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Capital Controls and International Capital Market Segmentation: The Evidence from the Japanese and American Stock Markets

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ABSTRACT

The paper focuses on two countries, Japan and the U.S., to test the integration of capital markets. In Japan, the enactment of the Foreign Exchange and Foreign Trade Control Law in December of 1980 amounted to a true regime switch that virtually eliminated capital controls. Using multifactor asset pricing models, we show that the price of risk in the U.S. and Japanese stock markets was different before, but not after, the liberalization. This evidence supports the view that governments are the source of international capital market segmentation.

THE NOTION THAT WORLD capital markets are perfectly integrated is central to most theoretical developments in both international finance and macroeconomics; indeed, it also lies behind much of the recent international policy discussion. Capital markets are integrated if assets with perfectly correlated rates of return have the same price regardless of the location in which they are traded.[1] Segmentation may arise either because of government impediments to capital movements or because of individuals' attitudes or irrationality. If the hypothesis that the international capital markets are integrated is rejected, there is still an important empirical question left unanswered which is related to the source of segmentation. While there is a rich body of theoretical research on international market integration, only a few studies have tried to investigate this important question empirically.

A first group of empirical studies on market integration has generally adopted an international single index asset pricing model à la Sharpe-Lintner to test

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[1] This seems to be the natural way of defining capital market integration. (See Stulz (1981b).)

whether a purely domestic factor—usually the part of the return on the domestic market portfolio that is orthogonal to the world portfolio—has explanatory power in a regression of stock returns on a world market index.[2] The finding that the domestic factor is often priced is then taken to support market segmentation.

The interpretation of this empirical evidence, however, is not so clear-cut because a single index international capital asset pricing model can only be obtained under the restrictive assumptions of either a universal logarithmic utility function (Adler and Dumas (1983)) or purchasing power parity (Grauer, Litzenberger, and Stehle (1976)) or no correlation between exchange rate movements and stock returns as well as deterministic domestic consumption deflators (Solnik (1974a)). With deviations from purchasing power parity and more risk aversion than the logarithmic investor, for example, the equilibrium rate of return on an asset would also depend on its correlation with the inflation rates in various countries (Adler and Dumas (1983) and Stulz (1985)) so that the pricing of purely domestic factors would not necessarily indicate lack of integration.

Single index models also do not address the question of whether segmentation arises from government policies or from market inefficiency. Theoretically, a powerful test could be devised by specifying two asset pricing models—one without and one with barriers to international investments—and then verifying whether the additional restrictions implied by capital controls are supported by the data. Unfortunately, the pricing models will be different depending on the type of barrier imposed by governments (Stulz (1981a) and Eun and Janakiramanan (1986)), and, realistically, one cannot hope to capture the characteristics of the various countries' capital control regimes with just one model specification.

A second group of studies have used an international version of the Arbitrage Pricing Theory (APT).[3] Because the pricing in this model is based on an arbitrage condition of nominal returns, APT has the advantage of eluding the problem of purchasing power parity deviations. These studies generally apply APT to a large set of countries, and they typically point to market segmentation although they do not provide information as to the sources of segmentation, namely government policies versus market failures. Still another approach is to use consumption asset pricing models. Obstfeld (1986) tests the equality of the marginal rate of substitution in consumption in different countries. He assumes that agents have access to the same risk-free interest rate and cannot reject the null hypothesis of integration. Wheatley (1988) tests international equity market integration using a simple version of the consumption asset pricing model. His tests provide little evidence against the joint hypothesis that the U.S. and the seventeen non-U.S. equity markets are integrated and that the consumption-based asset pricing model holds.

Given the current status of international asset pricing models and the complexity of governments' impediments to capital movements, we believe that

[2] This approach is used by Stehle (1977) and Jorion and Schwartz (1986). A similar framework is used in the empirical studies of Solnik (1974b) and Errunza and Losq (1985).

[3] Cho, Eun, and Senbet (1986) and Korajczyk and Viallet (1986) use an APT framework to test international capital markets integration. See also Berges (1981) for an empirical study using an international APT.

generalized tests of capital market integration are likely to be uninformative.[4] Therefore, in this paper, we try a different approach which is similar in spirit to an event study.[5] We focus on the experience of just one country, Japan, at the time of the implementation of its new Foreign Exchange and Foreign Trade Control Law in December 1980. This law amounted to a true regime switch that virtually eliminated most capital controls in an economy where the financial markets had been highly regulated. If the controls were the only source of segmentation, we should be able to reject the hypothesis that the price of risk in any currency—defined in terms of various multifactor asset pricing models—was the same in the Tokyo and in the New York stock markets before the end of 1980 but not after that date. The results that we obtain thus lack generality, but, we hope, are more conclusive than previous research in the field.

In the four years after liberalization, we cannot find any sign of segmentation between the Japanese and U.S. security markets. Arbitrage has thus ensured that risk carries the same price in the two countries when expressed in a common numeraire. By contrast, we often reject the equality of both the risk premia and the return on the risk-free asset before the liberalization of December 1980. Although international asset pricing models perform less well after the liberalization, thus reducing the power of the test, the data support the view that government policies, rather than investor behavior, are the source of international capital market segmentation.

Our paper is organized in the following order: Section I reviews the Japanese experience with capital controls; Section II describes the model and the hypotheses being tested; Section III explains the data; Sections IV and V present empirical tests utilizing prespecified risk factors for individual securities and portfolios; Section VI presents tests utilizing risk factors estimated with factor analysis; and Section VII concludes the paper.

I. The Japanese Experience with Capital Controls: A Brief Account

Until 1974, no Japanese security firm could buy foreign financial assets and no foreign company could buy Japanese securities. Domestically, there was no free market for short-term assets. A steady liberalization process of domestic financial markets began in 1974, although interest rates in the money markets were completely deregulated only in 1978–1979.[6] Controls on international capital flows were instead effectively maintained even in the second part of the seventies as a policy tool to manage the exchange rate (Otani and Tiwari (1981) and Ito (1986)). In an attempt to check the sharp appreciation of the yen from mid-1977 to the end of 1978, capital outflows were encouraged and inflows discouraged: nonresidents were prohibited from purchasing Japanese securities with maturities of less than five years and one month; the marginal reserve requirement on yen

[4] The difficulties associated with testing international capital asset pricing models have long been recognized in the literature. Solnik (1977) contains an early discussion of the issue.

[5] An "event study" approach to capital market segmentation is also used by Alexander, Eun, and Janakiramanan (1987, 1988) who look at the dual listing of companies in the Canadian and U.S. markets.

[6] See Pigott (1983) for a detailed account of the Japanese experience with financial deregulation.

accounts of nonresidents was increased in steps to 100 percent; and Japanese institutions were for the first time allowed to purchase foreign securities. After a brief relaxation of the controls in 1979, Japanese authorities resorted again to their use in 1980, this time to support a falling yen: the reserve requirements on nonresident deposits were rolled back to zero; Japanese banks were encouraged to raise funds in London and to transfer them to Tokyo; and foreigners were allowed to trade in the domestic money market.

A sharp change in the capital control regime has occurred since December 1980, when the enactment of the Foreign Exchange and Foreign Trade Control Law completely liberalized short-term capital movements. Since then, the Japanese government has also supported the "internationalization" of the yen and improved foreigners' access to all Japanese security markets.

The dramatic impact of Japanese controls between 1977 and 1980, and of the subsequent liberalization, is clearly visualized in Figure 1. This figure depicts the interest rate differential between three-month Euroyen deposits traded in London and three-month repurchase agreements (Gensaki) traded in Tokyo—the market that has been least affected by administrative controls. Before 1981, the interest rate differential was substantial—it was negative until 1979, when capital was prevented from flowing into Japan, and positive in 1980, when the opposite occurred. However, the differential has been practically zero since 1981, indicating that any government-induced segmentation of the market had ceased its effects on market prices. The interest rate differential is only a measure of money market segmentation. However, data on stock-related capital movements in and out of Japan (Table I) reveal a sharp increase in two-way flows after 1980.

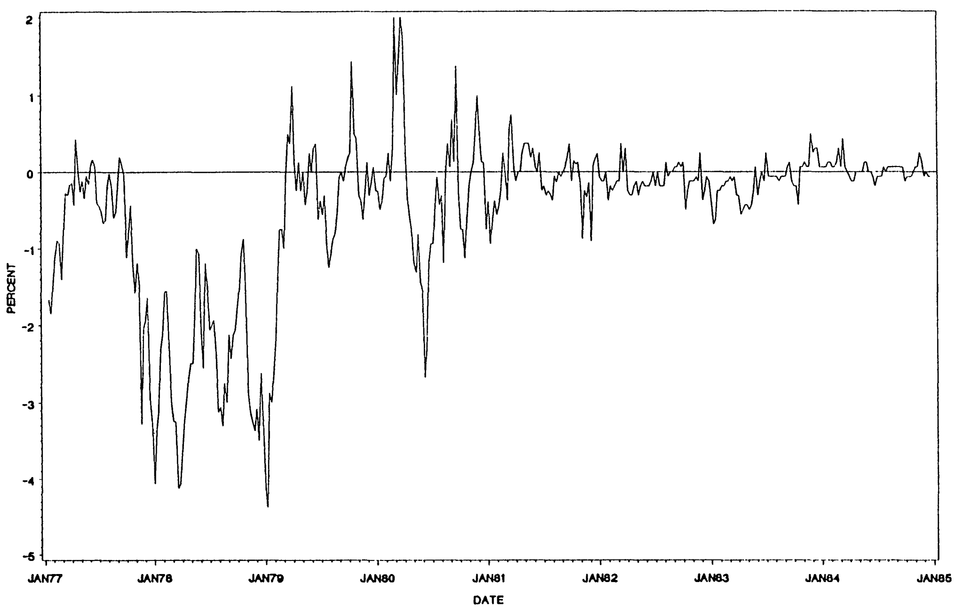

Figure 1. Spread between three-month Euroyen in London and three-month Gensaki in Tokyo. Data were from the Board of Governors of the Federal Reserve and Data Resources Inc.

Table I

Gross Equity Related Capital Movements in Japan

Gross equity trade is measured as the sum of stock purchases and sales for the fiscal year. Fiscal years in Japan begin on April 1. Trades are in million dollars. Source: Bank of Japan, Annual Report, 1987.

Fiscal Year	Japanese Trade in Foreign Stocks	Foreign Trade in Japanese Stocks
1977	466	6,131
1978	653	11,385
1979	1,030	12,034
1980	1,132	25,383
1981	1,533	44,699
1982	2,663	35,905
1983	3,356	70,025
1984	4,503	77,825

Given the depth of the controls and the extent of the subsequent liberalization program, the Japanese experience constitutes a unique opportunity for a semi-controlled experiment to test international capital market segmentation. If government policies, as opposed to investor attitudes, were the only source of segmentation, the price of risk in the U.S. and Japanese stock markets (expressed in the same currency) should be different before 1981, but the difference should disappear afterwards. This is the hypothesis that we test in the following sections.

II. Testing Market Segmentation with a Multifactor Asset Pricing Model

In order to test the hypothesis of market integration, namely that the price of risk in the two markets is the same, we need a model that defines risk. We choose an international version of a multifactor asset pricing model in which equilibrium returns are set by arbitrage (APT) for two reasons. First, earlier empirical evidence indicates that more than one factor may affect security prices in an international setting, but it is difficult to derive the exact form of the pricing equation. Second, the pricing of assets can be based on an arbitrage relation of nominal rates of return so that one need not worry about deviations from purchasing power parity. It is worthwhile to note, however, that this paper is by no means a test of APT or any other model; rather the multifactor specification and the arbitrage pricing framework is used to generate a benchmark which is used to compare pre- and post-liberalization rates of return.

In our two-country case we have $N + 2$ securities traded in the U.S. and Japan, which are indexed by i. We assume that security 0 is the risk-free asset in U.S. dollars and security 1 is the Japanese risk-free asset, which, however, is risky in dollar terms. The remaining N assets are Japanese and American stocks. If the U.S. dollar is used as the numeraire, APT postulates that the return-generating function for the $N + 1$ risky securities is linear:

$$\boldsymbol{r}_t = \boldsymbol{E}_t + \boldsymbol{B}\delta_t + \boldsymbol{e}_t, \tag{1}$$

where $\boldsymbol{r}$ is the $(N + 1) \times 1$ vector of nominal rates of return (in dollars) at time t, $\boldsymbol{E}$ is the vector of expected returns, δ is the $k \times 1$ vector of risk factors with mean zero and covariance matrix of unity, $\boldsymbol{B}$ is the $(N + 1) \times k$ matrix of the sensitivity coefficients to the risk factors, and $\boldsymbol{e}$ is the vector of idiosyncratic terms with a diagonal covariance matrix $\boldsymbol{\Sigma}$.

It is well known that, in the absence of arbitrage, expected returns (in dollars) should be linear in $\boldsymbol{B}$; that is,

$$\boldsymbol{E}_t = \lambda_{0t}\iota + \boldsymbol{B}\lambda_t, \tag{2}$$

where λ is the vector of risk premia associated with risky factors, λ_0 is the risk-free rate, and ι is an $(N + 1) \times 1$ vector of ones.

Solnik (1983) has shown that the same arbitrage portfolio will hold for the Japanese investor who calculates rates of return in yen provided that the dollar/yen rate follows the linear return-generating function given in (1). The empirical counterpart of the equilibrium pricing equation is given at any point in time t by substituting equation (2) into (1):

$$\boldsymbol{r}_t^c = \lambda_{0t}^c\iota + \boldsymbol{B}^c\lambda_t^c + \boldsymbol{v}_t^c, \quad \mathrm{c} = US,\ JA, \tag{3}$$

where $\boldsymbol{v}$ is the vector of error terms, $\boldsymbol{B}\delta_t + \boldsymbol{e}_t$, and the superscript c indicates whether it is a Japanese or an American stock.

Based on the specification in (3), our empirical tests of market segmentation are based on a well known two-stage estimation approach. In the first stage the elements of $\boldsymbol{B}$ are estimated for each security by using the time series of returns.[7] In the second stage estimated matrix $\boldsymbol{B}$ is used as an independent variable to estimate λ_0 and vector λ for each country both for the sample period 1977–1980, during which the capital controls were in effect, and for the subsequent four years 1981–1984, which were characterized by free capital movements. Then, the composite null hypothesis that

$$\hat{\lambda}^{US} \neq \hat{\lambda}^{JA}, \quad \hat{\lambda}_0^{US} \neq \hat{\lambda}_0^{JA} \quad \text{in 1977–1980}$$

and

$$\hat{\lambda}^{US} = \hat{\lambda}^{JA}, \quad \hat{\lambda}_0^{US} = \hat{\lambda}_0^{JA} \quad \text{in 1981–1984}$$

is tested formally.

Because there is no agreement on the empirical implementation and testing of multifactor models, we used several approaches—prespecified factors and factor analysis with individual securities and portfolios of securities—as well as different multivariate testing techniques. If the impact of the capital market liberalization had a sufficiently strong impact on rates of return around the time of the regime switch, as we conjecture, then we should expect the results to be robust with

[7] In the case of the factor analytic approach, the number of factors k and the elements of $\boldsymbol{B}$ are jointly estimated in the first stage.

respect to the alternative testing procedures and inevitable model specification errors.

III. Data

We used weekly stock returns calculated from daily prices quoted at the close of the markets on Wednesdays from January 1, 1977, to December 31, 1984. As indicated before, we focus on two sample periods, January 1977 to December 1980 and January 1981 to December 1984. We selected two four-year periods around the December 1980 liberalization measures because four years was the longest period for which we felt comfortable with the assumption of risk factor stationarity.[8] In addition, the Bank of Japan did not officially admit a free money market until 1977, even though one started to develop as early as 1971.

We selected two sets of Japanese and U.S. securities, each consisting of 110 stocks, which reproduced the industry composition of the two countries' market portfolios as closely as data availability permitted us. The choice of an equal number of stocks and the attention paid to their industrial classification were dictated by our concern regarding the introduction of spurious risky factors in either set of securities. In Table II we show the proportion of the companies in the sample relative to the total number of companies in each industrial sector for two broad market indexes: the total of the publicly traded companies in Japan as reported by *Toyo Keizai Shinposha* (1985) and the largest 1,000 corporations in the U.S. as ranked by *Business Week* (1986).[9]

We converted Japanese stock returns into dollars. We tried to match the timing of stock transactions with that of the exchange transactions as closely as possible. An investor who wants to get the Wednesday dollar rates of return at the closing prices of the Japanese stock market will buy the stock on Wednesday at 3 p.m. Tokyo time and sell it at the same time one week later. Given that it takes three business days to settle stock transactions in Tokyo, he or she needs yen on Monday, while he or she will receive the yen the next Friday. Two business days are needed for settlement in foreign exchange markets. Therefore, the foreign exchange transaction must occur on Monday in Tokyo. By taking the dollar/yen exchange rates quoted at 1 p.m. New York time on Thursdays, we lead the opening of the Tokyo foreign exchange market by four hours (1 p.m. in New York on Thursday is 4 a.m. in Tokyo on Friday). A far more serious problem, albeit unsolvable, is that of nonsynchronous trading given that trading hours in Tokyo and New York markets never overlap.[10] The sources of all the data are given in the Appendix.

[8] Nomura Research Institute kindly computed the stock returns for us which had to be adjusted for frequent stock splits in Japan.

[9] Because the original data on Japanese stocks were classified into twenty industry categories, we regrouped the U.S. companies according to that classification as close as our data allowed us.

[10] Nonsynchronous trading may bias not only risk premia but also the returns covariance, and therefore, the estimation of risk factors. (See Shanken (1987).)

Table II

Industry Composition of Sample Portfolio and Market Portfolio

The industry compositions of the Japanese market portfolio and the U.S. market portfolio are obtained from the *Japan Company Handbook,* Toyo Keizai Shinposha (1985) and the special issue, "The Top 1000", of *Business Week* (1986), respectively.

	Number of Companies in the Sample as Percentage of Country Market Portfolio	
Sectors[a]	Japan	U.S.
Fishery	14.3	—
Construction	5.3	10.0
Foods	9.6	11.1
Textiles	6.5	13.3
Chemicals	16.1	11.4
Natural Resources	20.0	11.1
Rubber Goods	11.1	16.6
Iron and Steel	12.8	11.1
Non-Ferrous Metals	8.7	9.1
Machinery	4.9	10.0
Electronics and Computers	16.8	11.1
Transport Equipment	14.3	13.3
Precision Machinery	23.5	9.4
Miscellaneous Manufacturing	13.0	11.8
Commerce	8.6	11.0
Banking	6.1	11.3
Non-Banking Financial	28.0	11.1
Real Estate	15.8	20.0
Transportation	10.0	10.3
Utilities	14.3	10.8
Tobacco, Paper and Forestry	—	10.0
Conglomerates	—	8.3

[a] Because the Japanese stocks in our data were classified into twenty industry categories, we regrouped the U.S. companies according to that classification as close as the data allowed us.

IV. Prespecified Factors with Individual Stock Returns

The first approach was to identify the sources of risk explicitly by "guessing" the economic variables that have a systematic effect on stock prices (Chen, Roll, and Ross (1986) and Hamao (1988)). The approach can be criticized for its arbitrariness, but it was important to check the sensitivity of the results to all possible alternative implementations of the multifactor model. In addition, this approach subsumes early empirical work in international finance (Solnik (1974b) and Stehle (1977)).

With weekly data, the choice of economic variables entering the stock return function is limited. We selected six variables. *DIJ* and *DIU* measure the change in the short-term interest rate in Japan and the U.S., respectively. The three-month Gensaki rate is used to calculate *DIJ*, and the three-month Eurodollar deposit rate is used to calculate *DIU*. These variables capture movements in the term structure of interest rates in the spirit of recent single-state-variable theoretical models of the equilibrium structure. The third variable, the percentage change in the Dow Jones Commodity Futures Index (*FUT*)—an index of twelve commodities including foods, metals, and woods—was chosen to proxy for world demand pressures and inflationary expectations.

In principle, stock market indexes should not enter the regressions if all relevant economic variables are identified and properly measured; in practice, however, misspecifications are bound to occur, especially when the choice of variables is limited due to the weekly data frequency. Therefore, we added the percentage change in the world stock market (*WINDEX*). *WINDEX* was calculated by weighing the indexes of the stock markets of France, Germany, Switzerland, the United Kingdom, Japan, Australia, Canada, and the U.S. by their capitalizations at the end of 1980, the midpoint of our sample period. Table III reveals that the stock markets of these countries represent 93.4 percent of total world stock market capitalization. Thus, our proxy for the world stock market index is adequate.

Finally, we use two purely domestic market factors in the regressions, namely the parts of the Japanese (*JRES*) and U.S. (*USRES*) stock indexes that are orthogonal to the world stock index. The last three variables typically have been used in many studies of international capital market integration. Changes in the exchange rate, by themselves, are not a source of systematic risk for stocks since the stockholder can hedge against it perfectly by borrowing the foreign currency. The changes, however, may be an indirect source of risk to the extent that they affect the real exchange rate and, consequently, the profitability of the export sector. The empirical evidence presented in Hamao (1988) shows that neither the exchange rate nor the terms of trade are priced factors in the Japanese stock

Table III

Equity Markets Capitalization

Equity market capitalization is measured in U.S. dollars as of the end of 1980. Data are from Ibbotson, Carr, and Robinson (1982).

Countries Included in the World Index	$ Billions	% World Total	% World Index
France	53	2.2	2.3
Germany	71	2.9	3.1
Switzerland	46	1.9	2.0
United Kingdom	190	7.8	8.4
Japan	357	14.7	15.7
Australia	60	2.5	2.7
Canada	113	4.6	5.0
United States	1381	56.8	60.8
Total World Index	2271	93.4	100.0
Other Countries	159	6.6	—
Total World Equities	2430	100.0	—

market. In addition, any direct effect of foreign exchange movements should be already captured by the two domestic factors *JRES* and *USRES*.

The model specification imposes that the vector of risk factors, δ, be mean zero and serially uncorrelated, although they may be contemporaneously correlated. Because all the variables are first differences of weekly data, we do not find strong serial correlation, with one exception—the Japanese interest rates which moved in a stepwise fashion throughout most of the sample period. The estimated means are comfortably close to zero, with the exception of the world stock index (Table IV).[11]

The initial step is the estimation of the $\boldsymbol{B}$ matrix. In this section we use the time series of individual stock returns from the entire sample period 1977–1984. We thereby assume that the risk characteristics of the securities remain unchanged throughout the liberalization process even though the price of the risk factors changed between the two subperiods. There are two problems with this approach: the consistency of the risk premia estimates—because both $\boldsymbol{B}$ and λ are estimated from the same sample period—and the stationarity of $\boldsymbol{B}$. Both problems will be dealt with in the next section by following the usual approach of forming portfolios and using out-of-sample betas; in this section, however, we do not want to lose the information that the grouping of securities would involve.

Some descriptive statistics for betas from individual stock returns are reported in Table V. For the entire sample period, all the risk factors had a strong impact on U.S. stock returns, with the exception of Japanese interest rates. For the Japanese stocks, only the world stock market index and Japanese variables seem to matter. The U.S. betas seem remarkably stable. The estimated betas are similar for the two subperiods, 1977–1980 and 1981–1984, although the t-statistics suggest rejection of the equality of the world index parameters. As for the Japanese stocks, the betas are quite unstable—the Hotelling T^2 rejects the null of equal $\boldsymbol{B}$'s at any conventional level of significance. In particular, commodity futures and Japanese interest rates appear to have different effects in the two subperiods.[12]

The next step is to test the equality of the vector of risk premia for the U.S. and Japanese securities by using cross-sectional regressions of stock returns on the estimated betas. We use three testing procedures. First, we use mean rates of stock returns in each of the subperiods as the dependent variables in the cross-sections and estimate the vector of U.S. and Japanese premia jointly with a seemingly unrelated regression. Using our notations, we estimate the system of two equations:

$$\left.\begin{aligned} \bar{r}_j^{US} &= \lambda_0^{US}\iota + \hat{\boldsymbol{B}}_j^{US}\hat{\lambda}^{US} + v_j^{US} \\ \bar{r}_j^{JA} &= \lambda_0^{JA}\iota + \hat{\boldsymbol{B}}_j^{JA}\hat{\lambda}^{JA} + v_j^{JA} \end{aligned}\right\}, \quad j = \begin{matrix} 1977\text{–}1980, \\ 1981\text{–}1984 \end{matrix}$$

[11] An alternative would have been to filter the data with univariate models or vector autoregressions, but we were afraid of the potential bias caused by a wrong specification of the processes for the time series.

[12] To determine whether the liberalization of capital movements could account for the changes in the Japanese betas, in addition to the risk premia, we used the differential between domestic yen and Euroyen deposits of the same maturity as a proxy variable. This variable is plotted in Figure 1. Perhaps the proxy was inappropriate; when we re-estimated the betas with the capital control proxy added on the right-hand side, its coefficient was never significant.

with seemingly unrelated regression for each of the two subperiods, where $\bar{r}_j$ stands for the mean return over the period j. The equality of the premia is then tested with a standard F-test of the linear constraint (Theil (1971), p. 313).

Second, we follow the multivariate approach of Shanken and Weinstein (1985), which takes into account the estimation errors of the betas in the first step of the procedure.[13] The test statistic, which we denote SW, is given by the ratio of two quadratic forms and is distributed T^2:

$$SW = \frac{P(\hat{\lambda}^{US} - \hat{\lambda}^{JA})'(\boldsymbol{H}\Sigma^{-1}\boldsymbol{H}')^{-1}(\hat{\lambda}^{US} - \hat{\lambda}^{JA})}{1 + \hat{\lambda}'\Omega^{-1}\hat{\lambda}} \approx T^2(k+1, P-k-1), \quad (4)$$

where P is the number of periods used to estimate the betas in the first step, Ω is the covariance matrix of risky factors, λ is the vector of premia obtained from the regression in which the equality of the premia had been imposed, and $\boldsymbol{H}$ is defined as

$$(\boldsymbol{H} = (\boldsymbol{A}^{US} : -\boldsymbol{A}^{JA}), \text{where} \qquad \boldsymbol{A} = (\hat{\boldsymbol{B}}'^{c}\hat{\boldsymbol{B}}^{c})^{-1}\hat{\boldsymbol{B}}'^{c}, \quad c = US, JA.$$

Third, we adopt the Fama and MacBeth (1973) procedure and estimate the risk premia for both countries in each week of the two subperiods, or 208 weeks for each subperiod. In this way we obtain two time series of premia differentials which should have a zero mean if capital markets were perfectly integrated. The Hotelling T^2 statistics are then used to test the vector of mean differentials.

The results are presented in Table VI. Cross-sectional estimates of risk premia, obtained by ordinary least squares, are shown at the tops of Panels A and B. More sources of risk seem to be priced before the liberalization of capital controls than after; this is particularly true for the Japanese securities. At the bottom sections of Panels A and B, we present the statistics and procedures testing the equality of premia with and without the intercept term. Without the intercept, the SW and T^2 statistics indicate that capital market integration is supported by the data in both subperiods, even though the absolute values of the statistics decline substantially after 1980. The F-statistics suggest instead that the hypothesis of segmentation cannot be rejected at the 0.05 significance level in the 1977–1980 period but can in the subsequent years.

The results are substantially the same when the equality of the intercept term is tested together with the equality of the premia. This time, however, even the F-statistics do not detect segmentation before 1981, which is somewhat surprising given that we have imposed an additional restriction. Based on the analysis of individual securities, we mostly find that the price of risk has been successfully arbitraged between the U.S. and Japanese stock markets notwithstanding the capital controls existing before 1981. These controls appear to have segmented only the Japanese money market, as clearly illustrated in Figure 1. As we pointed out, lack of stationarity and consistency of the parameter estimates may mar the analysis of individual stock returns. In the next section we repeat the tests by grouping securities into portfolios, a procedure commonly used to minimize the two sources of bias (Black, Jensen, and Scholes (1972)).

[13] The test draws on the work of Shanken (1985). See also Bodurtha (1986) for an application of this test to an international asset pricing model.

Table IV

Risk Measures: Descriptive Statistics

Descriptive statistics are computed using weekly data for the full sample period from January 5, 1977 to December 26, 1984. The risk measures are:

WINDEX: Percentage change in World stock market index.
JRES: Purely domestic component of Japanese stock market index.
USRES: Purely domestic component of U.S. stock market index.
DIU: Change in short-term interest rate in the U.S.
DIJ: Change in short-term interest rate in Japan.
FUT: Percentage change in the Dow Jones commodity futures index.

World stock market index is a value-weighted portfolio of the eight countries' stock indices. The list of the countries and the weights are given in Table III. Individual stock market indices are obtained from the Financial Times, London. Short-term interest rates are obtained from the Board Governors of the Federal Reserve, Washington, D.C. The Dow Jones commodity futures index is obtained from the Dow Jones Educational Services, Princeton.

Mean and First-Order Autocorrelation of Risk Measures

	Mean (%)	t	ρ	SE (ρ)
WINDEX	0.19805	2.15	0.048	0.05
JRES	0.00000	0.00	0.032	0.05
USRES	0.00000	0.00	0.025	0.05
DIU	0.00017	0.31	0.076	0.05
DIJ	0.00001	0.00	−0.502	0.05
FUT	0.03195	0.31	0.127	0.05

Correlation Matrix of Risk Measures

	WINDEX	*JRES*	*USRES*	*DIU*	*DIJ*	*FUT*
WINDEX	1.00	0.00	0.00	−0.40	−0.15	0.37
JRES		1.00	−0.78	−0.08	−0.49	0.03
USRES			1.00	−0.17	0.46	−0.17
DIU				1.00	0.13	−0.18
DIJ					1.00	−0.10
FUT						1.00

V. Prespecified Factors with Portfolios of Securities

Given the small number of securities in each country sample, precisely 110, we can only form twenty-two portfolios of five securities. In each subperiod, 1977–1980 and 1981–1984, we took the first two years, 1977–1978 and 1981–1982, and estimated the beta of each security with respect to its domestic market index. This beta was used to rank securities and form the portfolios. We then estimated the portfolio betas with respect to all six sources of risk specified in the previous section using again the returns from the first two years only. The portfolio betas estimated in this way were used as independent variables to obtain estimates of

Table V

Security Betas: Descriptive Statistics

OLS coefficients of risk measures are estimated for each security with weekly data. The Full period covers from January 5, 1977 to December 26, 1984. The first subperiod, Period 1, is from January 5, 1977 to December 31, 1980, and the second subperiod, Period 2, is from January 7, 1981 to December 26, 1984. Means and standard deviations of coefficients are computed over 110 companies in each country. The coefficients correspond to risk measures:

WINDEX: Percentage change in World stock market index.
JRES: Purely domestic component of Japanese stock market index.
USRES: Purely domestic component of U.S. stock market index.
DIU: Change in short-term interest rate of the U.S.
DIJ: Change in short-term interest rate in Japan.
FUT: Percentage change in the Dow Jones commodity futures index.

World stock market index is a value-weighted portfolio of the eight countries' stock indices. The list of the countries and the weights are given in Table III. Individual stock market indices are obtained from the Financial Times, London. Short-term interest rates are obtained from the Board of Governors of the Federal Reserve, Washington, D.C. The Dow Jones commodity futures index is obtained from the Dow Jones Educational Services, Princeton.

Estimation Period		U.S. Securities: Coefficients of Risk Measures					
		WINDEX	*JRES*	*USRES*	*DIU*	*DIJ*	*FUT*
Full Period:	Mean	1.08	0.10	1.18	−21.70	0.0003	−0.052
	Std. Dev.	0.396	0.195	0.678	37.833	0.112	0.108
Period 1:	Mean	1.18	0.10	1.19	−21.60	0.005	−0.052
	Std. Dev.	0.496	0.273	0.756	39.191	0.159	0.116
Period 2:	Mean	1.02	0.09	1.17	−21.30	−0.004	−0.052
	Std. Dev.	0.391	0.261	0.856	46.748	0.178	0.187

Estimation Period		Japanese Securities: Coefficients of Risk Measures					
		WINDEX	*JRES*	*USRES*	*DIU*	*DIJ*	*FUT*
Full Period:	Mean	0.73	1.03	0.01	−0.37	−0.027	−0.02
	Std. Dev.	0.261	0.251	0.479	14.519	0.117	0.083
Period 1:	Mean	0.31	1.05	0.09	−0.66	0.02	−0.04
	Std. Dev.	0.237	0.313	0.745	21.464	0.147	0.102
Period 2:	Mean	0.92	1.01	0.01	0.43	−0.07	0.02
	Std. Dev.	0.326	0.322	0.645	28.488	0.177	0.139

	Tests of Equality of Betas between Subperiods (t-Values)						Hotelling[a] T^2
U.S. Securities	4.78	0.31	0.36	0.09	0.40	0.004	4.52 (0.0005)
Japanese Securities	19.3	1.37	0.78	0.28	4.53	3.45	107.0 (0.000)

[a] p-Values in parentheses.

the risk premia in the second half of the two subperiods, that is, 1979–1980 and 1983–1984.

There are no theoretical reasons for using the domestic market beta to rank securities; it simply turned out to be an effective way to spread securities returns out of sample.[14] Due to lack of dispersion in firm size in both the U.S. and Japanese samples, the popular method of ranking on firm size could not be used.

We are also aware of the bias that may be introduced when individual security betas, used to form the portfolios, and portfolio betas are estimated from the same sample period. We did not have much choice, however, because our entire sample period consists of only four years, two of which were needed for estimating the risk premia.[15]

Table VII reports the estimated portfolio betas for the two subperiods. As for individual securities, U.S. betas seem more stable than Japanese betas. Surprisingly, changes in Japanese interest rates are correlated with U.S. portfolio returns but not with Japanese portfolio returns.

In Table VIII we show the same three test statistics that were introduced in the previous section and the risk premia estimates for portfolios. The results are clear-cut. With or without the intercept term included, the hypothesis of integration is rejected at the 0.05 significance level by both SW and F-statistics for the 1979–1980 period. In contrast, the same hypothesis cannot be rejected at any meaningful significance level for the years 1983–1984. Only the T^2 statistic, with a significance level of 0.18, does not provide evidence of segmentation before the liberalization process. The significance of the results, however, is somewhat reduced by the lack of precision with which the risk premia are estimated in the second subperiod, both with the individual securities and with portfolios.

VI. Factor Analytic Approach

Factor analysis is an alternative methodology where one determines the number of factors k and estimates the elements of $\boldsymbol{B}$ jointly. We extracted the factor loadings from the full covariance matrix of U.S. and Japanese stock returns, as well as the dollar return on the Japanese risk-free rate, which we assumed to be equal to the three-month interest rate on Euroyen deposits. The matrix was calculated by using the time series of returns for the entire sample period, from 1977 to 1984, thereby assuming stationarity of the individual stocks' "riskiness." In this respect the analysis is similar to that of Section IV.

A well known issue in the estimation of APT models is the decision about the number of factors to be extracted. Because the number of factors required grows with the number of securities in the sample, the model soon loses its empirical content unless one arbitrarily fixes the number of priced factors (Roll and Ross (1980) and Dhrymes, Friend, and Gultekin (1984)). In our case, the covariance matrix of returns contains 221 securities, and the number of factors needed to explain the covariance structure of security returns would exceed thirty. We limited the number of factors to five, ten, and twenty factors, as is done in most current empirical tests of the APT.

Once the security factor loadings were obtained, we estimated and tested the

[14] Multivariate tests reject the equality of portfolio mean returns in each of the subperiods. Thus, we can safely rule out the probability of results being an artifact of the experimental design.

[15] We also formed portfolios by grouping securities according to their gross returns, but the results did not change significantly.

Table VI

Risk Premia Based on Individual Stock Returns

The risk premia are estimated cross-sectionally using the coefficients of risk measures, betas, as the independent variables. The dependent variables are the mean returns of 110 U.S. securities and the Japanese securities in each of the subperiods. The first subperiod, Period 1, is from January 5, 1977 to December 31, 1980, and the second subperiod, Period 2, is from January 7, 1981 to December 26, 1984. The risk premia estimates correspond to risk measures:

WINDEX: Percentage change in World stock market index.
JRES: Purely domestic component of Japanese stock market index.
USRES: Purely domestic component of U.S. stock market index.
DIU: Change in short-term interest rate in the U.S.
DIJ: Change in short-term interest rate in Japan.
FUT: Percentage change in the Dow Jones commodity futures index.

The F-statistic tests the equality of the premia for Japanese and U.S. securities estimated by seemingly unrelated regressions. The SW statistic is the ratio of two quadratic forms and is distributed T^2 (Shanken and Weinstein (1985)). The T^2 statistic tests the equality of premia using the Fama and MacBeth (1973) procedure.

Estimation Period A: 1/5/77–12/31/80	Risk Premia							
	INTERCEPT	*WINDEX*	*JRES*	*USRES*	*DIU*	*DIJ*	*FUT*	R^2
U.S. Securities	0.078	0.315	0.174	−0.084	0.003	0.421	−0.087	0.33
(t-value)	(0.86)	(3.25)	(0.85)	(1.18)	(3.66)	(1.71)	(0.34)	
Japanese Securities	−0.010	0.097	0.328	−0.197	0.0008	−0.018	−0.300	0.24
(t-value)	(0.09)	(0.82)	(2.52)	(3.68)	(0.67)	(0.11)	(1.38)	

	Tests of Equality of Risk Premia[a]		
	F	SW	T^2
Excluding Intercept	2.17	1.01	0.802
	(0.047)	(0.413)	(0.569)
Including Intercept	1.86	0.950	0.686
	(0.076)	(0.467)	(0.683)

Estimation Period B: 1/7/81–12/26/84	Risk Premia							
	INTERCEPT	*WINDEX*	*JRES*	*USRES*	*DIU*	*DIJ*	*FUT*	R^2
U.S. Securities	0.405	−0.176	−0.161	0.112	−0.001	−0.097	−0.207	0.24
(t-value)	(6.40)	(2.60)	(1.12)	(2.24)	(2.46)	(0.56)	(1.15)	
Japanese Securities	0.020	0.015	0.213	−0.0032	−0.0017	0.014	−0.127	0.08
(t-value)	(0.15)	(0.12)	(1.32)	(0.04)	(1.41)	(0.07)	(0.47)	

	Tests of Equality of Risk Premia[a]		
	F	SW	T^2
Excluding Intercept	1.49	1.32	0.390
	(0.183)	(0.243)	(0.885)
Including Intercept	1.83	1.10	0.470
	(0.082)	(0.361)	(0.855)

[a] p-Values in parentheses.

Table VII

Portfolio Betas: Descriptive Statistics

OLS coefficients of risk measures are estimated for each portfolio with weekly data. Portfolios are formed by market model beta rankings. Means and standard deviations of coefficients are computed over twenty-two portfolios in each country. The coefficients correspond to risk measures:

WINDEX: Percentage change in World stock market index.
JRES: Purely domestic component of Japanese stock market index.
USRES: Purely domestic component of U.S. stock market index.
DIU: Change in short-term interest rate in the U.S.
DIJ: Change in short-term interest rate in Japan.
FUT: Percentage change in the Dow Jones commodity futures index.

World stock market index is a value-weighted portfolio of the eight countries' stock indices. The list of the countries and the weights are given in Table III. Individual stock market indices are obtained from the Financial Times, London. Short-term intrest rates are obtained from the Board of Governors of the Federal Reserve, Washington, D.C. The Dow Jones commodity futures index is obtained from the Dow Jones Educational Services, Princeton.

Estimation Period		U.S. Portfolios: Coefficients of Risk Measures					
		WINDEX	*JRES*	*USRES*	*DIU*	*DIJ*	*FUT*
1/5/77–	Mean	1.484	0.072	1.198	−21.612	0.020	−0.014
12/27/78	Std. Dev.	0.623	0.198	0.724	38.000	0.067	0.066
1/7/81–	Mean	0.999	0.115	1.025	−18.868	0.050	−0.037
12/29/82	Std. Dev.	0.285	0.132	0.694	21.168	0.111	0.107

Estimation Period		Japanese Portfolios: Coefficients of Risk Measures					
		WINDEX	*JRES*	*USRES*	*DIU*	*DIJ*	*FUT*
1/5/77–	Mean	0.213	1.077	0.348	7.327	0.0003	−0.036
12/27/78	Std. Dev.	0.102	0.286	0.462	42.189	0.098	0.064
1/7/81–	Mean	0.979	1.107	−0.304	6.097	0.004	0.020
12/29/82	Std. Dev.	0.178	0.230	0.320	18.753	0.135	0.082

equality of the risk premia for the two countries in each of the two subperiods. We estimated weekly cross-sections of the premia with generalized least squares. Formally, for each week of the sample period, the risk premia were estimated by

$$\lambda = (\hat{\boldsymbol{B}}^{*\prime}\hat{\boldsymbol{W}}^{-1}\hat{\boldsymbol{B}}^{*})^{-1}\hat{\boldsymbol{B}}^{*}\hat{\boldsymbol{W}}^{-1}r_t \tag{5}$$

and

$$\hat{\boldsymbol{W}} = \hat{\boldsymbol{B}}\hat{\boldsymbol{B}}' + \hat{\Sigma},$$

where $\boldsymbol{B}^*$ are now the estimated factor loadings augmented with a column of ones, $[\iota\colon \boldsymbol{B}]$, and Σ is the specific variance or idiosyncratic risk associated with individual securities. We thus obtained a time series of premia differentials, $\hat{\Pi}_t = \hat{\lambda}_t^{US} - \hat{\lambda}_t^{JA}$, with covariance matrix Φ. Under the null hypothesis that the mean differential is equal to zero, the statistic

$$T\bar{\Pi}'\Phi^{-1}\bar{\Pi} \sim \chi_k^2 \tag{6}$$

is distributed chi-square with k degrees of freedom, where

$$\bar{\Pi} = \frac{1}{T}\sum_{t=1}^{T}\hat{\Pi}_t \tag{7}$$

Table VIII

Risk Premia Based on Portfolio Returns

The risk premia are estimated cross-sectionally using the coefficients of risk measures, betas, as the independent variables. Betas are estimated from periods 1/5/77–12/27/78 and 1/7/81–12/29/82, respectively. Dependent variables are the mean returns of twenty-two portfolios of U.S. and Japanese stocks, respectively, during the period indicated. Each portfolio contained five stocks. The risk premia estimates correspond to risk measures:

WINDEX: Percentage change in World stock market index.
JRES: Purely domestic component of Japanese stock market index.
USRES: Purely domestic component of U.S. stock market index.
DIU: Change in short-term interest rate in the U.S.
DIJ: Change in short-term interest rate in Japan.
FUT: Percentage change in the Dow Jones commodity futures index.

The F-statistic tests the equality of the premia for Japanese and U.S. portfolios estimated by seemingly unrelated regressions. The SW statistic is the ratio of two quadratic forms and is distributed T^2 (Shanken and Weinstein (1985)). The T^2 statistic tests the equality of premia using the Fama and MacBeth (1973) procedure.

Estimation Period A: 1/3/79–12/31/80	Risk Premia							
	INTERCEPT	*WINDEX*	*JRES*	*USRES*	*DIU*	*DIJ*	*FUT*	R^2
U.S. Portfolios	0.104	0.362	0.198	−0.184	0.0007	−0.118	1.08	0.61
(t-value)	(1.31)	(3.05)	(0.73)	(1.40)	(0.91)	(0.23)	(2.02)	
Japanese Portfolios	0.706	−0.529	−0.460	0.077	0.0004	1.44	−1.17	0.59
(t-value)	(4.84)	(1.46)	(3.27)	(1.25)	(0.61)	(3.84)	(2.39)	

	Tests of Equality of Risk Premia[a]		
	F	SW	T^2
Excluding Intercept	2.42	2.58	1.52
	(0.049)	(0.023)	(0.18)
Including Intercept	3.52	2.90	1.41
	(0.007)	(0.008)	(0.20)

Estimation Period B: 1/5/83–12/26/84	Risk Premia							
	INTERCEPT	*WINDEX*	*JRES*	*USRES*	*DIU*	*DIJ*	*FUT*	R^2
U.S. Securities	0.569	−0.235	0.083	−0.017	0.0015	−0.75	−0.045	0.41
(t-value)	(4.03)	(1.27)	(0.32)	(0.19)	(0.93)	(0.25)	(0.20)	
Japanese Securities	−0.212	0.567	−0.022	0.027	−0.005	−0.771	0.804	0.29
(t-value)	(0.70)	(1.09)	(0.06)	(0.17)	(1.48)	(2.14)	(1.39)	

	Tests of Equality of Risk Premia[a]		
	F	SW	T^2
Excluding Intercept	0.71	1.08	0.76
	(0.642)	(0.380)	(0.60)
Including Intercept	1.02	0.74	0.676
	(0.435)	(0.639)	(0.69)

[a] p-Values in parentheses.

Table IX

Risk Premia Based on Factor Analysis

Factor loadings and the number of factors are jointly estimated by maximum-likelihood factor analysis based on the returns of 221 securities: 110 U.S. securities, 110 Japanese securities, and the dollar return on the Japanese risk-free asset with weekly data for the full period covering 1/5/77–12/26/84. Risk premia are estimated with cross-sectional GLS in each of the subperiods, 1/5/77–12/31/80 and 1/7/81–12/26/84, respectively. The overall χ^2 tests that the vector of risk premia estimated with generalized least squares is significantly different from zero. The χ^2 statistics tests that the vector of differences in the two countries' premia estimated by generalized least squares weekly is equal to zero.[a]

	Period			
	1/5/77–12/31/80		1/7/81–12/26/84	
	US	*JA*	*US*	*JA*
A. Five-Factor Model				
Overall χ^2	21.62	1.85	2.31	10.77
	(0.0006)	(0.869)	(0.805)	(0.056)
	Test of Equality of Risk Premia			
Excluding Intercept	$\chi^2 =$	11.92	$\chi^2 =$	3.43
		(0.035)		(0.559)
Including Intercept	$\chi^2 =$	15.06	$\chi^2 =$	5.32
		(0.019)		(0.503)
B. Ten-Factor Model				
Overall χ^2	29.89	6.60	8.09	12.55
	(0.0008)	(0.762)	(0.620)	(0.249)
	Test of Equality of Risk Premia			
Excluding Intercept	$\chi^2 =$	17.20	$\chi^2 =$	12.97
		(0.070)		(0.225)
Including Intercept	$\chi^2 =$	20.97	$\chi^2 =$	15.08
		(0.034)		(0.178)
C. Twenty-Factor Model				
Overall χ^2	45.15	12.05	16.08	23.95
	(0.001)	(0.914)	(0.711)	(0.245)
	Test of Equality of Risk Premia			
Excluding Intercept	$\chi^2 =$	36.83	$\chi^2 =$	25.58
		(0.012)		(0.180)
Including Intercept	$\chi^2 =$	39.02	$\chi^2 =$	28.97
		(0.009)		(0.114)

p-Values in parentheses.

and

$$\Phi = \frac{1}{T-1} \sum_{t=1}^{T} (\hat{\Pi}_t - \bar{\Pi})(\hat{\Pi}_t - \bar{\Pi})'.$$

The results for the three models are shown in Table IX. The first statistic for each model is an overall chi-sequare test of significance for the vector of estimated risk premia. This is an important piece of information because the null hypothesis of equal premia in the two countries could not be statistically rejected even with perfectly segmented capital markets if no factors were priced. For all three models we find that at least one factor is indeed priced.[16]

The chi-square tests of equality of risk premia are consistent with the findings of the previous section. There are significant differences in risk premia based on five-, ten-, and twenty-factor models during the regime of capital controls in Japan before 1981. We do not find any difference in risk premia for the U.S. and Japanese stocks after 1981. The results of this section support, more strongly than those of the previous section, the proposition that government controls are the source of segmentation. This is because risk premia are estimated with individual stock returns rather than with portfolio returns and because the pricing model performs equally well in the second subperiod, thus raising the power of the test.

VII. Conclusions

General tests of international capital market integration are likely to be inconclusive both because it is difficult to specify a testable capital asset pricing model in an open economy and because it is difficult to distinguish between segmentation due to objective restrictions to trade in financial assets and that arising from individuals' attitudes and irrationality. In this paper, therefore, we decided to focus exclusively on perhaps the most important recent episode of capital market liberalization in an industrialized country, and elimination of capital controls in Japan at the end of 1980, and analyze the dollar price of risk in the Japanese stock market at the time of the liberalization.

If governments, rather than individuals, are the source of segmentation, we should observe a price differential for risk between the Japanese and the U.S. capital markets before, but not after, the liberalization. The empirical evidence presented in the paper supports this view. The data were examined using different model specifications and testing procedures, but in the vast majority of cases we were unable to reject the hypothesis of perfect integration after 1980. Before that date, integration is instead rejected most of the time.

Appendix: Data Sources

Japanese Security Prices: Nomura Research Institute, Tokyo, Data Tape.

U.S. Security Prices: Center for Research on Security Prices, Chicago, Data Tape.

Dow Jones Commodity Futures Index: Dow Jones Educational Services, Princeton, NJ.

Exchange Rates: Board of Governors of the Federal Reserve, Washington, D.C., International Finance Division, Data Tape.

Gensaki Interest Rates: Board of Governors of the Federal Reserve, Washington, D.C., International Finance Division, Data Tape.

[16] The individual risk premia are not reported, however, since they are identified only up to an orthogonal transformation and no economic significance can be attributed to them (Dhrymes, Friend, and Gultekin (1984)).

3-Month Eurodollar and Euroyen Rates: Data Resources Incorporated, Lexington, MA, Macro Data Bank.
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