"AN EMPIRICAL INVESTIGATION OF INTERNATIONAL ASSET PRICING"

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An Empirical Investigation of International Asset Pricing

Abstract

We investigate several asset pricing models in an international setting. We use data on a large number of assets traded in the United States, Japan, the United Kingdom, and France. The models together with the hypothesis of capital market integration imply testable restrictions on multivariate regressions relating asset returns to various benchmark portfolios. We find that multifactor models tend to outperform single-index models in both domestic and international forms especially in terms of explaining seasonality in asset returns. We also find that the behavior of the models is affected by changes in the regulatory environment in international markets. The evidence regarding the performance of models assuming segmented versus integrated markets is mixed.

In this paper we present evidence on the pricing performance of alternative domestic and international asset pricing models. The models are compared when pricing assets within national economies and, in their international versions, when pricing assets across economies. The pricing models together with the hypothesis of capital market integration imply testable restrictions on multivariate regression models relating asset returns to various benchmark portfolios. Conditional on capital market integration, the tests provide information on the validity of the model. Conversely, given that the assumed type of pricing model is correct, the tests provide information about integration across markets. We compare domestic and international versions of the Capital Asset Pricing Model (CAPM) and Arbitrage Pricing Theory (APT) based multifactor models in which the pervasive factors are estimated by an asymptotic principal components technique.

We focus on three questions. First, we investigate whether the factor models have explanatory power over the CAPM in a domestic as well as in an international setting. Secondly, we ask whether international versions of the asset pricing models outperform or underperform single-economy versions. Finally, we look for the influence of changes in the regulation of international financial markets on the deviations of returns from the predicted asset pricing relations.

Our study covers the period 1969-1983. We utilize a large number of securities from four countries both for factor estimation and hypothesis testing. The countries are the United States, Japan, the United Kingdom, and France. The number of assets available for each time period and other information about the data are contained in Table 1.

Asset pricing theories are commonly tested in a closed economy setting in which assets are priced relative to benchmark portfolios constructed from

assets trading in the same economy. Fama and MacBeth (1973) and Roll and Ross (1980) are well known examples of single economy tests of the CAPM and APT, respectively. A variety of asset pricing anomalies have been uncovered by single-economy studies. In particular, seasonal, firm size, and dividend yield related mispricing have been documented.[1] Single-economy applications of the APT have had some success in explaining pricing anomalies.[2]

Two related studies are Cho, Eun, and Senbet (1986) and Gultekin, Gultekin, and Penati (1987). Cho Eun and Senbet (1986) reject an international version of the APT. Using a two-country version of the APT, Gultekin, Gultekin, and Penati (1987) find that the performance of the model is affected by changes in capital controls. We find that rejection of the international APT is sensitive to inclusion of sample periods with strict capital controls. Our study covers more countries than Gultekin et al (1987) but fewer than Cho et al (1986). However, for the countries studied here, we utilize many more securities.[3] The large number of cross-sections allows more precise estimation of the factors. Also, the above studies do not address the issue of comparative performance across models (e.g., CAPM versus APT or international versus domestic).

The next section of the paper contains a brief description of the alternative asset pricing models. In Section II we describe the data. The techniques used to estimate the pervasive factors and test the alternative models are described in Section III while the empirical results are given in Section IV. A summary and conclusions are given in section V.

I. Alternative Asset Pricing Models

We investigate the pricing performance of domestic and international versions of the CAPM and APT. The economic content of the CAPM [pointed out

by Roll (1977)] or the APT [Huberman and Kandel (1987)] is that a particular benchmark portfolio or linear combination of a group of benchmark portfolios lies on the minimum variance boundary of risky assets. The domestic and international versions of the models differ in that only securities traded on the local exchange are included in the benchmark portfolios for the former model while the benchmarks for the international versions include all the assets in the sample. Since the basic models are rather well known we will merely state the implications of the models and concentrate our discussion on the problems associated with implementing the models.

The standard version of the CAPM postulates that the market portfolio is on the mean/variance efficient frontier which, in turn, implies that the expected return on each asset is linearly related to its beta [$\beta_{iM} = \mathrm{cov}(\tilde{r}_i, \tilde{r}_M)/\mathrm{var}(\tilde{r}_M)$]. Assuming the existence of a real riskless asset with return r_F, we have:

$$E(\tilde{R}_i) = E(\tilde{r}_i) - r_F = \beta_{iM}[E(\tilde{r}_M) - r_F] = \beta_{iM}E(\tilde{R}_M) \tag{1}$$

where $\tilde{r}_i$ and $\tilde{r}_M$ are the real returns on asset i and the market portfolio ($\tilde{R}_i$ and $\tilde{R}_M$ are returns in excess of the riskless return, r_F). In a closed economy setting the market portfolio, M, is the portfolio of all domestic assets weighted by their respective proportionate values. Extending (1) to an international setting generally involves more than replacing the domestic market portfolio with an international market portfolio. Exchange rate uncertainty and, particularly, potential deviations from strict purchasing power parity can lead to incremental hedging demands for assets (hedging against shifts in the consumption-investment opportunity set is not peculiar to international models). Under admittedly restrictive conditions there is no

excess demand for hedging exchange risks and we can proceed with a relation like (1).[4] Note that in both domestic and international applications one is never able to obtain the true market portfolio relevant for the particular model. Thus, tests of the pricing relation (1) for different proxies, M, are tests of mean/variance efficiency for these proxies.

As in many empirical investigations, we use a proxy for the riskless rate of interest. We use the return on short-term US treasury bills. Since these returns are not strictly riskless in real terms we also test the restrictions implied by the Black (1972) zero-beta CAPM, assuming that the difference between the expected returns on the zero-beta portfolio and treasury bills is a constant, λ. It follows that expected returns, in excess of the T-bill return, are determined by

$$E(\tilde{R}_i) = (1-\beta_{iM})\cdot\lambda + \beta_{iM}E(\tilde{R}_M). \tag{2}$$

A value of λ equal to zero is consistent with the pricing relation (1).

An assumption underlying the APT is that asset returns follow a factor model:

$$\tilde{r}_i = \mu_i + b_{i1}\tilde{f}_1 + b_{i2}\tilde{f}_2 + \ldots + b_{ik}\tilde{f}_k + \tilde{\epsilon}_i$$

where b_{ij} is the sensitivity (beta) of asset i relative to factor j and $E(\tilde{f}_j) = E(\tilde{\epsilon}_i) = E(\tilde{f}_j\tilde{\epsilon}_i) = 0$ for all i and j. The number of assets in the economy is assumed to be sufficiently large and the correlation across the idiosyncratic returns (ϵ_i's) is assumed to be sufficiently small that the idiosyncratic risk can be diversified away in large portfolios.[5] Lack of arbitrage opportunities and existence of a riskless asset imply that

$$E(\bar{R}_i) \approx b_{i1}\gamma_1 + b_{i2}\gamma_2 + \ldots + b_{ik}\gamma_k. \qquad (3)$$

Additional equilibrium conditions [as in Connor (1984)] can lead to the pricing relation (3) holding as an equality rather than an approximation. Our empirical work below tests (3) as an equality. Ross and Walsh (1983) and Solnik (1983) extend the APT to an international setting. With the assumption that exchange rates follow the same factor model as asset returns, they find that the standard APT pricing relation (3) can be applied directly in an international setting. Thus, exchange rate uncertainty is priced to the extent that it represents pervasive factor risk. We also estimate a zero-beta version of (3) which we discuss in more detail below.

Table 2 presents the particular models we investigate. Two of them, the CAPM-EW and the CAPM-VW, are models where the benchmark portfolios are equal-weighted and value-weighted portfolios of common stock, respectively. The last two, the APT-5 and APT-10 factor models, use statistically estimated factors. Each of the four models (and their zero-beta alternatives) are tested in three versions. In the first version we test the mean/variance efficiency of domestic benchmark portfolios relative to domestic assets. In the second and third versions we test the mean/variance efficiency of international benchmark portfolios relative to both domestic assets (for each economy separately) and relative to an international set of assets.

II. Data Sources

The selected countries, markets, and data sources are presented in Table 1. They are the result of a compromise between the data requirement of the asymptotic principal component analysis - a large number of securities in each country - and the availability of reliable data.

We were able to obtain monthly stock returns data for four countries spanning fifteen years from January 1969 through December 1983. Our sample includes three major markets: the New York and American Stock Exchanges, the Tokyo Stock Exchange, and the London Stock Exchange. For these three countries our sample includes all assets traded on the exchanges. The Paris Bourse was added in order to introduce a country with severe foreign exchange controls. Unlike the major markets, our sample from this market includes only a subsample of the number of traded assets (approximately 20%). The four markets represented nearly 65% of the world equity market capitalization at the end of 1983. Returns from France, Japan, and the United Kingdom, adjusted for dividends and stock splits, were transformed into dollar returns using end-of-month exchange rates from the Data Resources Incorporated data file. Excess returns were computed using the short term US treasury bill return from Ibbotson Associates (1985).[6] We perform our tests on both nominal and real returns. Nominal dollar returns are converted into real returns using the inflation series from Ibbotson Associates (1985).

III. Estimation of Pervasive Economic Factors and Hypothesis Tests

A. Estimation of Pervasive Factors

Our tests of the CAPM amount to prespecifying the benchmark portfolios whose mean-variance efficiency is being tested. However, the assumed linear factor structure which underlies the APT lends itself naturally to direct statistical estimation of the factors. In fact, most empirical tests of the APT to date use standard factor analytic techniques to estimate either the betas of assets or the factor realizations. For our factor models we use the asymptotic principal components technique of Connor and Korajczyk (1986, 1988b). An advantage of their procedure is that it can utilize very large

cross-sections to estimate the pervasive factors. Also, while the number of time periods, T, has to be larger than the number of assumed factors, k, it does not have to be larger than the number of assets, n. While maximum likelihood factor analysis is, in theory, more efficient than principal components, standard factor analysis packages cannot handle the number of cross-sections analyzed here (e.g., the international APT uses between 4211 and 6692 securities to estimate the factors). A brief outline of the asymptotic principal components technique is presented below.

Assume that asset returns follow an approximate k-factor model [in the sense of Chamberlain and Rothschild (1983)], that exact multifactor pricing holds [i.e., (3) holds as an equality], and that we observe the returns on n risky assets and the riskless interest rate over T time periods. Let R^n be the $n \times T$ matrix of excess returns; F be the $k \times T$ matrix of realized factors plus risk premia (i.e., $F_{jt} = f_{jt} + \gamma_{jt}$); and B^n be the $n \times k$ matrix of factor loadings. The estimation procedure allows the risk premia, γ_{jt}, to vary through time. Exact multifactor pricing implies that

$$R^n = B^n F + \varepsilon^n \tag{4}$$

where: $E(F\varepsilon^{n\prime}) = 0$, $E(\varepsilon^n) = 0$, and $E(\varepsilon^n\varepsilon^{n\prime}/T) = V^n$.

Let Ω^n be the $T \times T$ matrix defined by $\Omega^n = R^{n\prime}R^n/n$ and G^n be the $k \times T$ matrix consisting of the first k eigenvectors of Ω^n. Theorem 2 of Connor and Korajczyk (1986) shows that G^n approaches a non-singular transformation of F as $n \to \infty$. That is, $G^n = L^n F + \phi^n$ where plim $\phi^n = 0$. The transformation L^n reflects the standard rotational indeterminacy of factor models but is irrelevant for testing the multifactor pricing model. We assume that our sample size is sufficiently large so that ϕ^n is the null matrix.

Note that, while we are working with cross-sections as large as 6692, the factor estimation method only requires the calculation of the first k eigenvectors of a $T \times T$ matrix. In our work T is equal to 180 (fifteen years with monthly data). For the domestic versions of the APT Ω^n is calculated over the assets traded on the domestic stock exchange and over the entire sample for the international versions. We use the extension of the principal components technique [from Connor and Korajczyk (1988b)] which does not require that assets have a continuous time series of returns. Because of this, our factor estimates are not contaminated through any survivorship bias.

A difficulty which arises in any application of the APT is choosing the appropriate number of factors. A common approach, found in the factor analysis literature, tests whether V^n is diagonal after extracting k factors. This test is inappropriate when asset returns follow an approximate rather than a strict factor model since V^n need not be diagonal in the latter case. We report the results of two tests for the appropriate number of factors.

The first one, suggested by Kandel and Stambaugh (1987), tests a k_1 factor model versus a k_2 factor model ($k_2 - k_1 = p > 0$) by examining whether the mean/variance efficient portfolio formed from k_2 benchmark portfolios (plus the riskless asset) can be formed from the first k_1 benchmark portfolios (plus the riskless asset). Let $\mathbf{P}_{1t}$ denote the $k_1 \times 1$ vector of period t excess returns on the first k_1 factors and $\mathbf{P}_{2t}$ denote the $p \times 1$ vector of period t excess returns on the remaining factors. The null hypothesis that k_1 factors are sufficient implies that the $p \times 1$ vector of intercepts in a multivariate regression of $\mathbf{P}_2$ on $\mathbf{P}_1$ are equal to zero. That is, $\mathbf{a} = 0$ in

$$\mathbf{P}_{2t} = \mathbf{a} + \beta \mathbf{P}_{1t} + \eta_t. \tag{5}$$

To test $a = 0$ we use a modified likelihood ratio (MLR) statistic [see Rao (1973, p. 555)]. The MLR statistic for our hypotheses is given by

$$[(|\hat{\Sigma}_N|/|\hat{\Sigma}_A|) - 1] \cdot (T - k_1 + 1 - p)/p \qquad (6)$$

where T is the number of time series observations, $|\cdot|$ denotes determinant, and $\hat{\Sigma}_N$ ($\hat{\Sigma}_A$) are the maximum likelihood estimates of $E[\bar{\eta}_t\bar{\eta}_t']$ assuming the null, $a = 0$ (the alternative, $a \neq 0$). Under the null the MLR statistic has an F distribution with degrees of freedom equal to p and $T-k_1+1-p$. An advantage of the MLR over alternative test statistics (such as the Wald or unmodified LR statistics) is that its exact small sample distribution is known (when $\bar{\eta}$ has a multivariate normal distribution).[7] We apply the above test to the factors estimated by the asymptotic principal components technique. The results are reported in Table 3. The tests do not seem to provide much power to discriminate against any hypothesized number of factors. In only two out of five cases are we able to reject the null of no factors in favor of the alternative of one factor.

We suggest an alternative test which under certain conditions will give us asymptotically (as the number of assets, n, increases) valid inferences regarding the number of factors for both strict or approximate factor structures. It relies on a result from Ingersoll (1984) which states the cross-sectional mean-square of assets' betas relative to a non-pervasive factor must approach zero as n approaches infinity. That is, if the kth factor is non-pervasive then $B^{n\prime}_{\cdot k}B^{n}_{\cdot k}/n \to 0$ as $n \to \infty$, where $B^{n}_{\cdot k}$ is the kth column of B^n in (4). A necessary condition for the mean-square beta to approach zero is that the average beta must also approach zero since $B^{n\prime}_{\cdot k}B^{n}_{\cdot k}/n \to \sigma^2 + (\bar{B}_k)^2$

where σ^2 is the cross-sectional variance in betas and $\bar{B}_k$ is the average beta. We can estimate the average beta by regressing the excess return of an equal-weighted portfolio on the factors. Non-pervasive factors should have coefficients that are asymptotically zero as the number of assets in the equal-weighted portfolio approaches infinity. Thus, we can use a simple t-test for the null hypothesis that the equal-weighted portfolio has zero sensitivity to the kth factor. There is a reason to expect this test will extract too few factors and a reason to expect the test will extract too many factors. The test might indicate to few factors because it is possible for the mean beta to be zero while the limiting variance is not zero (for example the betas could alternate between 1 and -1). In small samples the test might indicate too many factors since the mean beta relative to a non-pervasive factor is only zero in the limit. The results of this test are reported in Table 4. They are diametrically opposed to the previous one. Only for the United Kingdom do we accept less than fifteen factors. Given that the restriction being tested is only strictly valid for n equal to infinity, the apparently large number of factors should be interpreted with some caution.

Discrepancies in the number of factors is a common feature of the empirical literature on the APT [e.g., Roll and Ross (1980), Dhrymes, Friend, and Gultekin (1984), and Trzcinka (1986)]. Since our tests also do not provide a consistent picture we use other grounds to choose the number of factors. One criterion is that the number of factors should be greater than or equal to the number of countries investigated. This leads to a lower bound of four factors. A second criterion is than we not use too many degrees of freedom through inclusion of too many factors. In some of our empirical work we allow for seasonality in betas and in mispricing. Since we have fifteen

years of data, use of more than fourteen factors is not feasible. We have chosen to estimate multifactor models with five factors and with ten factors.

B. Tests of the Asset Pricing Models

The alternative asset pricing models (1) and (4) each place testable restrictions on the relation between asset returns and the returns on the benchmark portfolios. If we let P denote the vector of excess returns on a generic benchmark proxy (i.e., the return on some market index for the CAPM or the return on either prespecified or estimated factors for the APT) then the intercept in the regression of any asset's excess returns on P should be zero. Thus, given a sample of m assets and the regressions

$$R_{it} = \alpha_i + b_i P_t + \varepsilon_{it} \qquad i = 1, 2, \ldots, m; \; t = 1, 2, \ldots T \tag{7}$$

the pricing models imply the restriction

$$\alpha_1 = \alpha_2 = \ldots = \alpha_m = 0. \tag{8}$$

We will refer to α_i as the mispricing of asset i relative to the benchmark P. We first test whether mispricing is non-zero across assets for each of our alternative benchmarks. This is a test for unconditional mean/variance efficiency of some linear combination of the benchmark portfolios, P.

Because of the well-documented January seasonal patterns in asset returns we also allow the mispricing of assets to differ in January from the mispricing common to all months.[8] This is done by estimating the regression

$$R_{it} = \alpha_{iNJ} + \alpha_{iJ} D_{Jt} + b_i P_t + \varepsilon_{it} \tag{9}$$

where D_{Jt} is a dummy variable equal to 1 in January and zero otherwise. Mispricing specific to January is measured by α_{iJ} while mispricing which is

not specific to January is reflected in α_{iNJ}.[9] The hypotheses regarding α_i, α_{iNJ}, and α_{iJ} [e.g., as in (8)] are tested using the MLR statistic described above. Under the null, the test statistic has a central F distribution with degrees of freedom equal to m (the number of assets in the sample) and T-k+1-m (where k is the number of regressors, excluding the constant).

As discussed above, rejection of the null hypothesis in (8) might be attributable to a difference between the expected return on the true zero beta asset and the return on our proxy for r_{Ft}. We allow for this by testing the restrictions implied by zero-beta forms of the models. We assume that the expected excess zero-beta return, $\lambda = E(\bar{R}_{zt}) - r_{Ft}$, is constant through time.

The restrictions implied by the zero-beta CAPM in (2) on the multivariate regression (7) are given by

$$\alpha_i = (1 - b_i)\cdot\lambda \qquad i = 1, 2, \ldots, m \tag{10}$$

We test the nonlinear cross-equation restriction (10) using the likelihood ratio (LR) test [see Gallant (1987, pp. 457-8)] which has a χ^2 distribution, asymptotically, with degrees of freedom equal to m - 1. Unlike the MLR statistic for the linear multivariate regression case, we do not know the exact small sample distribution of the LR test of the restriction (10).

The restrictions implied by the zero-beta version of the APT are less complicated if we are willing to assume that our proxy for the riskless asset bears only factor related risk. In the Appendix A we show that our estimates of the factors converge to the true factors plus risk premia relative to the true zero-beta return as long as the return, r_{Ft}, is well diversified. This implies that the intercepts in the regression (7) are equal to λ when the benchmark portfolio proxies, P, are derived from the asymptotic principal

components technique.

The analysis of Hansen and Richard (1987) provides a framework for interpreting our tests of unconditional and conditional mean/variance efficiency. Note that our conditional tests use only a subset of information available to economic agents. Thus, we need to make the distinction between unconditional efficiency, efficiency conditional on a coarse (the econometricians') information set, and efficiency conditional on the full information set.[10] Hansen and Richard (1987) show that unconditional efficiency implies conditional efficiency but that the converse is not true.[11] Thus, failing to reject unconditional or limited conditional efficiency is consistent with the hypothesis of conditional efficiency. On the other hand, rejecting unconditional or limited conditional efficiency does not imply rejection of conditional efficiency. Rejection of limited conditional efficiency implies rejection of unconditional efficiency.

Panel A of Table 5 summarizes the parameter restrictions implied by the various models described above.

C. Tests of the Effects of Capital Controls

The regulatory environment of international financial markets is likely to be an important factor in capital market integration and asset pricing. In Appendix B we give a brief description of changes in capital controls over our sample period. There is a general trend towards deregulation marked by two major periods of change. The first took place at the beginning of 1974 when the Interest Equalization Tax was eliminated in the United States (January) while other countries loosened restrictions on capital inflows (January-February). Also, the early 1974 period marks the completion of the transition from a regime of fixed exchange rates to one of floating rates. The second is

in 1979 when the United Kingdom and Japan dismantled a number of controls.[12]

We investigate whether periods of more strict controls (ending in January 1974 and November 1979, respectively) are associated with deviations from the predictions of the asset pricing models. This is done by testing whether the size of mispricing is different during these periods. We construct two dummy variables D_{74t} and D_{79t} such that D_{74t} is equal to 1.0 before February 1974 and 0.0 afterwards while D_{79t} is equal to 1.0 before December 1979 and 0.0 otherwise. We then test $\alpha_{i74} = 0$ and $\alpha_{i79} = 0$, for all i, in the regression

$$R_{it} = \alpha_i + \alpha_{i74}D_{74t} + \alpha_{i79}D_{79t} + b_iP_t + \varepsilon_{it} \tag{11}$$

$$i = 1, 2, \ldots, m; \; t = 1, 2, \ldots T$$

We also estimate variants of (11) which allow for a January seasonal in α_i and b_i as well as the zero-beta forms of the models. Mispricing which is invariant over the entire fifteen year period is measured by α_i.

In panel B of Table 5 we summarize our tests for the influence of capital controls.

D. Choice of Dependent Variables for Hypothesis Tests

As discussed above the asset pricing models imply restrictions on the coefficients of a multivariate regression of asset returns on particular benchmark portfolios. One would normally proceed in testing the hypothesis of zero mispricing by estimating the restricted null model [e.g., equation (7) with the constraint $\alpha_i = 0$] and the unrestricted version [e.g., (7) with the intercepts allowed to be non-zero]. Standard approaches to hypothesis testing involve investigating the increase in the generalized variance (determinant) of the residual covariance matrix, $\hat{V}$, due to additional restrictions (as in likelihood ratio tests) or calculating quadratic forms relative to $\hat{V}$. Large

values of m [i.e., many assets on the left hand side (LHS) of the regression] present some difficulties in hypothesis testing. In particular, when m is larger than T the generalized variances are uniformly zero and the estimated residual covariance matrix is singular. There are several alternative techniques designed to overcome this problem.

A common approach, which we adopt, is to group assets into portfolios on the basis of some instrumental variables. Thus, rather than having m individual assets on the LHS of the regressions we have p portfolios (with p << m). This makes testing feasible, allows more precise estimates of the parameters but also runs the risk of masking mispricing if the values of α_i are uncorrelated with the instruments. Thus, there is a tradeoff between increased precision of our estimates and decreased heterogeneity in the sample.[13] The instrument used to form portfolios should be chosen to insure heterogeneity across portfolios. The instrumental variable chosen here is the "size" of the firm.[14] We form five sets of size portfolios - one set per country plus a set which includes all assets. For each set we rank firms on the basis of market value of equity at the beginning of the period (December 1968) and form ten equal-weighted portfolios (the first portfolio containing the smallest 10% of the firms, etc.). A firm remains in its portfolio as long as there are observed returns for this asset. Assets are reallocated to size portfolios at five year intervals (i.e., December 1973 and December 1978).

IV. Empirical Results

The results reported below are robust to a variety of permutations in estimating the models. We estimate each model using both nominal and real returns. The inferences we draw about the model are not dependent on whether real or nominal returns are used. Because of this, we report our results

using nominal returns. Since we are assuming that various parameters are constant over our fifteen year sample period we check whether our results are robust to allowing changes in the parameters. We do this by estimating the models over three five year subperiods and aggregating the subperiod results. The aggregated results did not yield different inferences from the entire period. We report the results from the entire period.

A. The Correlation Structure of Asset and Factor Returns

Before we proceed with the formal hypothesis tests we present some evidence on the covariance structure of asset returns across countries and evidence on the relation between market indices and our estimated factors. Table 6 is the sample correlation matrix of several market indices over the 1969-1983 period. The correlations across national indices range from 0.20 to 0.47 and are consistent with previous evidence. While there are important common movements in the various indices, there also appear to be substantial country specific components to the return series. The correlation between equal-weighted and value-weighted indices in the same country are, as one would expect, high (from 0.87 to 0.98).

Table 7 provides some evidence on the relation between the country indices and our international factors estimated by the asymptotic principal components procedure which allows us to use the returns on every available asset to calculate Ω^n. When we estimate international factors across all four economies our sample includes between 4211 and 6692 firms. Over our 180 month period, the monthly average number of firms with returns data is 5596. The eigenvectors of Ω^n are our estimates of the pervasive factors. Regressions of the excess returns of the national indices on the first five international factors are reported in Table 7. The coefficients of determination (R^2),

reported in the last column, indicate a very strong relation between the estimated factors and the indices for every country except France. Also, each of the five factors generally has significant explanatory power across all countries. The results in Table 7, and some extensive canonical correlation analysis not reported here, indicate that there are several common international factors across countries. The estimated mispricing of each index relative to these five benchmark portfolios (in % per annum) is listed in the third column. The values of $\hat{\alpha}_{FR}$ are (economically) very negative but are not measured with much precision. The estimated mispricing relative to the 10-factor APT is generally smaller (in absolute value). They are not reported in detail here in order to conserve space.

B. Multi-Index versus Single-Index Models

In this section we compare the performance of multi-index and single index models using the tests described in Section III.B. We use two criteria (1) whether or not we reject the restrictions implied by the models and (2) the relative magnitudes of the estimates of mispricing across models. We find the second criterion useful given the non-nested nature of the models. For example, rejection of the restriction (8) for one model and failure to reject (8) for a second model does not imply that the second model fits better. The mispricing parameters (α_i) of the first model might be closer to zero but measured with more precision. Thus a combination of the two criteria is more informative than either one alone.

When we assume that the US T-bill return is the appropriate riskless return, the asset pricing models imply that the intercepts (α_i) are zero in a multivariate regression of size portfolios excess returns on particular benchmark portfolios. When we allow mispricing to be seasonal, both the

seasonal and non-seasonal components of mispricing (α_{iJ} and α_{iNJ}) should be zero. In Table 8 we present the results of the tests.

There is no model without at least one rejection, at the 5% level of significance, of the null that non-seasonal mispricing is zero ($\alpha_i = 0$ or $\alpha_{iNJ} = 0$). The CAPM-VW model has the fewest rejections while the APT-10 model has the most. The null is always rejected for UK and for international size portfolios but never rejected for Japanese portfolios. The hypothesis that January specific mispricing is zero ($\alpha_{iJ} = 0$) is never rejected by the APT-10 model and more often rejected by the CAPM than by the APT-5 model. Considering the three null hypotheses together: $\alpha_i = 0$, $\alpha_{iNJ} = 0$ and $\alpha_{iJ} = 0$, it appears that there is some evidence against all of the models (with the exception of the APT for Japan). Also, it appears that the CAPM does better in explaining returns that are not specific to January and the APT does better in explaining January specific returns.

Test results for the zero-beta specifications of the models are presented in Table 9. With few exceptions (CAPM-EW model for the UK and the domestic APT-10 for France), whenever the null is rejected with the US T-Bill rate as the zero-beta return, we also reject the zero-beta variant of the model. On the other hand, there are a few cases where we reject the zero-beta model but do not reject the model using the US T-bill return. This is due to the fact that the p-values in Table 8 are valid in small samples (assuming normality of ε) while the p-values for the restrictions in Table 9 are valid asymptotically. The likelihood ratio tests in Table 9 will tend to reject too often in small samples.

The tests reported in Tables 8 and 9 provide us with our first criterion for model evaluation. However, sole reliance on the p-values in those tables

may be misleading because, among other things, the power of the tests may be different across models. The power of the above tests increases with the precision of our estimates of mispricing, *ceteris paribus*. Holding the level of mispricing constant, we would expect more precise estimates of mispricing for portfolios with larger numbers of securities (by diversification). The number of assets included in our size portfolios vary greatly across economies. For example, each of the ten international size portfolios have 457 assets, on average, while the French size portfolios have only 12. Thus a simple comparison of the test statistics may be insufficient to estimate the relative performance of each model and of each of their various versions across countries.

Hence, we present evidence of the relative magnitude of the mispricing of the alternative models. Space limitations prevent us from showing the entire set of graphs corresponding to each possible permutation. We chose only two graphs which, along with Table 10, best illustrate the most important findings from a detailed comparisons of the models.

Figure 1 shows the mispricing for the four models using international size portfolios with international benchmarks. The graph plots the mispricing for each size portfolio, from the smallest (S1) to the largest (S10). Mispricing for small size portfolios is larger than for large size portfolios, whatever the model: actually none of the four model seems to fully explain the size related anomaly. This finding holds for each of the four countries individually, using domestic as well as international benchmarks, with the UK showing the strongest size effect and France the weakest.[15] In Table 10 we present the average absolute mispricing of the size portfolios for the models as an estimate of the extent to which they deviate from zero mispricing.

Mispricing is relatively large (in economic terms) for the CAPM-VW model and is systematically larger than any of the three other models whatever the version. Differences in mispricing between the factor models and the CAPM-EW model are minimal. There is a striking contrast between the frequency of rejection based on the test statistics and the level of mispricing. The CAPM-VW has the fewest number of rejections but the largest estimates of mispricing. Similarly the APT has more frequent rejections of the restrictions but fits the data better than the CAPM.

January mispricing for the same models and size portfolios are shown on Figure 2. As with mispricing, January mispricing of small size international portfolios is larger than for large size ones. Again, this finding holds at the country level with the US showing the strongest effect and France the weakest. However, the effect is clearly more pronounced for the CAPM, a finding which confirms the results of the statistical tests. This is also true for each country using domestic as well as international benchmarks. In other words, the APT models seems to include seasonal factors not "picked up" by the alternative models. From Table 10, the CAPM-VW model, again, shows the largest average absolute mispricing of the four models, but contrary to the previous finding, the APT models and specially the APT-10 model show a much smaller mispricing than the CAPM-EW model, except in some cases for France.

To summarize, although the size effect is present when estimating each of the four models, the APT models tend to perform better than the CAPM models especially when comparing the the magnitude of the January mispricing. The difference in performance between the two factor models is minimal. In particular, both seem to include seasonal factors which "explain" January-specific asset return behavior. Our results for domestic benchmarks are

consistent with the single-economy applications of the APT cited in footnote 2. We know of no previous study which directly compares the international APT to the international CAPM.

C. Domestic versus International Benchmarks

When comparing domestic versus international models we focus on the mispricing of the domestic size portfolios relative to the domestic and international benchmark portfolios, respectively. From Tables 8 and 9 the models with international benchmarks are rejected slightly less often than the domestic models. However, when comparing the magnitude of the mispricing estimates in Table 10, it appears that the domestic versions marginally outperform the international versions, except for the CAPM-VW model.

D. Impact of Regulatory Changes in International Financial Markets

D.1 Impact on Model Performance

In Section III.C, we describe a test of the impact of changes in the regulatory environment of the international financial markets on the performance of the models which allows mispricing to vary across regulatory regimes. If the changes in regulations do not influence international asset pricing, then we would expect our regime shift coefficients [α_{i74} and α_{i79} in (11)] to equal zero.

Table 11 contains our test results when we allow mispricing to be regime dependent (Table 11 is comparable to Table 8). The hypotheses $\alpha = 0$ and $\alpha_{NJ} = 0$ are not rejected whatever the model, except for France where they are always rejected. The hypothesis $\alpha_J = 0$ is overwhelmingly rejected for the CAPM while it is seldom rejected for the APT (especially the ten factor model). These results are quite different from those shown in Table 8 when no adjustment was made for changes in the international financial markets. They are consistent

with international regulatory influences on asset pricing.

CAPM and APT models differ on the significance of α_{i74}: it tends to be significant for the APT model but not for the CAPM. However, in the case of international size portfolios, α_{i74} is always statistically significant. The hypothesis that $\alpha_{i79} = 0$ is never rejected, except for France. These findings tend to show that the asset pricing models are sensitive to the changes around early 1974 which include switching from fixed to floating exchange rates, the elimination of the interest equalization tax in the US, and liberalization of capital controls on the part of the other countries. The performance of the models does not seem to be affected by the changes in 1979.

We present in Table 12 the tests of the zero-beta variants of the models. They mimic those obtained when the US T-Bill return is used as the zero-beta return. The restrictions implied by the zero-beta models cannot be rejected except for France and they are also in sharp contrast with the results of the tests presented in Table 9 which do not allow for regime shifts in mispricing.

D.2 Multi-Index versus Single-Index Models Adjusting for Regime Shifts

The above results indicate that the APT tends to be rejected less often, especially relative to January mispricing. Figures 3 and 4 show our estimates of mispricing and January mispricing as a function of portfolio size when we include dummy variables for the 1974 and 1979 changes in the regulatory environment. The graphs show patterns that are similar to those in Figures 1 and 2. There is still a size effect for all models and a strong January effect for the CAPM models. However, the size effect is much less pronounced when the models are adapted for the regime changes. At the country level, the US exhibits the strongest size effect and Japan and France the weakest. In none of the four countries there is any noticeable January effect for the APT.

Average absolute mispricing sizes are presented in Table 13. As before, CAPM-VW shows by far the largest mispricing and January mispricing of the four models (except for France). Differences are minimal between the CAPM-EW and the APT models except for the January mispricing which, again, is much lower with the APT. When compared to the estimates presented in Table 10, the APT's mispricing is systematically lower, except for France, while their January mispricing is comparable, again, except for France.

D.3 Domestic versus International Models Adjusting for Regime Shifts

The international version of the CAPM seems to outperform the domestic version in terms of both of our criteria. We reject the CAPM restrictions slightly more often for the domestic versions than the international versions (see Table 11). We also find that the CAPM has smaller pricing errors in its international version than in its domestic version (see Table 13).

We generally find the opposite results for the APT. Using domestic size portfolios we reject the APT restrictions slightly more often for the international benchmarks. From the levels of absolute mispricing in Table 13 the domestic versions of the APT seems to outperform the international versions. However, when we use international size portfolios, the ten factor APT is the only model which does not reject the absence of a January seasonal effect in pricing.

In summary, the analysis of the size of the mispricing confirms the finding of the statistical analysis: the four asset pricing models seem to be sensitive to changes in the regulatory environment of the international financial markets. The period from January 1969 (the beginning of our sample) to January 1974 seems to cause many of the rejections of the model. Abstracting from the period prior to February 1974 we find that multi-index

models continue to outperform the single-index models. Also, International versions of the CAPM outperform domestic versions while the opposite is generally true for the APT.

V. Conclusions

We compare domestic and international versions of a several alternative asset pricing models. The empirical results indicate that: (1) There is some evidence against all of the models, especially in terms of pricing common stock of small market value firms. (2) Multifactor models tend to outperform single-index CAPM-type models in both domestic and international forms. The value-weighted CAPM has much larger pricing errors than the APT. The equal-weighted CAPM performs about as well as the APT except in terms of explaining seasonality in asset returns. (3) There is strong evidence that the behavior of the models in the period from January 1969 to January 1974 is different from their behavior after January 1974. We interpret this evidence as being consistent with a scenario in which some combination of capital control deregulation and the break down of the fixed exchange rate regime lead to pricing effects that are not well captured by models of either completely segmented or completely integrated markets. (4) Controlling for regime shifts in the level of capital controls, international versions of the CAPM outperform domestic versions while the opposite is true for the APT. The evidence is generally consistent with non-trivial international influences in asset pricing.

Appendix A

In some asset pricing models (particularly international models) the US treasury bill return is not the appropriate zero-beta return, i.e., $r_{Ft} \neq E(\tilde{r}_{zt})$. In this appendix we show that we need not assume that $r_{Ft} = E(\tilde{R}_{zt})$ or even that r_{Ft} is riskless in order to obtain valid estimates of the pervasive factors and their associated risk premia $(f_{jt} + \gamma_{jt})$. Although r_{Ft} is riskless in nominal US dollar returns is easy to see that it may not be riskless in real terms or relative to another currency. Under certain conditions we can use excess returns relative to any well diversified asset or portfolio. Let $\tilde{R}_{it} = \tilde{r}_{it} - \tilde{r}_{\delta t}$ (i.e., we are calculating excess returns relative to asset δ). We assume that (a) $r_{\delta t}$ is well diversified and (b) taking excess returns with respect to $r_{\delta t}$ does not alter the basic nature of the factor structure (i.e., taking excess returns with respect to $r_{\delta t}$ does not turn a k-factor model into a q-factor model with q<k). That is,

a) $r_{\delta t} = \mu_{\delta t} + b_{\delta} f_t$

b) $\|(B^{*n\prime} B^{*n})^{-1}/n\| \leq c < \infty$ for all n

where: $B^{*n} = B^n - \iota b_{\delta}'$;

ι = an $n \times 1$ vector of 1's;

and B^n is as defined in (4).

Given these conditions, all of the assumptions required by Connor and Korajczyk (1986) hold and their Theorem 2 can be applied to show that $G^n = L^n F + \phi^n$ with plim $\phi^n = 0$. Now the pricing model implies that $E(\tilde{r}_{it}) - E(\tilde{r}_{zt}) = b_i \gamma$ and, hence, that $E(\tilde{r}_{it}) - E(\tilde{r}_{\delta t}) = [E(\tilde{r}_{zt}) - E(\tilde{r}_{\delta t})] + b_i \gamma = \lambda_t + b_i \gamma$. Under the assumption that $\lambda_t = \lambda$, the intercept terms in (7) should all equal λ as was stated in the text. If condition (b) does not hold then λ will include the risk premia for the k-q factors that were eliminated.

Appendix B[16]

Significant Changes in Capital Flows Regulations France, Japan, United Kingdom, and United States 1969-1983

Chronology:

Situation end 1968

US: Purchases of foreign securities are subject to the Interest Equalization Act. Capital exports are subject to mandatory controls.

FR: Capital transfers are subject to exchange control approval.

UK: All capital transfers by residents outside the Sterling area require approval under the Exchange Control Act.

JP: Capital transactions are permitted unless specifically restricted.

1969

All four countries ease their control on capital flows through marginal changes in regulations.

1970

Overall, relaxation of controls in all countries within the framework of existing regulations.

1971

"In the turbulent exchange market conditions prevailing in mid-71, measures to control capital movements were widespread among industrial countries in mid-71."

1972

Some control measures implemented in 1971 were revoked after the Smithsonian Agreement (Dec 18, 1971). But reliance on capital controls increased after mid-72 and intensified in late 1972. No significant exceptions among the

four countries.

1973

Measures to curb capital flows were not relaxed until mid-73. At that time all four countries initiated a series of measures towards liberalization of capital flows.

US: from June 1963 to December 1973 a series of measures were taken to relax controls on capital flows.

FR: in August a series of measures were announced which dismantled or relaxed most of exchange restrictions taken in 68/69 but several were restarted in January 1974.

UK: May 22, relaxation of restrictions on Sterling lending to non-residents. Oct 19, borrowing abroad facilitated.

JP: In Nov and Dec, series of measures to relax controls on capital outflows.

1974

US: in January 1974, controls on capital outflows were terminated.

Rest "Industrial countries of Europe and Japan had, by February, partially dismantled the complex system of control over capital inflows. A few tightened controls on outflows.

1975

"For the first time in a number of years, industrialized countries adopted relatively few major measures affecting capital movements."

1976

"Measures taken did not, in general, signify any fundamental changes in politics on capital flows."

1977

"Measures bearing on capital movements during the period represented in general the continuation of policies already in place." However, in Nov 77 JP undertook measures to restrict capital inflows for portfolio investment.

1978

"Measures...represented in general an adaptation of existing policy trends."

1979

UK: Historical decision to dismantle the system of exchange controls through measures taken from June to October.

JP: A series of liberalization measures were announced in January and, mostly, in May to take effect in March 1980.

1980

JP: Further liberalization.

1981

FR: May 1981, severe exchange controls established.

1982

"On balance, measures were taken to facilitate the international mobility of capital."

1983

"On the whole, changes in the regulations governing capital flows do not appear to have had an important impact on international capital flows."..."The long term trend toward liberalization of capital control regulations was sustained in the industrial countries." However, capital transfers between France and all other countries are still subject to exchange controls. They are permitted in Japan "unless specifically restricted."

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Endnotes

1. See, for example, Banz (1981) and Keim (1983) for evidence on US exchanges, Kato and Schallheim (1985) for the Tokyo stock exchange, Corhay, Hawawini, and Michel (1987) for the London stock exchange, and Hawawini and Viallet (1987) for the Paris stock exchange.
2. In the US, Chen (1983) finds that the size anomaly becomes insignificant when the APT is used while Lehmann and Modest (1988) and Connor and Korajczyk (1988a) find a significant size effect remaining. Lehmann and Modest (1988) do find that the dividend yield anomaly is no longer significant. In the UK, Beenstock and Chan (undated) find that the APT does significantly better than the CAPM in explaining asset returns, as do Dumontier (1986) using French stocks and Hamao (1986) using Japanese stocks.
3. Compare the numbers in Table 1 to the number of securities used in Cho, Eun, and Senbet (1986) -- US (60), Japan (55), UK (48), and France (24) and the numbers used in Gultekin, Gultekin, and Penati (1987) -- US (110) and Japan (110).
4. In addition to the usual assumptions needed for the CAPM [see Constantinides (1980)], assuming that strict purchasing power parity (PPP) holds (i.e., the law of one price must hold across national boundaries) would be sufficient for (1) to hold internationally. Exchange rate uncertainty is not priced separately from market risk because the PPP assumption implies that changes in exchange rates do not change relative prices. See, for example, Solnik (1974, pp. 514-18) and Stulz (1985 pp. 77-79).

5. Ross (1976) assumes that $E(\bar{\varepsilon}_i \bar{\varepsilon}_j) = 0$ (a strict factor model). Chamberlain and Rothschild (1983) and Ingersoll (1984) show that the APT can be derived under the weaker condition that the eigenvalues of the cross-sectional idiosyncratic covariance matrix are bounded as the number of assets grows large (an approximate factor model).

6. We run nominal versions of the models using the other three currencies: franc, pound, and yen and the short term interest rates prevailing in each of the country's currency. Interest rates were taken from the International Financial Statistics tables. We find that the test results are not significantly affected by the currency chosen. As a result, and to save space, we choose to present our results using US dollar as numeraire only.

7. Geometric interpretations of this test are provided in Gibbons, Ross, and Shanken (1986) and Kandel and Stambaugh (1987).

8. Some previous studies have reported January and April seasonality in stock returns in the United Kingdom [Corhay, Hawawini, and Michel (1987)]. Our tests show no significant April mispricing for the UK over our sample period. These results are not necessarily inconsistent since seasonality in risk premia need not imply seasonality in mispricing.

9. The specification in (9) incorporates variation in conditional means but assumes that conditional betas are constant. We also estimate a specification which incorporates variation in conditional betas by letting b_i be seasonal:

$$R_{it} = \alpha_{iNJ} + \alpha_{iJ}D_{Jt} + b_{iNJ}P_t + b_{iJ}D_{Jt}P_t + \varepsilon_{it}.$$

We find no substantive difference in the estimated mispricing between this specification and that of (9). For this reason we only report the results from (9).

10. We will refer to these as unconditional, limited conditional, and conditional efficiency, respectively.

11. The same logic shows that unconditional efficiency implies limited conditional efficiency but that the converse is also not true.

12. In Japan, deregulation measures, announced in early 1979, were implemented in 1980.

13. The implications of this tradeoff, in terms of the power of the tests, is analyzed in Gibbons, Ross, and Shanken (1986).

14. The papers cited in footnote 1 indicate that size is a reasonable instrument in terms of insuring heterogeneity across portfolios.

15. It is not surprising that France shows the weakest size effect. The 126 firms in the French sample are only a fraction of the firms traded on the Paris Bourse and represent the most frequently traded shares. As a consequence the sample is comprised of firms which are rather homogeneous in size.

16. Reference: "Main Developments in Restrictive Practices, Measures Affecting Capital", IMF Annual Report on Exchange Arrangements and Exchange Restrictions, 1970-1984.

Table 1

Exchange Market Data and Sample Data Summary

COUNTRY	UNITED STATES	JAPAN	UNITED KINGDOM	FRANCE	TOTAL
Exchange Market Data	NYSE & AMEX	TOKYO	LONDON	PARIS	
Market Capitalization (12/83)[a] World Capitalization	43%	15%	6.1%	1%	65.1%
Number of listed firms (12/83)[a]	2274	1441	2217	518	6450
Sample Data					
Sample source	CRSP	Japanese Research Institute (JSRI)	London Share Price Data Base	Compagnie des Agents de Change	
Frequency of returns	Monthly	Monthly	Monthly	Monthly	
Number of sample firms:					
Minimum	2187	672	1138	112	4211
Maximum	2706	1420	2555	126	6692
Average	2457	1144	1874	124	5596

a: source - International Federation of Stock Exchange Statistics, 1983

Table 2

Models and Versions Tested

MODELS[a]			
CAPM-EW	CAPM-VW	APT-5	APT-10

VERSIONS

Domestic/Domestic		Domestic/Int'al		Int'al/Int'al	
R_i	Benchmark	R_i	Benchmark	R_i	Benchmark
US[b]	US	US	Int'al[c]	Int'al	Int'al
JP	JP	JP	Int'al		
UK	UK	UK	Int'al		
FR	FR	FR	Int'al		

[a] Zero-beta variants of each model are also tested.
[b] US: United States, JP: Japan, UK: United Kingdom, FR: France.
[c] These benchmark portfolios include all four countries' assets.

Table 3

Tests of the null hypothesis of k_1 factors versus the alternative of k_2 factors. The null hypothesis implies that the intercepts in a multivariate regression of the last $k_2 - k_1$ factors on the first k_1 factors equals zero.

		P-values				
k_1	k_2	US	UK	Japan	France	World
0	1	0.366	0.074*	0.003*	0.368	0.180
1	2	0.349	0.014*	0.894	0.400	0.103
2	3	0.328	0.452	0.745	0.103	0.032*
3	4	0.771	0.464	0.730	0.725	0.292
4	5	0.636	0.138	0.336	0.623	0.531
5	6	0.445	0.801	0.932	0.797	0.511
1	5	0.711	0.050*	0.882	0.088	0.068
5	10	0.973	0.986	0.989	0.414	0.839
10	15	0.872	0.398	0.588	0.485	0.937

Note: Factors estimated by asymptotic principal components using monthly data from January 1969 through December 1983. P-values are the right tail area of the MLR statistic for restriction that intercepts equal zero. See Kandel and Stambaugh (1987).

* denotes significance at the 5% level.

Table 4

Test of k_1 factor model versus k_2 factor model. The null hypothesis implies that the betas of an equal-weighted portfolio relative to factor k_1+1 through factor k_2 are equal to zero asymptotically (as the number of assets in the equal-weighted portfolio increase).

		P-values				
k_1	k_2	US	UK	Japan	France	World
1	5	<0.001*	<0.001*	<0.001*	<0.001*	<0.001*
5	10	0.010*	0.139	0.001*	0.005*	<0.001*
10	15	<0.001*	0.060	<0.001*	<0.001*	0.003*

Note: Factors estimated by asymptotic principal components using monthly data from January 1969 through December 1983. P-values are the right tail area of the MLR statistic for restriction that the betas of the equal-weighted portfolio relative to factors k_1+1 through k_2 are jointly zero.

* denotes significance at the 5% level.

Table 5

Panel A: Tests of the Asset Pricing Models

Summary[a]

Null	Regression	Model
$\alpha_i=0$	$R_i=\alpha_i+b_iP+\varepsilon_i$	CAPM, APT
$\alpha_{iJ}=0$; $\alpha_{iNJ}=0$	$R_i=\alpha_{iNJ}+\alpha_{iJ}D_J+b_iP+\varepsilon_i$	CAPM, APT
$\alpha_{iJ}=0$; $\alpha_{iNJ}=0$	$R_i=\alpha_{iNJ}+\alpha_{iJ}D_J+b_{iJ}D_JP+b_{iNJ}P+\varepsilon_i$	CAPM, APT
$\alpha_i=(1-b_i)\lambda$	$R_i=\alpha_i+b_iP+\varepsilon_i$	CAPM Zero-b
$\alpha_i=\lambda$	$R_i=\alpha_i+b_iP+\varepsilon_i$	APT Zero-b
$\alpha_{iNJ}=(1-b_i)\lambda$	$R_i=\alpha_{iNJ}+\alpha_{iJ}D_J+b_iP+\varepsilon_i$	CAPM Zero-b
$\alpha_{iNJ}=\lambda$	$R_i=\alpha_{iNJ}+\alpha_{iJ}D_J+b_iP+\varepsilon_i$	APT Zero-b
$\alpha_{iNJ}=(1-b_i)\lambda$	$R_i=\alpha_{iNJ}+\alpha_{iJ}D_J+b_{iJ}D_JP+b_{iNJ}P+\varepsilon_i$	CAPM Zero-b
$\alpha_{iNJ}=\lambda$	$R_i=\alpha_{iNJ}+\alpha_{iJ}D_J+b_{iJ}D_JP+b_{iNJ}P+\varepsilon_i$	APT Zero-b

[a] $i=1,..m$ refers to the dependent variables; P to the benchmarks; D_J to a dummy variable with $D_J=1$ in January and zero otherwise.

Table 5 (Cont'd)

Panel B: Tests of the Effects of Capital Controls

Summary[b]

Null	Regression	Model
$\alpha_i=0$; $\alpha_{i74}=0$; $\alpha_{i79}=0$	$R_i=\alpha_i+\alpha_{i74}D_{74}+\alpha_{i79}D_{79}+b_iP+\varepsilon_i$	CAPM,APT
$\alpha_{iNJ}=0$; $\alpha_{iNJ74}=0$; $\alpha_{iNJ79}=0$; $\alpha_{iJ}=0$	$R_i=\alpha_{iNJ}+\alpha_{iNJ74}D_{74}+\alpha_{iNJ79}D_{79}+\alpha_{iJ}D_J+b_iP+\varepsilon_i$	CAPM,APT
$\alpha_{iNJ}=0$; $\alpha_{iNJ74}=0$; $\alpha_{iNJ79}=0$; $\alpha_{iJ}=0$	$R_i=\alpha_{iNJ}+\alpha_{iNJ74}D_{74}+\alpha_{iNJ79}D_{79}+\alpha_{iJ}D_J+b_{iJ}D_JP+b_{iNJ}P+\varepsilon_i$	CAPM,APT
$\alpha_i=(1-b_i)\lambda$	$R_i=\alpha_i+\alpha_{iNJ74}D_{74}+\alpha_{iNJ79}D_{79}+b_iP+\varepsilon_i$	CAPM Zero-b
$\alpha_i=\lambda$	$R_i=\alpha_i+\alpha_{iNJ74}D_{74}+\alpha_{iNJ79}D_{79}+b_iP+\varepsilon_i$	APT Zero-b
$\alpha_{iNJ}=(1-b_i)\lambda$	$R_i=\alpha_{iNJ}+\alpha_{iNJ74}D_{74}+\alpha_{iNJ79}D_{79}+\alpha_{iJ}D_J+b_{iNJ}P+\varepsilon_i$	CAPM Zero-b
$\alpha_{iNJ}=\lambda$	$R_i=\alpha_{iNJ}+\alpha_{iNJ74}D_{74}+\alpha_{iNJ79}D_{79}+\alpha_{iJ}D_J+b_{iNJ}P+\varepsilon_i$	APT Zero-b
$\alpha_{iNJ}=(1-b_i)\lambda$	$R_i=\alpha_{iNJ}+\alpha_{iNJ74}D_{74}+\alpha_{iNJ79}D_{79}+\alpha_{iJ}D_J+b_{iJ}D_JP+b_{iNJ}P+\varepsilon_i$	CAPM Zero-b
$\alpha_{iNJ}=\lambda$	$R_i=\alpha_{iNJ}+\alpha_{iNJ74}D_{74}+\alpha_{iNJ79}D_{79}+\alpha_{iJ}D_J+b_{iJ}D_JP+b_{iNJ}P+\varepsilon_i$	APT Zero-b

[b] $i=1,..,m$, P and D_J as in Panel A; D_{74}, D_{79} refer to dummy variables with $D_{74}=1$ until January 1974 and zero afterwards, $D_{79}=1$ until November 1979 and zero afterwards.

Table 6

Sample Correlation of Equal-Weighted and Value-Weighted Domestic and International Market Index Portfolios 1969-1983

	US-VW	UK-EW	UK-VW	JP-EW	JP-VW	FR-EW	FR-VW	W-EW	W-VW
US-EW	0.87	0.44	0.48	0.22	0.27	0.31	0.33	0.85	0.85
US-VW		0.38	0.46	0.20	0.29	0.35	0.37	0.74	0.96
UK-EW			0.93	0.36	0.38	0.44	0.44	0.80	0.55
UK-VW				0.31	0.35	0.46	0.47	0.78	0.62
JP-EW					0.90	0.38	0.36	0.52	0.41
JP-VW						0.38	0.38	0.54	0.51
FR-EW							0.98	0.49	0.46
FR-VW								0.49	0.48
W-EW									0.86

Note: US, UK, JP, FR, and W denote United States, United Kingdom, Japan, France and International portfolios, respectively. EW denotes equal-weighted and VW denotes value-weighted.

Table 7

Regression of Market Index Excess Returns on Estimated International Factors

Five Factor Model

$$R_{it} = \alpha_i + \beta_{i1}P_{1t} + \ldots + \beta_{i5}P_{5t} + \varepsilon_{it}$$

INDEX	$\alpha_i \times 1200$	$\beta_{i1} \times 10$	$\beta_{i2} \times 10$	$\beta_{i3} \times 10$	$\beta_{i4} \times 10$	$\beta_{i5} \times 10$	R^2
US-EW	0.76	8.51	3.13	-0.02	0.29	0.45	0.990[a]
	(1.19)	(121.91)	(44.76)	(-0.30)	(4.11)	(6.51)	0.990[b]
US-VW	-3.60	5.09	1.72	0.01	1.68	2.45	0.916
	(-2.78)	(36.05)	(12.16)	(0.05)	(11.93)	(17.47)	0.913
UK-EW	1.78	6.76	-6.27	-1.25	0.26	-0.30	0.988
	(2.52)	(87.90)	(-81.38)	(-16.19)	(3.40)	(-3.87)	0.988
UK-VW	-6.59	7.78	-6.19	-2.01	0.29	1.34	0.920
	(-3.15)	(34.02)	(-27.02)	(-8.73)	(1.27)	(5.89)	0.917
JP-EW	1.69	2.75	-2.54	6.29	-0.46	0.66	0.964
	(1.75)	(26.13)	(-24.10)	(59.31)	(-4.41)	(6.26)	0.963
JP-VW	-1.57	2.97	-2.12	5.22	-0.15	1.69	0.827
	(-0.77)	(13.26)	(-9.45)	(23.13)	(-0.69)	(7.59)	0.822
FR-EW	-6.91	3.96	-2.56	1.62	0.16	2.23	0.349
	(-1.35)	(7.10)	(-4.58)	(2.88)	(0.29)	(4.01)	0.330
FR-VW	-9.63	4.25	-2.57	1.55	0.35	2.36	0.346
	(-1.78)	(7.22)	(-4.34)	(2.61)	(0.59)	(4.02)	0.327
W-EW	0.03	6.70	-1.18	0.94	0.24	0.15	0.997
	(0.11)	(226.40)	(-40.06)	(31.40)	(8.00)	(5.04)	0.997
W-VW	-5.05	4.93	0.31	0.78	1.22	2.15	0.942
	(-5.31)	(47.40)	(3.02)	(7.49)	(11.82)	(20.73)	0.941

Note: US, UK, JP, FR, and W denote United States, United Kingdom, Japan, France and International portfolios, respectively. EW denotes equal-weighted and VW denotes value-weighted.

[a] Unadjusted R^2.

[b] Adjusted R^2.

Table 8

Tests of H_0: $\alpha_i = 0$ for size portfolios in

$$R_i = \alpha_i + b_i P + \epsilon_i$$

and $R_i = \alpha_{iNJ} + \alpha_{iJ} D_J + b_i P + \epsilon_i$

1969 - 1983[a]

		CAPM-EW			CAPM-VW		
R_i	P	α_i=0	α_{iNJ}=0	α_{iJ}=0	αi=0	α_{iNJ}=0	α_{iJ}=0
US	US	2.28*	2.20*	8.46*	1.65	1.58	11.06*
		(.016)	(.020)	(.020)	(.095)	(.118)	(<.001)
JP	JP	1.00	1.09	2.06*	1.64	1.37	2.56*
		(.447)	(.374)	(.031)	(.100)	(.198)	(.007)
UK	UK	4.39*	3.95*	.57	4.71*	3.96*	1.64
		(<.001)	(<.001)	(.838)	(<.001)	(<.001)	(.100)
FR	FR	1.58	1.89*	1.45	1.58	1.89*	1.63
		(.115)	(.049)	(.155)	(.115)	(.049)	(.101)
US	Int'l	1.76	1.80	8.62*	1.51	1.57	10.93*
		(.072)	(.065)	(<.001)	(.139)	(.118)	(<.001)
JP	Int'l	1.61	1.57	2.71*	1.88	1.59	2.57*
		(.108)	(.120)	(.004)	(.051)	(.112)	(.006)
UK	Int'l	4.23*	3.92*	.65	4.65*	3.89*	1.65
		(<.001)	(<.001)	(.770)	(<.001)	(<.001)	(.096)
FR	Int'l	1.69	1.92*	1.17	1.63	1.92*	1.87
		(.086)	(.045)	(.316)	(.101)	(.045)	(.121)
Int'l	Int'l	3.28*	3.14*	5.88*	3.64*	3.16*	8.31*
		(<.001)	(<.001)	(<.001)	(<.001)	(<.001)	(<.001)

Table 8 (Cont'd)

		APT-5			APT-10		
R_i	P	$\alpha_i=0$	$\alpha_{iNJ}=0$	$\alpha_{iJ}=0$	$\alpha i=0$	$\alpha_{iNJ}=0$	$\alpha_{iJ}=0$
US	US	5.60* (<.001)	4.32* (<.001)	1.48 (.152)	6.37* (<.001)	4.26* (<.001)	1.51 (.140)
JP	JP	1.16 (.323)	1.32 (.226)	1.23 (.274)	1.28 (.248)	1.18 (.307)	1.11 (.359)
UK	UK	4.03* (<.001)	3.73* (<.001)	.69 (.788)	3.93* (<.001)	3.66* (<.001)	.50 (.865)
FR	FR	1.81 (.063)	2.08* (.028)	1.31 (.228)	1.95* (.042)	1.93* (.047)	.88 (.898)
US	Int'l	2.69* (.004)	2.30* (.015)	2.18* (.021)	3.33* (.001)	2.92* (.002)	1.82 (.061)
JP	Int'l	1.14 (.336)	.88 (.551)	.98 (.464)	1.35 (.205)	1.58 (.304)	.38 (.952)
UK	Int'l	3.94* (<.001)	3.67* (<.001)	1.05 (.401)	4.01* (<.001)	3.03* (<.001)	.46 (.915)
FR	Int'l	1.56 (.123)	1.79 (.066)	1.65 (.096)	1.53 (.135)	1.76 (.073)	1.68 (.089)
Int'l	Int'l	5.31* (<.001)	5.12* (<.001)	2.34* (.013)	5.79* (<.001)	5.39* (.134)	1.53 (<.001)

[a] MLR test statistics with p-values in parentheses. Under the null they have a central F distribution (degrees of freedom equal to 10 and 171 - k, where k is the number of non-constant regressors).

* Indicates significance at the 5% level.

Table 9

Mispricing tests results of the Zero-Beta models

1969-1983[a]

Panel A: CAPM models

Tests of H_0: $\alpha_i = (1-b_i)\lambda$ and H_0: $\alpha_{iNJ} = (1-b_i)\lambda$

for i = 1,..., 10 size portfolios

in $R_i = \alpha_i + b_i P + \varepsilon_i$ and $R_i = \alpha_{iNJ} + \alpha_{iJ} D_J + b_{iP} + \varepsilon_i$

		CAPM-EW		CAPM-VW	
R_i	P	$\alpha_i=(1-b_i)\lambda$	$\alpha_{iNJ}=(1-b_i)\lambda$	$\alpha_i=(1-b_i)\lambda$	$\alpha_{iNJ}=(1-b_i)\lambda$
US	US	22.94* (.006)	19.49* (.021)	19.40* (.022)	16.16 (.064)
JP	JP	10.39 (.320)	11.19 (.263)	13.63 (.136)	11.91 (.218)
UK	UK	17.96* (.036)	14.55 (.104)	25.99* (.002)	16.97* (.049)
FR	FR	15.94 (.068)	19.13* (.024)	15.33 (.082)	18.29* (.032)
US	Int'l	18.47* (.030)	17.48* (.042)	16.06 (.066)	15.70 (.073)
JP	Int'l	11.10 (.269)	7.34 (.602)	6.54 (.685)	7.04 (.633)
UK	Int'l	18.20* (.033)	16.25 (.062)	20.68* (.014)	17.63* (.040)
FR	Int'l	16.93 (.050)	19.97* (.018)	15.71 (.073)	19.18* (.024)
Int'l	Int'l	30.25* (<.001)	33.98* (<.001)	32.76* (<.001)	28.67* (<.001)

Table 9 (Cont'd)

Panel B: APT models

Tests of H_0: α_i = constant and H_0: α_{iNJ} = constant

for i = 1,..10 size portfolios

in $R_i = \alpha_i + b_i P + \varepsilon_i$ and $Ri = \alpha_{iNJ} + \alpha_{iJ} D_J + b_i P + \varepsilon_i$

		APT-5		APT-10	
R_i	P	α_i=constant	α_{iNJ}=constant	α_i=constant	α_{iNJ}=constant
US	US	52.38* (<.001)	42.18* (<.001)	60.26* (<.001)	42.48* (<.001)
JP	JP	12.17 (.204)	13.18 (.155)	13.90 (.126)	12.53 (.185)
UK	UK	39.46* (<.001)	31.16* (<.001)	39.66* (<.001)	37.46* (<.001)
FR	FR	16.43 (.058)	18.42* (.031)	15.09 (.088)	20.65* (.014)
US	Int'l	25.78* (.002)	22.92* (.006)	34.33* (<.001)	30.54* (<.001)
JP	Int'l	11.15 (.265)	9.02 (.435)	12.71 (.176)	10.29 (.328)
UK	Int'l	34.12* (<.001)	33.99* (<.001)	36.95* (<.001)	29.48* (<.001)
FR	Int'l	14.23 (.114)	20.07* (.017)	15.10 (.088)	20.65* (.014)
Int'l	Int'l	50.78* (<.001)	47.69* (<.001)	56.48* (<.001)	53.00* (<.001)

a Likelihood ratio test (p-values in parentheses). Statistics are asymptotically χ^2 with 9 degrees of freedom.

* indicates significance at the 5% level.

Table 10

Mispricing size

Average absolute mispricing of 10 size portfolios in

$$R_i = \alpha_i + b_i P + \varepsilon_i$$

$$\text{and } R_i = \alpha_{iNJ} + \alpha_{iJ} D_J + b_i P + \varepsilon_i$$

$$A = \Sigma\ |\alpha_i|/10 \quad \text{and } AJ = \Sigma\ |\alpha_{iJ}|/10$$

(per cent per annum)

1969-1983

		CAPM-EW		CAPM-VW		APT-5		APT-10	
R_i	P	A	AJ	A	AJ	A	AJ	A	AJ
US	US	1.40	34.49	4.76	72.74	2.58	5.21	2.63	4.69
JP	JP	1.98	15.04	13.88	35.05	1.81	5.43	1.78	5.76
UK	UK	6.14	3.18	12.63	73.15	3.31	2.14	3.35	2.36
FR	FR	2.21	8.70	3.15	21.89	2.46	9.44	2.14	7.53
US	Int'l	3.61	35.67	3.38	74.10	2.62	8.33	2.84	7.80
JP	Int'l	10.60	19.35	14.79	33.80	2.14	10.34	2.29	4.87
UK	Int'l	6.08	4.58	11.91	69.44	5.23	8.70	4.30	3.69
FR	Int'l	2.70	16.03	2.59	17.90	2.16	14.98	2.07	16.60
Int'l	Int'l	4.31	23.94	7.50	64.42	4.64	7.04	4.58	5.28

Table 11

Mispricing test results with capital control dummies

Tests of $H_0: \alpha_i = 0$; $H_0: \alpha_{i74} = 0$; $H_0: \alpha_{i79} = 0$

for size portfolios in $R_i = \alpha_i - \alpha_{i74}D_{74} - \alpha_{i79}D_{79} - b_iP - \epsilon_i$

and

Tests of $H_0: \alpha_{iNJ} = 0$; $H_0: \alpha_{iJ} = 0$; $H_0: \alpha_{iNJ74} = 0$; $H_0: \alpha_{iNJ79} = 0$

for size portfolios in $R_i = \alpha_{iNJ} - \alpha_{iJ}D_J - \alpha_{iNJ74}D_{74} - \alpha_{iNJ79}D_{79} - b_iP - \epsilon_i$

1969 - 1983[a]

		CAPM-EW							CAPM-VW						
R_i	P	α_i =0	α_{i74} =0	α_{i79} =0	α_{iNJ} =0	α_{iJ} =0	α_{iNJ74} =0	α_{iNJ79} =0	α_i =0	α_{i74} =0	α_{i79} =0	α_{iNJ} =0	α_{iJ} =0	α_{iNJ74} =0	α_{iNJ79} =0
US	US	4.47	33.42*	5.14	3.20	109.52*	38.57*	5.16	4.92	33.78*	5.07	3.17	120.52*	38.74*	5.11
		(.924)	(<.001)	(.881)	(.976)	(<.001)	(<.001)	(.880)	(.898)	(<.001)	(.886)	(.978)	(<.001)	(<.001)	(.884)
JP	JP	3.41	7.97	3.47	4.15	25.81*	8.18	3.90	2.91	12.22	4.38	3.85	28.09*	12.26	4.66
		(.970)	(.632)	(.968)	(.940)	(.004)	(.611)	(.952)	(.983)	(.270)	(.929)	(.934)	(.002)	(.268)	(.931)
UK	UK	12.42	13.55	10.72	12.33	22.51*	14.44	11.21	16.25	14.36	14.51	14.44	21.19*	14.18	14.45
		(.258)	(.195)	(.380)	(.264)	(.013)	(.154)	(.342)	(.093)	(.157)	(.151)	(.154)	(.020)	(.165)	(.153)
FR	FR	33.45*	9.76	29.73*	38.37*	24.59*	9.56	30.67*	32.29*	9.78	29.24*	36.86*	23.76*	9.57	29.94*
		(.001)	(.462)	(<.001)	(<.001)	(<.001)	(.480)	(<.001)	(<.001)	(.460)	(<.001)	(<.001)	(.008)	(.476)	(<.001)
US	Int'l	4.14	32.68*	5.81	3.33	96.76*	35.86*	5.80	4.73	34.47*	5.00	3.05	119.80*	39.53*	5.03
		(.941)	(<.001)	(.831)	(.973)	(<.001)	(<.001)	(.831)	(.908)	(<.001)	(.891)	(.980)	(<.001)	(<.001)	(.889)
JP	Int'l	2.85	13.48	4.50	4.17	31.06*	15.30	5.25	2.87	10.87	4.03	3.90	27.91*	11.24	4.36
		(.985)	(.198)	(.922)	(.910)	(<.001)	(.121)	(.873)	(.984)	(.368)	(.946)	(.952)	(.002)	(.339)	(.929)
UK	Int'l	11.56	13.99	10.52	12.11	7.66	14.71	10.87	11.97	13.32	10.89	11.80	18.02	13.85	11.30
		(.315)	(.173)	(.396)	(.278)	(.662)	(.143)	(.368)	(.287)	(.206)	(.366)	(.299)	(.055)	(.180)	(.334)
FR	Int'l	35.22*	10.95	30.18*	39.28*	17.49	10.49	31.13*	35.27*	10.10	30.73*	40.18*	22.53*	9.84	31.48*
		(<.001)	(.361)	(<.001)	(<.001)	(.064)	(.398)	(<.001)	(<.001)	(.432)	(<.001)	(<.001)	(.013)	(.455)	(<.001)
Int'l	Int'l	4.50	31.63*	7.37	6.88	64.19*	32.26*	8.37	5.96	30.24	6.00	7.02	95.76	32.35	7.25
		(.922)	(<.001)	(.690)	(.737)	(<.001)	(<.001)	(.512)	(.818)	(<.001)	(.816)	(.724)	(<.001)	(<.001)	(.702)

		APT-5							APT-10						
R_i	P	α_i =0	α_{i74} =0	α_{i79} =0	α_{iNJ} =0	α_{iJ} =0	α_{iNJ74} =0	α_{iNJ79} =0	α_i =0	α_{i74} =0	α_{i79} =0	α_{iNJ} =0	α_{iJ} =0	α_{iNJ74} =0	α_{iNJ79} =0
US	US	6.72	26.90*	14.08	4.75	15.01	25.20*	13.79	7.14	28.78*	14.14	4.83	15.26	27.08*	14.16
		(.751)	(.003)	(.170)	(.907)	(.132)	(.005)	(.182)	(.712)	(.001)	(.167)	(.902)	(.123)	(.003)	(.166)
JP	JP	3.40	9.38	4.46	3.89	14.42	10.14	4.49	4.59	8.61	5.61	4.48	13.64	3.59	5.70
		(.970)	(.496)	(.924)	(.952)	(.155)	(.428)	(.922)	(.917)	(.570)	(.847)	(.923)	(.190)	(.477)	(.840)
UK	UK	13.61	23.20*	9.02	13.19	8.14	23.72*	9.06	14.10	24.25*	10.21	13.83	6.23	24.42*	10.21
		(.191)	(.010)	(.530)	(.213)	(.616)	(.008)	(.526)	(.169)	(.007)	(.423)	(.181)	(.796)	(.007)	(.423)
FR	FR	21.79*	12.03	25.42*	23.91*	14.72	11.96	25.49*	22.38*	10.10	25.37*	24.25*	10.10	10.11	25.39*
		(.016)	(.283)	(.005)	(.008)	(.143)	(.288)	(.004)	(.013)	(.432)	(.005)	(.007)	(.432)	(.430)	(.005)
US	Int'l	4.12	28.64*	7.41	3.13	22.52*	26.57*	7.12	5.09	33.17*	10.27	3.24	18.68*	30.46*	9.54
		(.942)	(.001)	(.686)	(.978)	(.013)	(.003)	(.714)	(.885)	(<.001)	(.417)	(.975)	(.045)	(<.001)	(.482)
JP	Int'l	3.34	26.29*	4.78	3.74	10.50	26.29*	5.06	2.91	20.12*	3.57	2.43	8.24	22.09*	3.58
		(.972)	(.003)	(.905)	(.959)	(.398)	(.003)	(.887)	(.983)	(.028)	(.965)	(.992)	(.795)	(.015)	(.964)
UK	Int'l	12.53	16.37	10.18	12.06	15.24	19.36*	10.03	13.59	22.98*	12.25	12.54	5.56	23.38*	12.42
		(.251)	(.089)	(.425)	(.281)	(.123)	(.036)	(.438)	(.193)	(.011)	(.269)	(.181)	(.796)	(.007)	(.423)
FR	Int'l	34.47*	11.87	29.90*	42.22*	20.20*	11.52	32.55*	31.71*	9.73	27.58*	39.30*	18.13	9.61	30.21*
		(<.001)	(.294)	(.001)	(<.001)	(.027)	(.318)	(<.001)	(<.001)	(.464)	(.002)	(.007)	(.098)	(.475)	(<.001)
Int'l	Int'l	5.73	33.49*	8.48	8.55	22.82*	30.81*	8.55	4.59	32.67*	10.82	6.12	15.12	30.19*	10.32
		(.837)	(<.001)	(.583)	(.575)	(.011)	(<.001)	(.575)	(.917)	(<.001)	(.372)	(.805)	(.128)	(<.001)	(.412)

[a] Wald test (p-values in parentheses). Statistics are asymptotically χ^2 with 10 degrees of freedom.

* Indicates significance at the 5% level.

Table 12

Mispricing tests results of the Zero-Beta models with capital control dummies

1969-1983[a]

Panel A: CAPM models

Tests of H_0: $\alpha_i=(1-b_i)\lambda$ and H_0: $\alpha_{iNJ}=(1-b_i)\lambda$

for i = 1,..,10 size portfolios

in $R_i = \alpha_i + \alpha_{i74}D_{74} + \alpha_{i79}D_{79} + b_iP + \varepsilon_i$

and $R_i = \alpha_{iNJ} + \alpha_{iJ}D_J + \alpha_{iNJ74}D_{74} + \alpha_{iNJ79}D_{79} + b_iP + \varepsilon_i$

		CAPM-EW		CAPM-VW	
R_i	P	$\alpha_i=(1-b_i)\lambda$	$\alpha_{iNJ}=(1-b_i)\lambda$	$\alpha_i=(1-b_i)\lambda$	$\alpha_{iNJ}=(1-b_i)\lambda$
US	US	8.44 (.490)	6.12 (.728)	5.12 (.824)	3.15 (.958)
JP	JP	2.01 (.991)	2.00 (.991)	2.07 (.990)	3.24 (.954)
UK	UK	10.37 (.321)	10.08 (.344)	15.61 (.075)	13.59 (.137)
FR	FR	31.89* (<.001)	29.34* (<.001)	32.07* (<.001)	36.01* (<.001)
US	Int'l	4.03 (.909)	3.09 (.961)	4.42 (.882)	2.96 (.966)
JP	Int'l	2.43 (.982)	3.48 (.942)	2.54 (.980)	3.70 (.930)
UK	Int'l	10.23 (.332)	10.86 (.285)	11.47 (.245)	11.33 (.254)
FR	Int'l	29.54* (<.001)	27.85* (.001)	32.41* (<.001)	35.44* (<.001)
Int'l	Int'l	4.29 (.891)	6.37 (.702)	5.83 (.756)	6.77 (.661)

Table 12 (Cont'd)

Panel B: APT models

Tests of H_0: α_i = constant and H_0: α_{iNJ} = constant

in $R_i = \alpha_i + \alpha_{i74}D_{74} + \alpha_{i79}D_{79} + b_iP + \varepsilon_i$

and $R_i = \alpha_{iNJ} + \alpha_{iJ}D_J + \alpha_{iNJ74}D_{74} + \alpha_{iNJ79}D_{79} + b_iP + \varepsilon_i$

		APT-5		APT-10	
R_i	P	α_i=constant	α_{iNJ}=constant	α_i=constant	α_{iNJ}=constant
US	US	5.52 (.787)	3.90 (.918)	6.65 (.673)	4.39 (.883)
JP	JP	3.36 (.948)	3.72 (.929)	4.51 (.875)	4.93 (.840)
UK	UK	13.10 (.158)	12.67 (.178)	13.54 (.140)	13.30 (.150)
FR	FR	20.46* (.015)	22.38* (.008)	20.90* (.013)	22.52* (.007)
US	Int'l	3.62 (.934)	2.84 (.256)	4.42 (.882)	2.98 (.965)
JP	Int'l	2.85 (.970)	3.30 (.951)	2.65 (.977)	2.39 (.984)
UK	Int'l	10.76 (.292)	10.69 (.297)	12.33 (.195)	11.72 (.230)
FR	Int'l	29.43* (<.001)	35.10* (<.001)	27.96* (<.001)	34.79* (<.001)
Int'l	Int'l	4.71 (.859)	6.28 (.712)	4.41 (.816)	5.85 (.799)

a Likelihood ratio test (p-values in parentheses). Statistics are asymptotically χ^2 with 9 degrees of freedom.

* indicates significance at the 5% level.

Table 13

Mispricing size with capital control dummies

Average absolute mispricing of 10 size portfolios in

$$R_i = \alpha_i + b_i P + \varepsilon_i$$

and $R_i = a_{iNJ} + \alpha_{iJ} D + \alpha_{iNJ74} D_{74} + \alpha_{iNJ79} D_{79} + b_i P + \varepsilon_i$

$$A = \Sigma\ |\alpha_i|/10 \text{ and } AJ = \Sigma\ |\alpha_{iJ}|/10$$

(per cent per annum)

1969-1983

		CAPM-EW		CAPM-VW		APT-5		APT-10	
R_i	P	A	AJ	A	AJ	A	AJ	A	AJ
US	US	12.71	77.41	14.83	86.35	1.30	4.95	1.23	4.72
JP	JP	1.22	29.02	6.56	33.94	1.58	5.43	1.54	5.90
UK	UK	3.19	21.57	6.17	79.41	2.27	2.17	2.29	2.42
FR	FR	12.82	23.24	13.18	21.82	4.77	9.15	4.48	7.47
US	Int'l	5.69	36.61	7.90	69.77	1.60	8.48	1.93	8.12
JP	Int'l	1.35	19.67	2.99	28.32	1.71	9.49	1.93	6.95
UK	Int'l	2.29	4.26	5.23	63.15	2.21	12.23	2.48	3.38
FR	Int'l	16.82	16.80	15.71	14.31	16.46	31.60	9.68	23.60
Int'l	Int'l	.90	23.68	3.33	58.56	1.54	6.28	1.33	3.32

Figure 1

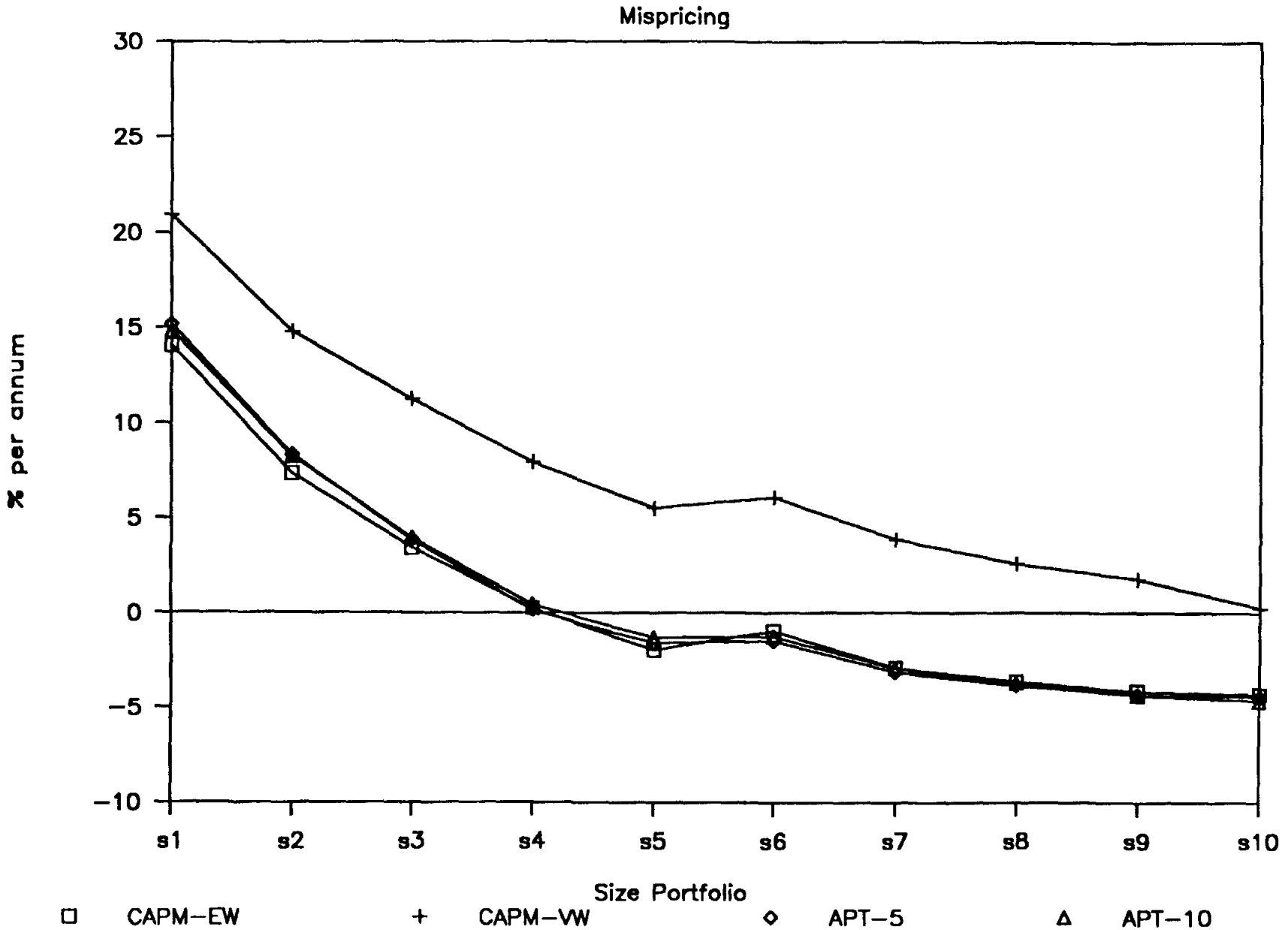

Mispricing
30
25
20
15
10
5
0
−5
−10
% per annum
s1
s2
s3
s4
s5
s6
s7
s8
s9
s10
Size Portfolio
CAPM−EW
CAPM−VW
APT−5
APT−10

Figure 2

January Mispricing

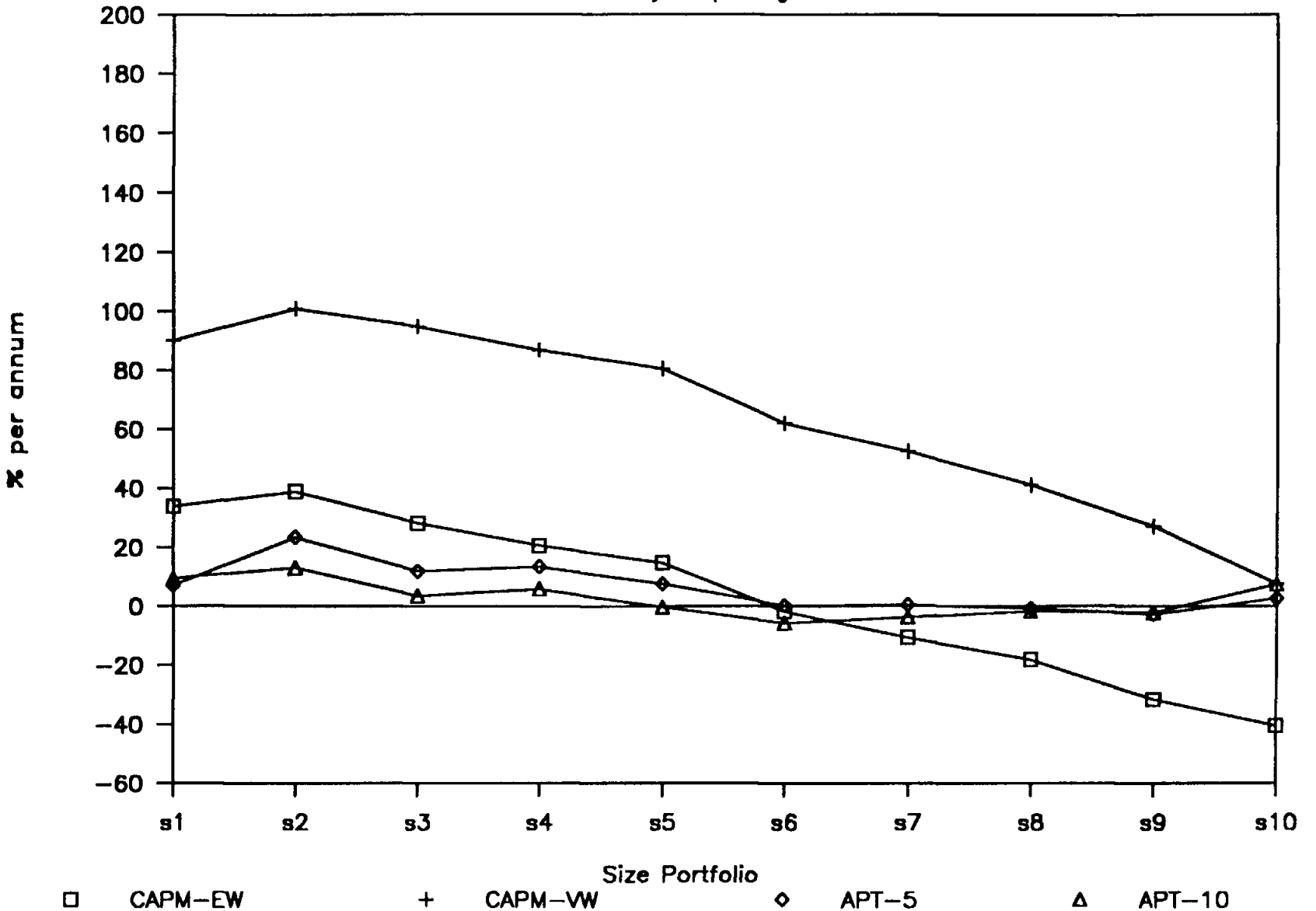


Figure 3

Mispricing -- Capital Controls

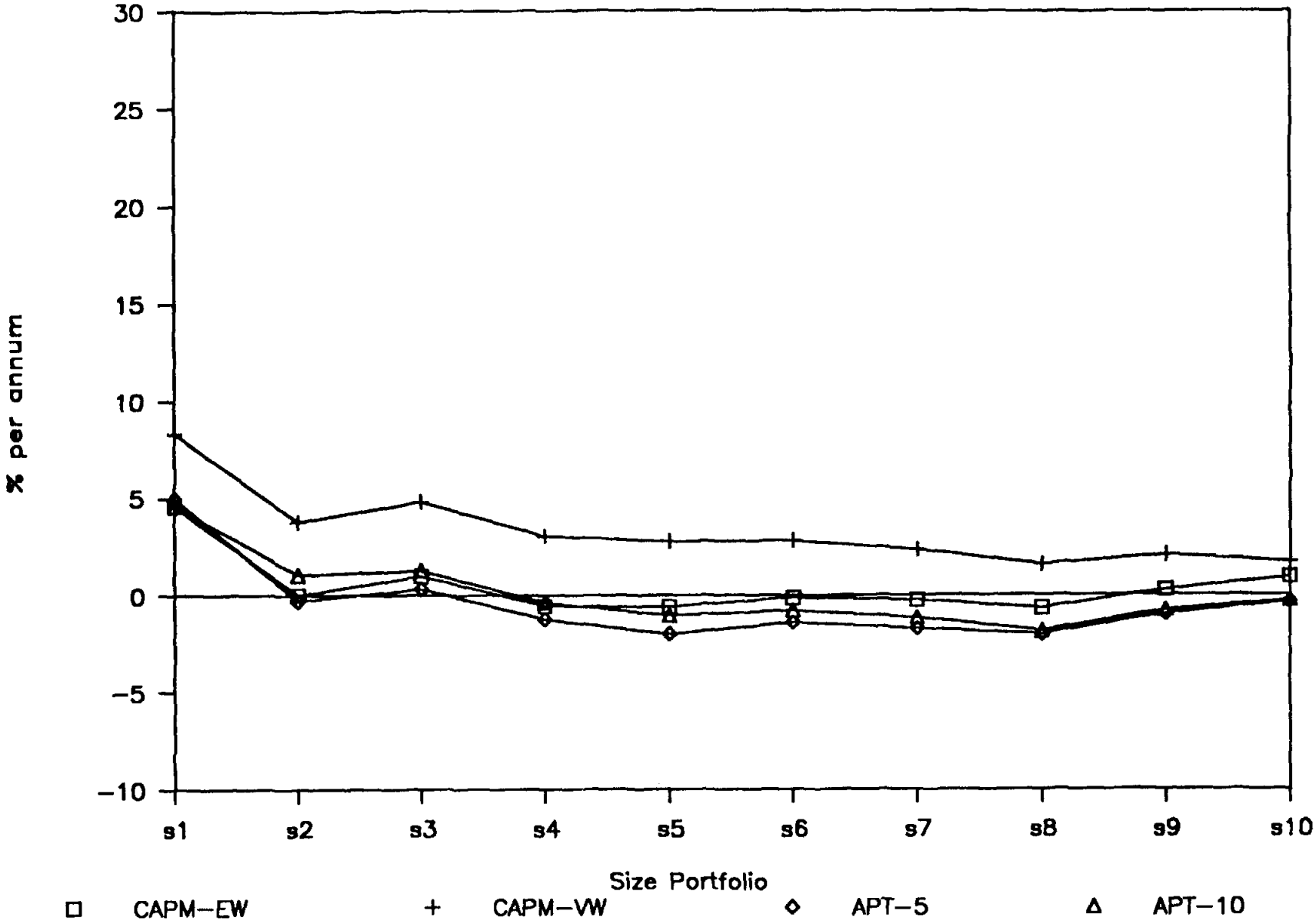
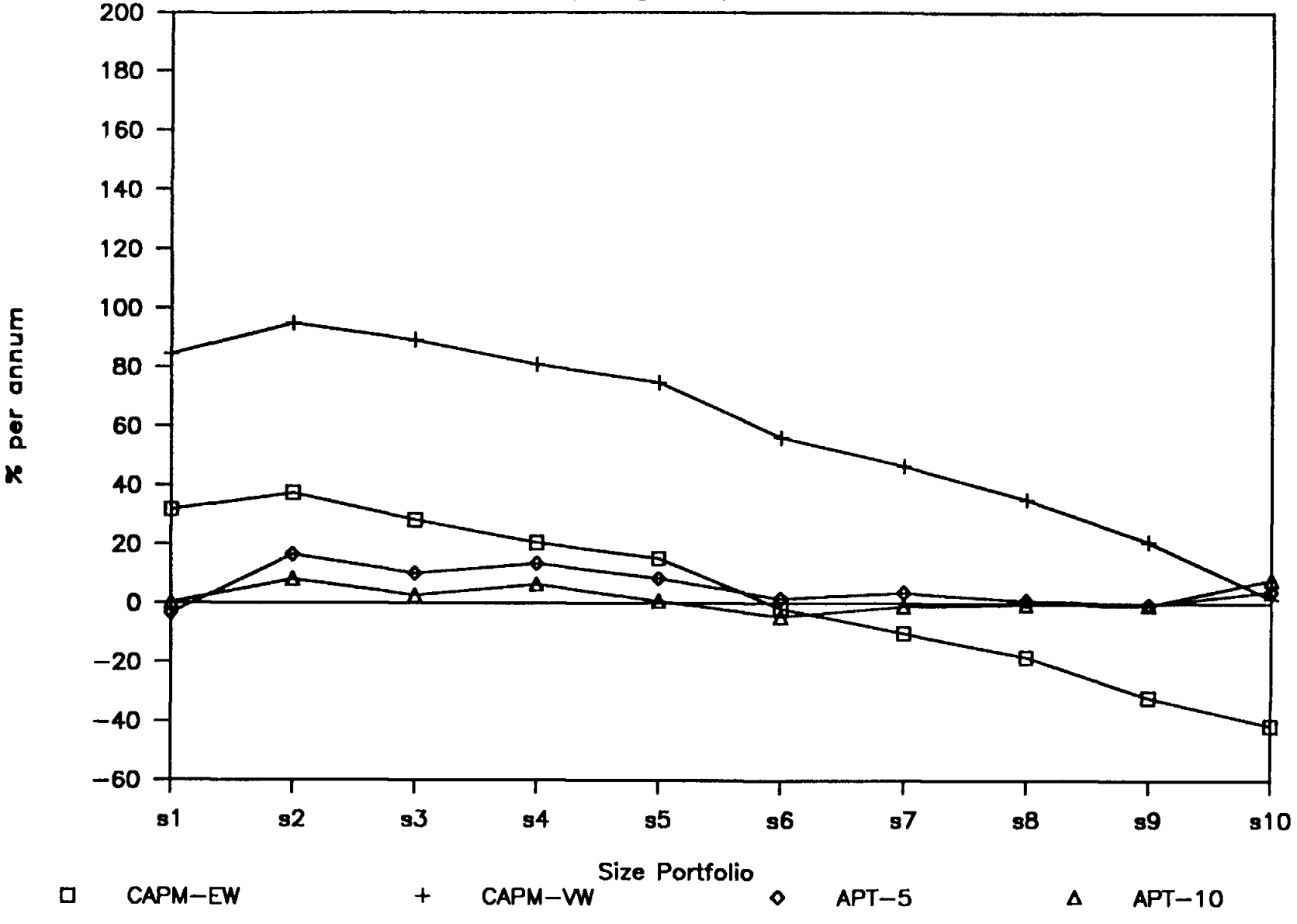


Figure 4

Jan Mispricing – Capital Controls



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