Predictable Stock Returns in the United States and Japan: A Study of Long-Term Capital Market Integration

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ABSTRACT

This paper uses the predictability of monthly excess returns on U.S. and Japanese equity portfolios over the U.S. Treasury bill rate to study the integration of long-term capital markets in these two countries. During the period 1971–1990 similar variables, including the dividend-price ratio and interest rate variables, help to forecast excess returns in each country. In addition, in the 1980's U.S. variables help to forecast excess Japanese stock returns. There is some evidence of common movement in expected excess returns across the two countries, which is suggestive of integration of long-term capital markets.

IF CAPITAL MARKETS ARE integrated, then financial assets traded in different markets, but with identical risk characteristics, will have identical expected returns. Alternatively, in segmented capital markets, barriers to arbitrage may allow assets traded in different markets to have different expected returns even when their risk characteristics are the same. This study explores the extent to which U.S. and Japanese stock markets can be described as integrated.

One obvious way to measure the extent of integration is to look for direct evidence of barriers to arbitrage across markets (legal restrictions on foreign share ownership, transactions taxes, and so forth), or for evidence that cross-border transactions in financial assets are limited in scale.[1] A problem with this straightforward approach is that legal barriers and taxes can often be circumvented, while a limited volume of cross-border trading might be sufficient to bring asset prices into line across markets.

Another approach is to test the hypothesis that assets with identical risk characteristics have the same expected returns in different markets, assum-

[1]For example, French and Poterba (1990) study the extent to which U.S. and Japanese investors make cross-border investments in common stocks.

ing that some mean-variance efficient benchmark portfolio is observable. If this assumption holds, then assets traded in integrated capital markets have expected returns that are determined by their observable betas with the benchmark return and by the observable mean benchmark return. Most commonly, these moments are assumed to be constant through time, but recent work has started to allow for moments that change with certain conditioning variables.[2]

A troublesome aspect of this research is the need to specify an observable benchmark portfolio a priori. The hypothesis of integration may be rejected merely because one has specified an inappropriate benchmark portfolio. In this paper we try to avoid the assumption that a benchmark return is observable. Without an observable benchmark, it is harder to measure assets' risk characteristics and harder to test the hypothesis of integration. However we can still make some progress if we are willing to use extra assumptions about the unobservable benchmark return. In particular, if assets have constant betas with the benchmark, but the conditional mean benchmark return is time-varying, then the returns on assets traded in integrated markets have a single-latent-variable representation. Expected returns on all such assets vary through time in a perfectly correlated fashion, because they are all being driven by the changing price of a single unobserved source of risk. In this paper we test a single-latent-variable model for U.S. and Japanese stock returns.

Our work is subject to some of the same difficulties as the observable-benchmark approach.[3] First, we may falsely reject the hypothesis of integrated capital markets if capital markets are in fact integrated but our assumptions about the unobservable benchmark fail to hold. For example, if Japanese and U.S. firms are exposed to different sources of risk, and if the prices of these risks move independently, then expected excess returns will move independently even if prices are set in a single world capital market. Second, there may be some alternatives against which the single-latent-variable test has no power. For example, national stock markets could be segmented but subject to common shocks that move expected returns in similar ways.

Nevertheless we believe that a finding of common movement is suggestive of integration. Common movement in expected returns implies that some force is affecting the equilibrium return in the U.S. and Japanese stock markets in the same way. We are agnostic about what this force might be. The possibilities include changes in volatility or some broader measure of "business cycle risk" (Fama and French (1989)), changes in the risk aversion of a representative agent as aggregate wealth rises and falls (Marcus (1989)), and exogenous shifts in the demand for stock of "noise traders" that must be

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[2]Cho, Eun, and Senbet (1986), Gultekin, Gultekin, and Penati (1989), Jorion and Schwartz (1986), and Stehle (1977) work with static models. Wheatley (1988) uses a dynamic asset pricing model but works with unconditional moments. Harvey (1991) uses a conditional approach, allowing moments to change through time.

[3]Wheatley (1989) emphasizes the problems with latent-variable modeling.

accommodated by utility-maximizing traders (Campbell and Kyle (1988), Shiller (1984)). But if market-clearing takes place in the U.S. and Japanese stock markets independently, then equilibrium returns would move together only by coincidence.[4]

Our work also has value as simple data description. To the extent that we find similar variables forecasting stock returns in the U.S. and Japan, this reinforces the large literature on predictable components of stock returns in the U.S. market.[5] Our single-latent-variable model generates estimates of the component of expected excess returns that is common to the two countries, and this is of some interest whether or not the model adequately describes all variation in expected excess returns.

The organization of our paper is as follows. In section I we describe the asset pricing framework that motivates our empirical work. In section II we describe our data set. In section III we present preliminary regressions that document the existence of predictable excess stock returns. In section IV we try to use the results from section III to characterize the extent to which U.S. and Japanese stock markets are integrated. We briefly discuss an observable-benchmark model using a world stock index as the benchmark. Then we estimate a single-latent-variable model that restricts expected excess stock returns in the U.S. and Japan to move together. Section V concludes.

I. The Asset Pricing Framework

The most general asset pricing model we consider is a K factor model of the following form:

$$\tilde{r}_{i,t+1} = E_t[\tilde{r}_{i,t+1}] + \sum_{k=1}^{K} \beta_{ik} \tilde{f}_{k,t+1} + \tilde{\varepsilon}_{i,t+1}. \tag{1}$$

[4]This argument is analogous to that of Feldstein and Horioka (1980). They argued that if international capital markets were perfectly integrated, then there would be no reason to expect savings and investment in a particular country to be correlated with one another. Evidence that these variables are correlated is suggestive that international capital markets are imperfectly integrated. Similarly, we argue that if international capital markets were entirely segmented, then there would be no reason to expect equilibrium returns in different countries to be correlated with one another.

[5]See for example Campbell (1987, 1990), Campbell and Shiller (1988), Fama and French (1988, 1989), Fama and Schwert (1977), and Keim and Stambaugh (1986). These papers find that excess U.S. stock returns are forecast by, among other variables, the dividend-price ratio on stock, the level of interest rates, the long-term yield spread, and the month of the year (the so-called "January effect"). There is a smaller recent literature on forecasting Japanese stock returns; Sentana and Wadhwani (1989) study the Japanese market in detail, while Bekaert and Hodrick (1990), Cumby (1990), Cutler, Poterba, and Summers (1990), Harvey (1990), and Solnik (1990) forecast Japanese stock returns as part of a multicountry study of stock price behavior. Gultekin and Gultekin (1983), Jaffe and Westerfield (1985), and Kato and Schallheim (1985) study the January effect in Japan. Gultekin (1983) and Solnik (1983) report international evidence on inflation (measured directly or using short-term interest rates) in relation to stock returns.

Here $\tilde{r}_{i,t+1}$ is the excess return on asset i held from time t to time $t+1$, the difference between the random real return on asset i and the risk-free real rate of interest. The excess return on asset i equals the expected excess return, plus the sum of K factor realizations $\tilde{f}_{k,t+1}$ times their betas or factor loadings β_{ik}, plus an idiosyncratic error $\tilde{\varepsilon}_{i,t+1}$. The asset pricing model is dynamic in the sense that the expected excess return can vary through time, but static in that the beta coefficients are assumed to be constant through time.

The expected excess return is restricted by the model as follows:

$$E_t[\tilde{r}_{i,t+1}] = \sum_{k=1}^{K} \beta_{ik}\lambda_{kt}, \tag{2}$$

where λ_{kt} is the "market price of risk" for the kth factor at time t. This type of restriction can be generated by any of a number of intertemporal asset pricing models.

Now suppose that the information set at time t consists of a vector of N forecasting variables X_{nt}, $n = 1, \ldots, N$ (where X_{1t} is a constant), and that conditional expectations are linear in these variables. Then the kth risk price can be written

$$\lambda_{kt} = \sum_{n=1}^{N} \theta_{kn}X_{nt}, \tag{3}$$

and equation (2) becomes

$$E_t[\tilde{r}_{i,t+1}] = \sum_{k=1}^{K} \beta_{ik} \sum_{n=1}^{N} \theta_{kn}X_{nt} = \sum_{n=1}^{N} \alpha_{in}X_{nt}. \tag{4}$$

Equation (4) says that the IN coefficients α_{in} obtained by regressing I excess returns on N forecasting variables can be written in terms of IK beta coefficients and KN coefficients which define market prices of risk.

There are two main ways in which this system can be used in empirical work. Either one can assume that certain factors are observable; or one can assume that factors are unobservable, but the number of factors is small relative to the number of assets and forecasting variables.

A. Observable Factors

Suppose that we observe a portfolio whose return has a beta of one on the first factor, and zero on the other factors. Suppose further that the return on this portfolio has zero idiosyncratic risk. Call the return on this portfolio $\tilde{r}_{1,t+1}$. Then we have

$$\tilde{r}_{i,t+1} = \beta_{i1}\tilde{r}_{1,t+1} + \sum_{k=2}^{K} \beta_{ik} \sum_{n=1}^{N} \theta_{kn}X_{nt} + \sum_{k=2}^{K} \beta_{ik}\tilde{f}_{k,t+1} + \tilde{\varepsilon}_{i,t+1}$$

$$= \beta_{i1}\tilde{r}_{1,t+1} + \sum_{n=1}^{N} \alpha_{in}^{*}X_{nt} + \tilde{u}_{i,t+1}. \tag{5}$$

In a regression of excess return i on excess return 1 and the information variables X_{nt}, the inclusion of excess return 1 "soaks up" the time variation in the risk price for factor 1. The coefficients on X_{nt}, α_{in}^*, now reflect only the time variation in the risk prices for factors 2 through K. If these risk prices are zero, then all coefficients α_{in}^* will be zero; if these risk prices are constant, then the intercept α_{i1}^* will be nonzero, but the other coefficients α_{in}^* for $n = 2, \ldots, N$ will be zero.

This approach can be applied to the international context as follows. Suppose we think that the Japanese stock market obeys a multi-factor model, where the first factor is an international factor and the other factors are domestic Japanese factors. Suppose that the international factor is well proxied by another stock market return, say the return on a world stock index. Then we can regress the Japanese market return on the world index return and a set of forecasting variables. The variance of $\sum \alpha_{in}^* X_{nt}$, relative to the variance of $\sum \alpha_{in} X_{nt}$ (the fitted value when the Japanese market is regressed only on X_{nt}), is a measure of the variation in risk prices of domestic factors relative to the variation in the risk prices of all factors. In the extreme case where only the international factor is priced, the coefficients α_{in}^* will all be zero. (This is the model discussed in the introduction to the paper, in which the international factor is an observable benchmark portfolio.) In the case where only the risk price for the international factor varies through time, the coefficients α_{in}^* will be zero apart from the intercept.

B. Unobservable Factors

One objection to the above procedure is that it assumes that a particular world stock index is an adequate proxy for the international factor in the asset pricing model. This may not be appropriate. An alternative approach is to assume that there is a single priced international factor which is unobservable, and no priced domestic factors in either the U.S. or Japan. If we work with two stock returns, one from each country, and N forecasting variables, then equation (4) imposes that $\alpha_{in} = \beta_i \theta_n$, where the k subscript has been dropped since there is only one factor. The underlying parameters β_i and θ_n are only identified up to a normalization; if we normalize $\beta_1 = 1$, the restricted system can be written as

$$
\begin{bmatrix} \tilde{r}_{1,\,t+1} \\ \tilde{r}_{2,\,t+1} \end{bmatrix} = \begin{bmatrix} \theta_1 & \theta_2 & \cdots & \theta_N \\ \beta_2\theta_1 & \beta_2\theta_2 & \cdots & \beta_2\theta_N \end{bmatrix} \begin{bmatrix} X_{1t} \\ \cdot \\ \cdot \\ X_{Nt} \end{bmatrix} + \begin{bmatrix} v_{1,\,t+1} \\ v_{2,\,t+1} \end{bmatrix} \tag{6}
$$

The first row of the coefficient matrix in (6) identifies the θ_n coefficients, the first column identifies the coefficient β_2, and the remaining $N - 1$ coefficients are restricted. These restrictions enforce a perfect correlation between the expected excess return in the U.S. market, and the expected excess return in the Japanese market. The restricted specification is sometimes

called a single-latent-variable model.[6] It can be estimated and tested using Hansen's (1982) Generalized Method of Moments, which allows for conditional heteroskedasticity in the variance-covariance matrix of returns.

The model (6) can be generalized to allow for unobserved domestic factors whose risk prices are constant or depend only on a subset of the X variables (say the first L variables, X_{nt} for $n = 1, \ldots, L$). When such factors are present, the restrictions in (6) apply only to the coefficients on the remaining X variables (X_{nt} for $n = L + 1, \ldots, N$). Unfortunately, we cannot allow for arbitrary domestic factors because the model then becomes unidentified.

Even if the overidentifying restrictions of equation (6) are rejected, the estimated coefficients may still be of interest. The fitted values from (6) are the best possible forecasts of stock returns in the two countries subject to the restriction that the forecasts be perfectly correlated with one another; thus they can be interpreted as estimates of a common component in expected stock returns. Below we will compare these estimates with unrestricted regression forecasts of stock returns in the two countries.

C. Omitted Information Variables and Other Problems

In our empirical work we use forecasting variables X_{nt} which are known to the market at time t. Generally, we do not wish to assume that we have included all the relevant variables. Fortunately, the methods described above are robust to omitted information. By taking conditional expectations of equations (5) and (6), it is straightforward to show that the various restrictions hold in the same form when a subset of the relevant information is used. Thus if the coefficients α_{in}^* in equation (5) are zero for the true information vector used by the market, they will also be zero if a subset of this vector is included in (5). Similarly, if the market's forecasts of excess returns in the two countries are perfectly correlated, then forecasts using a subset of the market's information must also be perfectly correlated.

The single-latent-variable approach does depend critically on the maintained assumption that assets have constant betas with the unobserved benchmark portfolio. If this assumption is false, then the single-latent-variable model will fail to describe the data even if U.S. and Japanese equity markets are integrated. Unfortunately, it is hard to generalize the approach to deal with violations of this assumption. Structural change in the θ coefficients of equation (6) can be handled by estimating a system with fixed β coefficients and randomly or deterministically changing θ coefficients. (Below we estimate a system of this type with a single change in the θ coefficients in the middle of the sample.) Structural change in the β coefficients is harder to deal with because the β's are identified only by the normalization that $\beta_1 = 1$. This normalization will not be appropriate if all assets' β coefficients are changing through time. Thus the results reported in

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[6]For more details on this model, see Hansen and Hodrick (1983), Gibbons and Ferson (1985), and Campbell (1987).

this paper must be interpreted conditionally on the maintained assumption that assets have fixed betas on the unobserved benchmark portfolio.

II. Data and Sample Period

The comparative approach of this paper requires that the data be comparable across the two countries to the greatest extent possible. The last month for which we are able to obtain complete data in both countries is March 1990.

A. Data Sources

For the U.S., we use standard publicly available data. Stock prices and dividends are taken from the Center for Research on Security Prices (CRSP) monthly stock tape. We study a value-weighted index of New York Stock Exchange stocks, and also a set of equally weighted portfolios, organized by firm size.[7] We use a 1-month Treasury bill yield as our short-term interest rate, and a long-term (approximately 20-year) government bond yield to compute the long-short yield spread. These series are from Ibbotson Associates (1990).

For Japan, the most commonly used and readily available stock price indexes are the Nikkei 225 and the Tokyo Stock Exchange Price Index (TOPIX). These indexes, however, are not comparable with the CRSP value-weighted New York Stock Exchange index. The Nikkei index is a price-weighted index of only 225 stocks out of more than 1500 stocks listed currently on the Tokyo Stock Exchange, representing about 50% of total capitalization. The TOPIX is a value-weighted index constructed from all the stocks traded on the first section of the Tokyo Stock Exchange with 97% of the total (first and second section) capitalization, but neither TOPIX nor Nikkei properly account for dividend payments.

We therefore constructed a value-weighted index, as well as a set of equally weighted size portfolios, from data on individual stock returns including and excluding dividends.[8] The universe of stocks is the Tokyo Stock Exchange, first and second sections; foreign firms listed on the TSE are excluded from the sample.[9] Our database is an extension of the one presented in detail in Hamao (1988, 1991) and Hamao and Ibbotson (1990), and it starts in January

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[7]The size portfolios are rebalanced monthly according to capitalization at the end of the previous month.

[8]Our Japanese individual stock returns data are compiled from the raw data on prices kept by Daiwa Securities and adjusted for dividend payments, stock splits, etc. This database is comparable to the CRSP files. Like the TOPIX, we use market value weights in our index. These weights differ from the ideal weights when there is cross-holding of firms' equity by other firms, which is an important phenomenon in Japan (McDonald (1989), French and Poterba (1991)). However we obtain similar results for equally weighted size portfolios of stocks, so it does not appear that our results are sensitive to the details of the weighting scheme.

[9]Our U.S. sample does include a few Japanese firms in the form of American Depositary Receipts, but overall there is minimal cross-listing.

1970. Since we need one year's lag in order to construct a 1-year moving average dividend-price ratio, our sample period starts in January 1971.

Japanese bond markets did not develop until the 1970's, and data are therefore not available before 1970. There is no equivalent of Treasury bills in Japan; thus the short-term interest rate used here is a combined series of the call money rate (1971:1–1977:11) and the Gensaki rate (1977:12–1990:3). The Gensaki rate, an interest rate applied to bond repurchase agreements, is less subject to regulation than the call money rate, but it became available only after 1977. The call money rate is the "unconditional" rate, which is applied to transactions maturing in less than one month, and we use a Gensaki rate with one month maturity. For the long-term Japanese government yield, we use a value-weighted index of yields on bonds with 9 to 10 years to maturity.[10]

We also use one piece of data from outside the national financial markets of the U.S. and Japan. This is the monthly return on the Morgan Stanley Capital International (MSCI) World Index, a market-value-weighted index covering just under 1500 companies listed on the stock exchanges of 20 countries. Together, these companies account for about 60% of the total market capitalization of the countries included in the index. At the end of September 1990 the U.S. market had a weight of 35%, the Japanese market had a weight of 30%, and the European stock markets had a combined weight of 28% in the index.[11] The MSCI world index is measured in dollar terms, inclusive of dividends.

Finally, we note that in forming excess return series, we measure both U.S. and Japanese stock returns in dollars, relative to the U.S. Treasury bill rate. In earlier versions of this paper, we measured the Japanese stock return in yen, relative to the Japanese short-term interest rate.[12] Excess yen returns on Japanese stocks are slightly more predictable than excess dollar returns on Japanese stocks, but the difference is small and does not affect the qualitative results of the paper. We also measure all returns in continuously compounded (log) form. This is common practice in empirical work on asset pricing, and it has the advantage that it enables us to use excess returns without measuring a dollar or yen price deflator. However it may introduce some approximation error in that asset pricing models generally apply to simple rather than log returns.[13]

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[10]This is the longest consistently available maturity. The yield index normally includes a benchmark issue, which has a large capitalization and therefore a high weight in the index.

[11]Earlier in our sample period the U.S. had a higher weight and Japan had a lower weight. The prevalence of cross-holding in the Japanese market tends to give it too high a weight in international value-weighted indices.

[12]The excess Japanese stock return, measured in yen, approximates the return that U.S. investors will receive if they finance their investment by borrowing yen. The approximation becomes accurate in the limit of continuous time if the yen return on Japanese stocks is uncorrelated with the dollar price of the yen (Stulz (1981)).

[13]As Bekaert and Hodrick (1990) point out, there can be systematic differences in expected log returns on different assets even if investors are risk-neutral; these result from Jensen's Inequality. Thus it is strictly speaking incorrect to refer to expected excess log returns as "risk premia."

B. Sample Period

Limitations on the availability of Japanese data, discussed above, confine us to the sample period 1971:1–1990:3. Within this period, financial markets in both countries have undergone some institutional changes. The system of financial regulation in the U.S. has changed gradually through the period we study, but Japanese capital markets have experienced a more radical deregulation.[14] Before 1970, there was virtually no free short-term interest rate. Although the Gensaki market grew substantially in the 1970's, it was not until 1978 that the authorities completely lifted restrictions in the short-term market. After the first issue of government bonds in 1966, financial institutions, which were the major bondholders, were not allowed to sell government bonds in a secondary market until 1977.

More recently a major deregulation occurred with the revision of the Foreign Exchange Law in December 1980. The old Foreign Exchange Law prohibited all transactions with foreign countries in principle, whereas the new law removed controls over many types of capital flow. For example, it is now possible for a foreigner to invest in up to 10% of the equity of a Japanese company without the permission of the Ministry of Finance.

Japanese deregulation took another step forward in May 1984 with the "Yen-Dollar Agreement". At this time interest rates were further deregulated, limitations on exchanging foreign currency into yen were abolished, yen-denominated foreign loans were deregulated, foreign brokers were allowed to obtain membership of the Tokyo Stock Exchange, and the Euroyen bond and loan markets were enlarged.[15] In 1987, bond markets were further liberalized as it became possible to short bonds in Japan for the first time.

This history, and the steady development of the secondary bond market in Japan, suggest that we ought to divide our sample period to see whether deregulation and financial innovation have had noticeable effects on stock market behavior. We choose to divide the whole period 1971:1–1990:3 (231 observations) into two subsamples, 1971:1–1980:12 (120 observations) and 1981:1–1990:3 (111 observations). One could argue for break points later in the sample, notably in 1984 and 1987, but the one we use has the advantage that it is close to a mid-sample split.[16]

III. Forecasting Excess Stock Returns in the United States and Japan

Table I reports basic statistics that summarize the behavior of some of the most important variables we study. For each variable we report the mean,

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[14]See Pigott (1983), Japanese Ministry of Finance (1987), and Suzuki (1987) for a description of Japanese financial deregulation.

[15]Frankel (1984) gives further details.

[16]We have also checked that our results are not sensitive to the inclusion of data from the fixed exchange rate period by dropping the observations before 1974.

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Table I

Summary Statistics for U.S. and Japanese Data

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The sample periods for this table are 1971:1–1990:3, 1971:1–1980:12, and 1981:1–1990:3, with 231, 120, and 111 observations, respectively. Units are percentage points at an annualized rate. S.D. is the standard deviation and ρ_1 is the first autocorrelation of the series. Excess stock returns are measured in dollars, relative to the U.S. one-month Treasury bill rate. The dividend-price ratio is the average dividend over the previous year, divided by the current stock price. The short-term interest rate is a one-month Treasury bill rate for the U.S., and a one-month Gensaki rate for Japan. The relative short rate is the difference between the current short rate and a one-year backward moving average. The long-short yield spread is the difference between a long-term bond yield (maturity approximately 20 years for the U.S., 9–10 years for Japan) and the short-term interest rate.

Sample Period	U.S. Mean	U.S. S.D.	U.S. ρ_1	Japan Mean	Japan S.D.	Japan ρ_1	U.S./Japan Correlation
			Excess Value-Weighted Stock Return				
71–90	3.66	56.7	0.051	12.5	69.3	0.213	0.274
71–80	2.11	56.8	−0.002	13.3	63.5	0.228	0.279
81–90	5.33	56.8	0.098	11.7	75.4	0.201	0.272
			Value-Weighted Dividend-Price Ratio				
71–90	4.17	0.812	0.999	1.69	0.855	0.990	−0.138
71–80	4.05	0.902	0.999	2.29	0.661	0.991	−0.461
81–90	4.30	0.683	0.999	1.04	0.484	0.998	0.802
			Short-Term Interest Rate				
71–90	7.39	2.75	0.992	6.57	2.39	0.997	0.207
71–80	6.58	2.53	0.973	7.47	2.80	0.995	0.284
81–90	8.28	2.70	0.989	5.60	1.27	0.990	0.677
			Relative Short-Term Interest Rate				
71–90	0.048	1.63	0.747	−0.079	1.49	0.936	0.093
71–80	0.260	1.55	0.674	0.118	1.89	0.945	0.233
81–90	−0.181	1.68	0.792	−0.291	0.834	0.915	−0.259
			Long-Short Spread				
71–90	1.82	1.69	0.918	0.501	1.84	0.960	0.196
71–80	1.38	1.58	0.884	0.142	2.42	0.967	0.169
81–90	2.29	1.69	0.941	0.890	0.667	0.943	0.193

standard deviation, and first autocorrelation of the U.S. and Japanese series, and the correlation between the U.S. and Japanese series, over the full sample and both subsamples.

At the top of the table we give statistics for the excess dollar returns on the U.S. and Japanese value-weighted indexes over the U.S. Treasury bill rate. Monthly returns are measured in percentage points at an annual rate. Japanese stocks have a higher mean return than U.S. stocks in both the 1970's and the 1980's, but also a higher standard deviation. In addition the

value-weighted Japanese stock index has a surprisingly high first-order autocorrelation coefficient of just over 0.2; this is stable across the two decades in the sample.[17] The correlation between U.S. and Japanese stock returns is also very stable at about 0.3.

Next we look at the behavior of dividend-price ratios on the two stock indexes (where the dividend is the average over the previous year, and the price is the current price). Dividend-price ratios have been found to predict excess returns in the U.S. (Campbell and Shiller (1988), Fama and French (1988)), and they will be important explanatory variables in our regression analysis. We find that the Japanese dividend-price ratio has a lower mean than the U.S. dividend-price ratio (in fact, it has been lower than the U.S. in every month since the mid-1970's). The Japanese dividend-price ratio is lower in the second half of our sample, reflecting the sustained rise in Japanese stock prices during the 1980's.[18] The U.S. and Japanese series are both extremely persistent, with first-order autocorrelations very close to one. They are negatively correlated in the 1970's, but highly positively correlated in the 1980's as both countries' dividend-price ratios drifted downward.

We repeat the exercise for the U.S. Treasury bill rate and the Japanese short rate, again measured at an annual rate. Short-term nominal interest rates have been found to forecast excess stock returns in U.S. data (Fama and Schwert (1977), Campbell (1987)). U.S. interest rates tend to rise slightly over the full sample period, while Japanese rates fall; however the medium-run movements of the two interest rates are positively correlated. For this reason the rates have higher correlations over the subsamples than over the whole sample period.

We also report summary statistics for the "relative short rate," defined as the difference between the current short-term interest rate and a 1-year backward moving average.[19] This variable is used to forecast stock returns in Campbell (1990) and Hodrick (1990). It removes the low-frequency variation from the interest rate series, and accordingly has a lower first-order autocorrelation coefficient than the raw interest rate. In the 1970's, the relative short rate is more variable in Japan and is positively correlated across the two countries, but in the 1980's this pattern reverses. The relative short rate becomes more variable in the U.S. and negatively correlated across the two countries.

Finally, we report summary statistics for the long-short yield spread. This variable also has been used to predict excess U.S. stock returns (Fama and

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[17]This high autocorrelation is due partly to the fact that we are measuring Japanese excess returns in dollar terms. Even in yen terms, however, the autocorrelation is between 0.10 and 0.15.

[18]Our sample period includes the first three months of 1990, during which the Japanese market reversed its rise. The cumulative excess log return on Japanese stocks in this period was about − 35%. For a detailed analysis of Japanese dividend yields and the level of the Japanese market, see French and Poterba (1991).

[19]This is equivalent to a triangularly weighted moving average of changes in short rates, so it is stationary if the short rate is stationary in first differences.

French (1989)). The U.S. and Japanese yield spreads are weakly positively correlated, with a higher mean in the U.S.

A. *Forecasting Excess Stock Returns with Own-Country Variables*

In Table II we regress excess returns in the U.S. and Japan on a variety of forecasting variables. U.S. results appear on the left-hand side of the table, and Japanese results on the right-hand side. For each country we use forecasting variables specific to that country. We report coefficients, with heteroskedasticity-consistent standard errors in parentheses, for the whole sample and each subsample.[20] For each regression, we also report the adjusted R^2 statistic, the joint significance of the coefficients (excluding a constant term and the January dummy), and the significance level for a test of stability of the coefficients across subsamples.

Panel A of Table II reports a simple forecasting equation, the "basic specification," that has been used for U.S. stock returns over a longer sample period by Campbell (1990) and Hodrick (1990). Three variables are included: a January dummy, the dividend-price ratio, and the relative short rate. In U.S. data the sign pattern of the variables is the same in the full sample and both subsamples. The January dummy has a positive sign, the dividend-price ratio also has a positive sign, while the relative short rate has a negative sign.[21] The dividend-price ratio and relative short rate are jointly significant at the 0.2% level over the full sample, and at the 0.1% level in the 1970's. The 1980's provide little evidence about the forecastability of returns; one cannot reject the hypothesis that the coefficients are stable from the 1970's to the 1980's, but one also cannot reject the hypothesis that the coefficients are zero in the 1980's. In the Japanese data the pattern is much the same; the point estimate of the dividend-price ratio coefficient actually switches sign from the 1970's to the 1980's, but again it is very imprecisely estimated in the 1980's, so that one cannot reject the hypothesis of coefficient stability across the two decades.

Panel B of Table II reports an "augmented specification," adding two other variables that are often thought to be relevant for forecasting returns: the lagged excess stock return and the long-short yield spread from the term structure of interest rates. As noted above there is some evidence of serial correlation in Japanese excess stock returns, and this improves the forecasting power of the model for Japanese returns. In the U.S. the augmented model does no better than the basic model in forecasting returns. Once again the 1980's add rather little to the evidence, since one cannot reject the null of coefficient stability or the null of zero coefficients in this part of the sample period.

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[20] The heteroskedasticity-consistent standard errors are generally quite similar to the ordinary standard errors in these regressions.

[21] Campbell (1990) and Hodrick (1990) show that this pattern also appears over the entire postwar period in the U.S.

Table II

Forecasting Excess Stock Returns with Own-Country Variables

The sample periods for this table are 1971:1–1990:3, 1971:1–1980:12, and 1981:1–1990:3, with 231, 120, and 111 observations, respectively. Dependent variables for all regressions are excess U.S. or Japanese stock returns, measured in dollars relative to the U.S. one-month treasury bill rate. The own-country dividend-price ratio is the previous year's average dividend divided by the current stock price. The relative short rate is the difference between the current short rate (one-month Treasury bill rate for U.S. and one month Gensaki rate for Japan) and a one year backward moving average. The own-country long-short yield spread is the difference between a U.S. long-term bond yield of 20 years maturity, or a 9–10 year maturity for Japan, and the respective short rate. Coefficients on the regressors are reported, with heteroskedasticity-consistent standard errors in parentheses. "Significance" is the joint significance of all the coefficients in the regression other than on the constant and January dummy. "Stability" is the rejection significance level for the hypothesis that all coefficients are constant across the two subsamples. A constant term is included in all regressions.

	U.S. Stock Returns			Japanese Stock Returns		
	71–90	71–80	81–90	71–90	71–80	81–90
Panel A: Basic Specification						
January	1.421	2.338	0.608	1.474	0.971	1.841
dummy	(1.296)	(1.722)	(1.839)	(1.090)	(0.818)	(1.980)
Dividend-	0.761	1.109	0.368	0.528	1.445	−0.488
price ratio	(0.403)	(0.459)	(0.823)	(0.454)	(0.658)	(1.306)
Relative	−0.645	−0.912	−0.491	−0.600	−0.519	−0.724
short rate	(0.200)	(0.267)	(0.262)	(0.232)	(0.237)	(0.803)
Adjusted R^2	0.067	0.116	0.013	0.025	0.066	−0.014
Significance	0.002	0.001	0.178	0.012	0.003	0.664
Stability	0.667			0.629		
Panel B: Augmented Specification						
January	1.422	2.500	0.667	1.189	0.673	1.110
dummy	(1.296)	(1.745)	(1.773)	(1.117)	(0.955)	(2.122)
Dividend-	0.846	1.193	0.647	0.643	1.877	0.231
price ratio	(0.385)	(0.452)	(0.855)	(0.460)	(0.667)	(1.299)
Relative	−0.431	−0.714	−0.256	−0.368	−0.122	−0.653
short rate	(0.309)	(0.396)	(0.409)	(0.268)	(0.276)	(0.772)
Lagged	0.021	−0.090	0.104	0.168	0.116	0.154
excess return	(0.083)	(0.107)	(0.108)	(0.066)	(0.080)	(0.105)
Long-short	0.269	0.374	0.259	0.183	0.412	−0.956
spread	(0.277)	(0.390)	(0.359)	(0.236)	(0.241)	(0.868)
Adjusted R^2	0.063	0.116	0.008	0.044	0.091	0.005
Significance	0.004	0.003	0.189	0.005	0.002	0.281
Stability	0.711			0.375		

In summary, Table II provides considerable evidence that U.S. and Japanese stock returns can be forecast using similar types of domestic variables. The major qualification to this statement is that the predictability of returns is tenuous in the 1980's, although this decade does not contradict the evidence from the 1970's.

B. *Forecasting Excess Stock Returns with Both Countries' Variables*

In Table III we push the investigation one stage further. We regress U.S. and Japanese excess returns on a common set of forecasting variables taken from both countries. This enables us to see whether foreign-country variables have any ability to predict excess returns when they are added to domestic variables. The basic set of forecasting variables in Panel A of Table III combines the two countries' basic forecasting variables from Panel A of Table II; it includes a January dummy, and U.S. and Japanese dividend-price ratios and relative short rates. The augmented set of forecasting variables in Panel

Table III
Forecasting Excess Stock Returns with Both Countries' Variables

The sample periods for this table are 1971:1–1990:3, 1971:1–1980:12, and 1981:1–1990:3, with 231, 120, and 111 observations, respectively. Dependent variables for all regressions are excess U.S. or Japanese stock returns, measured in dollars relative to the U.S. one-month treasury bill rate. The dividend-price ratio is the previous year's average dividend divided by the current stock price. The relative short rate is the difference between the current short rate (one-month Treasury bill rate for U.S. and one month Gensaki rate for Japan) and a one-year backward moving average. Panel B, in addition to one-month lags of U.S. and Japanese excess returns, includes the long-short yield spread as the difference between a U.S. long-term bond yield of 20 years maturity, or a 9–10 year maturity for Japan, and the respective short rate. Coefficients are reported, with heteroskedasticity-consistent standard errors in parentheses. "Significance (All)" is the joint significance of all the coefficients in the regression other than on the constant, and the January dummy. "Significance (U.S.)" and "Significance (Japan)" are the joint significance levels of both countries' variables. "Stability" is the rejection significance level for the hypothesis that all coefficients in the subsample are constant across the two subsamples. A constant term is included in all regressions.

	U.S. Stock Returns			Japanese Stock Returns		
	71–90	71–80	81–90	71–90	71–80	81–90
Panel A: Basic Specification						
January	1.452	2.281	0.602	1.321	1.050	1.388
dummy	(1.278)	(1.670)	(1.770)	(1.086)	(0.854)	(2.063)
U.S.	0.797	1.253	1.493	−1.176	−0.495	−5.030
dividend-price ratio	(0.427)	(0.574)	(1.300)	(0.477)	(0.671)	(1.207)
U.S.	−0.622	−0.792	−0.515	−0.373	−0.310	−0.364
relative short rate	(0.197)	(0.279)	(0.302)	(0.217)	(0.339)	(0.299)
Japanese	−0.052	0.090	−2.003	0.305	0.860	5.355
dividend-price ratio	(0.382)	(0.751)	(1.557)	(0.465)	(0.844)	(1.808)
Japanese	−0.297	−0.405	−0.023	−0.523	−0.474	−0.073
relative short rate	(0.209)	(0.231)	(0.576)	(0.237)	(0.250)	(0.799)
Adjusted R^2	0.067	0.126	0.010	0.053	0.061	0.076
Significance (All)	0.010	0.004	0.213	0.000	0.007	0.002
Significance (U.S.)	0.003	0.001	0.213	0.012	0.510	0.000
Significance (Japan)	0.360	0.171	0.265	0.065	0.105	0.007
Stability	0.256			0.047		

Table III-*Continued*

	U.S. Stock Returns			Japanese Stock Returns		
	71–90	71–80	81–90	71–90	71–80	81–90
Panel B: Augmented Specification						
January	1.432	2.337	0.013	1.077	0.278	1.089
dummy	(1.294)	(1.688)	(1.877)	(1.140)	(1.180)	(2.106)
U.S.	0.861	1.290	2.830	−0.766	0.378	−3.531
dividend-price ratio	(0.441)	(0.578)	(1.810)	(0.526)	(0.725)	(1.314)
U.S.	−0.506	−1.440	−0.096	0.076	0.875	−0.005
relative short rate	(0.337)	(0.604)	(0.364)	(0.403)	(0.607)	(0.515)
U.S.	0.011	−0.168	0.182	0.241	0.306	0.194
lagged return	(0.082)	(0.101)	(0.112)	(0.088)	(0.118)	(0.127)
U.S.	0.153	−0.542	0.692	0.372	1.377	0.473
long-short spread	(0.300)	(0.608)	(0.370)	(0.389)	(0.616)	(0.607)
Japanese	0.060	0.015	−2.699	0.667	2.641	4.228
dividend-price ratio	(0.384)	(0.833)	(1.733)	(0.533)	(0.939)	(1.615)
Japanese	−0.203	−0.342	−0.929	−0.212	0.523	−0.630
relative short rate	(0.249)	(0.364)	(0.705)	(0.276)	(0.361)	(0.902)
Japanese	0.007	0.079	−0.119	0.074	0.028	0.019
lagged return	(0.051)	(0.071)	(0.056)	(0.074)	(0.083)	(0.109)
Japanese	0.068	0.350	−2.161	0.171	0.440	−0.999
long-short spread	(0.238)	(0.230)	(1.298)	(0.218)	(0.216)	(1.007)
Adjusted R^2	0.053	0.139	0.053	0.088	0.142	0.068
Significance (All)	0.052	0.007	0.075	0.000	0.000	0.008
Significance (U.S.)	0.007	0.001	0.110	0.015	0.065	0.005
Significance (Japan)	0.819	0.173	0.047	0.219	0.026	0.048
Stability	0.027			0.046		

B of Table III similarly combines the two countries' augmented forecasting variables from Panel B of Table II; it includes a January dummy, U.S. and Japanese dividend-price ratios, relative short rates, lagged excess returns, and long-short yield spreads.

In Table III we find only weak evidence that Japanese variables help to forecast U.S. stock returns. The Japanese variables are jointly significant only in the augmented specification in the 1980's. Here the lagged Japanese excess return adds forecasting power so that the Japanese variables are jointly significant at the 4.7% level. The overall forecastability of U.S. excess returns is not much stronger in Table III than in Table II.

The addition of U.S. variables to the Japanese forecasting equation has a much more important effect, particularly in the 1980's subsample.[22] In Table II, we were unable to forecast Japanese excess returns in the 1980's; but in

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[22] This finding seems to be consistent with the results of Hamao, Masulis, and Ng (1990) for high-frequency data. They find that the Japanese stock market is more sensitive to foreign shocks than are the American or British stock markets.

Table III, the adjusted R^2 statistics for this decade rise from -0.01 to 0.08 when the U.S. variables are added to the basic specification, and from 0.01 to 0.07 when the U.S. variables are added to the augmented specification. The U.S. variables are jointly significant at the 0.5% level or better in both specifications. This improvement in 1980's forecasting power for Japan is accompanied by evidence of instability in the coefficients between the 1970's and the 1980's, as we can now reject the hypothesis of constant coefficients at the 5% level for both specifications.

A closer look at the pattern of coefficients in Table III reveals that many of the forecasting variables have parallel effects on the two countries' excess stock returns. The January dummy coefficients are positive for both countries and all sample periods, while the coefficients on U.S. and Japanese interest rates are negative for both countries and all sample periods. The dividend-price ratio effects are less consistent, however; the U.S. return is forecast by its own dividend yield with little contribution from the Japanese dividend yield, whereas the Japanese return seems to be forecast by the difference between the Japanese and U.S. dividend yields.[23] Overall, the fitted values from the Table III regressions have a positive correlation in the 1970's of about 0.4 in the basic specification and 0.2 in the augmented specification.[24] In the 1980's the correlation is zero or even negative, but one should not make too much of this since the overall forecastability of U.S. stock returns is quite weak in the 1980's.[25]

IV. Some Evidence on Capital Market Integration

We have found evidence that similar types of variables help to predict stock returns in the U.S. and Japan. The evidence is particularly strong in the 1970's, when stock returns in both countries are forecast by dividend yields (positively) and by the level of domestic short-term interest rates relative to their recent past (negatively). In the 1980's, there is little evidence for predictability of excess returns using own-country forecasting variables alone. But in this period there is an interesting cross-country effect: when U.S. variables are added to the forecasting equation, it becomes possible to predict Japanese excess returns with an adjusted R^2 of 7 or 8%. The next question we consider is whether these facts are consistent with any

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[23]Bekaert and Hodrick (1990) obtain a similar pattern of coefficients. They detrend the Japanese dividend yield in order to remove some of its low-frequency variation. The similarity of findings suggests that the downward drift in the Japanese dividend yield is not driving the results.

[24]The correlation of fitted values may be lower in the augmented specification merely because the addition of irrelevant forecasting variables has added noise to the forecasts.

[25]Bekaert and Hodrick (1990) report somewhat stronger results for forecasting U.S. stock returns in the 1980's using a system which also includes the differential between U.S. and Japanese interest rates (the forward premium). The pattern of their results is qualitatively similar to ours. Below we check the effect of adding the forward premium to our single-latent-variable model.

of the simple models of an integrated world capital market that we presented in section I.

A. An Observable Factor Model

In Table IV we estimate a regression in the form of equation (5). We add a world stock index excess return to the regressions of the U.S. and Japanese excess stock returns on forecasting variables. If the predictability of domestic returns is due merely to the changing risk price of an international factor, which is adequately proxied by the world index return, then the inclusion of the world index in the regression should destroy the significance of the forecasting variables.

Table IV
An Observable Factor Model for Excess Stock Returns

The sample periods for this table are 1971:1–1980:12 and 1981:1–1990:3, with 120 and 111 observations, respectively. Dependent variables for all regressions are excess U.S. or Japanese stock returns, measured in dollars relative to the U.S. one-month treasury bill rate. The dividend-price ratio is the previous year's average dividend divided by the current stock price. The relative short rate is the difference between the current short rate (one-month Treasury bill rate for U.S. and one month Gensaki rate for Japan) and a one-year backward moving average. Panel B in addition to one-month lags of U.S. and Japanese excess returns, includes the long-short yield spread as the difference between a U.S. long-term bond yield of 20 years maturity, or a 9–10 year maturity for Japan, and the respective short rate. The regressions also include a constant and the excess dollar return on the Morgan Stanley Capital International world index, the only coefficient reported, with heteroskedasticity-consistent standard errors in parentheses. "Significance (All)" is the joint significance of all the coefficients in the regression other than on the constant, the world index return and the January dummy. "Significance (U.S.)" and "Significance (Japan)" are the joint significance levels of both countries' variables without the constant, the world index return, and the January dummy.

	U.S. Stock Returns		Japanese Stock Returns	
	71–80	81–90	71–80	81–90
Panel A: Basic Specification				
World index return	1.036	0.886	0.747	1.046
	(0.052)	(0.083)	(0.125)	(0.138)
Adjusted R^2	0.815	0.661	0.342	0.590
Significance (All)	0.367	0.011	0.013	0.000
Significance (U.S.)	0.254	0.002	0.142	0.000
Significance (Japan)	0.820	0.024	0.449	0.000
Panel B: Augmented Specification				
World index return	1.035	0.869	0.740	1.066
	(0.053)	(0.077)	(0.111)	(0.126)
Adjusted R^2	0.822	0.657	0.416	0.583
Significance (All)	0.198	0.030	0.000	0.001
Significance (U.S.)	0.086	0.051	0.002	0.003
Significance (Japan)	0.553	0.034	0.033	0.003

In fact the addition of a world index generally has little effect on the other coefficients in the regression. The U.S. market has a beta of just over 1 in the 1970's, and a beta of just under 0.9 in the 1980's; this reflects the high but declining weight of the U.S. market itself in the world stock index. The Japanese market has a beta of about 0.75 in the 1970's, rising just above 1 in the 1980's. The other forecasting variables remain significant except for the U.S. regression in the 1970's, which is close to being a regression of the U.S. market on itself.[26]

B. An Unobservable Factor Model

We next ask whether predictable excess stock returns in the U.S. and Japan move together through time. As discussed in section I, if international capital markets are integrated and predictable excess returns are due to changes in the price of risk of a single world factor, then one would expect to find common movement in expected excess returns in the U.S. and Japan.

Common movement of fitted values can occur even when only own-country variables are significant for forecasting returns. If U.S. and Japanese forecasting variables are correlated, then own-country forecasts of excess returns can be highly correlated. This point is important for understanding the 1970's in our data. Table V shows that during the 1970's the forecasts of excess returns from Table III had correlations of 0.45 (basic specification, Panel A) and 0.23 (augmented specification, Panel B), even though we found very little evidence that foreign-country variables add to the forecasting power of own-country variables in this period. These correlations are somewhat increased by the presence of the January effect; if one looks at deseasonalized fitted values, the correlations fall slightly to 0.41 and 0.20.

Of course, it is essential to take into account the sampling error in the coefficients of Table III. In the 1980's, for example, the forecasts of U.S. excess returns are not statistically significant, so it is unlikely that their correlation with other forecasts can be estimated with any precision. In order to deal with sampling error properly, in Table V we estimate a single-latent-variable model of the form (6). This model imposes the testable restriction that expected excess stock returns are perfectly correlated across countries. We work with the raw data at the left of the table, and also with demeaned stock returns and with demeaned and deseasonalized returns (the residuals from a regression of returns on a constant and January dummy). This enables us to see whether any rejection of the latent-variable specification is due solely to the behavior of unconditional mean returns, or to the behavior of mean returns and January effects. The forecasting variables are the same

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[26] Given the instability of the Japanese regression coefficients, we present only subsample results. Full sample results are similar. We also ran regressions using the foreign-country excess return as the international factor. The Japanese return gets a coefficient of about 0.25 when it is added to the U.S. regression (this is the "beta" of the U.S. index on the Japanese index), while the U.S. return has a coefficient of 0.3 or 0.4 (the beta of the Japanese index on the U.S. index), but the other variables remain just as significant as they were before.

Table V

An Unobservable Factor Model for Excess Stock Returns

The sample periods for this table are 1971:1–1980:12 and 1981:1–1990:3, with 120 and 111 observations, respectively. The table reports the results of estimating a single-latent-variable model,

$$\begin{bmatrix} \tilde{r}_{1,t+1} \\ \tilde{r}_{2,t+1} \end{bmatrix} = \begin{bmatrix} \theta_1 & \theta_2 & \cdots & \theta_N \\ \beta_2\theta_1 & \beta_2\theta_2 & \cdots & \beta_2\,\theta_N \end{bmatrix} \begin{bmatrix} X_{1t} \\ \cdot \\ \cdot \\ X_{Nt} \end{bmatrix} + \begin{bmatrix} v_{1,t+1} \\ v_{2,t+1} \end{bmatrix}$$

where $r_{1,t+1}$ is the excess U.S. value-weighted stock index return and $r_{2,t+1}$ is the excess Japanese value-weighted stock index return, and both excess returns are measured in dollars relative to the U.S. one-month Treasury bill rate. The variables X_{it}, $i = 1, \ldots, N$ include in Panel A, the U.S. and Japanese dividend-price ratios (the previous year's average dividend divided by the current stock price), and the U.S. and Japanese relative short rates computed as the difference between the current short rate (one-month Treasury bill rate for U.S. and one month Gensaki rate for Japan) and a one-year backward moving average. Panel B includes in addition, one-month lags of the U.S. and Japanese excess returns, and U.S. and Japanese long-short yield spreads (the difference between a U.S. long-term bond yield of 20 years maturity, or a 9–10 year maturity for Japan, and the respective short rate). The system is estimated on raw data with a constant term and January dummy included, on demeaned data with a January dummy included, and on demeaned and deseasonalized data (the residuals from a first-stage regression of returns on a constant and January dummy). θ_i are free coefficients of the model not reported. "Japanese beta" is the estimated coefficient β_2 with an asymptotic standard error in parentheses. "Model restrictions" is the significance level for a test of the overidentifying restrictions of the single-latent-variable model. "R^2 ratio" is the ratio of the restricted model forecast variance to the unrestricted one obtained from regressing the excess stock return on the variables X_{it}. "Unrestricted correlation" is the correlation of the unrestricted regression forecasts of stock returns in the two countries.

	Raw Data		Demeaned		Demeaned and Deseasonalized	
	71–80	81–90	71–80	81–90	71–80	81–90
Panel A: Basic Specification						
Japanese beta	0.703	−9.511	0.822	−5.191	0.860	−4.944
	(0.233)	(17.716)	(0.291)	(5.812)	(0.311)	(5.465)
Model restrictions	0.015	0.264	0.055	0.305	0.028	0.211
R^2 ratio (U.S.)	1.016	0.041	0.711	0.125	0.730	0.147
R^2 ratio (Japan)	0.650	0.986	0.621	0.890	0.640	0.921
Unrestricted correlation	0.450	−0.309	0.450	−0.309	0.408	−0.384
Panel B: Augmented Specification						
Japanese beta	0.431	−1.038	0.473	−1.908	0.832	−1.548
	(0.163)	(0.631)	(0.189)	(1.127)	(0.263)	(0.903)
Model restrictions	0.002	0.041	0.015	0.183	0.015	0.127
R^2 ratio (U.S.)	1.070	0.424	0.852	0.276	0.726	0.358
R^2 ratio (Japan)	0.157	0.235	0.151	0.517	0.365	0.445
Unrestricted correlation	0.232	0.068	0.232	0.068	0.196	0.045

ones used in the basic and augmented specifications of Table III, Panels A and B. Given the evidence of coefficient instability, we estimate the system separately for the 1970's and the 1980's.

The first excess return in the system is the U.S. excess stock return; therefore we normalize the β for the U.S. to equal one. The free coefficients of the model are then the θ_n, $n = 1, \ldots, N$, and the β coefficient for the Japanese excess return. In Table V we report the Japanese β with an asymptotic standard error in parentheses. (To save space, the θ_n coefficients are not reported.)[27]

Table V shows that a single-latent-variable model for U.S. and Japanese excess returns can be rejected at the 0.2% to 5.5% level in the 1970's (depending on the specification). The β coefficient of the Japanese return on the unobserved common factor is estimated to be between 0.4 and 0.9. In the 1980's, the single-latent-variable specification is rejected at the 5% level when the augmented specification is used with the raw data, but otherwise is not rejected at even the 10% level. The Japanese β becomes large and negative, and very imprecisely estimated. The reason for these results is that U.S. excess returns are not reliably forecastable in the 1980's subsample, so the model is free to fit Japanese excess returns and its estimates of the θ and β coefficients become highly colinear.

Another way to evaluate the performance of the model with a single unobservable factor is to compare the variance of the restricted forecast with the variance of the unrestricted forecast from Table III. If the restricted variance is much smaller than the unrestricted variance, then the restrictions are causing a serious deterioration in forecast power.[28] In Table V we report the ratio of the two variances for the U.S. and Japanese markets. In the 1970's the single-latent-variable model fits 70% to 85% of the variance of the unrestricted forecast of U.S. returns, and 15% to 65% of the variance of the unrestricted forecast of Japanese returns. Even though the model is rejected statistically, the estimated common component of returns is clearly important. In the 1980's, the single-latent-variable model fits 10% to 35% of the variance of the unrestricted forecast of U.S. returns and 45% to 90% of the variance of the unrestricted forecast of Japanese returns. This reflects the fact that in this decade the unrestricted Japanese coefficients are statistically significant while the U.S. coefficients are not, so the latent-variable model fits the former at the expense of the latter.

A visual impression of these results is given in Figures 1 through 4. These figures plot the unrestricted and restricted fitted values, using solid lines and dashed lines, respectively, over the 1970's (Figures 1 and 2) and the 1980's

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[27] We use a two-stage version of Hansen's (1982) Generalized Method of Moments for estimation. We obtain starting values by regressing the first excess return in the system onto the forecasting variables to estimate the θ coefficients, and then regressing the second excess return onto the first to obtain an initial β estimate. Our qualitative results are robust to the choice of starting values and to the use of the U.S. or Japanese return as the first return in the system. Ferson and Foerster (1990) report Monte Carlo evidence that the two-stage method of moments estimator behaves well in finite samples when the number of assets in the system is small.

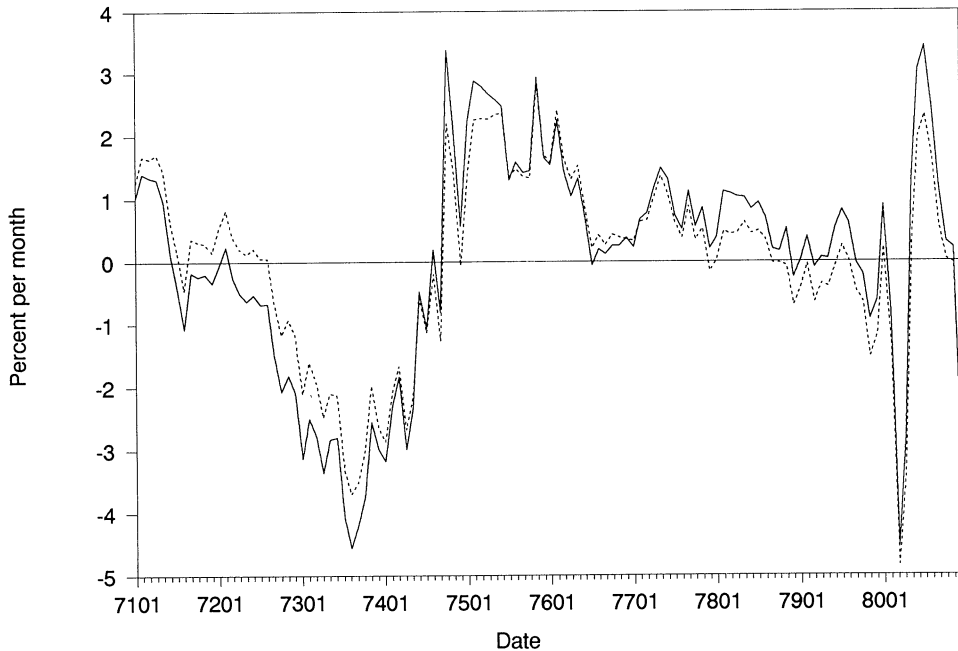[28] See Gibbons and Ferson (1985) for further discussion of this statistic.

Figure 1. Forecast U.S. Excess Stock Return, 1970–1980. The solid line is the unrestricted forecast from the basic specification estimated in Panel A of Table III. The dashed line is the restricted forecast from the single-latent-variable model estimated in Table V, again using the basic specification.

(Figures 3 and 4).[29] Figures 1 and 3 show the fitted values for the U.S. market, while Figures 2 and 4 show the fitted values for the Japanese market.

Figures 1 and 2 show an impressive degree of common movement of expected excess returns in the 1970's, despite the statistical rejection of the single-latent-variable model. In both countries the 1970's were characterized by large low-frequency swings in expected returns, with a decline from 1971 to 1974, a rise from 1974 to 1978, and a second decline from 1978 to 1980. In the 1980's the expected excess return in the U.S. is much less variable and there is no clear pattern of common movement, although a peak in the expected excess return occurred in early 1983 for each country.

C. How Robust Are the Results?

We have tried several alternative specifications in order to check the robustness of the results reported above. First, we have tried estimating the single-latent-variable model using the forward premium (the difference

[29]The fitted values are demeaned and deseasonalized, and the basic specification is used. The plots look similar when the January effect is included, and when the augmented specification is used (in the latter case the fitted values are less smooth, reflecting the higher-frequency movements of the forecasting variables included in the augmented specification).
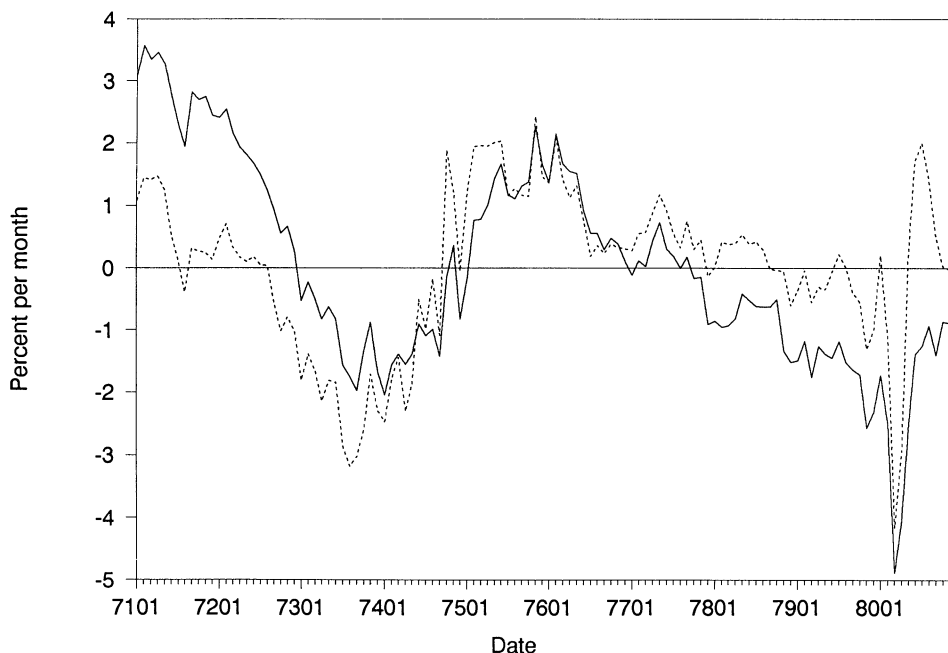
Figure 2. Forecast Japanese Excess Stock Return, 1970–1980. The solid line is the unrestricted forecast from the basic specification estimated in Panel A of Table III. The dashed line is the restricted forecast from the single-latent-variable model estimated in Table V, again using the basic specification.

between U.S. and Japanese short-term interest rates) as an additional forecasting variable. Bekaert and Hodrick (1990) find that this variable has forecasting power for U.S. stock returns in the 1980's. We also obtain this result, but find little forecasting power in the 1970's. Accordingly the inclusion of the forward premium has almost no effect on the single-latent-variable results in the 1970's; in the 1980's, the single-latent-variable model is rejected at about the 1% level when the forward premium is included.

Second, we have tried starting our sample period in 1974:1 in order to remove the period of fixed exchange rates from the sample. This has very little effect on the forecasting equations for U.S. stock returns, but it decreases the predictability of Japanese stock returns. Japanese returns are forecastable in 1974–1980 and 1974–1990 only when the augmented specification is used with both countries' forecasting variables. As one would expect from this, the single-latent-variable model is less strongly rejected when the sample period starts in 1974. It is not rejected at even the 10% level in 1974–1980, and is rejected at about the 5% level in 1974–1990.

Third, we have checked that our results are not sensitive to the use of a value-weighted stock index in each country. Single-latent-variable models applied to equally weighted portfolios of stocks in the first, third, and fifth
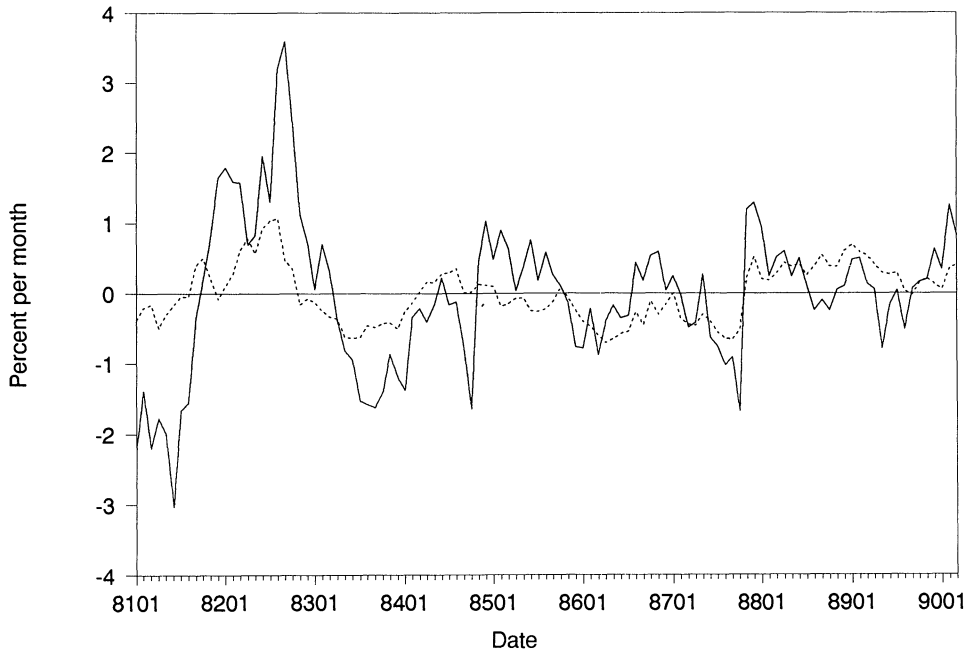
Figure 3. Forecast U.S. Excess Stock Return, 1981–1990. The solid line is the unrestricted forecast from the basic specification estimated in Panel A of Table III. The dashed line is the restricted forecast from the single-latent-variable model estimated in Table V, again using the basic specification.

quintiles of market value give results similar to those reported for value-weighted indexes.

Fourth, we have estimated a single-latent-variable model over the full 1971–1990 sample period, but allowing a change in the θ coefficients at the end of 1980. This model imposes perfect correlation in the forecasts for the U.S. and Japan, but allows these forecasts to shift in relation to the regressors in the middle of the sample period. This model is rejected at significance levels ranging from 8.3% to 0.03%, which is what one would expect given the rejections of the basic model in the 1970's.

V. Conclusion

In this paper we have studied international capital market integration by comparing the predictable components of excess stock returns in the U.S. and Japan. Our main results are as follows.

First, in both countries it is generally possible to forecast excess stock returns relative to the U.S. Treasury bill rate using similar sets of domestic variables. The domestic dividend-price ratio has a generally positive effect on excess stock returns, while the relative short rate (the difference between the
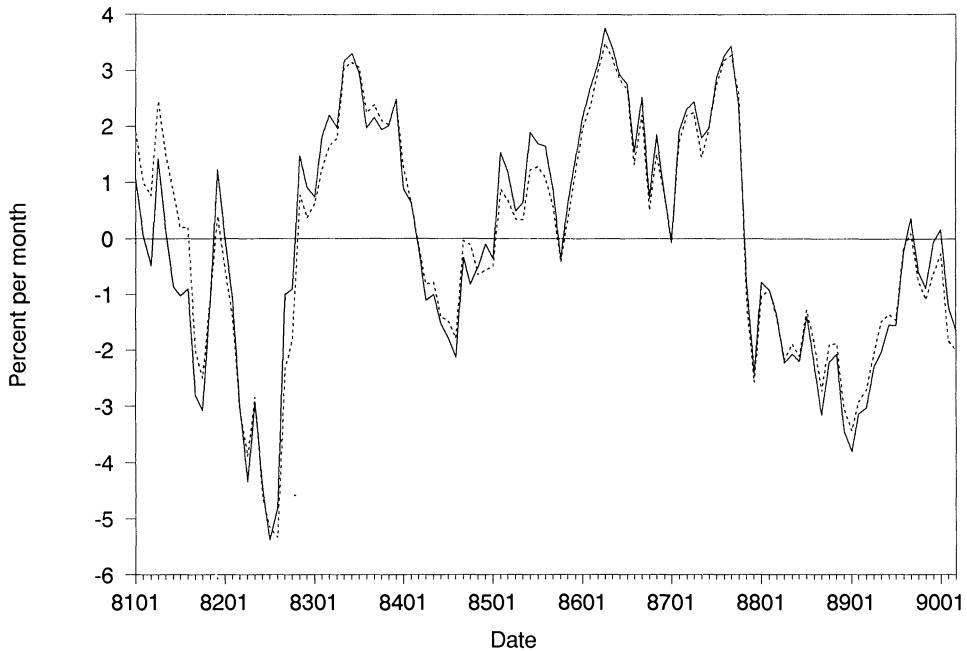
Figure 4. Forecast Japanese Excess Stock Return, 1981–1990. The solid line is the unrestricted forecast from the basic specification estimated in Panel A of Table III. The dashed line is the restricted forecast from the single-latent-variable model estimated in Table V, again using the basic specification.

current short rate and its 1-year backward moving average) has a negative effect. The main evidence for these effects comes from the 1970's in both countries. The 1980's add little to the evidence, because we cannot reject that the forecasting coefficients in this decade are the same as in the 1970's, but equally we cannot reject that they are zero.

Second, U.S. variables help to forecast Japanese excess stock returns in the 1980's. The level of the Japanese dividend-price ratio relative to the U.S. dividend-price ratio is a powerful forecasting variable for Japanese returns. There is weaker evidence that Japanese variables help to explain U.S. excess stock returns in the 1980's.

Third, the movements of expected excess returns on the U.S. and Japanese markets are not well explained by a model where assets have constant betas on a single "international factor," proxied by a world stock index return, whose risk price changes over time.

Fourth, in the 1970's expected excess stock returns in the U.S. and Japan are positively correlated. We can reject at the 5% level the hypothesis that expected excess stock returns in the two countries are perfectly correlated, but our estimates of the common "international" component of expected excess returns explain more than 70% of the variance of expected returns in the U.S., and as much as 60% of the variance of expected returns in Japan.

This common movement of expected excess returns is suggestive of at least partial integration of U.S. and Japanese stock markets.

We would like to be able to compare the common movement of expected excess returns in the 1970's with that in the deregulated 1980's. Unfortunately it is hard for us to measure the correlation of expected excess returns in the 1980's, because we have only weak forecasting power for excess U.S. stock returns in this decade.

These results are consistent with the view that an important determinant of expected stock returns is the changing price of risk of a single common factor in a world capital market. However we do not wish to overstate the strength of the evidence. In the 1980's we cannot precisely measure common movement of expected excess returns. In the 1970's our results are stronger, but it is of course possible that the common movement of expected returns results from common shocks affecting segmented markets, rather than from the operation of an integrated world capital market. In any event, our results should help to guide research on the causes of changing expected stock returns in the United States. Whatever these causes are, they cannot be entirely local but must have the potential to move expected stock returns in other countries as well.

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