"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 their ability to explain seasonality in asset returns. We also find that the behavior of the models is affected by changes in the regulatory environment in international markets.

In this paper we evaluate 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 the Arbitrage Pricing Theory (APT) where the pervasive factors are estimated by an asymptotic principal components technique.

We focus on three questions. First, we investigate whether the APT has greater explanatory power than 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, which covers the period 1969-1983, uses a large number of securities from the United States, Japan, the United Kingdom, and France both for factor estimation and hypothesis testing.

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]

In related work 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, Gultekin, and Penati (1987) but fewer than Cho, Eun, and Senbet (1986). However, for the countries we study, 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 2 we describe the data. The techniques used to estimate the pervasive factors and test the alternative models are described in Section 3, and the empirical results are given in Section 4. Section 5 comprises a summary and conclusions.

1. Alternative Asset Pricing Models

We investigate the pricing performance of domestic and international versions of the CAPM and APT. The CAPM or the APT imply that a particular benchmark portfolio or linear combination of a group of benchmark portfolios lies on the minimum variance boundary of risky assets [e.g., see Roll (1977) or Huberman and Kandel (1987)]. 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 domestic 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 them.

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:

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

where $\tilde{r}_i$ and $\tilde{r}_M$ are the real returns on asset i and the market portfolio and $\tilde{R}_i$ 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 (although hedging against shifts in the consumption-investment opportunity set is not peculiar to international models). Under admittedly restrictive conditions there will be 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, amount to tests of mean/variance efficiency for these proxies.

As in many empirical investigations, we use the return on short-term U.S. treasury bills as a proxy for the riskless rate of interest. 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

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

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

As an alternative to the CAPM we consider the APT. An assumption underlying the APT is that asset returns follow a factor model:

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

where b_{ij} is the sensitivity (beta) of asset i relative to factor j and $E(\tilde{f}_j) = E(\tilde{\varepsilon}_i) = E(\tilde{f}_j\tilde{\varepsilon}_i) = 0$ 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 eliminated in large portfolios.[5] Lack of arbitrage opportunities and existence of a riskless asset imply that

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

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. Also, the pervasive components of exchange rate risks will be reflected in the returns on our factor mimicking portfolios. We also estimate a zero-beta version of (3) which we discuss in more detail below.

In table 1 we present the particular models investigated. Two of them, the CAPM-EW and the CAPM-VW, are models in which the benchmark portfolios are equal-weighted and value-weighted portfolios of common stocks, 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.

2. Data Sources

The selected countries, markets, and data sources are presented in Table 2. We were able to obtain monthly stock return 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 is 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, are 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 U.S. treasury bill return.[6] We perform our tests on both nominal and real returns. Nominal dollar returns are converted into real returns using inflation calculated at the percentage change in the U.S consumer price index. The treasury bill returns and inflation series are from Ibbotson Associates (1985).

3. Estimation of Pervasive Economic Factors and Hypothesis Tests

3.1 Estimation of Pervasive Factors

Our tests of the CAPM amount to specifying 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 this procedure is its ability to 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 securities 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.

We 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. We assume that our sample size is sufficiently large 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 of monthly data). For the domestic versions of the APT Ω^n is calculated over the assets traded on the domestic stock exchange and, for the international versions, over the entire sample. 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 by any survivorship bias.

A difficulty which arises in any application of the APT is the choice of 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 former case. We report the results of two tests, each of which takes a very different approach to the problem.

The asymptotic principal components procedure provides us with excess returns on factor mimicking portfolios. We will consider the problem of testing a k_1 factor model versus a k_2 factor model ($k_2 > k_1$). The first test is suggested by Kandel and Stambaugh (1987). It is based on the observation that if a k_1 factor model actually describes cross-sectional expected returns then the expected excess returns on the remaining $k_2 - k_1$ factor mimicking portfolios should be described by the APT pricing relation, (3), using the first k_1 factors. This test is more stringent than most approaches to determining the number of factors. While most tests only examine whether additional factors have explanatory power in time series, the approach of Kandel and Stambaugh (1987) tests whether the additional factors have risk which is not already priced by the first k_1 factors. Factors which have time series explanatory power but which have a zero risk premium (i.e., $\gamma_j = 0$ for that factor) will not be identified as factors by this test.

Let P_{1t} denote the $k_1 \times 1$ vector of period t excess returns on the first k_1 factors and P_{2t} denote the $p \times 1$ vector ($p = k_2 - k_1$) of period t excess returns on the remaining factors. The null hypothesis that k_1 factors are sufficient [from (3)] implies that the $p \times 1$ vector of intercepts in a multivariate regression of P_2 on P_1 are equal to zero. That is, $a = 0$ in

$$P_{2t} = a + \beta P_{1t} + \eta_t. \tag{5}$$

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

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

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 hypothesis, the MLR statistic has an F distribution with degrees of freedom equal to p and $T - k_1 - p$. 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 test results, which are reported in Table 3, 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 hypothesis 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. This test is different from the above approach in that it uses the usual criterion for pervasiveness, time series explanatory power, and does not rely on pricing restrictions. Our test relies on a result from Ingersoll (1984) which states that the cross-sectional mean-square of assets' betas relative to a <u>non-pervasive</u> factor must approach zero as n approaches infinity. That is, if the k^{th} 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 k^{th} 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 of 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 k^{th} factor. The test might indicate too few factors because it tests a necessary condition for non-pervasiveness. That is, 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). On the other hand, the test might indicate too many factors, in small samples, 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 results of the first test.

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.

Conflicting evidence about the number of factors is common in the empirical literature on the APT [e.g., Roll and Ross (1980), Dhrymes, Friend, and Gultekin (1984), and Trzcinka (1986)]. Since our tests do not provide an unambiguous picture, we use other grounds to choose the number of factors. If we wish to allow for the possibility that movements in exchange rates are a source of non-diversifiable risk, then it seems reasonable to allow for at least four factors. The four factors might represent general market risk plus the risks associated with shifts in the three relative prices of the four currencies.[8] 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 and ten factors.

3.2 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

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

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.[9] 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}.[10] The hypotheses regarding α_i, α_{iNJ}, and α_{iJ} [e.g., as in (8)] are tested using the MLR statistic described in (6) 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 - 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. Let $\tilde{r}_{zt}$ denote the return on a portfolio with zero covariance with the market. We assume that the expected excess zero-beta return, $\lambda = E(\tilde{r}_{zt}) - r_{Ft}$, is constant through time.

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

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

This implication of the zero-beta CAPM is discussed in Black, Jensen, and Scholes (1972, eqn. 14) in a single-equation context. Gibbons (1982) derives and tests the non-linear cross-equation restrictions implied by the zero-beta CAPM. Our restrictions in (10) are of the same form as those tested in Gibbons (1982), although the interpretation is slightly different.[11]

There are a variety of asymptotically equivalent test statistics for the hypothesis (10). We use the likelihood ratio (LR) test which has a χ^2 distribution, asymptotically, with degrees of freedom equal to m - 1 [see Gibbons (1982) and Gallant (1987, pp. 457-8)]. 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 LR test tends to reject the null hypothesis too often in small samples. This is particularly true when the number of cross-sections, m, is large relative to the number of time-series observation, T, as is shown in Stambaugh (1982), Shanken (1985), and Amsler and Schmidt (1985). We have also calculated the cross-sectional regression test (CSRT) statistic suggested in Shanken (1985) for the null hypothesis in (10). An approximate small sample distribution of this statistic is given by the Hotelling T^2 distribution. This approximation is better than the asymptotic χ^2 approximation [Shanken (1985) and Amsler and Schmidt (1985)]. For our sample, we find that there is very little difference between the inferences one would draw based on the LR test and the CSRT. Because of this, we only report the LR tests. The small difference between the two statistics is due mainly to the fact that our time-series sample of 180 observations is large relative to our cross-section of 10 assets (portfolios) and is consistent with the simulation results of Amsler and Schmidt (1985).

We also test the restrictions implied by the zero-beta version of the APT. We obtain particularly simple restrictions if we assume that our proxy for the riskless asset bears only factor related risk.[12] In the Appendix 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. Thus, mispricing relative to the zero-beta APT is measured by $(\alpha_i - \lambda)$.

Our hypotheses about α_i in (7) and (10) are tests of unconditional mean/variance efficiency of the benchmark portfolios. When we allow the mispricing parameters to be seasonal we are testing a particular form of conditional mean/variance efficiency. 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.[13] Hansen and Richard (1987) show that unconditional efficiency implies conditional efficiency but that the converse is not true.[14] 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. Therefore, it may be possible to reject unconditional or limited conditional efficiency when the benchmarks are efficient conditional on the full information set.

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

3.3 Tests of the Effects of Capital Controls

The regulatory environment of international financial markets is likely to be an important determinant of capital market integration and asset pricing. Over our sample period there is a general trend towards deregulation marked by two major periods of change.[15] The first period of change 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 important period is 1979 when the United Kingdom and Japan dismantled a number of controls.[16]

We investigate whether periods of more strict controls (ending in January 1974 and November 1979, respectively) are associated with greater deviations from the predictions of the asset pricing models than are periods of less stringent control. 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

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

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

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.[17] The use of dummy variables is an admittedly crude method of measuring the effects of capital controls. However, in the absence of a finer metric for the severity of controls, the dummy variable approach is a reasonable alternative. If the loosening of capital controls leads to a more integrated global market we would expect that the performance of purely domestic models would deteriorate and the performance of international models improve in the periods with fewer controls.

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

3.4 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.[18] The instrument used to form portfolios should be chosen to ensure heterogeneity across portfolios. The instrumental variable chosen here is the "size" of the firm.[19] We form five sets of ten 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).

4. 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 models 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.

4.1 The Structure of Factor Returns

Before we proceed with the formal hypothesis tests, we discuss some evidence on the covariance structure of asset returns across countries and present evidence on the relation between market indices and our estimated factors. The correlations across national common stock portfolios 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).

The international factors are estimated by the asymptotic principal components procedure from returns on every available asset. Over our 180 month period, the monthly average number of firms with returns data is 5596. Regressions of the excess returns of the national indices and percentage changes in exchange rates on the first five international factors are reported in Table 6. The results 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. These results and some extensive canonical correlation analysis not reported here, indicate that there are several common international factors. The estimated mispricing of each index relative to these five benchmark portfolios (in % per annum) is listed in the second column. The estimated values of mispricing for the French indices, $\hat{\alpha}_{FR}$, 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.

The regressions of changes in exchange rates on the factors indicate that exchange rates are related to the pervasive sources of risk in the equity markets. For each of the three exchange rates there is a statistically significant (at the 5% level) relation with four of the five factors. The factors explain between thirty and fifty three percent of exchange rate variability. Thus, exchange rate risk is, in part, pervasive and is reflected in the estimated factors.

4.2 Multi-Index versus Single-Index Models

In this section we compare the performance of multi-index and single index models using two criteria: first, whether or not the tests described in Section 3.2 reject the restrictions implied by the models and, secondly, whether the magnitudes of mispricing differ across models. The second criterion is useful in view of 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 U.S. 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 excess returns of size portfolios on particular benchmark portfolio returns. 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 7 we present the results of the tests.

Each model has 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 U.K. 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). 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. Also, the pattern of rejections for the U.S. sample with domestic benchmarks is basically the same as that found by Connor and Korajczyk (1988a) and Lehmann and Modest (1988).

Test results for the zero-beta specifications of the models are presented in Table 8. With few exceptions (CAPM-EW model for the U.K. and the domestic APT-10 for France), whenever the null is rejected with the U.S. T-Bill rate as the zero-beta return, we also reject the zero-beta variant of the model. Thus, the rejections do not seem to be driven by our choice of the U.S. T-bill return as the zero-beta return.

The tests reported in Tables 7 and 8 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 figures corresponding to each possible permutation. We chose only four graphs which, along with Table 9, 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. A comparison of mispricing across countries is given in Figure 2 (CAPM using the value-weighted international market) and in Figure 3 (five factor international APT). The U.K. shows the strongest size effect and France the weakest.[20] Again, the patterns of mispricing for the U.S. are similar to those shown in Figures 1-3 of Connor and Korajczyk (1988a), although the levels of mispricing are slightly smaller. This may be due to the fact that we include all assets in our size portfolios while Connor and Korajczyk use a sample in which firms are required to have a continuous trading history over five year intervals. In Table 9 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 in Figure 4. As with average mispricing, January mispricing of small size international portfolios is greater than for large size ones. This finding also holds at the country level with the U.S. 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 9, 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 especially the APT-10 model show a much smaller mispricing than the CAPM-EW model.

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 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.

4.3 Domestic versus International Benchmarks

Most empirical studies of asset pricing models use securities and benchmark portfolios from a single country. While there undoubtedly exist some barriers to international investing, one might expect that increasing global diversification would lead to a greater role of international factors in asset pricing. In order to compare domestic versus international models, we consider the mispricing of the domestic size portfolios relative to the domestic and international benchmark portfolios, respectively. In terms of the frequency of rejection, there is not a clear difference between the use of domestic versus international benchmark portfolios. From Tables 7 and 8 the models with international benchmarks are rejected slightly less often than the domestic models. This lower frequency of rejection of the models with international benchmarks could be due to smaller levels of mispricing or to lower power of the tests. In Table 9 we provide estimates of the average absolute pricing error across models. On the basis of the magnitude of the estimated mispricing it appears that the domestic versions marginally outperform the international versions, except for the CAPM-VW model. Thus, at this stage the evidence does not unambiguously support the use of domestic or international benchmarks.

4.4 Effect of Regulatory Changes in International Financial Markets

4.4.1 Effect on Model Performance

If the changes in regulations do not influence international asset pricing, then we would expect the regime shift coefficients [α_{i74} and α_{i79}] in equation (11) to equal zero.

Table 10, which is comparable to Table 7, reports the results of allowing mispricing to be regime dependent. The hypotheses $\alpha = 0$ and $\alpha_{NJ} = 0$ are not rejected for any 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, which are quite different from those shown in Table 7 when no adjustment was made for changes in the international financial markets, are consistent with international regulatory influences on asset pricing.

The statistical significance of α_{i74} differs between the CAPM and APT. The coefficient tends to be significant for the APT 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 model results 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 U.S., and liberalization of capital controls on the part of the other countries. The performance of the models does not seem to have been affected by the changes in 1979.

We present in Table 11 the tests of the zero-beta variants of the models. They are similar to those obtained when the U.S. 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 8 which do not allow for regime shifts in mispricing.

Because of the many different types of changes in international financial markets around 1974, it is not possible to attribute the apparent shift in pricing to particular changes in regulations or to the switch to floating from fixed exchange rates. Indeed, there may be other external reasons for the period-specific levels of mispricing since the purely domestic models also tend to perform well after adjusting for regime shifts. However, the fact that there are significant changes in the pricing of assets, relative to our benchmarks, around these regime shifts may indicate the importance of capital controls and exchange rate regimes for asset pricing.

4.4.2 Multi-Index versus Single-Index Models Adjusting for Regime Shifts

The above results indicate that the APT tends to be rejected less often than the CAPM, especially relative to January mispricing. Figures 5 and 6 show our estimates of mispricing and January mispricing of international size portfolios 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 U.S. exhibits the strongest size effect and Japan and

France the weakest. In none of the four countries is there any noticeable January effect for the APT. Estimates of average absolute mispricing are presented in Table 12. 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 9, the APT's mispricing is systematically lower, except for France, while their January mispricing is comparable, again, except for France.

4.4.3 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 10). We also find that the CAPM has smaller pricing errors in its international version than in its domestic version (see Table 12).

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 12 the domestic versions of the APT tend 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 causes 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.

5. 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

Let $\tilde{r}_{zt}$ denote the return on a portfolio which has zero covariance with the benchmark portfolios. In most cases (particularly international models) the U.S. treasury bill return is theoretically 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} + γ_{jt}). Although r_{Ft} is riskless in nominal U.S. 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 to obtain consistent estimates of $(f_{jt} + \gamma_{jt})$.

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 $r_{\delta t}$ is well diversified and that calculating excess returns with respect to $r_{\delta t}$ does not alter the basic nature of the factor structure (i.e., calculating 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).

Under 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.

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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 (1984) find that the APT performs significantly better than the CAPM in explaining asset returns. Similar results are found by Dumontier (1986) using French stocks and Hamao (1986) using Japanese stocks.
3. The minimum number of firms from our four countries is 4211 while the maximum number is 6692. The number of securities used in Cho, Eun, and Senbet (1986), by country, are US (60), Japan (55), UK (48), and France (24) while the numbers used in Gultekin, Gultekin, and Penati (1987) are 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 real relative prices. See, for example, Solnik (1974, pp. 514-18) and Stulz (1985 pp. 77-79).

5. Ross (1976) assumes that $E(\tilde{\epsilon}_i \tilde{\epsilon}_j) = 0$ (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 also test versions of the models denominated in each of the other three currencies. Excess returns are calculated relative to the short term interest rates prevailing in each of the countries' currencies which are obtained from the International Financial Statistics tables. We find that the test results are not significantly affected by the currency chosen. As a result we only present results using the US dollar as numeraire.

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

8. Of course this ignores the fact that exchange rate movements relative to currencies not in our sample might also be a source of pervasive risk. We are only suggesting a lower bound.

9. 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.

10. 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:

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

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).

11. Gibbons uses returns rather than excess returns. In his case λ would equal $E(\bar{r}_z)$. Since we use excess returns, λ should be interpreted as $E(\bar{r}_{zt}) - r_{Ft}$. Note that mispricing relative to the zero-beta CAPM is measured by $[\alpha_i - (1 - b_i)\cdot\lambda]$.
12. Given that our proxy for the riskless asset is only riskless in nominal \$US returns, it is not necessarily a zero-beta asset in real terms. Thus, it is likely that the return on a portfolio with zero betas with respect to the international factors will differ from our T-bill return.
13. We will refer to these as unconditional, limited conditional, and conditional efficiency, respectively.
14. The same logic shows that unconditional efficiency implies limited conditional efficiency but that the converse is also not true.
15. In an unpublished appendix (available from the authors) we give a brief description of changes in capital controls over our sample period.
16. In Japan, deregulation measures, announced in early 1979, were implemented in 1980.
17. We also have estimated models which allow the sensitivities, b_i, to be period dependent. The results are essentially the same as those with constant sensitivities.
18. The implications of this tradeoff, in terms of the power of the tests, is analyzed in Gibbons, Ross, and Shanken (1986).

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

20. 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.

Table 1
Models and versions tested.

MODELS			
CAPM-EW	CAPM-VW	APT-5	APT-10

VERSIONS					
Domestic/Domestic		Domestic/Int'al		Int'al/Int'al	
R_i	P	R_i	P	R_i	P
US	US	US	Int'al	Int'al	Int'al
JP	JP	JP	Int'al		
UK	UK	UK	Int'al		
FR	FR	FR	Int'al		

Models are tested by estimating the mispricing of size ranked portfolios relative to various benchmark portfolios. CAPM-EW and CAPM-VW correspond, respectively, to the use of equal-weighted and value-weighted equity portfolios as the benchmarks. APT-5 and APT-10 use 5 and 10 factor mimicking portfolios estimated by the asymptotic principal components procedure as benchmarks. Versions of the model are distinguished by the markets from which the size based portfolios and benchmark portfolios are constructed. R_i identifies the source of the size based portfolios while P identifies the source of the benchmark portfolios. US: United States, JP: Japan, UK: United Kingdom, FR: France, Int'al: all four countries. Zero-beta variants of each model are also tested as are variants which use nominal and real returns.

Table 2
Exchange market data and sample data summary.

COUNTRY	UNITED STATES	JAPAN	UNITED KINGDOM	FRANCE	TOTAL
Stock Exchange	NYSE & AMEX	TOKYO	LONDON	PARIS	
Market Capitalization (12/83)					
World Capitalization	43%	15%	6.1%	1%	65.1%
Number of listed firms (12/83)	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	121	5596

Source of market capitalization percentages and number of listed firms December 1983 - International Federation of Stock Exchange Statistics, 1983.

Table 3
Tests of k_1 factors versus the alternative of k_2 factors based on the mean/variance efficiency of k_1 factor portfolios relative to k_2 factor portfolios.

k_1	k_2	P-values US	UK	Japan	France	International
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

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 equal zero. Factors are estimated by asymptotic principal components using monthly data from January 1969 through December 1983. P-values are the right tail area of the modified likelihood ratio (MLR) statistic for restriction that intercepts equal zero. * denotes significance at the 5% level.

Table 4
Test of k_1 factors versus the alternative of k_2 factors based on the time-series explanatory power of the additional k_2-k_1 factor portfolios.

		P-values				
k_1	k_2	US	UK	Japan	France	International
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*

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). Factors are 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
Summary of parameter restrictions implied by asset pricing models.

Panel A: **Regressions not adjusting for changes in capital controls.**

Null	Regression	Model
1 $\alpha_i=0$	$R_i=\alpha_i+b_iP+\varepsilon_i$	CAPM, APT
2 $\alpha_{iJ}=0$; $\alpha_{iNJ}=0$	$R_i=\alpha_{iNJ}+\alpha_{iJ}D_J+b_iP+\varepsilon_i$	CAPM, APT
3 $\alpha_{iJ}=0$; $\alpha_{iNJ}=0$	$R_i=\alpha_{iNJ}+\alpha_{iJ}D_J+b_{iJ}D_JP+b_{iNJ}P+\varepsilon_i$	CAPM, APT
4 $\alpha_i=(1-b_i)\lambda$	$R_i=\alpha_i+b_iP+\varepsilon_i$	CAPM Zero-b
5 $\alpha_i=\lambda$	$R_i=\alpha_i+b_iP+\varepsilon_i$	APT Zero-b
6 $\alpha_{iNJ}=(1-b_i)\lambda$	$R_i=\alpha_{iNJ}+\alpha_{iJ}D_J+b_iP+\varepsilon_i$	CAPM Zero-b
7 $\alpha_{iNJ}=\lambda$	$R_i=\alpha_{iNJ}+\alpha_{iJ}D_J+b_iP+\varepsilon_i$	APT Zero-b
8 $\alpha_{iNJ}=(1-b_i)\lambda$	$R_i=\alpha_{iNJ}+\alpha_{iJ}D_J+b_{iJ}D_JP+b_{iNJ}P+\varepsilon_i$	CAPM Zero-b
9 $\alpha_{iNJ}=\lambda$	$R_i=\alpha_{iNJ}+\alpha_{iJ}D_J+b_{iJ}D_JP+b_{iNJ}P+\varepsilon_i$	APT Zero-b

Table 5 (Cont'd)

Panel B: Regressions adjusting for changes in capital controls.

	Null	Regression	Model
10	$\alpha_i=0$; $\alpha_{i74}=0$; $\alpha_{i79}=0$	$R_i=\alpha_i+\alpha_{i74}D_{74}+\alpha_{i79}D_{79}+b_iP+\varepsilon_i$	CAPM, APT
11	$\alpha_{iNJ}=0$; $\alpha_{iNJ74}=0$; $\alpha_{iNJ79}=0$; $\alpha_{iJ}=0$	$R_i=\alpha_{iNJ}+\alpha_{iNJ74}D_{74}+\alpha_{iNJ79}D_{79}+\alpha_{iJ}D_J+b_iP+\varepsilon_i$	CAPM, APT
12	$\alpha_{iNJ}=0$; $\alpha_{iNJ74}=0$; $\alpha_{iNJ79}=0$; $\alpha_{iJ}=0$	$R_i=\alpha_{iNJ}+\alpha_{iNJ74}D_{74}+\alpha_{iNJ79}D_{79}+\alpha_{iJ}D_J+b_{iJ}D_JP$ $+b_{iNJ}P+\varepsilon_i$	CAPM, APT
13	$\alpha_i=(1-b_i)\lambda$	$R_i=\alpha_i+\alpha_{iNJ74}D_{74}+\alpha_{iNJ79}D_{79}+b_iP+\varepsilon_i$	CAPM Zero-b
14	$\alpha_i=\lambda$	$R_i=\alpha_i+\alpha_{iNJ74}D_{74}+\alpha_{iNJ79}D_{79}+b_iP+\varepsilon_i$	APT Zero-b
15	$\alpha_{iNJ}=(1-b_i)\lambda$	$R_i=\alpha_{iNJ}+\alpha_{iNJ74}D_{74}+\alpha_{iNJ79}D_{79}+\alpha_{iJ}D_J$ $+b_{iNJ}P+\varepsilon_i$	CAPM Zero-b
16	$\alpha_{iNJ}=\lambda$	$R_i=\alpha_{iNJ}+\alpha_{iNJ74}D_{74}+\alpha_{iNJ79}D_{79}+\alpha_{iJ}D_J$ $+b_{iNJ}P+\varepsilon_i$	APT Zero-b
17	$\alpha_{iNJ}=(1-b_i)\lambda$	$R_i=\alpha_{iNJ}+\alpha_{iNJ74}D_{74}+\alpha_{iNJ79}D_{79}+\alpha_{iJ}D_J+b_{iJ}D_JP$ $+b_{iNJ}P+\varepsilon_i$	CAPM Zero-b
18	$\alpha_{iNJ}=\lambda$	$R_i=\alpha_{iNJ}+\alpha_{iNJ74}D_{74}+\alpha_{iNJ79}D_{79}+\alpha_{iJ}D_J+b_{iJ}D_JP$ $+b_{iNJ}P+\varepsilon_i$	APT Zero-b

The index i=1,..m refers to the dependent variables which are sized based portfolios; P refers to the benchmarks portfolios; D_J to a dummy variable with $D_J=1$ in January and zero otherwise; D_{74} and 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. λ represents the average excess zero-beta return.

Table 6
Regression of market index excess returns on five estimated international factors.

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

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	0.99
	(1.19)	(121.91)	(44.76)	(-0.30)	(4.11)	(6.51)	
US-VW	-3.60	5.09	1.72	0.01	1.68	2.45	0.92
	(-2.78)	(36.05)	(12.16)	(0.05)	(11.93)	(17.47)	
UK-EW	1.78	6.76	-6.27	-1.25	0.26	-0.30	0.99
	(2.52)	(87.90)	(-81.38)	(-16.19)	(3.40)	(-3.87)	
UK-VW	-6.59	7.78	-6.19	-2.01	0.29	1.34	0.92
	(-3.15)	(34.02)	(-27.02)	(-8.73)	(1.27)	(5.89)	
JP-EW	1.69	2.75	-2.54	6.29	-0.46	0.66	0.96
	(1.75)	(26.13)	(-24.10)	(59.31)	(-4.41)	(6.26)	
JP-VW	-1.57	2.97	-2.12	5.22	-0.15	1.69	0.83
	(-0.77)	(13.26)	(-9.45)	(23.13)	(-0.69)	(7.59)	
FR-EW	-6.91	3.96	-2.56	1.62	0.16	2.23	0.35
	(-1.35)	(7.10)	(-4.58)	(2.88)	(0.29)	(4.01)	
FR-VW	-9.63	4.25	-2.57	1.55	0.35	2.36	0.35
	(-1.78)	(7.22)	(-4.34)	(2.61)	(0.59)	(4.02)	
I-EW	0.03	6.70	-1.18	0.94	0.24	0.15	1.00
	(0.11)	(226.40)	(-40.06)	(31.40)	(8.00)	(5.04)	
I-VW	-5.05	4.93	0.31	0.78	1.22	2.15	0.94
	(-5.31)	(47.40)	(3.02)	(7.49)	(11.82)	(20.73)	
UK-X	7.39	-0.94	1.67	-0.44	-0.59	0.00	0.35
	(3.83)	(-4.45)	(7.92)	(-2.07)	(-2.80)	(0.01)	
JP-X	3.39	-0.70	1.50	-2.30	0.04	-0.66	0.53
	(1.83)	(-3.46)	(7.41)	(-11.27)	(0.18)	(-3.28)	
FR-X	8.71	-0.64	1.55	-1.28	-0.36	-0.88	0.31
	(3.71)	(-2.49)	(6.06)	(-4.95)	(-1.42)	(3.45)	

US, UK, JP, FR, and I denote United States, United Kingdom, Japan, France and International portfolios, respectively. EW denotes equal-weighted market portfolio, VW denotes value-weighted market portfolio, and X denotes the percentage change in the spot exchange rate (in units of foreign currency per dollar). R^2 denotes the coefficient of determination. T-statistics in parentheses. Parameters are estimated using monthly returns over the period 1969-1983.

Table 7
Modified Likelihood Ratio (MLR) tests of no mispricing for size ranked portfolios.
Panel A: CAPM

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

Table 7 (Cont'd)
Panel B: **APT**

R_i	P	APT-5 $\alpha_i=0$	APT-5 $\alpha_{iNJ}=0$	APT-5 $\alpha_{iJ}=0$	APT-10 $\alpha_i=0$	APT-10 $\alpha_{iNJ}=0$	APT-10 $\alpha_{iJ}=0$
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* (<.001)	1.53 (.134)

Modified Likelihood Ratio (MLR) test from equation (6) with p-values in parentheses. Under the null they have a central F distribution (degrees of freedom equal to 10 and 170 - k, where k is the number of non-constant regressors). * indicates significance at the 5% level. Tests are for zero mispricing across ten size ranked portfolios given by restrictions 1 and 2 in Table 5. Parameters are estimated using monthly returns over the period 1969-1983. α_i, α_{iNJ}, and α_{iJ} are the estimates of mispricing, non-January mispricing, and January-specific mispricing over the 1969-1983 period. R_i identifies the market from which the size portfolios are constructed. P identifies the market from which the benchmark portfolios are constructed.

Table 8
Likelihood Ratio tests of no **mispricing for size** ranked portfolios - zero-beta models.

Panel A: **CAPM**

		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)	17.40* (.043)	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	33.98* (<.001)	30.25* (<.001)	32.76* (<.001)	28.67* (<.001)

Table 8 (Cont'd)
Panel B: **APT**

		APT-5		APT-10	
R_i	P	$\alpha_i = \lambda$	$\alpha_{iNJ} = \lambda$	$\alpha_i = \lambda$	$\alpha_{iNJ} = \lambda$
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)

Table 8 (Cont'd)

Likelihood ratio test statistics with p-values in parentheses. Statistics are asymptotically χ^2 with 9 degrees of freedom. * indicates significance at the 5% level. Tests are for zero mispricing across ten size based portfolios given by restrictions 4-7 in Table 5. Parameters are estimated using monthly returns over the period 1969-1983. For the CAPM, the estimates of mispricing and non-January mispricing are $\alpha_i-(1-b_i)\lambda$ and $\alpha_{iNJ}-(1-b_i)\lambda$, respectively. For the APT, the estimates of mispricing and non-January mispricing are $\alpha_i-\lambda$ and $\alpha_{iNJ}-\lambda$, respectively. Estimated difference between the zero-beta return and r_F is given by λ. R_i identifies the market from which the size portfolios are constructed. P identifies the market from which the benchmark portfolios are constructed.

Table 9
Average absolute mispricing across size ranked portfolios.

		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

Average absolute mispricing across ten size ranked portfolios, in percent per annum, are given by $A = \Sigma\,|\alpha_i|/10$ and $AJ = \Sigma\,|\alpha_{iJ}|/10$. α_i and α_{iJ} are the estimates of mispricing and January-specific mispricing, respectively, from regressions 1 and 2 in Table 5. R_i identifies the market from which the size ranked portfolios are constructed. P identifies the market from which the benchmark portfolios are constructed.

Table 10
Modified Likelihood Ratio (MLR) tests of no mispricing with capital control dummies.

Panel A: CAPM

		CAPM-EW							CAPM-VW						
R_i	P	α_i	α_{i74}	α_{i79}	α_{iNJ}	α_{iJ}	α_{iNJ74}	α_{iNJ79}	α_i	α_{i74}	α_{i79}	α_{iNJ}	α_{iJ}	α_{iNJ74}	α_{iNJ79}
US	US	0.90	2.76*	0.57	0.79	8.87*	3.08*	0.57	0.49	3.10*	0.46	0.30	11.68*	3.54*	0.46
		(.532)	(.004)	(.836)	(.641)	(<.001)	(.001)	(.839)	(.895)	(.001)	(.911)	(.980)	(<.001)	(<.001)	(.912)
JP	JP	0.28	1.26	0.52	0.39	2.08*	1.29	0.55	0.27	1.13	0.41	0.36	2.59*	1.13	0.43
		(.984)	(.258)	(.871)	(.950)	(.029)	(.240)	(.854)	(.987)	(.340)	(.942)	(.964)	(.006)	(.342)	(.931)
UK	UK	1.24	1.79	1.04	1.22	0.58	1.80	1.05	1.51	1.33	1.35	1.33	1.95*	1.31	1.33
		(.267)	(.065)	(.408)	(.284)	(.830)	(.065)	(.403)	(.140)	(.217)	(.210)	(.217)	(.041)	(.230)	(.217)
FR	FR	3.05*	0.91	2.64*	3.42*	1.99*	0.89	2.68*	3.05*	0.91	2.68*	3.42*	1.98*	0.88	2.72*
		(.001)	(.522)	(.005)	(<.001)	(.037)	(.546)	(.005)	(.001)	(.529)	(.005)	(<.001)	(.039)	(.553)	(.004)
US	Int'l	0.38	3.03*	0.54	0.31	8.92*	3.31*	0.54	0.44	3.20*	0.46	0.28	11.05*	3.65*	0.46
		(.952)	(.001)	(.860)	(.979)	(<.001)	(.001)	(.863)	(.925)	(.001)	(.911)	(.985)	(<.001)	(<.001)	(.911)
JP	Int'l	0.26	1.25	0.42	0.38	2.86*	1.41	0.48	0.27	1.01	0.37	0.36	2.57*	1.04	0.40
		(.988)	(.263)	(.937)	(.952)	(.003)	(.179)	(.899)	(.988)	(.438)	(.957)	(.962)	(.006)	(.415)	(.944)
UK	Int'l	1.07	1.30	0.98	1.12	0.71	1.36	1.00	1.11	1.24	1.01	1.09	1.66	1.28	1.04
		(.386)	(.235)	(.466)	(.353)	(.717)	(.205)	(.444)	(.357)	(.271)	(.437)	(.374)	(.094)	(.247)	(.410)
FR	Int'l	3.27*	1.02	2.80*	3.62*	1.61	0.97	2.87*	3.27*	0.94	2.85*	3.71*	2.08*	0.91	2.90*
		(.001)	(.432)	(.003)	(<.001)	(.107)	(.473)	(.003)	(.001)	(.501)	(.003)	(<.001)	(.029)	(.528)	(.002)
Int'l	Int'l	0.42	2.93*	0.68	0.63	5.92*	2.98*	0.77	0.55	2.81*	0.56	0.65	8.83*	2.98*	0.67
		(.937)	(.002)	(.739)	(.783)	(<.001)	(.002)	(.655)	(.850)	(.003)	(.848)	(.772)	(<.001)	(.002)	(.753)

Table 10 (continued)
Panel B: APT

		APT-5							APT-10						
R_i	P	α_i	α_{i74}	α_{i79}	α_{iNJ}	α_{iJ}	α_{iNJ74}	α_{iNJ79}	α_i	α_{i74}	α_{i79}	α_{iNJ}	α_{iJ}	α_{iNJ74}	α_{iNJ79}
US	US	0.61	2.45*	1.28	0.43	1.36	2.28*	1.25	0.63	2.54*	1.25	0.42	1.34	2.38*	1.24
		(.802)	(.009)	(.244)	(.930)	(.203)	(.016)	(.264)	(.786)	(.007)	(.264)	(.933)	(.214)	(.012)	(.268)
JP	JP	0.31	0.85	0.41	0.35	1.31	0.92	0.41	0.41	0.76	0.50	0.39	1.20	0.84	0.50
		(.978)	(.577)	(.942)	(.965)	(.231)	(.518)	(.942)	(.943)	(.667)	(.891)	(.948)	(.297)	(.589)	(.888)
UK	UK	1.24	2.11*	0.82	1.19	0.74	2.15*	0.82	1.25	2.14*	0.90	1.21	0.55	2.14*	0.90
		(.269)	(.026)	(.608)	(.298)	(.689)	(.024)	(.609)	(.266)	(.024)	(.533)	(.286)	(.855)	(.024)	(.539)
FR	FR	1.98*	1.10	2.32*	2.16*	1.33	1.08	2.31*	1.98*	0.89	2.24*	2.13	0.89	0.89	2.23*
		(.038)	(.369)	(.014)	(.022)	(.217)	(.379)	(.015)	(.039)	(.543)	(.018)	(.025)	(.548)	(.547)	(.019)
US	Int'l	0.38	2.61*	0.68	0.28	2.04*	2.41*	0.64	0.45	2.93*	0.91	0.28	1.64	2.67*	0.84
		(.956)	(.006)	(.746)	(.984)	(.032)	(.011)	(.774)	(.919)	(.002)	(.528)	(.984)	(.100)	(.005)	(.593)
JP	Int'l	0.30	2.39*	0.44	0.34	0.95	2.38*	0.46	0.26	1.78	0.32	0.21	0.55	1.94*	0.31
		(.979)	(.011)	(.927)	(.969)	(.489)	(.012)	(.915)	(.989)	(.069)	(.976)	(.995)	(.854)	(.044)	(.977)
UK	Int'l	1.14	1.49	0.93	1.09	1.38	1.75	0.91	1.20	2.03*	1.08	1.10	0.49	2.05*	1.09
		(.334)	(.146)	(.509)	(.371)	(.194)	(.073)	(.527)	(.295)	(.034)	(.379)	(.364)	(.896)	(.031)	(.373)
FR	Int'l	3.14*	1.08	2.72*	3.82*	1.83	1.04	2.95*	2.80*	0.86	2.44*	3.45*	1.42	0.84	2.65*
		(.001)	(.379)	(.004)	(<.001)	(.059)	(.410)	(.002)	(.003)	(.572)	(.010)	(<.001)	(.178)	(.587)	(.005)
Int'l	Int'l	0.52	3.05*	0.77	0.77	2.07*	2.79*	0.77	0.41	2.89*	0.96	0.54	1.33	2.65*	0.91
		(.873)	(.001)	(.655)	(.653)	(.030)	(.003)	(.653)	(.943)	(.002)	(.484)	(.861)	(.220)	(.005)	(.529)

Table 10 (continued)

Modified Likelihood Ratio (MLR) test statistics from equation (6) with p-values in parentheses. Under the null they have a central F distribution (degrees of freedom equal to 10 and 170 - k, where k is the number of non-constant regressors). * indicates significance at the 5% level. Tests are for zero mispricing across ten size ranked portfolios given by restrictions 10 and 11 in Table 5. Parameters are estimated using monthly returns over the period 1969-1983. Estimated mispricing throughout the 1969-1983 period is given by α_i. Estimated non-January mispricing throughout the 1969-1983 period is given by α_{iNJ}. Estimated January-specific mispricing is given by α_{iJ}. Estimated mispricing specfic to the 1969-1974 and 1969-79 periods is given by α_{i74} and α_{i79}, respectively. R_i identifies the market from which the size portfolios are constructed. P identifies the market from which the benchmark portfolios are constructed.

Table 11
Likelihood Ratio tests of no **mispricing for** size ranked portfolios - zero-beta models with capital control dummies.

Panel A: CAPM

		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 11 (Cont'd)

Panel B: **APT**

		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)

Table 11 (Cont'd)

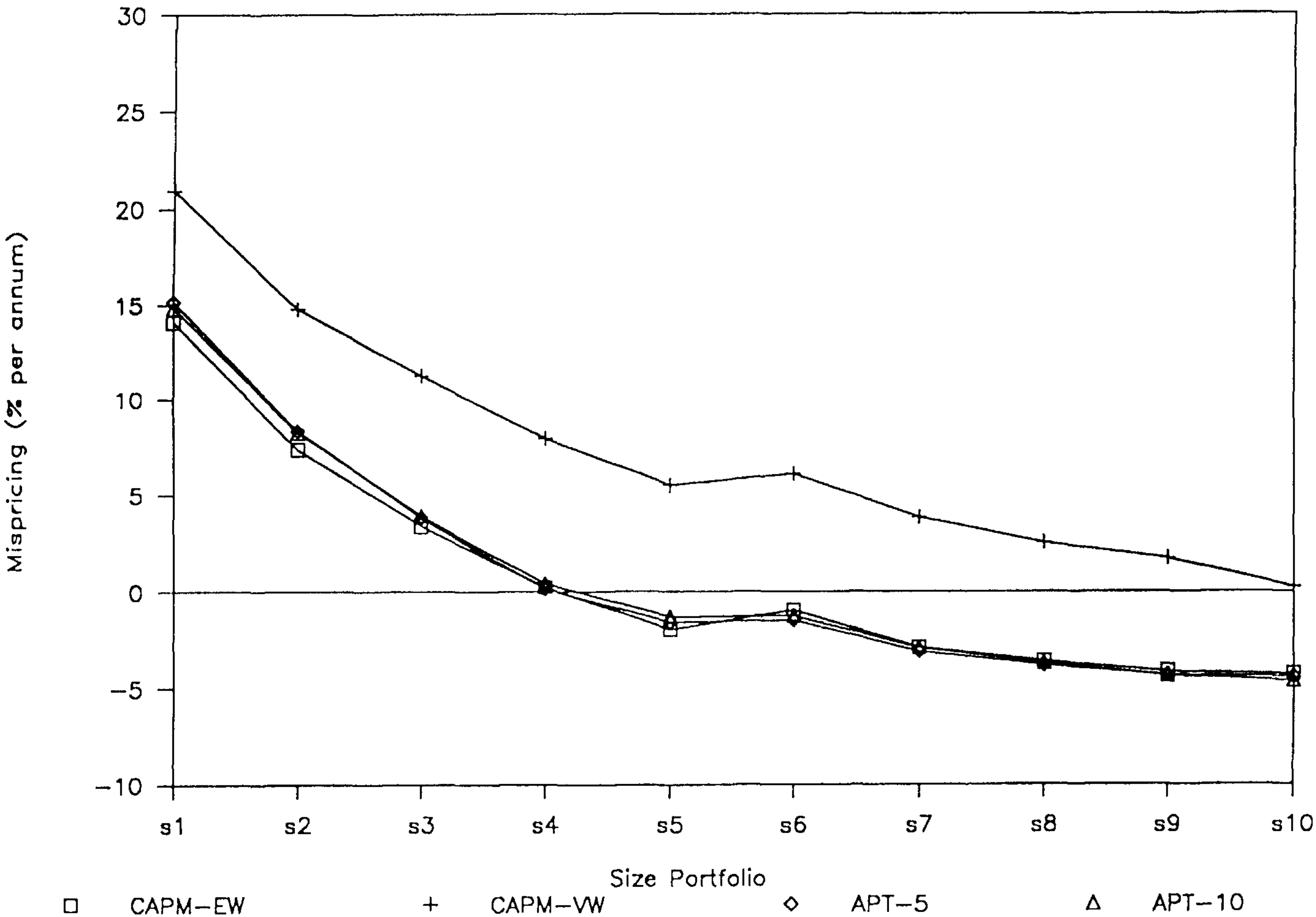
Likelihood ratio test statistics with p-values in parentheses. Statistics are asymptotically χ^2 with 9 degrees of freedom. * indicates significance at the 5% level. Tests are for zero mispricing across ten size based portfolios given by restrictions 13-16 in Table 5. Parameters are estimated using monthly returns over the period 1969-1983. For the CAPM, the estimates of mispricing and non-January mispricing are $\alpha_i-(1-b_i)\lambda$ and $\alpha_{iNJ}-(1-b_i)\lambda$, respectively. For the APT, the estimates of mispricing and non-January mispricing are $\alpha_i-\lambda$ and $\alpha_{iNJ}-\lambda$, respectively. Estimated difference between the zero-beta return and r_F is given by λ. R_i identifies the market from which the size portfolios are constructed. P identifies the market from which the benchmark portfolios are constructed.

Table 12
Average absolute mispricing across size ranked portfolios adjusting for changes in capital controls.

		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

Average absolute mispricing across ten size ranked portfolios, in percent per annum, are given by $A = \Sigma\ |\alpha_i|/10$ and $AJ = \Sigma\ |\alpha_{iJ}|/10$. α_i and α_{iJ} are the estimates of mispricing and January-specific mispricing, respectively, from regressions 10 and 11 in Table 5. Parameters are estimated using monthly returns over the period 1969-1983. R_i identifies the market from which the size ranked portfolios are constructed. P identifies the market from which the benchmark portfolios are constructed.

Figure 1. Mispricing, in percent per annum, for ten international portfolios formed by ranking on firm size. Mispricing is estimated by the intercept in the regression of monthly portfolio excess returns on a constant and (a) monthly excess returns on a value-weighted portfolio of international stocks, denoted CAPM-VW (plus signs) (b) monthly excess returns on an equal-weighted portfolio of international stocks, denoted CAPM-EW (squares); (c) first five international factor estimates from the asymptotic principal components procedure, denoted APT-5 (diamonds); and (d) first ten international factor estimates from the asymptotic principal components procedure, denoted APT-10 (triangles). Parameters are estimated using monthly returns over the period 1969-1983. S1 represents the portfolio of smallest firms while S10 represents the portfolio of largest firms. Size is defined as market value of common stock at the beginning of each five year subperiod.



30
25
20
15
10
5
0
-5
-10
Mispricing (% per annum)
s1
s2
s3
s4
s5
s6
s7
s8
s9
s10
Size Portfolio
□ CAPM-EW
+ CAPM-VW
◇ APT-5
Δ APT-10

Figure 2. Mispricing of international CAPM across countries, in percent per annum. Mispricing is estimated by the intercept in the regression of monthly portfolio excess returns on a constant and monthly excess returns on a value-weighted portfolio of international stocks. S1 represents the portfolio of smallest firms while S10 represents the portfolio of largest firms. Parameters are estimated using monthly returns over the period 1969-1983. UK size portfolios are denoted by squares, Japanese size portfolios by plus signs, US size portfolios by diamonds, and French size portfolios by triangles. Size is defined as market value of common stock at the beginning of each five year subperiod.

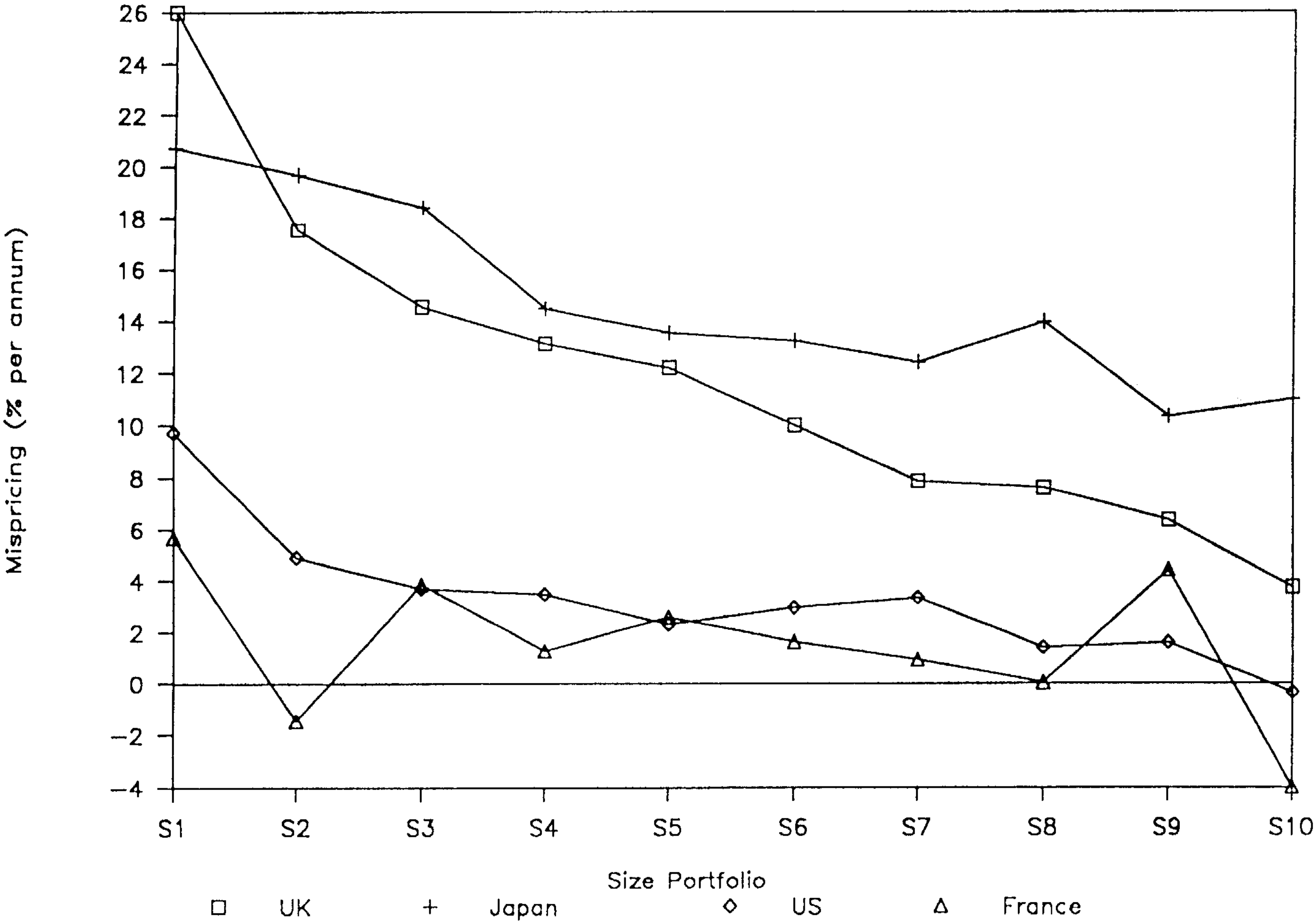
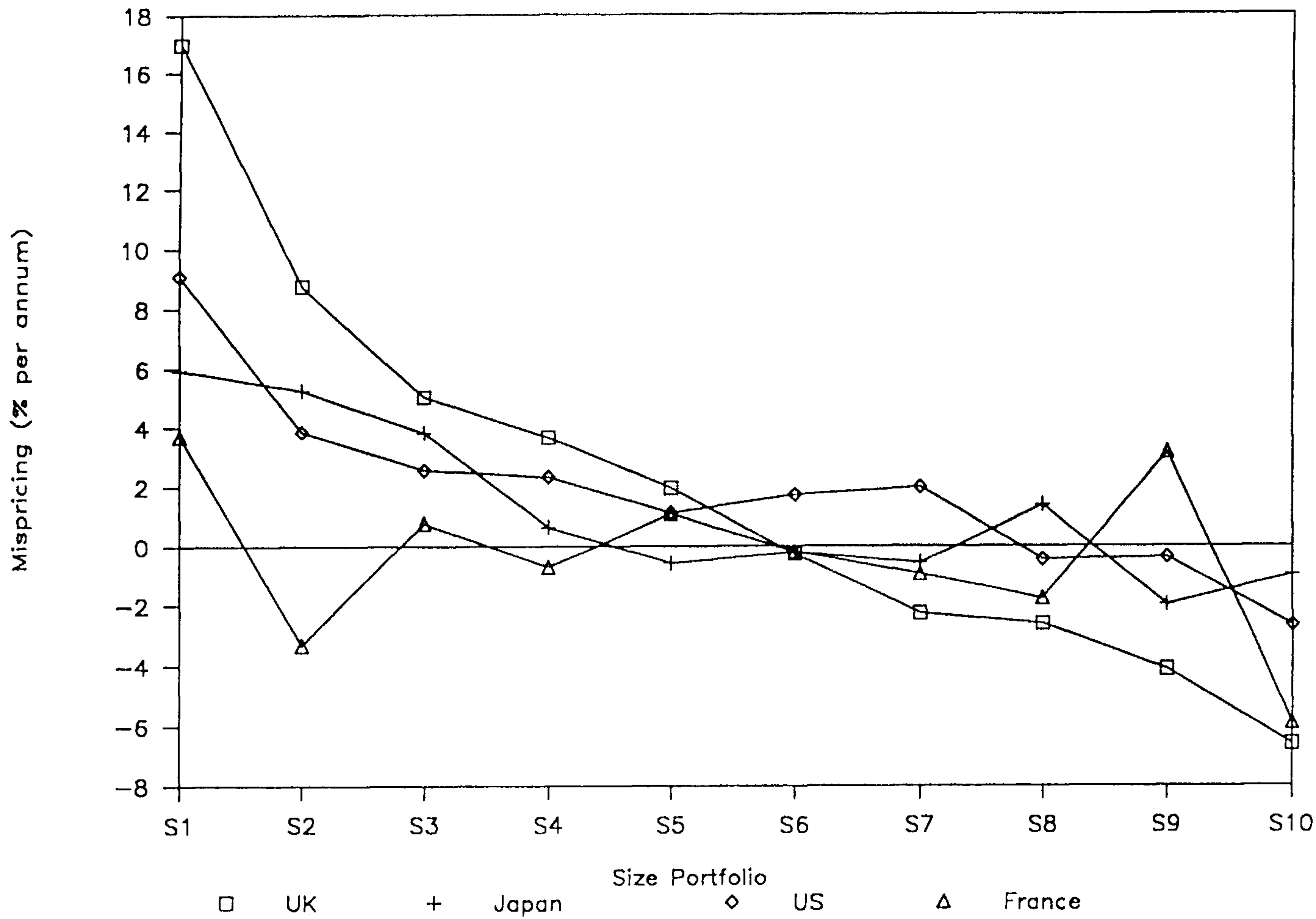
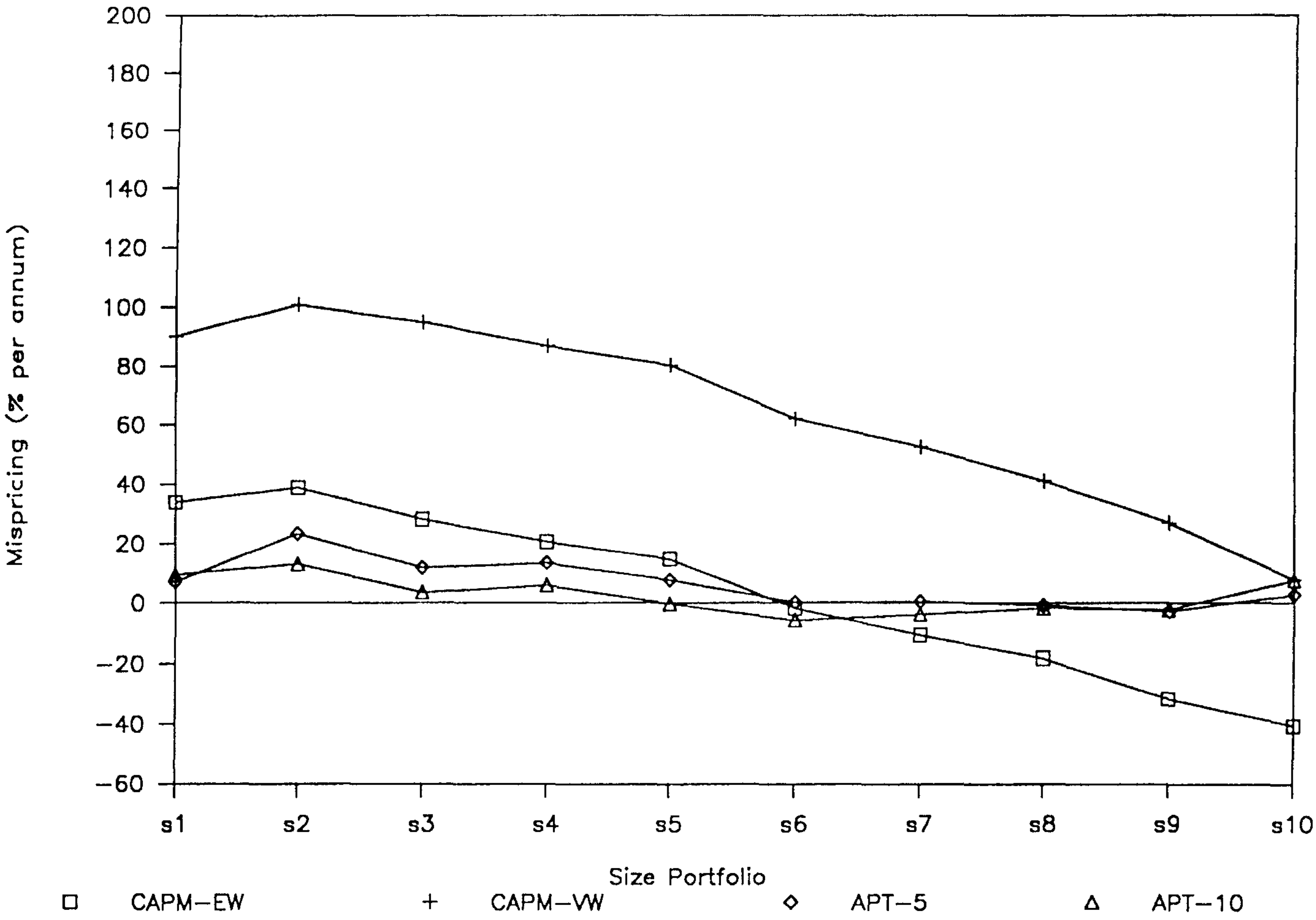

Mispricing (% per annum)
26
24
22
20
18
16
14
12
10
8
6
4
2
0
-2
-4
S1
S2
S3
S4
S5
S6
S7
S8
S9
S10
Size Portfolio
UK
Japan
US
France

Figure 3. Mispricing of international five factor APT, in percent per annum. Mispricing is estimated by the intercept in the regression of monthly portfolio excess returns on a constant and the first five international factor estimates from the asymptotic principal components procedure. Parameters are estimated using monthly returns over the period 1969-1983. S1 represents the portfolio of smallest firms while S10 represents the portfolio of largest firms. UK size portfolios are denoted by squares, Japanese size portfolios by plus signs, US size portfolios by diamonds, and French size portfolios by triangles. Size is defined as market value of common stock at the beginning of each five year subperiod.



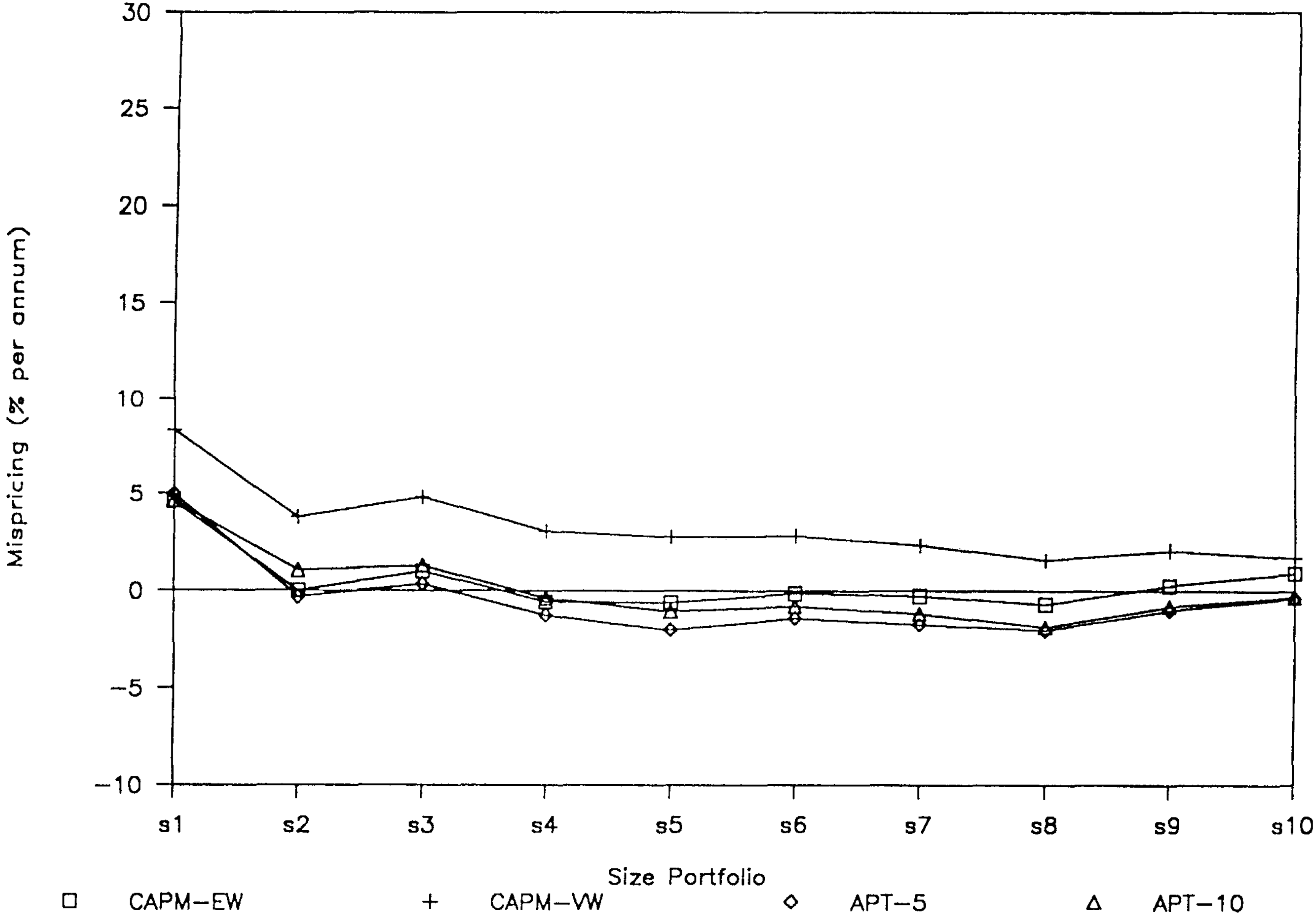
Mispricing (% per annum)
18
16
14
12
10
8
6
4
2
0
−2
−4
−6
−8
S1
S2
S3
S4
S5
S6
S7
S8
S9
S10
Size Portfolio
UK
Japan
US
France

Figure 4. January-specific mispricing, in percent per annum, for ten international portfolios formed by ranking on firm size. Mispricing is estimated by the slope coefficient on the January dummy variable in the regression of monthly portfolio excess returns on a constant, January dummy variable, and (a) monthly excess returns on a value-weighted portfolio of international stocks, denoted CAPM-VW (plus signs) (b) monthly excess returns on an equal-weighted portfolio of international stocks, denoted CAPM-EW (squares); (c) first five international factor estimates from the asymptotic principal components procedure, denoted APT-5 (diamonds); and (d) first ten international factor estimates from the asymptotic principal components procedure, denoted APT-10 (triangles). Parameters are estimated using monthly returns over the period 1969-1983. S1 represents the portfolio of smallest firms while S10 represents the portfolio of largest firms. Size is defined as market value of common stock at the beginning of each five year subperiod.



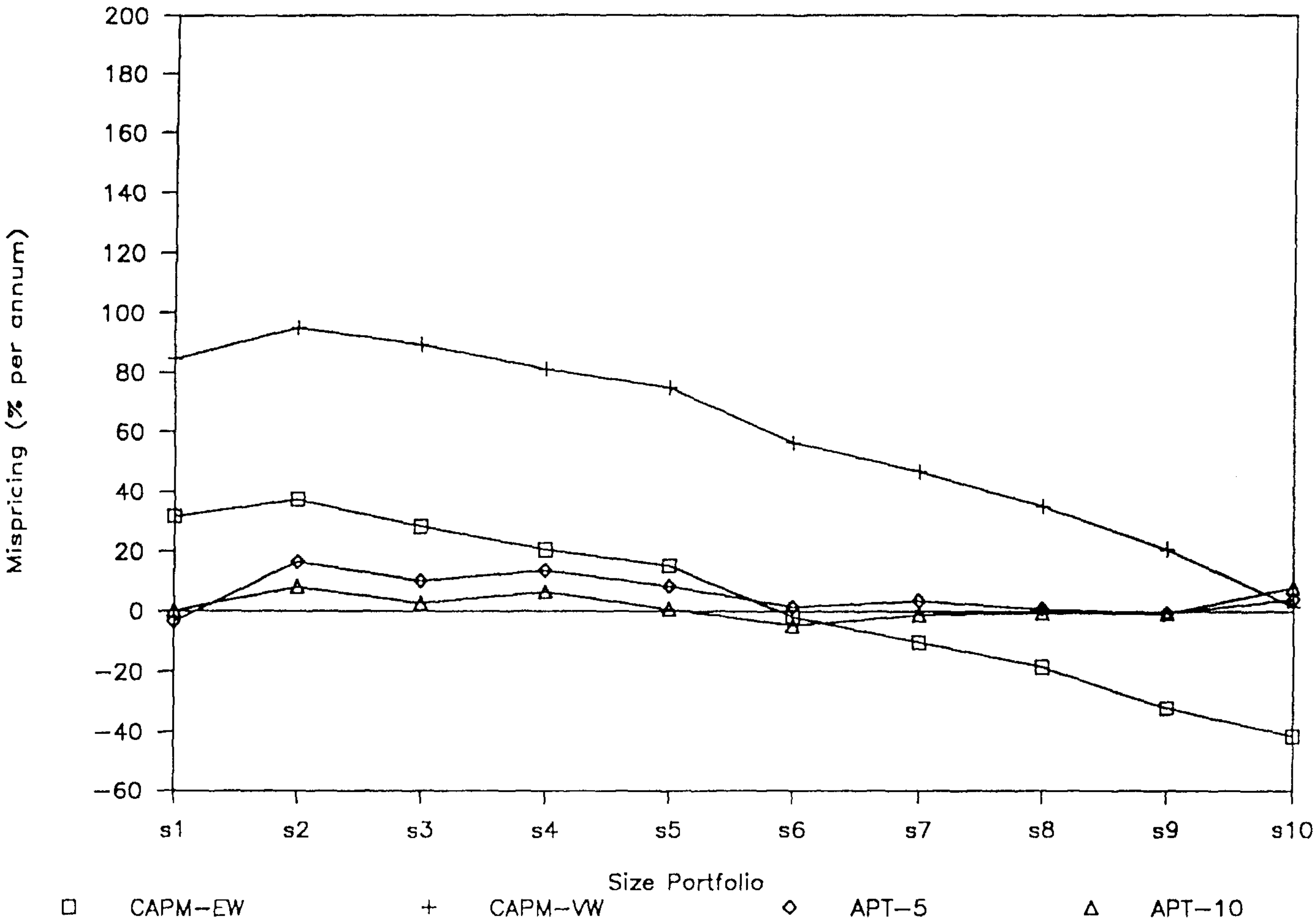
Mispricing (% per annum)
200
180
160
140
120
100
80
60
40
20
0
-20
-40
-60
s1
s2
s3
s4
s5
s6
s7
s8
s9
s10
Size Portfolio
CAPM-EW
CAPM-VW
APT-5
APT-10

Figure 5. Mispricing (controlling for changes in capital controls) in percent per annum, for ten international portfolios formed by ranking on firm size. Mispricing is estimated by the intercept in the regression of monthly portfolio excess returns on a constant, a dummy variable which is unity before February 1974, a dummy variable which is unity before December 1979, and (a) monthly excess returns on a value-weighted portfolio of international stocks, denoted CAPM-VW (plus signs) (b) monthly excess returns on an equal-weighted portfolio of international stocks, denoted CAPM-EW (squares); (c) first five international factor estimates from the asymptotic principal components procedure, denoted APT-5 (diamonds); and (d) first ten international factor estimates from the asymptotic principal components procedure, denoted APT-10 (triangles). Parameters are estimated using monthly returns over the period 1969-1983. S1 represents the portfolio of smallest firms while S10 represents the portfolio of largest firms. Size is defined as market value of common stock at the beginning of each five year subperiod.



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

Figure 6. January-specific mispricing (controlling for changes in capital controls) in percent per **annum**, for ten international portfolios formed by ranking on firm size. Mispricing is estimated by the slope coefficient on the January dummy variable in the regression of monthly portfolio excess returns on a constant, January dummy variable, a dummy variable which is unity before February 1974, a dummy variable which is unity before December 1979, and (a) monthly excess returns on a value-weighted portfolio of international stocks, denoted CAPM-VW (plus signs) (b) monthly excess returns on an equal-weighted portfolio of international stocks, denoted CAPM-EW (squares); (c) first five international factor estimates from the asymptotic principal components procedure, denoted APT-5 (diamonds); and (d) first ten international factor estimates from the asymptotic principal components procedure, denoted APT-10 (triangles). Parameters are estimated using monthly returns over the period 1969-1983. S1 represents the portfolio of smallest firms while S10 represents the portfolio of largest firms. Size is defined as market value of common stock at the beginning of each five year subperiod.



200
180
160
140
120
100
80
60
40
20
0
-20
-40
-60
Mispricing (% per annum)
s1
s2
s3
s4
s5
s6
s7
s8
s9
s10
Size Portfolio
CAPM-EW
CAPM-VW
APT-5
APT-10

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1989

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