country equity indices and the world portfolio are from MSCI (Morgan Stanley Capital International). The risk-free rate is the return on the U.S. T-Bill with maturity closest to 30 days from Center for Research in Security Prices (CRSP). Each mean equation relates the country index excess return R_{it}^e to its world covariance risk $\text{cov}(R_{it}; R_{mt}|\Im_{t-1})$:

$$R_{it}^e = \delta \text{ cov}(R_{it}; R_{mt}\Im_{t-1}) + \epsilon_{it}$$

where δ denotes the price of world covariance risk and $\epsilon_t|\Im_{t-1} \sim N(0, H_t)$. The conditional covariance matrix H_t is parameterized as follows:

$$H_t = H_0 * (\iota\iota' - aa' - bb') + aa' * \epsilon_{t-1}\epsilon_{t-1}' + bb' * H_{t-1}$$

where $*$ denotes the Hadamard matrix product, a and b are 9×1 vectors of constants and ι is an 9×1 unit vector. QML standard errors are reported in parentheses.

RMSE denotes Root Mean Squared Error. Pseudo R^2s are computed for each index as the ratio between the sum of squared fitted values of the risk premia and the sum of squared excess returns. Since the model restricts the price of risk to be constant across indices, individual pseudo R^2 are not guaranteed to be positive. B–J is the Bera–Jarque test statistic for normality. Q is the p-value of Ljung–Box test statistic of order 12 for standardized residuals (z) and standardized residuals squared (z^2). EN–S− and EN–S+ are Engle–Ng Lagrange multiplier tests for positive and negative size-bias in standardized residuals squared.

Panel A: Parameter Estimates

	Price of Covariance Risk
δ (const)	0.0237 (0.0161)

GARCH Process

	Canada	Japan	France	Germany	Italy	Switzerland	United Kingdom	United States	World
a_i	0.172	0.188	0.240	0.292	0.158	0.176	0.250	0.227	0.206
	(0.036)	(0.025)	(0.064)	(0.073)	(0.065)	(0.075)	(0.038)	(0.032)	(0.031)
b_i	0.962	0.966	0.872	0.703	0.895	0.922	0.947	0.961	0.965
	(0.016)	(0.009)	(0.044)	(0.501)	(0.066)	(0.027)	(0.019)	(0.012)	(0.010)

Panel B: Summary Statistics and Diagnostics for the Residuals

	Canada	Japan	France	Germany	Italy	Switzerland	United Kingdom	United States	World
Avg. Exc. Return	0.33%	1.00	0.64	0.57	0.23	0.63	0.72	0.38	0.45
Avg. pred. Error	−0.06%	0.54	0.20	0.23	−0.10	0.24	0.20	0.01	0.02
RMSE	5.50	6.70	6.96	6.06	7.76	5.61	7.45	4.46	4.22
Pseudo R^2	0.15%	−0.24	0.17	0.07	0.08	0.04	0.43	0.54	0.33
Kurtosis[a]	1.74**	0.64*	1.42**	0.70*	0.91**	1.28**	4.80**	2.05**	1.76**
B–J	41.72**	6.29*	23.57**	6.53*	13.72**	19.24**	290.91**	55.82**	48.99**
$Q_{12}(z)$	0.62	0.08	0.37	0.11	0.05	0.89	0.73	0.75	0.29
$Q_{12}(z^2)$	0.40	0.65	0.56	0.45	0.73	0.40	0.92	0.51	0.81
EN–S−	−0.38	−0.85	−1.32	−1.03	−0.52	−1.43	0.29	−0.85	−0.69
EN–S+	1.00	0.73	−0.58	1.13	1.98*	−0.37	−0.80	−0.58	−1.43
Likelihood function					−7508.60				

[a] Equal to zero for the normal distribution.

* and ** denote statistical significance at the 5 percent and 1 percent levels, respectively.

refer to as pseudo-R^2, is rather low and reaches its highest value of 0.54 percent for the U.S. market.[15] For most countries, the index of kurtosis for the standardized residuals is lower than the corresponding index for the excess returns. The same is true for the Bera-Jarque test statistic. Still, the hypothesis of normality is rejected in most cases. This suggests that, although the GARCH parameterization can accommodate some of the kurtosis in the data, the use of a fat-tailed conditional distribution might improve the performance of the model. This evidence against normality warrants the use of QML inferential procedures in our analysis.

We also construct the estimated standardized residuals ($\epsilon_t h_t^{-1/2}$) and the estimated standardized residuals squared ($\epsilon_t^2 h_t^{-1}$) and compute the Ljung-Box portmanteau statistic for each series to test the hypothesis of absence of autocorrelation up to order 12. The tests show that the null hypothesis is never rejected and that the GARCH (1, 1) parameterization that we adopt is satisfactory.[16]

Finally, we report test statistics for the presence of asymmetric responses to past innovations in the conditional covariance matrix. This phenomenon has been widely studied using univariate GARCH processes. For example, Engle and Ng (1993) find that bad news has a larger effect on volatility than good news. For this reason, we compute the test statistic for signed size-bias in the residuals that they propose. The test rejects the null hypothesis of no positive size-bias only for Italy. This result is not necessarily surprising, because most previous studies on this issue use higher frequency data. It is possible that volatility asymmetries are present in international markets, but they cannot be detected in monthly data.

Since our parameterization for the conditional covariance process has not yet been used in empirical applications, we also compare the estimation results of this specification to those from the Engle and Kroner (1995) diagonal specification. The latter model is widely used, but less parsimonious, because the intercepts of the conditional covariance process must be estimated. For this reason, we have to limit the comparison to a system of four countries plus the world portfolio.[17] The results show that the dynamics of the conditional second moments are essentially unaffected by the parameterization. Interestingly, the main difference between the two approaches is that the unconditional second moments implied by the Engle and Kroner parameterization are often much larger (sometimes more than double) than those implied by our parameterization or, more importantly, than the corresponding sample moments. Although this result may be sample specific, it suggests that an appealing feature of our parameterization is its ability to capture the dynamics of the conditional

[15] Since the model requires the coefficient δ to be equal across all markets, the pseudo-R^2 statistics are not guaranteed to be all nonnegative.

[16] We also find that this specification is selected by the Schwartz information criterion over a set of alternative specifications that go from a constant covariance matrix to a GARCH (2, 2) process.

[17] We include the four largest markets in the MSCI data set: United States, United Kingdom, Japan, and Germany. The results from this experiment are available upon request.

Table IV

Specification Tests of the Conditional International CAPM with Constant Price of Risk

This table reports robust Wald test statistics for two alternative specifications of the conditional international capital asset pricing model (CAPM). The first model (equation A) includes country specific risk and country specific intercepts in addition to market risk. The second model (equation B) includes asset specific intercepts and the instruments in z_{t-1} as independent conditioning variables, in addition to market risk. The instruments include a constant, a January dummy (JAN), the world index dividend yield in excess of the return on the 30-day U.S. T-Bill (XDPR), the change in the U.S. term premium (ΔUSTP), the U.S. default premium (USDP) and the change in the 30-day U.S. T-Bill return (ΔUSRF). Estimation is based on monthly dollar-denominated returns from January 1970 through December 1994. Data for the country equity indices and the world portfolio are from MSCI (Morgan Stanley Capital International). The risk-free rate is the return on the U.S. T-Bill with maturity closest to 30 days from the Center for Research in Security Prices (CRSP).

Eq. A

$$R_{it}^{e} = \alpha_{i} + \delta_{i}\mathrm{cov}(R_{it};\, R_{mt}|\Im_{t-1}) + \gamma_{i}\mathrm{var}(R_{it}|\Im_{t-1}) + \epsilon_{it}$$

Eq. B

$$R_{it}^{e} = \alpha_{i} + \lambda' z_{t-1} + \delta\ \mathrm{cov}(R_{it};\, R_{mt}|\Im_{t-1}) + \epsilon_{it}$$

where δ denotes the price of world covariance risk and $\epsilon_t|\Im_{t-1} \sim N(0, H_t)$. The conditional covariance matrix H_t is parameterized as follows:

$$H_t = H_0 * (\iota\iota' - aa' - bb') + aa' * \epsilon_{t-1}\epsilon_{t-1}' + bb' * H_{t-1}$$

where $*$ denotes the Hadamard matrix product, a and b are 9×1 vectors of constants and ι is an 9×1 unit vector.

Null Hypothesis	χ^2	df	p-level
Equation A			
Is the price of risk equal across countries?			
H_0: $\delta_i = \delta$, $\forall i$	4.767	8	0.782
Are country-specific intercepts jointly equal to zero?			
H_0: $\alpha_i = 0$, $\forall i$	6.534	9	0.685
Is the price of country-specific risk equal to zero?			
H_0: $\gamma_i = 0$, $\forall i$	4.231	8	0.765
Equation B			
Are the information variables z_{t-1} orthogonal to the risk-adjusted excess returns?			
H_0: $\lambda_k = 0$, $\forall k$	22.125	5	0.001

second moments, as well as other standard GARCH processes, while matching closely the sample second moments of the data.

Table IV contains the results of tests that compare the benchmark CAPM to alternative specifications. First, we estimate the partial market segmentation model specified in equation (10). In this model the risk premium for each country depends not only on the covariance with the world portfolio, but also on two country-specific factors: a constant and the conditional variance. The robust Wald tests support the conditional CAPM. Only one of the constant factors is individually significant at the 10 percent level and the p-value for the joint hypothesis that all the α_is are zero is equal to 0.69. None of the prices of country risk (γ_i) is individually significant either, and the p-value for the joint hypothesis that country risk is not priced in any market is equal to 0.77. We also consider an alternative form of market segmentation in which the price of covariance risk δ differs across countries. The null hypothesis of a common price cannot be rejected at any statistically reasonable level.

Second, we test the hypothesis that some of the variation in the excess returns can be explained by a number of information variables, even after accounting for the variation in the market risk premium of each asset. Formally, we test the hypothesis that the λ coefficients in equation (11) are jointly equal to zero. The robust Wald test statistic reported in the table shows that this hypothesis is strongly rejected by the data.

In summary, a number of diagnostic tests support the pricing restrictions of the traditional CAPM when applied to international financial markets. Expected returns are positively related to market risk and the price of covariance risk is equal across markets. In addition, neither idiosyncratic risk nor a country-specific constant factor add any explanatory power to the model. Nevertheless, two main issues emerge from the results discussed so far. First, although the estimates reveal a positive relation between excess returns and covariance risk, the link is statistically weak. Second, the model explains only a small amount of variation in the excess return series and some of the variation is still predictable after accounting for market risk. These results may be due, at least in part, to the auxiliary assumption on the constancy of the price of risk. If excess returns are considerably more variable than their conditional covariance with the market, a model with a constant price of risk may not have enough power to fully explain the dynamics of the risk premia. We analyze this issue in the next subsection by allowing δ to vary through time.

B. The Conditional CAPM with Time-Varying Price of Risk

Since the conditional CAPM is only a partial equilibrium model, the theory does not help identify the state variables that affect the price of market risk. Inevitably, any parameterization of the dynamics of δ_{t-1} can be criticized for being ad hoc. For example, many studies assume that the price of risk is linearly related to a given set of instruments, because this is an easy specification to estimate and test.[18] One limitation of this approach is that, in the absence of additional restrictions, a linear price of market risk can become negative, which is inconsistent with the theory. Merton (1980) points out that the nonnegativity restriction should be incorporated into the specification of the model to avoid biases in the estimation of the market premium. For this

[18] For example, Ferson (1989) and Ferson, Foerster, and Keim (1993) use the linearity assumption to test latent variable models.

reason, we assume that the price of market risk can be approximated by an exponential function of the instruments[19]

$$\delta_{t-1} = \exp(\kappa' z_{t-1}).$$

The instrument set z_{t-1} is described in the data section. To accommodate the January effect, we assume that the price of risk in the month of January is equal to $\delta_{t-1} = \kappa_J + \exp(\kappa' z_{t-1})$.

Panel A in Table V contains QML estimates of the parameters for the benchmark model with time-varying δ. The average price of market risk is equal to 3.96 and given a Newey-West consistent standard error of 0.38, is highly significant. The robust Wald test statistic for the hypothesis that the price of risk is constant (Panel B in the table) is equal to 22.59 and, therefore, the null hypothesis is strongly rejected at any conventional level. Individual test statistics reveal that the dynamics of δ are mostly driven by the excess dividend-yield, the change in the short-term interest rate and, in part, by the January dummy variable.[20]

These findings are very interesting, especially when compared to the earlier discussion of the conditional CAPM with constant price of risk. Obviously, the lack of statistical significance in the relation between expected returns and conditional market risk documented in Table III is due to the insufficient variation in the conditional covariances relative to the excess returns. The improved performance of the model is confirmed by the diagnostic statistics reported in Panel C. For example, the average value of the pseudo-R^2s increases from 0.16 percent in the case of constant price of risk to 2.52 percent when δ is time-varying. Most of the indices of kurtosis and the Bera-Jarque statistics are also improved relative to the model with constant δ. The standardized residuals, however, are still fat-tailed.

Figure 1 contains a plot of the estimated price of covariance risk.[21] Part of the volatility in the point estimates is inevitably due to estimation error. Since we believe that the most interesting component of our estimates is the trend in the series, we also include a plot of the price in which the high frequency components are filtered-out using the Hodrick and Prescott (HP) methodology (see Hodrick and Prescott (1996)). The filtered price reaches its highest values in the Seventies, becomes much lower during the Eighties and finally increases again in the early Nineties. The average value of δ_{t-1} between January 1980 and December 1994 is equal to 3.31 and significantly lower than its average value of 4.94 in the first part of the sample.[22]

[19] Bekaert and Harvey (1995) and De Santis and Gérard (1997) use a similar parameterization.

[20] We also investigate an alternative specification of the model in which the instruments include an aggregate measure of currency risk computed as the difference between the dollar and local currency returns on the world index. The coefficient of this variable is not statistically significant and none of the results are affected.

[21] To preserve readability, the plot does not include the January component of δ_t.

[22] Our choice to split the sample in 1980 is not completely arbitrary. In 1980 Japan, the country with the largest financial markets after the United States, enacted the Foreign Exchange and Foreign Control Law which lifted many restriction to international investments. Since then many other countries have liberalized their financial markets.

Another feature of the estimated price of risk is the presence of a pronounced January effect. The average estimated premia, excluding the January component, vary from a minimum of 0.45 percent per month (Germany) to a maximum of 0.72 percent (U.K.). They increase by an average of 1.47 percent in the month of January. This indicates that approximately 25 percent of the annual premium is earned during the first month of the year. Although our approach treats the January effect as a seasonal in the price of risk, rather than as a factor which affects risk premia linearly, our results confirm the findings of Gultekin and Gultekin (1983) and extend them to the Eighties and the first half of the Nineties.

Panel A in Table V shows that the structure of the conditional second moments is essentially unchanged when the price of market risk is allowed to vary. Panel C also includes some diagnostic tests on the squared standardized residuals to evaluate the specification of the conditional second moments. As before, the Ljung-Box statistics reveal that no significant autocorrelation is left in the series obtained from the GARCH (1, 1) parameterization and the Engle-Ng LM tests for signed size-bias suggest that, at least in our sample, there is no evidence of asymmetry in the response of conditional second moments to past innovations. Finally, we compute an extension of the LM test proposed by Engle and Ng (1993) to determine whether the instruments in z_{t-1} are orthogonal to the conditional variances.[23] In seven of nine cases we fail to reject the hypothesis that the instruments cannot predict conditional volatilities. The only exceptions are Canada and the world portfolio. Although this result opens the possibility for further studies on the specification of the GARCH process, we do not believe that the evidence is strong enough, in our context, to require a specification that would be much harder to estimate.

Having established a number of satisfactory features of the conditional CAPM with time-varying price of risk, we proceed to evaluate the two alternative specifications discussed at the beginning of this section. The results for these tests are contained in Table VI. First, we consider the model with partial market segmentation. In light of the previous testing results it is not surprising that neither the constants, nor the country specific volatilities are priced. In fact, the relaxation of the constancy hypothesis on δ_{t-1} should have no bearing on these results.

The second alternative specification is potentially more interesting. Since the model with time-varying price of risk explains a larger fraction of the variation in excess returns, this may eliminate the instruments' ability to predict excess returns, after accounting for market risk. The robust Wald test, however, shows that this is not the case and the null hypothesis that all the λs in equation (11) are equal to zero is strongly rejected again.

Given the persistence of this result, we believe it is worth exploring an additional hypothesis. The predictability may be driven by the nonnegativity restriction that we impose on the market risk premium. For example, Bou-

[23] For each country index, we run a regression of the squared standardized residuals on the variables included in z_{t-1}. Under the null hypothesis that the instruments have no predictive power, the regression R^2 multiplied by the number of observations is distributed as a chi-square variable with five degrees of freedom.

Table V

Quasi-Maximum Likelihood (QML) Estimates of the Conditional International CAPM with Time-Varying Price of Risk

Estimates are based on monthly dollar-denominated returns from January 1970 through December 1994. Data for the country equity indices and the world portfolio are from MSCI (Morgan Stanley Capital International). The risk-free rate is the return on the U.S. T-Bill with maturity closest to 30 days from Center for Research in Security Prices (CRSP). All returns are in percent per month. Each mean equation relates the country index excess return R^e_{it} to its world covariance risk $\text{cov}(R_{it}; R_{mt}|\Im_{t-1})$. The price of risk is function of a set of instruments, z_{t-1}, included in the investors' information set. The instruments include a constant, a January dummy (JAN), the world index dividend yield in excess of the return on the 30-day U.S. T-Bill (XDPR), the change in the U.S. term premium (ΔUSTP), the U.S. default premium (USDP) and the change in the 30-day U.S. T-Bill return (ΔUSRF).

$$R^e_{it} = \delta\text{cov}_{t-1}(R_{it};\, R_{mt}|\Im_{t-1}) + \epsilon_{it}$$

where $\delta_{t-1} = \exp(\kappa' z_{t-1})$ denotes the price of world covariance risk and $\epsilon_t|\Im_{t-1} \sim N(0, H_t)$. The conditional covariance matrix H_t is parameterized as follows:

$$H_t = H_0 * (\iota\iota' - aa' - bb') + aa' * \epsilon_{t-1}\epsilon'_{t-1} + bb' * H_{t-1}$$

where $*$ denotes the Hadamard matrix product, a and b are 9×1 vectors of constants and ι is an 9×1 unit vector. QML standard errors are reported in parentheses.

RMSE denotes Root Mean Squared Error. Pseudo R^2s are computed for each index as the ratio between the sum of squared fitted values of the risk premia and the sum of squared excess returns. Since the model restricts the price of risk to be constant across indices, individual pseudo R^2s are not guaranteed to be positive. B–J is the Bera–Jarque test statistic for normality. Q is the p-value of Ljung–Box test statistic of order 12 for standardized residuals (z) and standardized residuals squared (z^2). EN–S− and EN–S+ are Engle–Ng Lagrange multiplier tests for positive and negative size-bias in standardized residuals squared. EN-z denotes the Engle-Ng Lagrange multiplier test of the predictability of standardized residuals squared with the z variables.

Panel A: Parameter Estimates

Price of Covariance Risk

κ_0 (const)	κ_1 (JAN)	κ_2 (XDPR)	κ_3 (ΔUSTP)	κ_4 (USDP)	κ_5 (ΔUSRF)
−2.942	0.089	6.783	−0.652	0.414	−13.683
(2.570)	(0.054)	(2.483)	(2.263)	(2.288)	(5.517)

GARCH Process

	Canada	Japan	France	Germany	Italy	Switzerland	United Kingdom	United States	World
a_i	0.171	0.189	0.251	0.293	0.163	0.169	0.244	0.229	0.206
	(0.038)	(0.025)	(0.066)	(0.060)	(0.065)	(0.072)	(0.040)	(0.031)	(0.030)
b_i	0.961	0.966	0.868	0.671	0.888	0.921	0.950	0.960	0.965
	(0.018)	(0.009)	(0.044)	(0.438)	(0.070)	(0.033)	(0.018)	(0.012)	(0.010)

Table V—*Continued*

Panel B: Specification Test

Null Hypothesis	χ^2	df	p-level
Is the price of market risk constant? H$_0$: $\kappa_k = 0, \forall k > 1$	22.587	5	0.000

Panel C: Summary Statistics and Diagnostics for the Residuals

	Canada	Japan	France	Germany	Italy	Switzerland	United Kingdom	United States	World
Avg. exc. return	0.33%	1.00	0.64	0.57	0.23	0.63	0.72	0.38	0.45
Premia: exc. Jan.	0.51%	0.61	0.58	0.45	0.46	0.52	0.72	0.51	0.57
January	1.38%	1.66	1.59	1.27	1.26	1.41	1.80	1.33	1.53
Avg. pred. error	−0.29%	0.25	−0.08	0.01	−0.33	−0.00	−0.15	−0.23	−0.25
RMSE	5.40	6.66	6.93	6.06	7.77	5.56	7.32	4.33	4.10
Pseudo R^2 [a]	3.80%	1.06	1.03	0.01	−0.31	1.44	3.80	6.27	5.55
Kurtosis[a]	1.58**	0.58*	1.21**	0.60*	0.95**	1.03**	3.91**	1.83**	1.55**
B–J	35.91**	6.86*	17.20**	4.55	15.45**	12.43**	195.62**	45.64**	37.71**
$Q_{12}(z)$	0.52	0.22	0.50	0.10	0.02	0.93	0.73	0.68	0.28
$Q_{12}(z^2)$	0.67	0.60	0.52	0.25	0.76	0.38	0.78	0.71	0.89
EN–S−	−0.21	−0.95	−1.42	−0.72	−0.39	−1.26	0.38	−0.36	−0.02
EN–S+	1.41	0.37	−0.96	1.36	1.93	0.01	−0.67	−0.39	−1.07
EN–z	18.67**	2.11	5.83	4.08	6.27	10.40	8.01	7.01	16.34**
Likelihood function:					−7499.94				

[a] Equal to zero for the normal distribution.

* and ** denote statistical significance at the 5 percent and 1 percent levels, respectively.

doukh, Richardson, and Smith (1993) find that, for the U.S. market, negative risk premia are associated with periods of high expected inflation and downward sloping term structures. It is plausible that international financial markets display similar features, which obviously cannot be captured by the CAPM alone. To test this hypothesis we estimate the model in equation (11) assuming, this time, that the price of market risk is linear in the instruments, then we repeat the predictability test on the risk-adjusted returns. Figure 2 contains a plot of the estimated price of market risk with and without the nonnegativity restriction imposed. Although the two series have a correlation of 0.65, the difference is quite dramatic in some periods. Specifically, between the end of the Seventies and the early Eighties the unrestricted price of market risk is consistently negative. During those years, interest rates and inflation were unusually high and the slope of the yield curve was often negative. This shows that the findings of Boudoukh, Richardson, and Smith (1993) hold also in an international setting. Interestingly, when we allow risk premia to become negative, the joint test that all the λs in equation (11) are equal to zero has a p-value of 0.68, which implies that the predictability documented earlier is indeed driven by the inability of the CAPM to accommodate negative expected returns.

To summarize, we find that expected excess returns are positively related to their conditional covariance with a world-wide portfolio whereas country specific risk is not priced. Explicitly modeling the time-variation of the price of

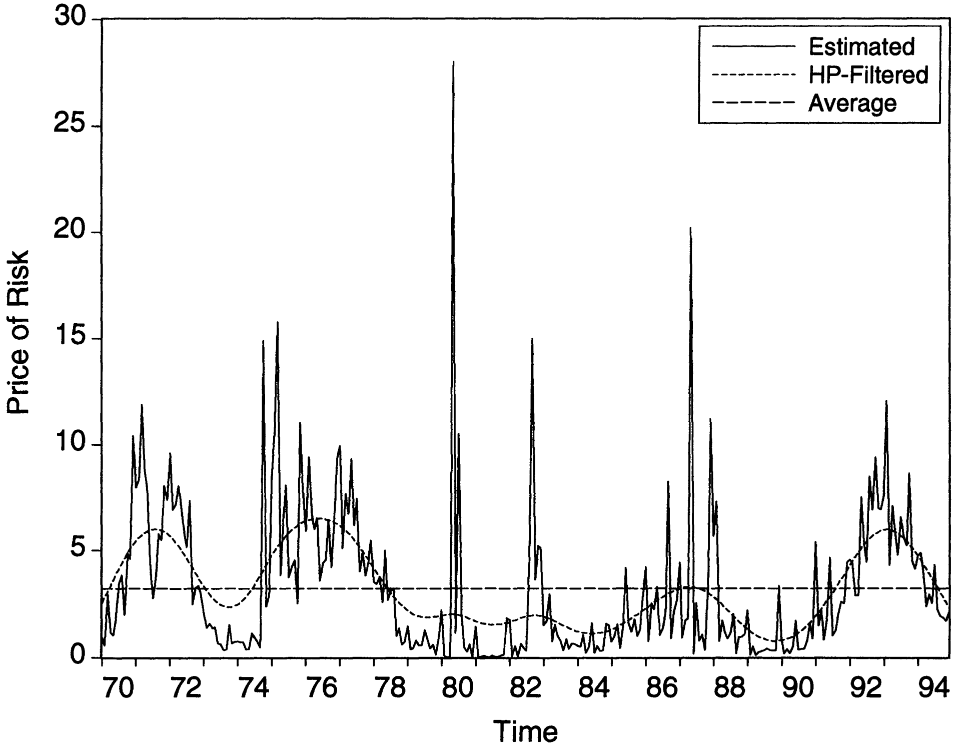

Figure 1. The price of world covariance risk. The figure contains plots of the estimated price of covariance risk, the corresponding HP-filtered series and a horizontal line that corresponds to the average price of risk. The estimates are obtained from the conditional version of the international CAPM. The plotted series do not include the January component.

market risk improves the ability to explain the dynamics of the risk premia. However expected excess returns remain predictable after accounting for world market risk. This is the most concerning result for advocates of the traditional CAPM as a model of international asset pricing: a one-factor model which imposes a nonnegativity constraint on the market risk premium cannot fully explain the dynamics of international expected returns.[24] If the premium for market risk has to be positive, an alternative model is needed to accommodate negative equity premia for some periods of time. For example, the recent findings of Dumas and Solnik (1995) and De Santis and Gérard (1997) indicate that currency risk is likely to be an important omitted factor for international assets. Or it is possible that national equity indices contain characteristics which are not accounted for by market risk.[25] Alternatively, a nonlinear pricing model like the one suggested by Bansal, Hsieh, and

[24] To verify the robustness of our results, we replicate all our tests using local currency returns in excess of the local risk free rate. This model assumes that investors use a unitary hedge ratio to offset their exposure to exchange rate risk (see, e.g., Chan, Karolyi, and Stulz (1992)). The evidence is essentially unaffected. All the results are available upon request.

Table VI

Specification Tests of the Conditional International CAPM with Time-Varying Price of Risk

This table reports robust Wald test statistics for two alternative specifications of the conditional international capital asset pricing model (CAPM). The first model (equation A) includes country specific risk and country specific intercepts in addition to market risk. The second model (equation B) includes asset specific intercepts and the instruments in z_{t-1} as independent conditioning variables, in addition to market risk. The instruments include a constant, a January dummy (JAN), the world index dividend yield in excess of the return on the 30-day U.S. T-Bill (XDPR), the change in the U.S. term premium (ΔUSTP), the U.S. default premium (USDP) and the change in the 30-day U.S. T-Bill return (ΔUSRF). Estimation is based on monthly dollar-denominated returns from January 1970 through December 1994. Data for the country equity indices and the world portfolio are from MSCI (Morgan Stanley Capital International). The risk-free rate is the return on the U.S. T-Bill with maturity closest to 30 days from the Center for Research in Security Prices (CRSP).

Eq. A $$R^e_{it} = \alpha_i + \delta\ \mathrm{cov}(R_{it};\ R_{mt}|\Im_{t-1}) + \gamma_i \mathrm{var}(R_{it}|\Im_{t-1}) + \epsilon_{it}$$

Eq. B $$R^e_{it} = \alpha_i + \lambda' z_{t-1} + \delta_{t-1}\mathrm{cov}(R_{it};\ R_{mt}|\Im_{t-1}) + \epsilon_{it}$$

where $\delta_{t-1} = \exp(\kappa' z_{t-1})$ denotes the price of world covariance risk and $\epsilon_t|\Im_{t-1} \sim N(0, H_t)$. The conditional covariance matrix H_t is parameterized as follows

$$H_t = H_0 * (\iota\iota' - aa' - bb') + aa' * \epsilon_{t-1}\epsilon'_{t-1} + bb' * H_{t-1}$$

where $*$ denotes the Hadamard matrix product, a and b are 9×1 vectors of constants and ι is an 9×1 unit vector.

Null Hypothesis	χ^2	df	p-level
Equation A			
Are country specific intercepts jointly equal to zero?			
H_0: $\alpha_i = 0, \quad \forall i$	7.566	9	0.578
Is the price of country specific risk equal to zero?			
H_0: $\gamma_i = 0, \quad \forall i$	4.182	8	0.840
Equation B			
Are the information variables z_{t-1} orthogonal to the risk-adjusted excess returns?			
H_0: $\lambda_k = 0, \quad \forall k$	25.096	5	0.000

Viswanathan (1993) may be needed to explain the cross-section of returns that we analyze.

IV. Should U.S. Investors Diversify Internationally?

The evidence discussed so far suggests that world risk is priced in international financial markets and that a number of restrictions of the traditional CAPM hold when the model is applied to an international setting. However,

[25] The possibility that the dynamics of expected returns can be explained by asset characteristics rather than risk has recently been suggested by Daniel and Titman (1997).

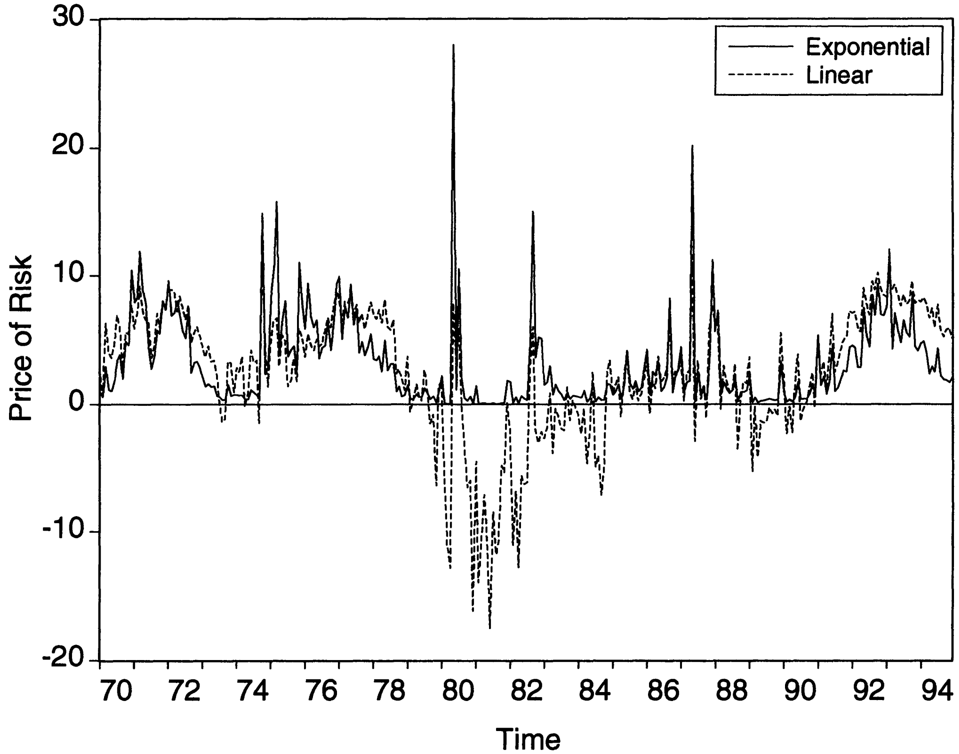

Figure 2. Restricted versus unrestricted price of world covariance risk. The figure contains plots of the estimated price of covariance risk with and without the nonnegativity restriction. The restricted value is obtained by assuming that the price is an exponential function of the instruments whereas the unrestricted value is obtained assuming linearity. The plotted series do not include the January component.

our results also indicate that a model with multiple sources of risk would probably perform better than a single factor model. This implies that the standard approach to portfolio choice proposed by Markowitz (1952) may not be completely adequate when evaluating the benefits of international diversification. Nevertheless, since it is common practice to use this framework also in international finance, in this section we discuss the implications of our findings for mean-variance analysis.[26]

International diversification is often regarded as the best way to improve portfolio performance. In fact, as long as financial markets are affected by country specific factors, correlations between asset returns from different countries are likely to be lower than correlations within the same country. Based on this reasoning, researchers and investment advisors alike have strongly promoted the idea of global diversification. Recently, however, a more

[26] Implicitly we make the assumption that all the markets in our study were financially integrated throughout the sample. Bekaert and Harvey (1995) analyze whether the degree of financial integration varies through time.

cautious attitude has been fostered by two arguments. First, international financial markets have become increasingly more integrated, due to the lifting of many restrictions to international investment during the Eighties. If the legal barriers were binding, cross-country correlations may have increased in recent years, thus reducing the benefits of diversification.[27] Second, recent studies suggest that global diversification may not be as good a safety net against bear markets in the United States as most people think. For example, Odier and Solnik (1993) and Lin, Engle, and Ito (1994) suggest that bear markets are contagious at the international level; when the U.S. market tumbles, cross-country correlations increase. In this sense, the benefits of holding an international portfolio may be reduced even further when investors need them the most.

To investigate those two issues, we first construct a measure of the expected gains from international diversification, then focus on its trend and cycle components. In particular, since the process of liberalization has taken place over several years, its effects on asset prices are better identified with the trend in the data. On the other hand, the behavior of international returns during U.S. bear and bull markets is mostly a short-term phenomenon and, therefore, should be reflected by the high frequency component in the data.

In practice, although correlations are the typical measure that analysts refer to, the model that we estimate suggests that the expected benefits of international diversification are driven by the interaction between conditional second moments and the time-varying price of market risk. Consider a U.S. investor who wants to evaluate the potential benefits of adding foreign assets to a portfolio which is fully invested in domestic stocks. The model predicts that the expected excess return on the U.S. portfolio should be equal to $\delta_{t-1}\text{cov}(R_{\text{US}t}, R_{mt}|\Im_{t-1})$. Obviously if the investor is not diversified, his risk exposure will be equal to the conditional volatility of the U.S. market. For the same level of volatility, the world portfolio of stocks, combined with the risk-free asset, has an expected excess return equal to $\delta_{t-1}\text{var}(R_{\text{US}t}|\Im_{t-1})$. Therefore, the expected gains of international diversification, for a level of risk equal to the volatility of the U.S. market, can be measured as

$$\text{E}(R_{dt} - R_{\text{US}t}|\Im_{t-1}) = \delta_{t-1}[\text{var}(R_{\text{US}t}|\Im_{t-1}) - \text{cov}(R_{\text{US}t}, R_{mt}|\Im_{t-1})] \qquad (12)$$

where R_{dt} is the return on an internationally diversified portfolio which includes R_m and R_f and has the same volatility as the U.S. portfolio. The term in brackets in equation (12) can be interpreted as a measure of the (time-varying) nonsystematic risk of the U.S. market, for which investors are not compensated. We discuss the benefits of international diversification by considering the dynamics of both the conditional second moments and the price of world risk.

[27] This argument is often found in the business press. However, an economic model that clearly predicts this result is not yet available. For example, in personal conversations with us, Fischer Black argued that the opposite effect could be expected.

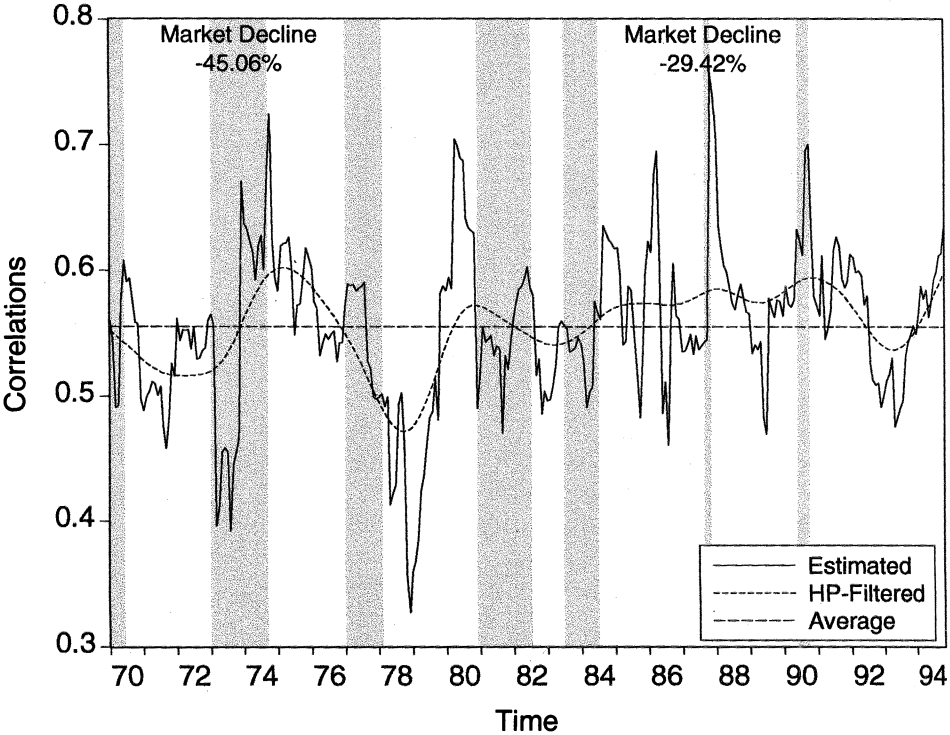

Figure 3. Portfolio correlation with the U.S. market. The figure contains plots of the correlation between the returns on the U.S. index and those of an equally-weighted portfolio which includes the remaining countries. The HP-filtered series and the average value of the correlation are also included. The shaded regions indicate periods of decline in the U.S. market, as identified by Forbes magazine. The two most severe declines in the sample (1973:01–1974:09 and 1987:09–1987:11) are explicitly labeled.

A. The Effects of Liberalization

First, consider the effect of market liberalization on the correlation of returns. Figure 3 contains plots of the time-varying correlations between the U.S. market and an equally weighted portfolio of the foreign assets. The estimated series reveals some interesting regularities. The twenty lowest correlations, with values ranging from 0.33 to 0.46, are all concentrated in the first part of the sample, prior to 1980. On the other hand, sixteen of the twenty largest correlations, with values between 0.68 and 0.76, correspond to the last fourteen years in the data.[28] These results suggest that the tendency of international markets to react simultaneously to innovations has increased with liberalization.

However, concluding that the reduction of legal barriers to international investments has significantly increased correlations would be premature. In

[28] All four remaining large correlations are associated with the bear market of 1973–1974.

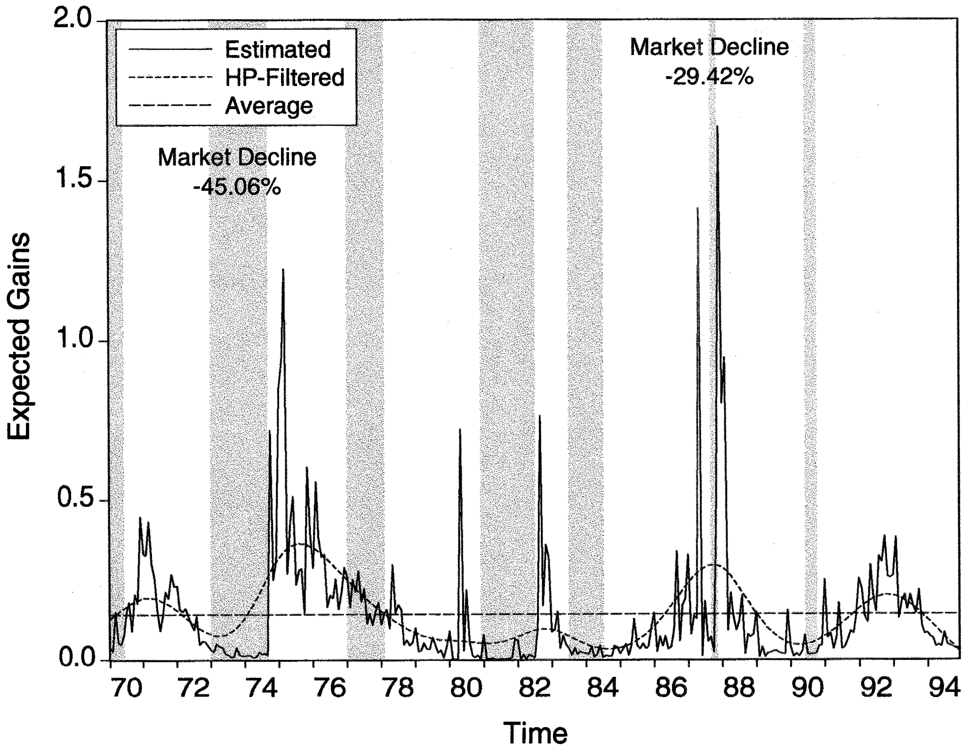

Figure 4. Expected gains from international diversification. The expected gains are measured as the increase in expected returns at the level of the U.S. market conditional volatility. The figure contains plots of the estimated gains, the HP-filtered series, and the average expected gains over the whole sample. The plotted series do not include the January component.

fact, the filtered series in the plot reveals that even recent values are well below the correlations that one would expect for U.S. portfolios. In this sense, the changes in correlations that we document can only be responsible for small changes in the expected gains from diversification.

Figure 4 contains plots of the expected gains from diversification, measured according to equation (12). Here, we focus on the low frequency component of the series to assess whether and how the expected benefits of diversification have been affected by market liberalization. Over the entire sample, the average premium for diversification is equal to 2.11 percent on an annual basis. The Newey-West consistent standard error for this estimate is equal to 0.284, which implies that the average expected gains are statistically significant. Our estimates indicate that the market risk premium for the U.S. index, including the January component, averages 7.45 percent per year. Therefore, for the same level of risk, the premium for an internationally diversified portfolio would be 28 percent higher. Interestingly, prior to 1980, the average gain is equal to 2.51 percent, whereas it is only equal to 1.84 percent for the last fourteen years. However the difference between the two means is not statistically significant. The plot of the filtered series reveals that for most of the Eighties the expected gains were lower than the average for the entire sample. In the early Nineties, however, the data again show relatively large expected benefits. Overall, although liberalization may have reduced the potential gains of diversifying internationally, the trend is rather weak, and the benefits remain economically significant.

B. U.S. Bear Markets and Portfolio Diversification

Next, we discuss whether bear markets are contagious across countries and how this may affect portfolio decisions. In Figures 3 and 4, we identify seven bear markets and seven bull markets for the United States, as defined by Forbes magazine.[29] The different periods are identified using the S&P 500 index as a benchmark. Some of the evidence seems to validate the claim that bear markets are contagious. In particular, we find that three of the largest correlations are associated with the most severe market declines in our sample. For example, between January 1973 and September 1974 the U.S. market dropped 45.06 percent. The estimated correlation at the end of that period is equal to 0.72, which is significantly larger than the sample average of 0.56. Similarly, after the 29.42 percent decline between September and November 1987, the correlation increases to 0.76. Finally, the model predicts another significant increase in correlation at the end of the market decline in the last part of 1990. In general, however, the plot shows that increases in international correlations are only obvious for large declines in the U.S. market and are often not very persistent.

When poor performance is spread across markets, investors have only limited opportunities to shield their portfolios. For example, our data show that the average benefits from international diversification are equal to only 0.08 percent per month during the seven bear markets in the sample, as opposed to 0.21 percent in the remaining periods. However, this result is not sufficient to discard the theory of international portfolio diversification. First, as long as diversification generates positive gains on average and investors are not able to predict bear and bull markets, it is still optimal to hold internationally diversified portfolios. Second, although the financial press often focuses on correlations, our measure in equation (12) suggests that a more appropriate assessment of the time-varying gains from diversification should be based on the dynamics of both nonsystematic risk and the price of market risk. For this reason, in Figure 5 we plot the difference between the conditional volatility of the U.S. market and the conditional covariance between the U.S. market and the World portfolio. It is interesting how, at the end of the 1974 and 1987 bear markets, the estimated difference increases rather dramatically. This implies that, after severe market declines, country specific volatility in the U.S. in-

[29] See recent issues of Forbes on mutual fund performance. The seven bear markets cover the following periods: January 1, 1970 to June 30, 1970; January 1, 1973 to September 30, 1974; January 1, 1977 to February 28, 1978; December 1, 1980 to July 31, 1982; July 1, 1983 to July 31, 1984; September 1, 1987 to November 30, 1987; June 1, 1990 to October 31, 1990.

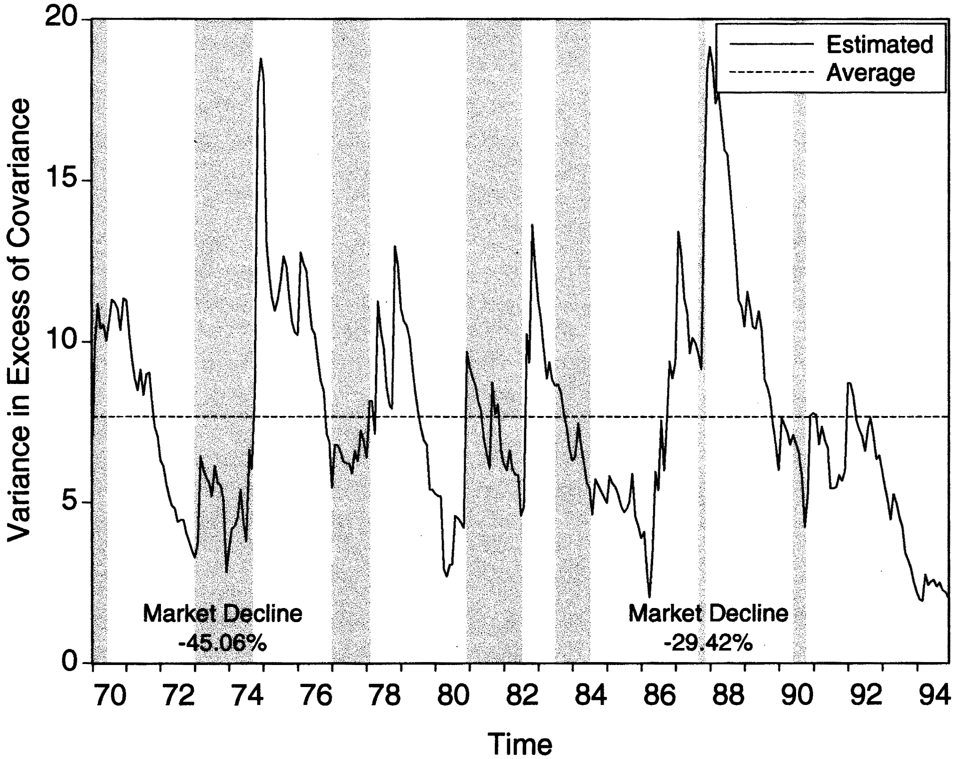

Figure 5. Variance risk in excess of world covariance risk for the U.S. stock market. The plot describes the dynamics of the estimated conditional volatility for the U.S. market, in excess of the conditional covariance between the U.S. market and the World portfolio. The shaded regions indicate periods of decline in the U.S. market, as identified by Forbes magazine. The two most severe declines in the sample (1973:01–1974:09 and 1987:09–1987:11) are explicitly labeled.

creases more than systematic risk. When combined with the estimated price of world risk plotted in Figure 1, this leads to the pattern in expected gains documented in Figure 4. Although the gains are often very low during declines of the U.S. market, they increase considerably when the market starts to recover. Given their inability to predict turns in market conditions, investors are better off holding internationally diversified portfolios even during periods of poor market performance.

V. Conclusions

In this article, we test the pricing restrictions of the conditional CAPM simultaneously for the world's eight largest equity markets. Since our approach fully specifies the dynamics of the conditional moments, we can also investigate how, from the perspective of a U.S. investor, the expected gains from international diversification respond to changing market conditions. In particular, we examine whether the expected benefits have been affected by the increasing integration of international financial markets and whether an internationally diversified portfolio provides a good hedge against large declines in the U.S. market.

Methodologically, we extend the multivariate GARCH process of Ding and Engle (1994) to accommodate GARCH-in-Mean effects and propose an alternative to the test of the CAPM of Bollerslev, Engle, and Wooldridge (1988). Our approach has several advantages. We can estimate the model and analyze international comovements for a large number of assets simultaneously without losing the flexibility of standard GARCH processes. The tests can be implemented on any subset of assets even when asset supplies are unobservable, as long as returns on a market-wide index are available. Finally, our specification can easily be extended to accommodate multiple risk factors.

The evidence supports most of the pricing restrictions of the conditional CAPM. Market risk, measured by the conditional covariance between the return on each asset and a world-index, is equally priced across countries. On the other hand, the price of country-specific risk, measured by the conditional variance of each country index, is not different from zero. Allowing for time-variation in the price of market risk significantly improves the performance of the model. However, when the price of market risk is restricted to be positive, as the theory implies, some of the variation in residual returns remains predictable. Interestingly, the predictability disappears when we relax the nonnegativity restriction. This result may be driven by the inability of the CAPM to accommodate negative risk premia during periods of high short-term interest rates. Even though the conditional version of the traditional CAPM provides useful information on the dynamics of market premia, a more adequate model of international asset pricing should probably include additional factors. Along these lines, recent work by Dumas and Solnik (1995) and De Santis and Gérard (1997) suggests that currency risk is priced in addition to market risk.

Our results have interesting implications for investors who want to diversify their portfolios internationally. We find that severe U.S. market declines are contagious at the international level, and often imply a significant reduction in the expected gains from international diversification. However, our estimates indicate that the ex ante gains for a U.S. investor average 2.11 percent a year and have not significantly declined over the last two and one-half decades. Hence, while holding an internationally diversified portfolio provides little protection against severe U.S. market declines, the long-term gains from international diversification remain economically attractive.

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