CAPITAL CONTROLS AND INTERNATIONAL CAPITAL MARKETS SEGMENTATION: THE EVIDENCE FROM THE JAPANESE AND AMERICAN STOCK MARKETS

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ABSTRACT

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

I. Introduction

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 governments impediments to capital movements or because of individual inhibitions. 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. Only a few empirical studies, however, have tried to test this proposition.

One set of studies have adopted an international single index asset pricing model à la Sharpe-Lintner to test 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 stocks 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 clearcut because a single index international capital asset pricing model can only be obtained under the restrictive assumptions of either 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, 1974). 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 various countries inflation rates (Adler and Dumas, 1983 and Stulz, 1985) so that the pricing of purely domestic factors would not necessarily indicate lack of integration.

In addition, a test should ideally discriminate segmentation due to government policies 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 verify 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, 1981, Eun and Janakiramanan, 1986) and, realistically, one cannot hope to capture the characteristics of the various countries' capital controls regimes with just one model specification. Furthermore, both the extent and form of the impediments to capital movements imposed by a country vary through time in response to balance of payments developments. Asset pricing will thus incorporate the anticipation of future governments' actions thus rendering the testability of any model nearly impossible.[3]

A different set of studies have used an international version of the Arbitrage Pricing Theory (APT).[4] Because the pricing in this model is based on an arbitrage condition of nominal returns, APT has the advantage of eluding the problem of PPP deviations. In addition, if this framework is applied to a large set of countries, it may not be used to discern the sources of markets segmentation, namely government policies versus market failures.[5]

Given the current status of the theory and the complexity of the problem, we believe that general test of capital markets integration are either not viable or deemed to be uninformative.[6] Therefore, in this paper, we try a different approach by focusing 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 until then. If the controls were the only source of segmentation, we should be able to reject the hypothesis that the price of risk--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.

In the four years after the liberalization, we cannot find any sign of segmentation between the Japanese and US 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. The data support the view that governments, rather than individuals' behavior, represents the only source of international capital markets segmentation.

Our paper is organized in the following order: Section II reviews the Japanese experience with capital controls; Section III describes the model and tests the hypothesis and the data; Sections IV and V produce empirical tests with prespecified risk factors for the case of individual securities and portfolios; Section VI presents tests when risk factors are estimated with factor analysis; and Section VII concludes the paper.

II. 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 after that date although interest rates in the money markets were completely deregulated only in 1978-1979.[7] 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 less than five years and one month, the marginal reserve requirement on Yen accounts by nonresidents was also increased in steps to 100%; while Japanese institutions were for the first time allowed to purchase foreign securities. After a brief relaxation of the controls in 1979, Japanese authorities resorted to their use in 1980, this time to support a falling Yen: the reserve requirements on nonresidents deposits was rolled back to zero, Japanese banks were encouraged to raise funds in London and to transfer them to Tokyo and, at the same time, 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 new 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 are clearly visualized in Figure 1, depicting the interest rate differential between 3-month euroyen deposits traded in London and 3-month repurchase agreements (Gensaki) in Tokyo--a market that has been least affected by administrative controls. Before 1981, the interest rate differential is substantial--it was negative until 1979 when capital were prevented from flowing into Japan and positive in 1980 when the opposite occurred--but has been practically zero after then, an indication that any government induced segmentation of the market had ceased its effects on market prices.

Given the depth of the controls and the extent of the subsequent liberalization program, the Japanese experience constitute a unique semi-controlled experiment for testing international capital markets segmentation. If government policies, as opposed to private attitudes, were the only source of segmentation, the price of risk in the US 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.

III. Testing Capital Controls with APT.

In order to test the equality between the price of risk in the two markets we need a model that defines what is systematic risk. We choose an international version of APT for two reasons: first, from the introductory remarks it is clear that more than one factor affect security prices in an international setting, but it is difficult to derive the exact form of the pricing equation; second, APT can be based on an arbitrage relation of nominal rates of returns so that one needs not to worry about deviations from PPP. It is worthwhile stressing, however, that this paper is by no means a test of APT; rather the APT 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 US and Japan, which are indexed by i. We assume that security 0 is the risk free rate in US 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 US dollar is used as numeraire, APT postulates that the return generating function for the N+1 risky securities is linear

$$r_t = E_t + \beta\delta_t + e_t \tag{1}$$

where r is the (N+1) $\times$ 1 vector of nominal rates of return (in dollar) at time t, E is the vector of expected return, δ is the k $\times$ 1 vector of risky factors with mean zero and variance-covariance matrix Ω, β is the (N+1) $\times$ k matrix of the sensitivity coefficients to the risky factors and e is the vector of idiosyncratic terms with variance-covariance Σ.

It is well-known from APT that there is an arbitrage portfolio constraining expected returns (in dollars) to be linear in β in equilibrium. Solnik (1983) has shown that the same arbitrage portfolio will hold for the Japanese investor who calculates rates of returns in Yen, providing the $/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

$$r_t^c = \Gamma_{0t}^c + \beta^c \Gamma_t^c + v_t^c \qquad c = US,\ JA \tag{2}$$

where Γ is the vector of risk premia associated with the risky factors, Γ_0 is the risk free rate, v is the vector of error terms and the superscript 'c' indicates whether it is a Japanese or an American stock.

Based on the specification in (2) our empirical tests of market segmentation are based upon a well-known two-stage estimation approach. In the first stage the elements of β are estimated for each security by using the time series of returns.[8] In the second stage $\hat{\beta}$ are used as independent variables, to estimate Γ_0 and vector Γ for each country, for the sample period 1977-1980 during which the capital controls were in effect, and for the subsequent four years, 1981-1984, that were characterized by free capital movements. Then, the composite null hypothesis that:

$$\hat{\Gamma}^{US} \neq \hat{\Gamma}^{JA} \qquad \hat{\Gamma}_0^{US} \neq \hat{\Gamma}_0^{JA} \qquad \text{in 1977-1980}$$

and

$$\hat{\Gamma}^{US} = \hat{\Gamma}^{JA} \qquad \hat{\Gamma}_0^{US} = \hat{\Gamma}_0^{JA} \qquad \text{in 1981-1984}$$

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--Hotelling T^2, the test proposed by Shanken and Weinstein (1985) (SW henceforth) and the classical F test of linear restrictions. 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 respect to the alternative testing procedures and inevitable model specification errors. The approach that we follow is thus similar in the spirit to an event study; as such, the results lacks generality but, we hope, they are more conclusive than previous empirical research in the field.[9]

Before discussing the various implementations of the test, we turn to a description of the data set used in the paper. We used weekly stock returns calculated from daily prices quoted at the closing of the markets on Wednesdays from January 1, 1977 to December 31, 1984. As indicated before, we focus on the two sample periods, January 1977 to December 1980 and January 1981 to December 1984. We do not consider data before 1977 because at that time Japan lacked truly free market interest rates. We do not consider data beyond 1984 because we want to match the length of the two sub-periods, the one with and the one without capital controls.

We selected two sets of Japanese and US securities, each consisting of 110 stocks, which reproduced the industry composition of the two countries market portfolios as closely as data availability allowed us. The choice of an equal number of stocks and the attention paid to their industrial classification were dictated by the concern of introducing spurious risky factors in either of the securities sets. In Table 1 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 and the largest 1,000 corporations in the US as ranked by Business Week.[10]

We converted Japanese stock returns into dollars. We were not able, however, to match exactly the timing of stock transactions with that of the exchange transactions. 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 needs Yen on Saturday, a working day in Japan, while he will receive Yen the next Saturday. Two business days are needed for settlement in foreign exchange markets; the American investor must then sell dollars on Thursday, Tokyo time, and sell Yen one week later. The time series for the $/Yen exchange rate that we have, however, are 1 p.m. New York quotes. There are two possibilities then: we either use New York Wednesday data, and we pick an exchange rate that is quoted five hours before the opening of the Tokyo exchange markets on Thursday, or we use New York Thursday data, and we get a rate quoted half-day after the closing of the market in Tokyo. For convenience we choose the second option.[11] The sources of all the data are given in the Appendix.

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). 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, 1974 and Stehle, 1977).

With weekly data, the choice of economic variables entering the stock return function is very limited. We selected six variables. We take the change in the short-term interest rate in both Japan--3-month Gensaki rate (DIJ)--and US--3-month eurodollar deposit rates (DIU). These variables are meant to capture movements in the term structure of interest rates, in the spirit of recent single-state-variable theoretical model of the equilibrium structure. The third variable was the percentage change in the Dow Jones Commodity Futures Index (FUT)--an index of twelve commodities including foods, metals and woods--as a proxy for world demand pressures and inflationary expectations.

In principle, stock market indexes should not enter the regressions if all relevant economic variables had been 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 index (WINDEX) which was calculated by weighing the stock markets of France, Germany, Switzerland, United Kingdom, Japan, Australia, Canada and the US--or 93.4% of total world stock markets, Table 2--with their capitalization at the end of 1980, the mid-point of our sample period.

Finally, we use two purely domestic market factors in the regressions, namely the part of the Japanese (JRES) and US (USRES) stock indexes that were orthogonal to the world stock indexes. The last three variables is what has been typically used in many studies of international capital markets 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. This secondary effect should be already captured by the two domestic factors JRES and USRES.

The model specification imposes that the variables, δ, measures of risk 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 the only exception of Japanese interest rates which moved in a stepwise fashion throughout most of the sample period--and the estimated means are comfortably close to zero, with the exception here of the world stock index (Table 3). 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. Our approach seems to conform to current practices in the empirical literature on APT with prespecified factors (Chan, Chen, and Hsieh, 1985 and Chen, Roll, and Ross, 1986).

The initial step is the estimation of the β matrix. In this section we use the time series of individual stock returns from the entire sample period, 1977 to 1984. We assume thereby that the risk characteristics of the securities remained unchanged throughout the liberalization process even though the price of the risk factors changed between the two sub-periods. There are two problems with this approach: the consistency of the risk premia estimates--because both β and Γ are estimated from the same sample period--and the stationarity of β. Both problems will be dealt 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.

The mean estimated betas from individual stock returns are reported in Table 4. For the entire sample period all the risk factors had a strong impact on US stock returns with the only exception of Japanese interest rates. For the Japanese stocks, instead, only Japanese variables seems to matter, in addition to the world stock market. The table also shows the betas estimated for the two sub-periods, 1977 to 1980 and 1981 to 1984. One can notice the remarkable stability of the betas for the US stocks over this interval, even though the t statistics suggests 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 βs at any significant level. In particular commodity futures and Japanese interest rates appear to have exerted different effects in the two sub-periods. We checked whether the liberalization of capital movements could account for the changes in the Japanese betas, in addition to the risk premia, by using the differential between domestic Yen and euroyen deposits of the same maturity as proxy variable (CAPC). This variable is plotted in Figure 1. Perhaps the proxy is inappropriate, but we do not find any correlation between betas and capital controls. The lower part of Table 3 shows that CAPC is only slightly correlated with the other sources of risk, suggesting no omitted variable problems; when we re-estimated the betas with the capital control proxy added at the right hand side, its coefficient was never significant.[12]

The next step is to test the equality of the vector of risk premia for the US and Japanese securities after they have been estimated by regressing cross sections of stock returns on the estimated betas. We use three testing procedures. First, we use the stocks mean rates of return in each of the sub-periods as the dependent variables in the cross sections and we estimate the vector of US and Japanese premia jointly with a seemingly unrelated estimator. The equality of the premia is then tested with a standard F test of linear constraint (Theil, 1971, p.313).

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

$$SW = \frac{T(\hat{\Gamma}^{US} - \hat{\Gamma}^{JA})'(H\hat{\Sigma}^{-1}H')^{-1}(\hat{\Gamma}^{US} - \hat{\Gamma}^{JA})}{1 + \hat{\Gamma}'\Omega^{-1}\hat{\Gamma}} = T^2(k + 1, T - k - 1)$$

where T is the number of periods used to estimate the betas in the first step, Γ is the vector of premia obtained by regression with equality restriction imposed, and H is defined as

$$H = (A^{US} : -A^{JA})$$

$$A^c = (\hat{\beta}^{c'}\hat{\beta}^c)^{-1}\hat{\beta}^{c'} \quad .$$

Finally, we adopt the Fama and MacBeth (1973) procedure and we estimate the risk premia for both countries in each week of the two sub-periods, or 208 weeks for each sub-period. 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 is then used to test the vector of mean differentials.

The results are presented in Table 5. More sources of risk seems to be priced before the liberalization of capital controls than after it; this is particularly true for the Japanese securities. Looking at the tests of the equality of premia, thus excluding the intercept, the SW and T^2 statistics indicate that perfect capital market integration is supported by the data in both sub-periods, even though the absolute value of the statistics declines substantially after 1980. The F statistics suggests instead that the hypothesis of segmentation cannot be rejected at the .05 significance level in the 1977-1980 period, but not 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 does 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 US 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 test by grouping securities into portfolios, which is the usual procedure to minimize these two sources of bias (Black, Jensen and Scholes, 1972).

V. Prespecified Factors with Portfolios of Securities.

Given the small number of securities in each country sample, precisely 110, we can only form 22 portfolios of 5 securities. In each sub-period, 1977-1980 and 1981-1984, we took the first two years, 1977-1978 and 1981-1982, and estimated the securities beta with respect to their domestic market index. This beta was used to rank securities and form the portfolios. We then estimated the portfolios betas with respect to all the six sources of risk specified in the previous section using again the returns from the first two years only. The portfolios betas estimated in this way were used as independent variables to obtain estimates of the risk premia in the second half of the two sub-periods, 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] We could not use firm size--a more popular choice in the literature--because both the US and Japanese samples comprised the largest companies in the two countries, which lacked sufficient dispersion in the data.

We are also aware of the bias that we may introduce by estimating the individual securities betas, with which we form the portfolios, and the portfolio betas from the same sample period. We did not have much of a choice, however, because our entire sample period consists of only four years and two of them were needed for estimating the risk premia with a sufficiently long series. Finally, we formed portfolios by grouping securities according to their gross returns but the results did not change in any significant way.

Table 6 reports the estimated portfolio betas for the two sub-periods. As for individual securities, US betas seem more stable than Japanese betas. Surprisingly, changes in Japanese interest rates are correlated with US portfolio returns but not with Japanese returns.

In Table 7 we show the same three test statistics that we introduced in the previous section and the estimated portfolio premia. The results are clearcut. With or without the intercept term included, the hypothesis of integration is rejected at the .05 significance level by both SW and F statistics for the 1979-1980 period. By contrast the same hypothesis cannot be rejected at any significance level for the years 1983-1984. Only the T^2 statistics, with a significance level of .18, does not provide evidence of segmentation before the liberalization process. The significance of the results is somewhat reduced, however, by the lack of precision with which the risk premia are estimated in the second sub-period.

VI. Factor Analytic Approach

Factor analysis is an alternative methodology where one determines the number of factors (k) and estimates the elements of β jointly. We extracted the factor loadings from the full variance-covariance matrix of US 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 fix 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 exceeds thirty. We decided to estimate models with five, ten and twenty factors.

Once we had obtained the securities factor loadings, we estimated and tested the equality of the risk premia in two different ways. First, we estimated the premia for the US and Japanese securities by using seemingly unrelated regressions with cross sections of mean stock returns in the two subperiods. The equality of the premia was then tested with the usual F test of linear constraints.

Second, we estimated weekly cross sections of the premia with generalized least square. In other words, for each week of the sample period the premia were estimated from

$$\hat{\Gamma}_t = (\hat{\beta}\hat{W}^{-1}\hat{\beta}')^{-1}\hat{\beta}\hat{W}^{-1}r_t'$$

and

$$\hat{W} = \hat{\beta}'\hat{\beta} + \hat{\Sigma}$$

where $\hat{\beta}$ are now the estimated factor loadings and the definition of W reflects the constraint of $\Omega = I$ which is imposed for the estimation of the factors. We thus obtained a time series of premia differentials, $\hat{\pi}_t = \hat{\Gamma}_t^{US} - \hat{\Gamma}_t^{JA}$, with variance-covariance matrix D. Under the null hypothesis that the mean differential is equal to zero, the statistics

$$T\bar{\pi}D^{-1}\bar{\pi}' \sim \chi_k^2$$

is distributed chi square with k degrees of freedom, where $\bar{\pi} = \frac{1}{T}\sum_{t=1}^{T}\hat{\pi}_t$.

The results for the three models are shown in Table 8. The first statistics for each model is an overall chi square test testing whether the vector of premia is significantly different from zero. 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 perfect capital markets segmentation if no factors were priced. For all three models we find that at least one factor is indeed priced. 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).

The chi square test strongly supports the findings of the previous section, namely, significant differentials in premia emerge before 1981 but they disappear after that year. The same conclusion can be reached by looking at the F test for the 5 and 10 factor models with the intercept term excluded. When the intercept is included, integration is instead rejected by the F test at the .05 significance level even for the years after 1980. The F test for the 20 factor model is the only instance in the paper in which the null hypothesis of equal price of risk in the Japanese and US markets is rejected for the post-liberalization period. A reasonable explanation is that the number of "firm specific" factors grows rapidly with the number of factor extracted. If that is the case, the F test should not be interpreted as evidence of segmentation in recent years.

VII. Conclusions

General tests of international capital markets integration are likely to be inconclusive both because it is difficult to specify a testable capital asset pricing model in the open economy and because it is difficult to distinguish between segmentation due to objective restrictions to trade in financial assets from that arising because of individuals ' inhibitions 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, 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 only source of segmentation, we should observe a price differential for risk between the Japanese and the US capital markets before the liberalization but not after it. On the whole the empirical evidence presented in the paper supports this view. The data was 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, N.J.

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

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

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

Countries' Stock Market Indexes: Financial Times, London.

FOOTNOTES

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

[2]This approach is used by Stehle, 1975 and Jorion and Schwartz, 1986. A similar framework is used in the empirical studies of Solnik, 1974 and Errunza and Losq, 1985.

[3]The recent experiences of France, Japan and Italy, among the major industrialized countries, is very illuminating on this point. The important relationship between asset pricing and government policies in an international context has been stressed in Stockman and Hernandez (1985) and Stockman and Dallas (1986).

[4]Cho, Eun and Senbet (1986) and Korajczyk and Viallet (1986) use an APT framework to test international capital markets integration. See also Cosset (1984) for an empirical study using an international APT.

[5]Still another approach is to test the equality of the marginal rate of substitution in consumption in different countries if agents have access to the same risk-free interest rate. This approach is adopted by Obstfeld (1986) who cannot reject the null hypothesis of integration. It would be difficult to extend this approach, however, to the pricing of risky assets. Some of the problems with consumption based asset pricing models are discussed in Cornell (1981).

[6]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.

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

[8]In the case of the factor analytic approach the number of factors (k) and the elements of β are jointly estimated in the first stage.

[9]An "event study" approach to capital market segmentation is also used by Gordon, Eun and Janakiramanan (1986) by looking at the dual listing of companies in the Canadian and US markets.

[10]Because the original data on Japanese stocks were classified into 20 industry categories, we found convenient to regroup the US companies according to that classification. There was no Japanese equivalent, however, for two of the US categories, Conglomerates and Tobacco, Paper and Forestry.

[11]There is also the complication of non synchronous trading given that trading hours in the Tokyo and New York markets never overlap. We believe, however, that none of the results depend in any significant way on these data problems.

[12]These regressions are not reported in the paper but are available upon request.

[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.

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

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Figure 1. DIFFERENTIAL BETWEEN 3-MONTH EUROYEN IN LONDON AND 3-MONTH GENSAKI IN TOKYO

(Daily Data: 1/1/1977 - 12/30/1984)

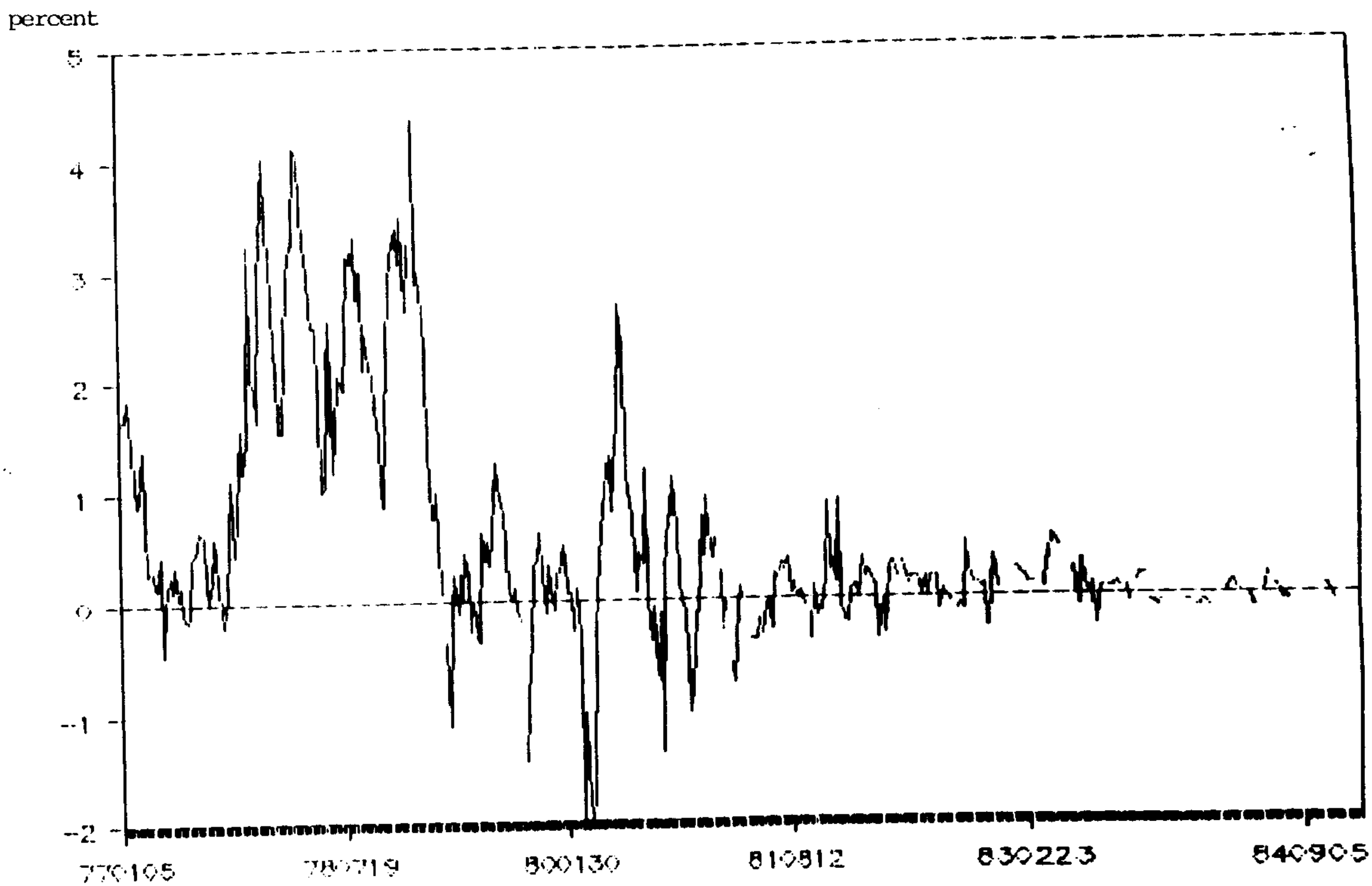

Table 1. Industry Composition of Sample Portfolio and Market Portfolio

	Number of Companies in the Sample as Percentage of Market Portfolio	
Sectors	Japan[1]	USA[2]
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 Mfg.	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

[1]The industry composition of the market portfolio is obtained from Toyo Keizai Shinposha (1985).

[2]The industry composition of the market portfolio is obtained from Business Week (1986).

Table 2. Equity Markets Capitalization[1]
(US dollars; end of 1980)

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	-

[1]Source: Ibbotson, Carr and Robinson (1982).

Table 3. Risk Measures: Descriptive Statistics

(1977 - 1984)

	Mean (%)	t	ρ	σ(ρ)
WINDEX	.19805	2.15	.048	.05
JRES	.00000	0.00	.032	.05
USRES	.00000	0.00	.025	.05
DIU	.00017	0.31	.076	.05
DIJ	.00001	0.00	-.502	.05
FUT	.03195	0.31	.127	.05

CORRELATION MATRIX

(1977 - 1984)

	WINDEX	JRES	USRES	DIU	DIJ	FUT
WINDEX	1.00	.00	.00	-.40	-.15	.37
JRES		1.00	-.78	-.08	-.49	.03
USRES			1.00	-.17	.46	-.17
DIU				1.00	.13	-.18
DIJ					1.00	-.103
FUT						1.00

CORRELATION WITH CAPC

	WINDEX	JRES	USRES	DIU	DIJ	FUT
1977-1980	.16	.11	-.13	-.14	-.09	.14
1981-1984	.06	.002	-.01	-.13	-.03	-.04

Table 4. Security Betas: Descriptive Statistics

Estimation Period	Estimated Betas of Risky Factors						
	WINDEX	JRES	USRES	DIU	DIJ	FUT	
US SECURITIES							
1977-1984							
Mean	1.08	0.10	1.18	-21.7	0.0003	-0.05	
(t value)	28.5	5.2	18.2	6.0	.03	5.0	
1977-1980							
Mean	1.18	0.10	1.19	-21.6	0.005	-0.05	
(t value)	25.0	3.9	16.6	5.8	0.37	4.7	
1981-1984							
Mean	1.02	0.09	1.17	-21.3	-0.004	-0.05	
(t value)	27.5	3.65	14.27	4.78	0.21	2.89	
JAPANESE SECURITIES							
1977-1984							
Mean	0.73	1.03	0.01	-0.37	-0.027	-0.02	
(t value)	29.5	42.9	0.2	0.27	2.6	1.93	
1977-1980							
Mean	0.31	1.05	0.09	-0.66	0.02	-0.04	
(t value)	13.60	35.33	1.22	0.32	1.8	3.60	
1981-1984							
Mean	0.92	1.01	0.01	0.43	-0.07	0.02	
(t value)	29.61	32.72	0.23	0.16	4.31	1.68	
	Equality of Betas between Sub-periods						Hotelling[1] T^2
	t values						
US SECURITIES	4.78	0.31	0.36	0.09	0.40	0.004	4.52 (.0005)
JAPANESE SECURITIES	19.3	1.37	0.78	0.28	4.53	3.45	107.0 (.000)

[1]p values in parenthesis.

Table 5. Risk Premia Based on Individual Stock Returns[1]

Risk Premia (t value)	1977-1980		1981-1984	
	US	JA	US	JA
INTERCEPT	.078 (.86)	-.010 (.09)	.405 (6.40)	.020 (.15)
WINDEX	.315 (3.25)	.097 (.82)	-.176 (2.60)	.015 (.12)
JRES	.174 (.85)	.328 (2.52)	-.161 (1.12)	.213 (1.32)
USRES	-.084 (1.18)	-.197 (3.68)	.112 (2.24)	-.0032 (.04)
DIU	.003 (3.66)	.0008 (.67)	-.001 (2.46)	-.0017 (1.41)
DIJ	.421 (1.71)	-.018 (.11)	-.097 (.56)	.014 (.07)
FUT	-.087 (.34)	-.300 (1.38)	-.207 (1.15)	-.127 (.47)
R^2	.33	.24	.24	.08

TEST OF EQUALITY OF RISK PREMIA

		1977-1980		1981-1984
Excluding Intercept	F = (p value)	2.17 (.047)	F = (p value)	1.49 (.183)
	SW = (p value)	1.01 .586	SW = (p value)	1.32 .756
	T^2 = (p value)	.802 (.569)	T^2 = (p value)	.390 (.885)
Including Intercept	F = (p value)	1.86 (.076)	F = (p value)	1.83 (.082)
	SW = (p value)	.950 (.529)	SW = (p value)	1.10 (.639)
	T^2 = (p value)	.686 (.683)	T^2 = (p value)	.470 (.855)

[1]Based on mean returns of 110 US securities and 110 Japanese securities. The F statistics tests the equality of the premia for Japanese and US securities estimated by seemingly unrelated regressions. The SW statistics is described in the text and is distributed F. The T^2 statistics is the Hotelling multivariate T^2.

Table 6. Portfolio Betas: Descriptive Statistics

Estimation Period	Estimated Betas of Risky Factors					
	WINDEX	JRES	USRES	DIU	DIJ	FUT
US PORTFOLIOS						
1977-1978						
Mean	1.48	0.07	1.19	-21.6	0.02	-0.01
(t value)	(11.16)	(1.71)	(7.76)	(2.67)	(1.42)	(1.00)
1981-1982						
Mean	0.99	0.11	1.02	-18.9	0.05	-0.04
(t value)	(16.42)	(4.04)	(6.92)	(4.10)	(2.12)	(1.64)
JAPANESE PORTFOLIOS						
1977-1978						
Mean	0.21	1.08	0.35	7.33	0.003	-0.04
(t value)	(9.73)	(17.7)	(3.53)	(0.81)	(0.01)	(2.64)
1981-1982						
Mean	0.97	1.11	-0.30	6.09	0.004	0.02
(t value)	(25.74)	(22.53)	(4.45)	(1.53)	(0.13)	(1.17)

Table 7. Risk Premia Based on Portfolio Returns[1]

Estimation Period	Risk Premia							
	INTERCEPT	WINDEX	JRES	USRES	DIU	DIJ	FUT	R^2
A. 1979-1980								
US Securities								
(t value)	.104	.362	.198	-.184	.0007	-.118	1.08	.61
	(1.31)	(3.05)	(0.73)	(1.40)	(0.91)	(0.23)	(2.02)	
Japanese Securities								
(t value)	.706	-.529	-.460	.077	.0004	1.44	-1.17	.59
	(4.84)	(1.46)	(3.27)	(1.25)	(0.61)	(3.84)	(2.39)	

TEST OF EQUALITY OF RISK PREMIA[2]

Excluding Intercept		Including Intercept	
F =	2.42	F =	3.52
(p value)	(0.49)	(p value)	(.007)
SW =	2.58	SW =	2.90
(p value)	(.023)	(p value)	(.008)
T^2 =	1.52	T^2 =	1.41
(p value)	(.18)	(p value)	(.20)

Estimation Period	INTERCEPT	WINDEX	JRES	USRES	DIU	DIJ	FUT	R^2
B. 1983-1984								
US Securities								
(t value)	.569	-.235	.083	-.017	.0015	-0.75	-.045	.41
	(4.03)	(1.27)	(0.32)	(0.19)	(0.93)	(0.25)	(0.20)	
Japanese Securities								
(t value)	-.212	.567	-.022	.027	-.005	-.771	.804	.29
	(.70)	(1.09)	(.06)	(.17)	(1.48)	(2.14)	(1.39)	

TEST OF EQUALITY OF RISK PREMIA[2]

Excluding Intercept		Including Intercept	
F =	0.71	F =	1.02
(p value)	(.642)	(p value)	(.435)
SW =	1.08	SW =	0.74
(p value)	(.380)	(p value)	(.639)
T^2 =	.76	T^2 =	.676
(p value)	(.60)	(p value)	(.69)

[1]The premia were obtained by regressing mean portfolio returns during the period indicated on the betas estimated from 1977-78 and 1981-82 respectively. The sample consisted of 22 portfolios of US and Japanese stocks respectively. Each portfolio contained 5 stocks.

[2]The F tests the equality of the premia for Japanese and US securities estimated by seemingly unrelated regression. The SW statistics is described in the text and is distributed F. The T^2 statistics is the Hotelling multivariate T^2.

Table 8. Risk Premia Based on Factor Analysis[1]

	1977-1980 US	1977-1980 JA	1981-1984 US	1981-1984 JA
A. FIVE FACTOR MODEL				
Overall χ^2	21.62	1.85	2.31	10.77
(p value)	(.0006)	(.869)	(.805)	(.056)
		TEST OF EQUALITY OF RISK PREMIA		
Excluding Intercept	F =	4.32	F =	1.45
	(p value)	(.001)	(p value)	(.208)
	χ^2 =	11.92	χ^2 =	3.43
	(p value)	(.035)	(p value)	(.559)
Including Intercept	F =	3.70	F =	1.26
	(p value)	(.001)	(p value)	(.277)
	χ^2 =	15.06	χ^2 =	5.32
	(p value)	(.019)	(p value)	(.503)
B. TEN FACTOR MODEL				
Overall χ^2	29.89	6.60	8.09	12.55
(p value)	(.0008)	(.762)	(.620)	(.249)
		TEST OF EQUALITY OF RISK PREMIA		
Excluding Intercept	F =	3.65	F =	1.67
	(p value)	(.0002)	(p value)	(.089)
	χ^2 =	17.20	χ^2 =	12.97
	(p value)	(.070)	(p value)	(.225)
Including Intercept	F =	3.73	F =	2.20
	(p value)	(.0001)	(p value)	(.015)
	χ^2 =	20.97	χ^2 =	15.08
	(p value)	(.034)	(p value)	(.178)
C. TWENTY FACTOR MODEL				
Overall χ^2	45.15	12.05	16.08	23.95
(p value)	(.001)	(.914)	(.711)	(.245)
		TEST OF EQUALITY OF RISK PREMIA		
Excluding Intercept	F =	1.56	F =	2.01
	(p value)	(.065)	(p value)	(.008)
	χ^2 =	36.83	χ^2 =	25.58
	(p value)	(.012)	(p value)	(.180)
Including Intercept	F =	1.64	F =	1.99
	(p value)	(.045)	(p value)	(.008)
	χ^2 =	39.02	χ^2 =	28.97
	(p value)	(.009)	(p value)	(.114)

[1] Factor analysis based on returns of 110 US securities, 110 Japanese securities and the dollar return on the Japanese risk free asset. The overall χ^2 tests that the vector of risk premia estimated with generalized least square is significantly different from zero. The F statistics tests the equality of the premia for Japanese and US securities estimated by seemingly unrelated. The χ^2 statistics tests that the vector of differences in the two countries premia estimated by generalized least square weekly is equal to zero.