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International Investment Performance: Evidence from Institutional Investors' Foreign Equity Holdings*

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Abstract

This paper analyzes the international equity holdings of a large panel of UK pension funds. We find considerable evidence of market timing activity, as illustrated by the funds' decision to scale back their investments in the US stock market during the 1990s. To explain this we jointly model time-varying conditional moments and portfolio weight dynamics. Past returns do not adequately explain the funds' international portfolio flows and only matter for the short run. Instead we find that a substantial part of the evolution in portfolio weights is explained by time-varying conditional expected returns, volatilities and covariances with domestic equity returns. However, once we control for publicly known state variables, there is no evidence of extra-market timing skills and most funds appear to have earned small but negative returns from international market timing.

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I. Introduction

Foreign investors' market timing activity has long been the subject of speculation. Reflecting on the large movements in international capital flows that occurred in the early 1990s, Lewis (1998) concludes that investors do not appear to follow passive buy-and-hold strategies in foreign markets. She conjectures that "... domestic investors may be trying to follow market timing strategies" (page 27). However, little is known about the factors influencing investors' market timing and strategic asset allocation decisions in international equity markets.

This paper investigates the extent of and rewards to institutional investors' market timing activity by analyzing a panel of 247 UK pension funds' foreign equity holdings in four regional markets (Japan, North America, Europe and Asia-Pacific) over the period 1991 to 1997. We find evidence of extensive attempts at market timing. At first this seems unsurprising since it is well-known that British funds persistently bet against the US stock market during the 1990s, not just by initially underweighting US stocks, but also by systematically reducing their US investments during a period when the global weight of the US stock market rose substantially. More surprising perhaps are the drivers behind international asset allocation decisions. We find strong evidence that a substantial portion of the market timing activity of individual funds can be explained by time-varying expected returns, volatilities and covariances in the four developed regions that we investigate. In particular, the observed decline in allocation to the North American stock market coincided with a systematic decline over the sample period in *expected* returns on North American assets relative to those from other developed markets, even though *ex post* the realized returns in this market were very high.

Studying individual funds' investment decisions turns out to have many advantages. Bohn & Tesar (1996) were the first to draw attention to the importance of expected returns for international capital flows. However, they studied aggregate flows for US investors and found only limited empirical support for the proposition that expected returns could explain portfolio flows. Since the composition of aggregate capital flows is unlikely to remain stable over time, such findings can be difficult to interpret. In fact we find that time-varying expected returns are more important at the level of individual institutional investors' asset allocation decisions than in the aggregate. In portfolio weight regressions, a remarkable 94 percent of all funds generated a positive and significant coefficient on expected returns for North America. In an attempt to capture international diversification effects, we include conditional own-market volatility and conditional covariances with domestic returns as additional explanatory variables. The percentage of funds that generated a negative and significant coefficient on either own-market volatility or conditional covariance with domestic stock returns varied from 53 percent (North America) to 94 percent (Europe). These results indicate that time-varying conditional moments are essential for explaining and evaluating institutional investors' international asset allocations.

The withdrawal from North America together with the fact that the US stock market paid substantially higher returns than the other major developed markets during the sample period might lead one to conclude that returns from international market timing were negative simply as a result of this one major market timing bet.[1] Compared with a strategy of using global market capitalization weights for their foreign equity portfolio, UK pension funds were 29 percentage points underweight in North America over the sample period. Since the average return on North America was more than 10 percentage points above the average international return for other developed markets, a negative mean return of around 3 percent per annum from this decision is suggested. However, this conclusion is premature since it confuses *ex post* returns with *ex ante* expected performance. The decision to withdraw from North America appears to have been the result of low or even negative expected returns resulting from rising stock valuations and low dividend yields in this market. So it is possible that UK funds possessed market timing skills over and above that which could have been inferred from a model of expected returns based on public information. We argue that a valid assessment of market timing skills has to be conducted in the context of a conditional analysis which allows for time-varying investment opportunities.

We investigate the market timing skills of the pension funds using a battery of tests that account for a time-varying investment opportunity set, as proposed in the recent literature on conditional performance measurement, c.f. Ferson & Schadt (1996) and Graham & Harvey (1996). Once we control for the effects of

[1]For example, in his *Financial Times* column of December 16, 1997 Barry Riley wrote "The latest revival on Wall Street ... has further inflamed the wounds of the overseas managers who have been so underweight in US equities all year". In his column of May 13, 1998, under the heading 'Wall Street Misread', Riley writes that "Last year's huge underweighting [of the US market] is being blamed on strategists' poor judgement" (our insert).

public information, there is no evidence of what Graham & Harvey (1996) refer to as extra-market timing, i.e. the ability to anticipate return movements beyond what could have been predicted using public information. Our estimates indicate that the median fund earned a negative return from international market timing of around -0.4 percent per annum.

Our work is related to a recent literature on the aggregate behavior and price impact of foreign investors. Froot, O'Connell & Seasholes (1999) use daily data on international trades over the period 1994 to 1998 to shed light on the relationship between foreign asset trades and stock returns. Their data comprises detailed records on aggregate holdings of pension, endowment and mutual funds and of governments. Equally impressive is the data set used by Choe, Kho & Stulz (1999) to examine the complete set of transactions of foreign investors on the Korea Stock Exchange over the period November 1996 to December 1997. Our data set is unique relative to those analyzed in previous studies in that it is organized by individual pension funds' asset holdings. While our data set is not well suited for studying the price impact of foreign investors in a particular domestic market, it is ideally suited for analyzing institutional investors' reallocations of funds across major developed capital markets and hence allows us to characterize and quantify the investment strategy of a key group of investors. Although our sample period is relatively short, this has certain advantages. As Tesar & Werner (1994) and Kang & Stulz (1997) point out, barriers to international investment have been declining over the last 20 to 30 years and our post-1980s data set is unlikely to be contaminated by this relaxation of capital controls.

The plan of the paper is as follows. Section II provides a description and initial characterization of our data set. To establish the extent of UK pension funds' international market timing activities and to assess the similarity between their investment strategies in foreign markets, Section III analyzes the evolution in both aggregate and individual fund portfolio weights and the extent to which these can be explained by time variations in the investment opportunity set. Section IV examines evidence on returns from international market timing and Section V concludes.

II. Data

Our data consists of monthly observations on 247 UK pension funds' investments in international equities over the period 1991:1 - 1997:12. It was provided to us by The WM Company of Edinburgh, UK. The sample is complete in the sense that it contains all of the funds that maintained the same single, externally-appointed fund management group throughout the period and which also reported their performance data continuously to WM over the period. Furthermore, UK pension funds face very few restrictions on their investment strategies.

The fact that we only consider funds with the same manager in place over the sample period raises the possibility that our sample is subject to survivor bias. Fortunately this bias is likely to be very small: a comparison of the mean return on the international equity portfolio of the full set of funds tracked by WM (12.50 percent per annum) with either the average value-weighted return (12.58 percent) or the average equal-weighted return (12.51 per cent) on our sample of funds reveals that the difference in mean returns is negligible. The finding of a seven basis point difference between the equal-weighted and value-weighted returns of pension funds also suggests that there is no significant difference between small and large funds' average performance. Further confirmation of the similarity between our sample and the full set of funds tracked by WM is provided by a time-series correlation of 0.998 between (value-weighted) returns on the two sets of funds.

For each fund, we have data on four regional constituents: Japan, North America, Europe (excluding the UK) and Asia-Pacific (excluding Japan).[2] For each region, every fund reports initial market value and net investment, income received, and return over the month. All asset holdings and returns are reported in pounds sterling.

There is already some evidence on international portfolio flows. Based on monthly recordings of transactions in long-term marketable securities reported to the US Treasury International Capital system, and using similar sources for other

[2]Some funds also held positions in a sector entitled 'other international equities' which largely consists of African, Middle Eastern and South American equities as well as mutual funds that could not be allocated exclusively to one of the four main categories. But these holdings were very small, less than 1 percent of total international equity holdings for much of the sample. Since the data records on this category were found to be incomplete, this sector was dropped entirely from the analysis and the weights rescaled for the four main regions.

countries, Tesar and Werner (1994, 1995) analyze the evolution in aggregate holdings of foreign assets in five major economies. Cooper & Kaplanis (1994) consider market capitalization data but do not discuss portfolio flows. Kang & Stulz (1997) examine foreign investors' aggregate holdings of individual firms' stocks. They find that foreign investors are cautious in their choice of assets and predominantly hold the equities of large firms in manufacturing industries as well as the equities of firms with good accounting performance. Choe et al. (1999) and Froot et al. (1999) use data sets that allow them to provide a detailed analysis of the short run, aggregate behavior and price impact of foreign investors.

A first impression of some key features of our data is provided in Figure 1 which plots aggregate portfolio weights in the four regions against the corresponding global market capitalization weights. The figure shows that UK pension funds' total international portfolio weights vary considerably over time. The aggregate weights in Japan, for example, increased by almost eight percentage points in 1991 only to more than halve from 25 to 11 percent between 1992 and 1994. They more than doubled in early 1994 and then drifted back again between 1995 and 1997. Over the full sample there is little overall change in the portfolio weight for Japan.

Turning to North America, a very different picture emerges. The weights decreased almost consistently throughout the sample from an initial level of 28 percent in early 1991 to around 10 percent at the end of 1996. Although there is a slight increase to 14 percent by the end of the sample, this cannot hide UK pension funds' massive withdrawal from North American equities at a rate in excess of 200 basis points per year.

Unsurprisingly, European equities account for around half of UK pension funds' international equity holdings. This weight increased over the sample period, particularly from 1996 to the end of 1997 when it rose from 39 to 57 per cent of total foreign equity holdings.[3] For the whole sample, the average annual increase in the weight in Europe amounted to 179 basis points. The weight in Asia-Pacific excluding Japan (Asia-Pacific for short) rose consistently over most of the period, almost trebling from 10 to 28 per cent of the total from 1991 to the end of 1995. In 1996

[3] Using an equilibrium model for global financial markets, Dumas (1998) estimates very small effects from European Economic and Monetary Union on equity and currency risk premiums and on international asset allocations. This suggests that anticipaton of EMU cannot provide a plausible explanation for UK funds' increased European exposure.

and 1997 investments in this region drop sharply to around 12 percent, however, as a result of the Asia-Pacific economic crisis.

Several interesting features emerge from comparing these weights with their global counterparts (rescaled to sum to 100 percent): (i) At the beginning of the sample, UK pension funds had less than half the global weighting in Japanese equities. However, following the drop in both the yen and Japanese stock prices over the decade, this difference had virtually disappeared by the end of 1997. (ii) UK pension funds were initially underweight in North America by about 15 percentage points (28 versus 43 percent) and this difference widened steadily during the 1990s. The global weight of the US equity market was close to 60 per cent by the end of 1997, while UK funds scaled their holdings of international equities in the North American market back to 14 per cent. Hence the global weight in North America was an astonishing *four* times higher than that held by UK funds by the end of the sample. (iii) UK pension funds were overweight in Europe and Asia-Pacific. They held three times the world weight in both Europe and Asia-Pacific both at the beginning and at the end of the sample. An important conclusion emerges from this behavior: there is no evidence of convergence to the global weights over the sample period (with the possible exception of Japan).

Figure 1 also offers the impression that the volatility of UK pension funds' aggregate portfolio weights exceeds those of the global weights (which represents the average global investor's portfolio). To test this formally, we computed for each month the variance of the portfolio weight changes across markets (in basis points) both for the global portfolio and for the value-weighted portfolio of UK pension funds. The average standard deviation of the pension fund portfolio weights was 104 basis points against 92 basis points for the world portfolio. A one-sided test of the null that these standard deviations are identical against the alternative that UK pension funds have greater volatility in portfolio weight changes could be rejected at the five percent critical level. This is consistent with Tesar & Werner (1995)'s finding that investors' turnover rate in foreign equity investments is high relative to their home market turnover rate.

III. Sources of Variation in UK Pension Funds' International Equity Holdings

Our next task is to study the factors that determine the evolution in our sample of pension funds' international equity holdings. Initially we summarize the aggregate evidence for the full set of funds. We subsequently consider the extent to which the strategies followed by individual funds reflect differential returns, rebalancing effects and time-varying investment opportunities.

A. Decompositions of Aggregate Portfolio Weight Changes

To help identify the factors causing portfolio weights to change over time, Table 1 reports the decomposition of portfolio weight changes recently proposed by Blake, Lehmann & Timmermann (1999). This decomposition recognizes that asset classes enjoying large positive relative returns also experience an increase in their allocations in the total portfolio unless fund managers deliberately rebalance portfolios as this occurs.

We first apply the decomposition to the pension funds' aggregate (value-weighted) holdings of international equities. Let W_{jt} be the total holding in region j at the end of month t across all funds in the sample while W_t is the total holding across all M regions. These holdings must satisfy the accounting identity:

$$W_{jt} \equiv W_{jt-1}(1 + r_{jt} + NCF_{jt}), \tag{1}$$

where r_{jt} is the rate of return on UK pension funds' holdings in region j and NCF_{jt} is the rate of net cash flows into that region during month t. Using this relation, the portfolio weight for region j (ω_{jt}) can be written as:

$$\omega_{jt} \equiv \frac{W_{jt}}{W_t} = \frac{\frac{W_{jt-1}}{W_{t-1}}(\frac{W_{jt}}{W_{jt-1}})}{W_t/W_{t-1}} = \omega_{jt-1}\frac{1 + r_{jt} + NCF_{jt}}{1 + \sum_{k=1}^{M}\omega_{kt}(r_{kt} + NCF_{kt})}. \tag{2}$$

Taking logarithms, it follows that:

$$\Delta \ln(\omega_{jt}) = \ln(1 + r_{jt} + NCF_{jt}) - \ln(1 + \sum_{k=1}^{M}\omega_{kt}(r_{kt} + NCF_{kt})), \tag{3}$$

so that, to a close approximation,

$$\Delta \ln(\omega_{jt}) \approx r_{jt} - r_{pt} + NCF_{jt} - NCF_{pt}, \tag{4}$$

where r_{pt} is the value-weighted total return and NCF_{pt} is the value-weighted net cash flow into the total portfolio during month t.

The decomposition in (4) allows us to measure the extent to which changes in aggregate portfolio weights are caused by differential returns across international regions, as indicated by r_{jt}-r_{pt}, or by shifts in the net cash flows across regions, as indicated by NCF_{jt}-NCF_{pt}. Variations arising from the first component represent the passive investment strategy of 'buy-and-hold', reinvesting asset income in the region whence it originated and distributing any net inflows into the pension fund according to the *ex post* asset allocation. Revisions in portfolio weights associated with the second component represent the active strategy of rebalancing the portfolio by redirecting cash flows across regions. The dramatic decrease over the sample period in the allocation to North American equities, for example, might either reflect low relative returns or the active withdrawal of funds.

Trends in international equity investments reflect very different factors in different regions, c.f. Panels A and B of Table 1. The small percentage decline over the sample period in the weight for Japan arises from a combination of very low relative returns and a very large inflow of funds. The large declining weight in North America, in comparison, arises for the exact opposite set of reasons: a very large cash outflow more than compensates for the very high relative returns in North America over the sample period. The increase in Europe is almost entirely due to the above-average returns in this region, while, conversely, the small increase in Asia-Pacific reflects mean returns that are marginally below average combined with relatively large net cash inflows.[4]

To determine how much return variability matters for monthly variations in portfolio weight changes, we adopted the following variance decomposition:

$$Var(\Delta \ln(\omega_{jt})) \approx Var(r_{jt} - r_{pt}) + Var(NCF_{jt} - NCF_{pt}) + \tag{5}$$

[4]We computed the mean annual percentage change in portfolio weights due to currency fluctuations and found this to be -2.88 per cent for Japan, using the Sterling-Yen exchange rate, -2.28 per cent for North America, using the Sterling-US dollar exchange rate, and 0.36 per cent for Europe, using the Sterling-Deutschemark exchange rate. These figures suggest that currency fluctuation is not the dominant factor in explaining the long-run variation in portfolio weights.

$$2Cov(r_{jt} - r_{pt}, NCF_{jt} - NCF_{pt}).$$

Panel C shows that the return component accounts for 32 (Japan), 24 (North America), 29 (Europe) and 39 (Asia-Pacific) per cent, respectively, of the monthly variance of aggregate portfolio weight changes. This indicates the presence of a significant short-run common component of the portfolio weight changes that is unrelated to actual returns.

B. Variations in Individual Funds' Portfolio Weights

To investigate the dispersion about the common component, we next study the dynamics of the cross-section of individual funds' portfolio weights. The various theories that have been suggested to explain home country bias have different implications for domestic investors' foreign asset holdings. Models of home country bias based on informational asymmetries such as those of Gehrig (1993) or Brennan & Cao (1997) assume that foreign investors are worse informed than domestic ones. This distinction only makes sense, however, if foreign investors follow similar investment strategies, i.e. if there is a strong common component in their portfolio holdings. The validity of this assumption has not previously been assessed.

To evaluate the level of heterogeneity in the pension funds' investment strategies, we first consider the cross-sectional distribution of portfolio weights across the four markets. Figure 2 presents histograms for the individual funds' portfolio weights computed as an average over the sample period. Only North America has a relatively symmetric and concentrated distribution, centered on 18 per cent, while two separate clusters, centered at 15 and 22 percent, can be found for Japan. Similarly, two clusters centered at 18 and 27 percent appear for Asia-Pacific, while clusters centered at 40 and 48 percent appear for UK pension funds' holdings of European equities. These figures indicate a wide dispersion in individual funds' investment strategies, but say nothing about the presence of any common dynamics.

To investigate the extent to which individual funds' portfolio weights revert towards a common mean, we consider the cross-sectional distribution of the individual funds' asset allocation dynamics. This has implications both for performance measurement and for our understanding of institutional investors' behavior in foreign markets. We sorted the individual funds into two groups according to whether or not their portfolio weights in a particular region exceeded the median holding in that region. The sort is performed every December. We then estimated the probability that a fund, i, with above-average weighting in region j ($\omega_{ijt} > \overline{\omega}_{jt}$) continues to have above-average weighting in the same region one year later:

$$p_{ij} = P(\omega_{ijt+12} > \overline{\omega}_{jt+12} | \omega_{ijt} > \overline{\omega}_{jt}) \tag{6}$$

This generates a time series of first-order Markov transition probability estimates for which we report the mean and standard error in Panel A of Table 2. Panel B of the table also reports the Markov transition probabilities based on initial and terminal portfolio weights in the four regions ($P(\omega_{ijT} > \overline{\omega}_{jT} | \omega_{ij1} > \overline{\omega}_{j1})$). If there was no persistent, fund-specific asset allocation component, the transition probabilities should be distributed randomly around 0.5. The year-to-year transition probabilities for the relative weights range from 0.65 to 0.78, indicating again that there is considerable persistence in deviations from the median asset allocation at the annual horizon. Even at the seven year horizon there is no tendency to mean revert towards the average international asset allocation in Europe and Asia. This evidence is consistent with Figure 2 and suggests that individual funds follow sufficiently different strategies in their international asset allocation that substantial information is lost by considering only aggregate equity holdings.

C. Cash Flow Dynamics and Past Returns

Recent research into the portfolio flows of international investors has focused on the extent to which these flows are driven by past returns. There are good theoretical reasons why past returns should help to determine portfolio weights: Brennan & Cao (1997) construct an asymmetric information model in which foreign investors increase their holdings in a particular market if that market experiences high returns. The intuition behind this result is easy to understand. In a noisy rational expectations equilibrium, the worse-informed (foreign) investors will revise their expected returns on foreign assets by more than the better-informed (local) investors who have established more precise estimates of the return generating process. If a positive shock occurs in a particular market, foreign investors revise upwards their return estimates by more than domestic investors and hence will, in equilibrium, demand a larger share of the assets in that market. Brennan and Cao find some evidence that US investors' total net purchases of foreign assets are positively correlated with both concurrent returns on the foreign market indices and the previous quarter's return.

We can test this prediction by analyzing whether the individual funds' short-run investment flows reflect differential returns in the four regions. Since there has been an upward drift in the value of UK funds' international equity holdings and since we are interested in explaining the relative holdings in different foreign markets, we will examine the portfolio weight changes after correcting for the automatic rebalancing resulting from differential returns. Applying (4) to the individual funds gives:

$$\Delta \ln(\omega_{ijt}) - (r_{ijt} - r_{ipt}) \approx NCF_{ijt} - NCF_{ipt}, \tag{7}$$

where i indexes individual pension funds, while j is the region and t the time period. Equation (7) allows a direct test of Brennan and Cao's model which hypothesizes that foreign investors' cash flows into regions with high relative returns should increase while cash flow rates into regions with low relative returns should decline. To test this proposition we regress net cash flow rates on a constant and the returns in each region:[5]

$$NCF_{ijt} - NCF_{ipt} = \alpha_{ij} + \sum_{k=1}^{M} \beta_{ik} r_{kt}^{h} + \varepsilon_{ijt} \tag{8}$$

where M is the number of regions and r_{kt} is the market index return in region k; $r_{kt}^{h} = \sum_{l=0}^{h-1} r_{k,t-l}$ is the corresponding cumulative return over the $h-$month horizon ending in period t.[6] For each region, we use Financial Times/Standard & Poors market indices as external benchmarks.[7] We also regress the net cash flow rates in each region on the return in that region relative to the average world market return (ex-UK):

[5]This provides a stricter test of the Brennan-Cao 'trend chasing' hypothesis than a regression of portfolio weight changes on return differentials. The latter regression would be dominated by the automatic rebalancing effect due to the increase in the portfolio weight of an asset whose return was relatively high. In fact, in a regression of portfolio weight changes on a constant and own-market returns, we found that 100, 94, 89 and 99 percent of the funds in Japan, North America, Europe and Asia-Pacific, respectively, had positive and statistically significant coefficients on own-market returns.

[6]Our tests are consistent with Brennan and Cao's set-up since the investment flows in their model can be scaled by total wealth to derive changes in portfolio weights.

[7]Section IV provides a range of statistics on the adequacy of these return benchmarks in the regional regressions.

$$NCF_{ijt} - NCF_{ipt} = \alpha_{ij} + \beta_{ij} xr_{jt}^h + \varepsilon_{ijt}, \tag{9}$$

where $xr_{jt}^h = (r_{jt}^h - r_{\omega t}^h)$ is the excess return in region j and $r_{wt}^h = \sum_{l=0}^{h-1} r_{w,t-l}$ is the world market return over the $h-$month period ending in period t. To summarize the outcome from the cross-section of 247 time series regressions, Table 3 reports the proportion of regression coefficients that are significant at the 5% critical level as well as the median values of the regression coefficients. Since we have no information on the funds' actual portfolio adjustment horizon, we present short-run results based on returns over the most recent month ($h = 1$) as well as longer-term results that use cumulated returns over the most recent six-month period ($h = 6$).

The Brennan-Cao model predicts positive coefficients on own-market returns and negative coefficients on returns in other markets. This pattern emerges from Table 3 in the case of a one-month horizon. With the exception of cash flows into North America, the highest proportion of positive and significant return coefficients in the multiple regression model (8) relate to own-market returns (Panel A) and the median value of these coefficient estimates is positive for each market (Panel C). In contrast, cross-market return coefficients are generally negative. In the simpler model (9), the proportion of positive and significant regression coefficients exceeds the proportion of negative coefficients for each region (Panel B).

The very large proportion of funds with a significantly negative intercept term in the North American net cash flow regressions and a negative coefficient on returns in the single regression model for this region (Panel D) again indicate that UK pension funds scaled down their North American equity investments well beyond what would have resulted if investors simply revised their future return estimates based on current returns as assumed by the information updating story of Brennan and Cao. The other regressions indicate that investors systematically raise their holdings of foreign assets (in excess of that due to passive rebalancing) in regions that have experienced above-average returns, while they scale back their investments in regions with below-average returns. In addition, a large proportion of funds reduce their weight in a particular market whenever returns in *other* markets rise.

Turning next to the six-month horizon, Panels E-H in Table 3 show very different results. The median own-market return coefficient is now large and negative both for Japan and North America (Panels G and H). In Europe and Asia-Pacific, there is no longer evidence of significant positive coefficients on own-market returns.

These findings bear an interesting comparison with the results reported by Bohn & Tesar (1996) who consider a related decomposition of net purchases into changes in desired portfolio weights and portfolio rebalancing terms. They find that net purchases are significantly positively correlated with short-run local capital gains and excess returns. Rather than selling assets in countries with above-average returns to rebalance their international portfolios, US investors appear to increase investments in markets that have recently experienced high relative returns. Our findings suggest that, although in the very short run our sample of investors tended to be trend-chasers in foreign markets and a momentum effect thus is important, over the longer term they appear to behave like 'reverse trend-chasers', pulling out of regions whose returns over the previous six months had been relatively high. It is difficult to explain this behavior in terms of differential information effects between domestic and foreign investors. Clearly state variables other than past returns are required to track investors' expectations of future returns.

D. Portfolio Weights and Time-varying Investment Opportunities

The analysis in the previous section suggests that past returns do not fully explain UK pension funds' international asset allocation decisions. In this section, we address the issue of whether the dynamics of portfolio changes reflect time variation in the conditional moments of stock returns in the four regions. To quote Brennan, Schwartz & Lagnado (1997): "A *sine qua non* of tactical asset allocation is time variation or predictability in expected asset returns" (page 1378). The theory of mean-variance optimizing investors' portfolio behavior implies that optimal portfolio weights should reflect the conditional correlation structure of international asset returns, conditional expected asset returns and a set of hedge factors, c.f. Solnik (1974), Stulz (1981) and Adler & Dumas (1983). This suggests that UK pension funds' decisions may, at least in part, have been driven by time-varying expected moments of returns.

While there is no consensus on how best to model the conditional moments of asset returns, there is now strong evidence that the investment opportunity set in most countries displays considerable time variation, c.f. Harvey (1991), Bekaert & Hodrick (1992), Campbell & Hamao (1992) and Ferson & Harvey (1993). Return correlations also appear to increase in bear markets, c.f. Erb, Harvey & Viskanta (1994), Lin, Engle & Ito (1994) and Longin & Solnik (1995). Some studies suggest that changes in the investment frontier can be permanent. For example, Erb et al. (1994) report an upward trend in the conditional correlation between UK stock returns and those of Germany, Italy, France and the US. Studies such as Dumas & Solnik (1995) and De Santis & Gerard (1998) find that foreign exchange risk is priced in equilibrium and may vary substantially over time.

We follow the literature, most notably Harvey (1991) and Bohn & Tesar (1996), and model expected returns in each region as a function of a set of commonly used state variables. As instruments we use an intercept term, the quality spread or default premium (Def_t) on US bonds computed as the differential yield on Baa and Aaa rated bonds, the 1-month US T-bill rate (I_t^{us}) and the US-UK T-bill spread ($I_t^{us} - I_t^{uk}$). Finally, we include the local dividend yield in each region ($Yield_{jt}$). These instruments are very similar to those adopted by Harvey, with the exception of the T-bill spread between the US and UK markets which is included to reflect a key information variable from the perspective of UK investors.[8] All returns are denominated in Sterling to reflect the objectives of a UK pension fund. Hence the specification of the conditional mean in our regressions is:

$$r_{jt+1} = \gamma_{0j} + \gamma_{1j} Yield_{jt} + \gamma_{2j} Def_t + \gamma_{3j} I_t^{us} + \gamma_{4j}(I_t^{us} - I_t^{uk}) + \eta_{jt+1} \tag{10}$$

To capture possible time variations in conditional volatilities and covariances, we model returns in the context of a bivariate generalized ARCH model. Since UK pension funds hold upwards of 70 percent of their funds in domestic assets and this component is dominated by UK equities, the contribution of foreign equity holdings to UK pension funds' total volatility is determined in part by their own volatility and in part by their covariance with the UK stock market. Let $\mathbf{r}_{t+1} = (r_{jt+1}\ r_{ukt+1})'$, where r_{jt+1} and r_{ukt+1} are region j and UK returns in month $t+1$, respectively. We follow Bollerslev (1990) and model returns as follows:

$$\begin{aligned} \mathbf{r}_{t+1} &= \Gamma \mathbf{Z}_t + \boldsymbol{\eta}_{t+1}, \\ \sigma_{kk,t} &= \alpha_{kk} + \beta_{k0}\eta_{kt}^2 + \beta_{k1}\sigma_{kk,t-1}, \\ \sigma_{kl,t} &= \psi_{kl}\sqrt{\sigma_{kk,t}\sigma_{ll,t}},\ k, l = j, uk \end{aligned} \tag{11}$$

[8] Returns and dividend yields were obtained from Morgan Stanley Capital International. The quality spread is based on data from DRI, the US T-bill rate is from the CRSP tapes, while the UK T-bill rate is from DataSTREAM.

where $\boldsymbol{\eta}_{t+1} = (\eta_{jt+1}, \eta_{ukt+1})'$ is the set of heteroskedastic return innovations defined as $\boldsymbol{\eta}_{kt+1} = \sqrt{\sigma_{kk,t}}\varepsilon_{kt+1}$, where $\boldsymbol{\varepsilon}_{t+1} = (\varepsilon_{jt+1}, \varepsilon_{ukt+1})'$ are normal, independent and identically distributed residuals so that $\boldsymbol{\eta}_{t+1} \sim N(0, \Sigma_t)$. $\Sigma_t = [\sigma_{kl,t}]$ is the conditional covariance matrix and ψ_{kl} is the conditional correlation coefficient which is assumed to be constant. Finally $\mathbf{Z}_t = \{Yield_{jt},\ Def_t,\ I_t^{us}, I_t^{us} - I_u^{uk}\}$ is a vector of instruments while Γ is a conformable matrix of regression coefficients. This model generates an estimate of the expected returns and conditional volatility of returns in each region as well as its conditional covariance with UK returns.

Table 4 reports the outcome from these regressions estimated on data over the sample 1970:1 to 1997:12. The default premium variable is highly significant with a positive coefficient in all regions, while the 1-month T-bill rate has a negative and significant coefficient in all regions. Local dividend yields seem to be important only for Asia-Pacific. ARCH effects are strong and volatility persistent in Japan, Asia-Pacific and the UK.

Armed with these measures we next investigate the relation between portfolio weights and the expected return in each region, the conditional volatility of the return within the region as well as the region's conditional covariance with returns on the UK stock market. This analysis extends the work by Bohn & Tesar (1996) which focused on conditional means but did not include an estimate of conditional volatility and covariance. Each of the conditional moments was computed in excess of the corresponding average 'world ex-UK' moment computed as a capitalization-weighted average across the four regions.[9] We include expected own-market excess returns rather than the separate expected returns for all markets to reduce the number of parameters to be estimated.

We do not experiment with different specifications but simply use linear projections of portfolio weights on first and second conditional moments as an approximation to a relationship between portfolio weights and conditional moments that could be both more complex and vary over time.[10] As a result we may underestimate the predictive power of these moments over portfolio weights. Consistent with

[9] Let $\mathbf{e}_j$ be a 4×1 vector with a one in the jth row and zeros elsewhere, let $\boldsymbol{\mu}_t$ be the vector of expected returns for period t, and let $\boldsymbol{\omega}_t$ be the vector of world capitalization weights for the four regions under investigation (rescaled to sum to unity). Then expected excess returns in region j were computed as $\rho_{jt}^e = (\mathbf{e}_j' - \boldsymbol{\omega}_t')\boldsymbol{\mu}_t$.

[10] Notice also that, unlike our data, the stochastic process assumed in the return generating process in Brennan et al. (1997) assumes constant volatility.

theoretical models of intertemporal asset allocation (e.g. Brennan et. al (1997)), we use portfolio weights as dependent variables. For each fund and each region we estimated a set of time-series regressions:

$$\omega_{ijt} = \alpha_{ij} + \beta_{1ij}\widehat{\rho}_{jt} + \beta_{2ij}\widehat{\sigma}_{jj,t}^{1/2} + \beta_{3ij}\widehat{\sigma}_{juk,t}^{1/2} + \varepsilon_{ijt}, \tag{12}$$

where $\widehat{\rho}_{jt}$ is the expected excess return in region j, while $\widehat{\sigma}_{jj,t}^{1/2}$ is the conditional return volatility and $\widehat{\sigma}_{juk,t}^{1/2}$ is the conditional covariance with UK equity returns, all estimated from (11). These moments are based on information at time $t-1$ and hence are known by the time the fund decides on ω_{ijt}.

Table 5 summarizes the empirical results. Panel A projects portfolio weights onto expected returns. More than 90 percent of all funds generated a positive coefficent on expected returns in North America and Europe, while only 29 percent did so for Japan and none for Asia-Pacific. The result for Asia-Pacific may initially seem puzzling, but is related to the importance of the dividend yield in the bivariate GARCH model for this region and its behavior during the Asian crisis of 1997. During 1997, Asian stock prices plummeted and the dividend yield rose sharply: the outcome was both a sharp fall in the Asia-Pacific portfolio weight and an increase in expected returns. Most remarkable perhaps is the fact that 94 and 80 percent of the expected return coefficients are statistically significant and positive for North America and Europe, respectively. Since these two regions account for around 75 percent of the total foreign equity holdings of our sample of pension funds, we can conclude that expected return variation is a significant determining factor of the international asset allocation of the vast majority of funds.

We next included the conditional own-market volatility as a regressor in the portfolio weight equation. Panel B shows that the vast majority of funds - in excess of 92 percent in all regions - generated a negative coefficient on volatility, indicating that the funds decreased their allocation towards regions whose volatility was expected to go up. Between 6 percent (Europe) and 82 percent (Japan) of the volatility coefficients were negative and statistically significant.

Panel C reports the outcome from using expected returns and conditional covariance with UK stock returns as regressors. Conditional covariances seem relatively less important than own-market volatility with one important exception: in the case of European equity weights, 99.6 percent of the funds produced a negative coefficient on the conditional covariance. For 94 percent of the funds this regressor was statistically significant. Furthermore, the median R^2 of the portfolio weight regression doubles for the European portfolio weight regressions when conditional covariance is added to expected returns as a regressor.[11,12] For completeness, Panel D shows the results from regressions that include all three explanatory variables. The sign of the separate coefficients on the volatility and covariance are now difficult to interpret since the two series are driven by a common component and hence strongly correlated.

To summarize the importance to the evolution in portfolio weights of the time variation in conditionally expected returns, volatilities and covariances, Figure 3 shows the cross-sectional distribution of R^2 values from these regressions performed for each fund. The R^2 values are high in all four regions with medians of 0.59 in North America, 0.70 in Asia, 0.48 in Japan and 0.29 in Europe.[13]

Once again it is interesting to compare these findings with the results in Bohn & Tesar (1996). In regressions of net purchases of foreign equity on predicted excess returns in a cross-section of countries, Bohn and Tesar found that expected excess returns were statistically significant and positively correlated with net purchases in roughly a third of the countries they examined. It is difficult to compare directly their findings on aggregate flows with our results on individual funds since they do not report R^2 statistics and also do not include time-varying second moments.

[11]In a panel analysis of equity flows, Portes & Rey (1999) find that equity flows between pairs of countries do not seem to be determined by the correlation between equity returns in the two countries, while volatility of returns in the two markets does matter. Our findings suggest that conditional covariances between returns in the host and foreign country do influence portfolio holdings. The difference between these findings may be explained by the use of time-varying conditional moments.

[12]We also estimated a specification that included the conditional covariance with the world index instead of the conditional covariance with UK stock returns. However, we found that the world covariance did not possess the same predictive power over portfolio weights as the UK covariance.

[13]Brennan & Cao (1997) conclude that their model based on lagged returns "is able to explain only a small portion of the variance of international equity flows" (page 1876). To explore the relative importance to portfolio weights of time variations in expected returns and second moments relative to past returns, we also estimated regressions that include the most recent returns (namely eqn. (12) with lagged returns added). The median R^2 increased only marginally from between 0.00 to 0.02 for the four regions, thereby suggesting that time-varying conditional moments, rather than lagged returns, are more important for explaining individual funds' asset allocation decisions.

Nevertheless, the fact that expected returns matter to almost all funds in at least one market (North America) suggests that time-varying expected moments may be more important at the level of individual institutional investors' asset allocation decisions than in the aggregate.

IV. Returns from International Market Timing

A. Unconditional Return Performance

To assess the performance of the funds in its stable, WM uses a range of value-weighted asset-class benchmarks. As external benchmarks it employs Financial Times/Standard & Poor indices, all of which assume that income is reinvested (gross of tax). However, the choice of external index to represent the benchmark returns is far from obvious: Kang & Stulz (1997), for example, show that foreign investors' holdings of Japanese equities are concentrated in the largest firms.

Figure 4 plots time series of monthly returns on the value-weighted portfolio of funds included in our sample. Also shown in the figure are returns on the corresponding FT/S&P indices. The pairs of return series are clearly strongly correlated. This impression is confirmed by the sample correlations reported in the last row of Table 6. Estimated time-series correlations between the FT/S&P indices and the returns on the value-weighted portfolio of pension funds exceed 0.97 and are as high as 0.99. In light of these high correlations, we do not consider alternative candidates for benchmark returns.

Table 6 also reveals large variations across markets in the individual funds' mean returns and volatilities relative to the external indices. First consider the mean returns. For Japan, the value-weighted sample mean return was 2.85 percent over the period, while the corresponding FT/S&P index paid an average of -0.73 percent per annum.[14] However, this region is the only one in which a typical UK pension

[14] The apparent exceptional outperformance in the Japanese stock market can be explained as follows. Although there are no legal constraints on foreign holdings in Japanese bank stocks, UK pension funds were underweight in the Japanese banking sector as a result of the small percentage of this sector's stock available for public purchase (a consequence of the high degree of cross-holdings in Japanese banking sector equities). This matters because Japanese banks paid exceptionally low returns over the sample period. In a two-factor regression of the funds' excess returns in Japan on the Japanese stock market index and the banking sector index, only 0.4 percent of funds had significantly positive Jensen alpha estimates. The results may also be

fund earned a higher mean return than the benchmark index. In the other regions, the pension funds underperformed the FT/S&P index on a raw return basis by an average of 0.43 (North America), 0.50 (Europe) and 2.06 (Asia-Pacific) percentage points per annum.[15] For the total international equity portfolio, UK pension funds underperformed by 0.70 percentage points per annum an index formed by weighting the four regions' FT/S&P returns by their international market shares: only 13 percent of the funds outperformed the passive world market portfolio.

Figure 5 presents histograms of the individual funds' annualized mean returns. In line with the clusters found in the cross-sectional distribution of portfolio weights, the cross-sectional distribution of mean return performance in Japan is bimodal with clusters of funds centered around 1 and 5 percent per annum. A similar bimodality is observed for Asia-Pacific, although not for North America or Europe. Such differences can only reflect disparities in choice of assets within particular markets. Hence funds differ not only in their strategic asset allocation towards major international regions, as shown in Figure 2, but also in their choice of equities within each region.

B. Conditional Market Timing Tests

To test whether UK pension funds possess market timing skills after controlling for public information, we ran a battery of tests inspired by Graham & Harvey (1996). Their regressions were designed to measure the market timing skills of newsletters recommending stocks versus cash and hence assume the existence of a single risky asset. Since we consider the allocation between four risky assets, we have to modify these measures. Initially we regress returns in each region in excess of the World (ex-UK) return, $\rho_{jt+1} = r_{jt+1} - r_{\omega t+1}$, on the previous period's portfolio weight change and the vector of instruments. Excess returns relative to the average foreign market are used in the regression since return-maximizing funds

explained by the fact, documented by Kang and Stulz (1997), that foreigners in the Japanese stock market tend to hold the equities of large firms. During the 1990s, large Japanese firms paid higher returns than small firms. When we controlled for a capitalization factor, once again we found that only 0.6 percent of the funds generated statistically significant outperformance.

[15]These differences show up clearly in the proportion of funds that outperformed the indices on a raw return basis: 97 percent of funds outperformed the index in Japanese equities, while only 20, 21 and 9 percent of the funds outperformed the FT/S&P indices in North America, Europe and Asia-Pacific, respectively.

ought to increase allocations to regions with above-average foreign asset returns:

$$\rho_{jt+1} = c_j + \beta_{1j}\Delta\omega_{jt} + \beta_j'\mathbf{Z}_t + \varepsilon_{jt+1}. \tag{13}$$

This regression tests whether funds successfully change their portfolio weights in anticipation of future relative returns in the various markets, after controlling for the publicly known state variables, $\mathbf{Z}_t$, listed in Section 3. Market timing skills should show up in the form of a positive coefficient estimate, $\widehat{\beta}_{1j}$.

Panel A of Table 7 shows that on this measure there is some evidence that UK investors possessed market timing skills: the median estimate of β_{1j}, computed across individual funds, is positive for three out of four regions, the exception not surprisingly being North America. Furthermore, the percentage of funds with positive estimates of market timing skills is very high in Japan (91 percent of all funds), Asia (89 percent) and relatively high in Europe (61 percent). In contrast, only 30 percent of funds obtained a positive market timing coefficient for North America. However, the percentage of funds with estimates of β_{1j} that are statistically significant and positive at the 5% level is quite low (below 6 percent in all regions).

In the presence of multiple risky assets, it is possible that investors do not simply increase their allocation towards the asset with the highest return and instead choose the asset with the highest expected return per unit of risk. We investigated this possibility by normalizing the future returns either by the *ex-ante* expected own-market volatility ($\widehat{\sigma}_{jj,t+1}^{1/2}$) or by the conditional covariance with returns on UK stocks ($\widehat{\sigma}_{juk,t+1}^{1/2}$), both obtained from the bivariate GARCH model (11). The results, reported in Panels B and C of Table 7, do not change very much, suggesting that the evidence on market timing is robust in the presence of time-varying risk.

We next conducted a Merton-style market timing regression based on indicator functions for the sign of future returns. Let $I_{\{\rho_{jt+1}\geq 0\}}$ be an indicator function for the event that future realized excess returns in region j relative to the world market average is positive or zero, while $I_{\{\rho_{jt+1}<0\}}$ is an indicator for the situation where future excess returns are negative. We estimated regressions

$$\Delta\omega_{jt} = \beta_{1j}I_{\{\rho_{jt+1}\geq 0\}} + \beta_{2j}I_{\{\rho_{jt+1}<0\}} + \varepsilon_{jt}. \tag{14}$$

A measure of market timing skills, proposed in this context by Graham & Harvey (1996), is whether $\beta_{1j} > 0$, in which case the asset allocation to markets with positive future excess returns is increased and $\beta_{1j} < 0$, in which case the allocation to markets with negative future excess returns is decreased. Panel A of Table 8 shows that over 90% of all funds generated positive estimates of β_{1j} for Japan, Europe and Asia; again the exception was North America for which only eight percent of funds obtained a positive estimate of β_{1j}. Likewise, there is strong evidence that the funds successfully timed periods with negative excess returns, the proportion of negative coefficient estimates of β_{2j} ranging from 65 to 99 percent.

These regressions have to be interpreted with caution, however. For instance, the large percentage of funds generating a negative estimate of β_{2j} for North America is likely to reflect the long-run strategic asset allocation decision of the funds to pull out of North America. This is different from tactical asset allocation skills as reflected in the ability to successfully switch in and out of markets in the short run according to the anticipated sign of future returns. A test of the tactical asset allocation skills based on the independence between the sign of the portfolio weight change and the sign of future returns, was proposed by Henriksson & Merton (1981) and generalized to account for sampling variation in the estimated 'hit rate' by Pesaran & Timmermann (1992). We report the outcome of this test in Panel B of Table 8. When applied to the four regions, we find very weak evidence of market timing skills. Only for Europe did more than 5 percent of the funds generate a positive and significant value for this test statistic.

Equation (14) is also subject to the criticism that any market timing skills reflected in the funds' portfolio weight changes might simply reflect publicly available information. To see if the funds possessed market timing skills over and above that contained in public information, we follow Graham & Harvey (1996) and regress the current portfolio weight change on indicators for the sign of the unanticipated future return component, $\rho^u_{jt+1} = \rho_{jt+1} - \rho^e_{jt+1}$, as well as the anticipated part, ρ^e_{jt+1}, computed from the regression of excess returns (11) on the lagged instruments:[16]

$$\Delta\omega_{jt} = \beta_{1j} I_{\{\rho^u_{jt+1} \geq 0\}} + \beta_{2j} I_{\{\rho^u_{jt+1} < 0\}} + \beta_{3j} I_{\{\rho^e_{jt+1} \geq 0\}} + \varepsilon_{jt}, \tag{15}$$

If funds can predict the part of future differential returns unaccounted for by current public information, β_{1j} should be positive and β_{2j} should be negative.

Table 9 shows very little evidence of extra-market timing skills. While 80 and

[16] We do not include a fourth indicator $I_{\{\rho^e_{jt+1} < 0\}}$ since in our application the pair of indicator functions $I_{\{\rho^u_{jt+1} \geq 0\}}$ and $I_{\{\rho^u_{jt+1} < 0\}}$ always sum to unity. Adding both $I_{\{\rho^e_{jt+1} \geq 0\}}$ and $I_{\{\rho^e_{jt+1} < 0\}}$ would hence make the regression perfectly collinear.

92 percent of the funds generated positive estimates of β_{1j} for Japan and Asia-Pacific, only 0 and 2 percent of the funds did so for North America and Europe, respectively. Even weaker evidence emerges for the market timing skills in down markets. Here there is only evidence of market timing skills in North America (95 percent) but negative evidence for Japan (15 percent), Europe (2 percent) and Asia (5 percent). Furthermore, some of the market timing ability shown in Table 8 appears to reflect publicly available information as evidenced by the many positive estimates of β_{3j} for Japan and Europe.[17]

Finally, we tested whether the funds correctly increase their portfolio weights the most for the region whose return next period is highest, or conversely whether they correctly decrease their weights the most for the region with the smallest future return. We conducted this test using a simple χ^2-test based on the diagonal cells in the 4x4 contingency table matching realized returns, r_{jt+1}, against weight changes, $\Delta\omega_{jt}$, in each of the four regions. We found that, using a 5% critical level, no fund showed ability to consistently anticipate the market with the highest return, while only one out of 247 funds seemed able to anticipate which market would pay the lowest return.

The evidence so far suggests that there is only very weak evidence of genuine market timing skills. However, it also suggests that UK pension funds may simply have followed the predictions from standard models of expected returns in revising their portfolio weights. To measure the total returns from the extra-market timing activities, we compute for each region the return from the part of the portfolio weights that is orthgonal to time-varying moments, $\omega_{ijt}^{u} = \omega_{ijt} - \widehat{\omega}_{ijt}$, where $\widehat{\omega}_{ijt}$ is the projection of ω_{ijt} on the conditional mean, variance and covariance from equation (12), rescaled to sum to unity. For each fund (i) the ω_{ijt}^{u} sum to zero (across j) and these weights therefore represent a zero-investment portfolio. Summing across regions gives a measure of the total return to the zero-investment portfolio that tracks extra-market timing skills:

[17] We also applied the Henriksson-Merton test to the relationship between the sign of the portfolio weight change and the unexpected future excess return. Compared with the results for the total future excess returns, the results, which are shown in Panel C of Table 8, are even weaker. Only 2, 0, 4 and 2 percent of the funds generated a significant value of this market timing test in the four regional markets.

$$\sum_{j=1}^{4} \omega_{ijt}^{u} r_{jt}. \tag{16}$$

The median of the time-series average of this measure is -0.39 percent per annum. Figure 6 provides a histogram of the statistic, demonstrating that there are two clusters of funds, one with a mean return from extra-market timing of around -0.90 percent per year, and another with mean returns of around -0.4 percent per year. These reflect the clusters in regional portfolio weights and raw performance that we reported earlier. Only 18 out of 247 or seven percent of the funds generated positive mean returns from extra-market timing. None of these time-series means was individually statistically significant, however.

V. Conclusion

Several new insights into institutional investors' behavior and performance in foreign equity markets have resulted from this study. Most of the existing literature has attempted to model international investment flows in terms of *past* returns and we find evidence of short-run return chasing. However, we also find that return chasing does not appear to be a significant determinant of our sample of institutional investors' longer-term strategic asset allocation decisions. Instead we find that portfolio weights are much more highly correlated with time-varying expected returns, volatilities and conditional covariances with domestic equity returns.

We decomposed the investors' market timing activity into two parts: that due to time-varying moments and that due to extra-market timing. The decision by UK pension funds to withdraw from North America and increase their allocation towards Europe appears to reflect the time-series of expected returns in these markets in excess of expected returns in other foreign markets. Since the *ex post* realized returns in the North American market were very high during the sample, a Bayesian learning model of the kind advanced by Brennan and Cao (1997) cannot be used to explain this strategy.

Our findings should be seen in the light of Barry Riley's comments in the *Financial Times* of 16 December 1997: "It is worth setting out once again why London's global managers have been so sceptical of Wall Street in the 1990s. But more fundamentally, what 1997 has shown so clearly, not just in the US, is that when value managers lose control of their markets they can flounder for extended periods. Investor attitudes were formed in the 1980s when the US stock market was the worst performing of the major global markets. The US was written off by many foreigners as slow-growing and inefficient, albeit with an interesting technology sector." Taken at face value, these comments support the argument that by not holding the global weight in the US stock market, UK pension funds' market timing activity cost them an average of 3 per cent per annum. However, we show that when we orthogonalize portfolio movements with respect to time-varying moments, the extra-market timing performance was -0.4 per cent per annum. There is still a loss from international market timing, but it is much lower than unconditional estimates indicate.

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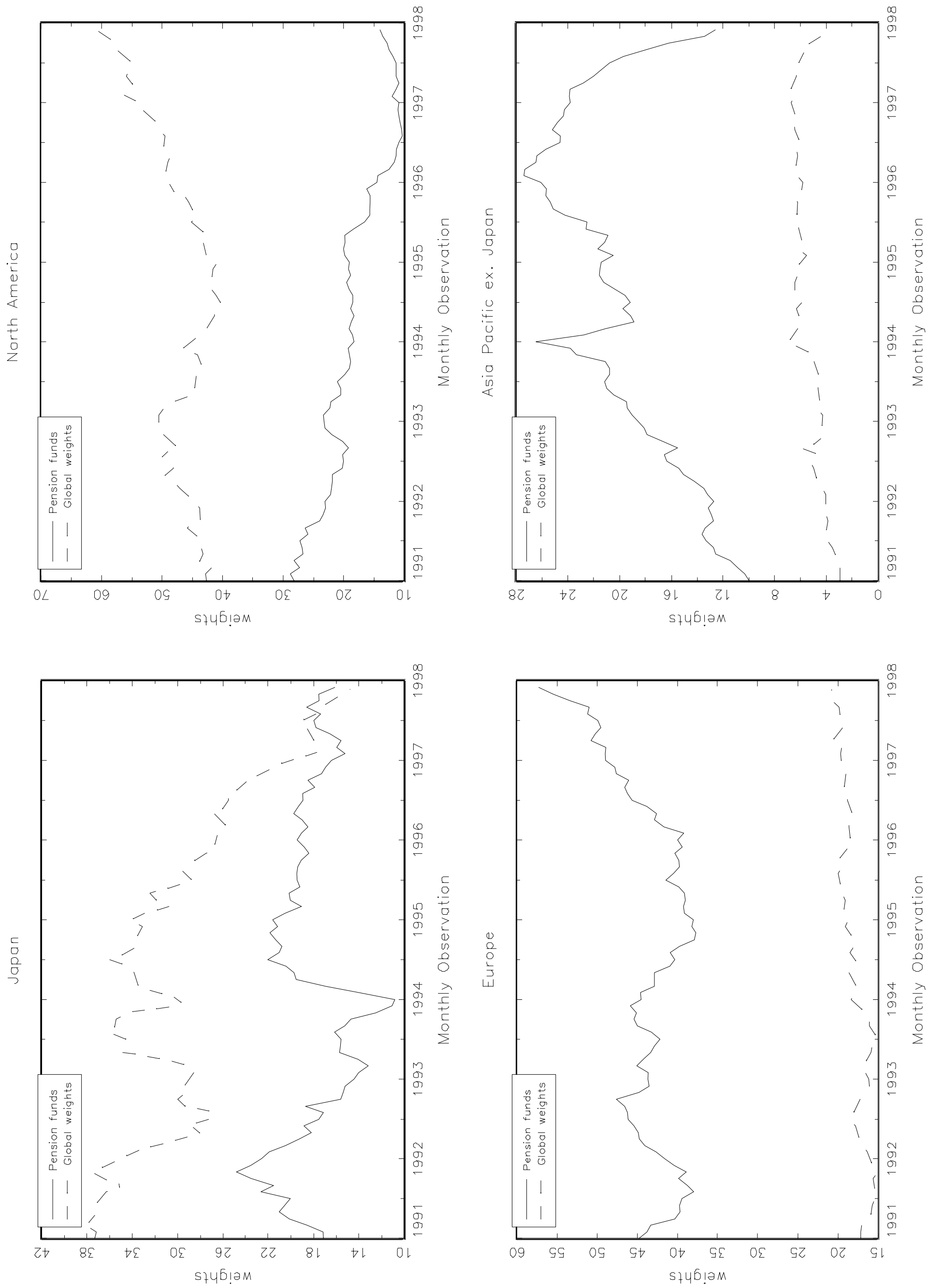

Figure 1: UK pension funds' portfolio weights and global market weights

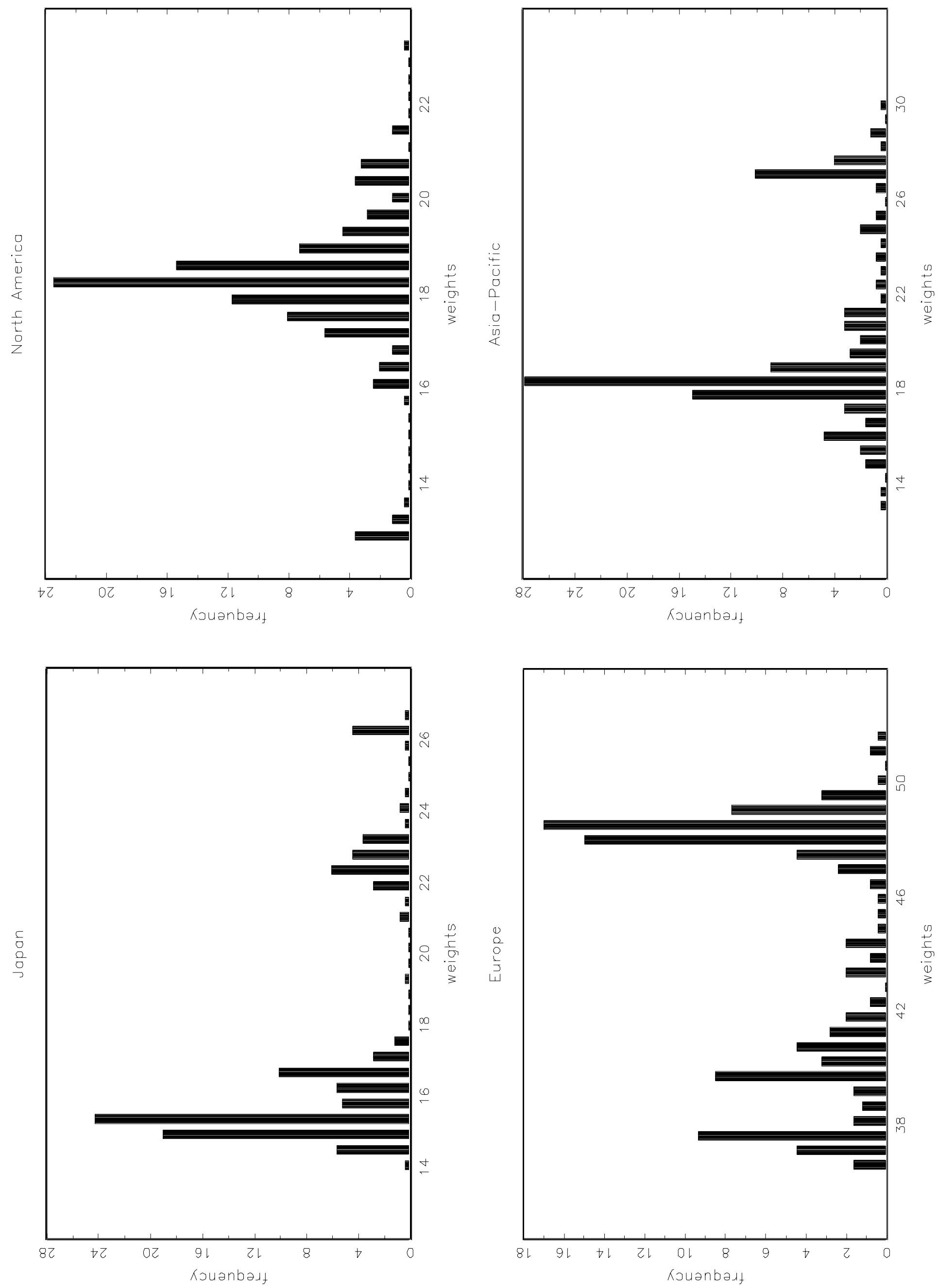

Figure 2: Histogram of individual funds' average portfolio weights

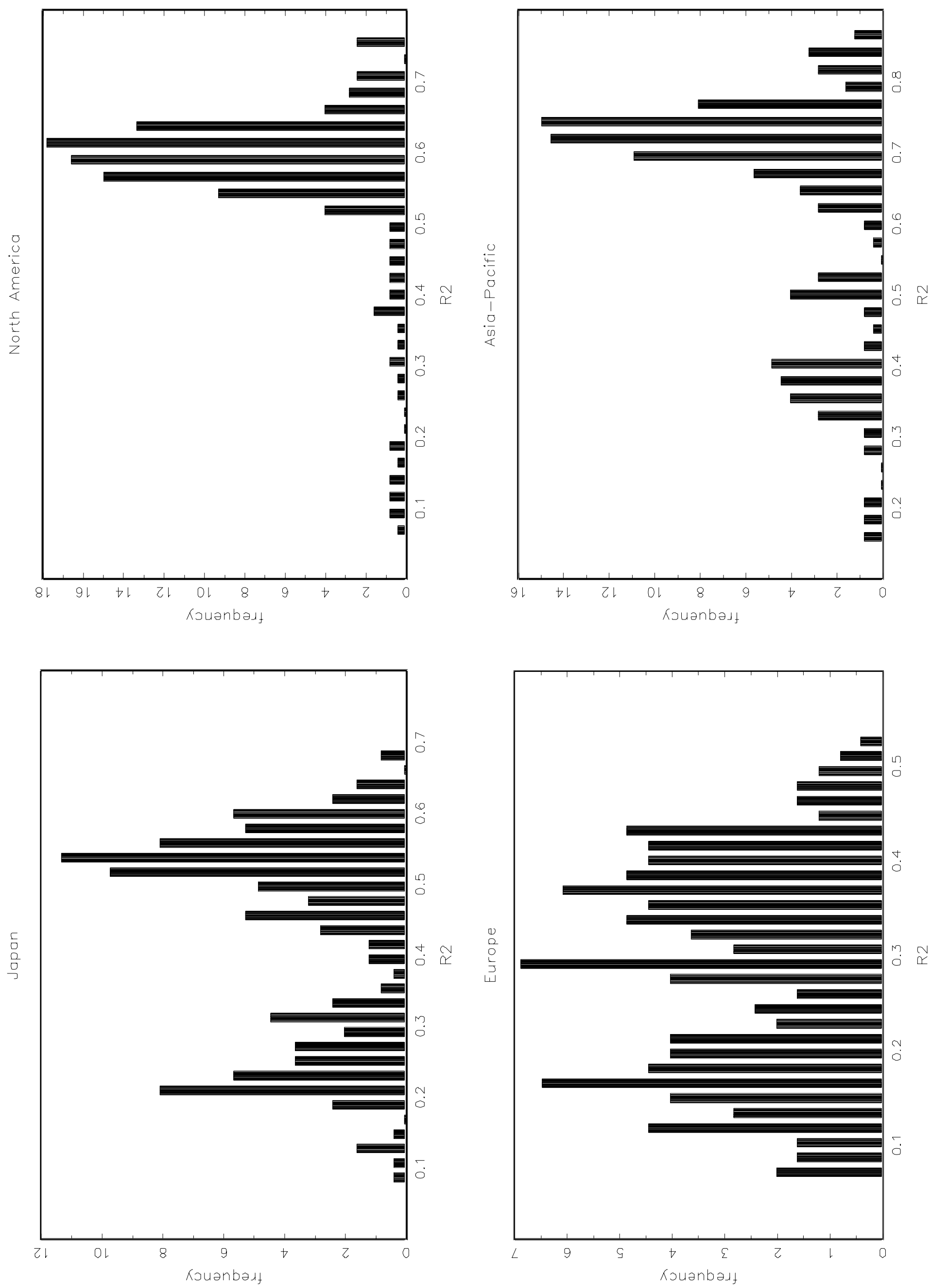

Figure 3: Histogram of R^2 from regressions of portfolio weights on conditional moments

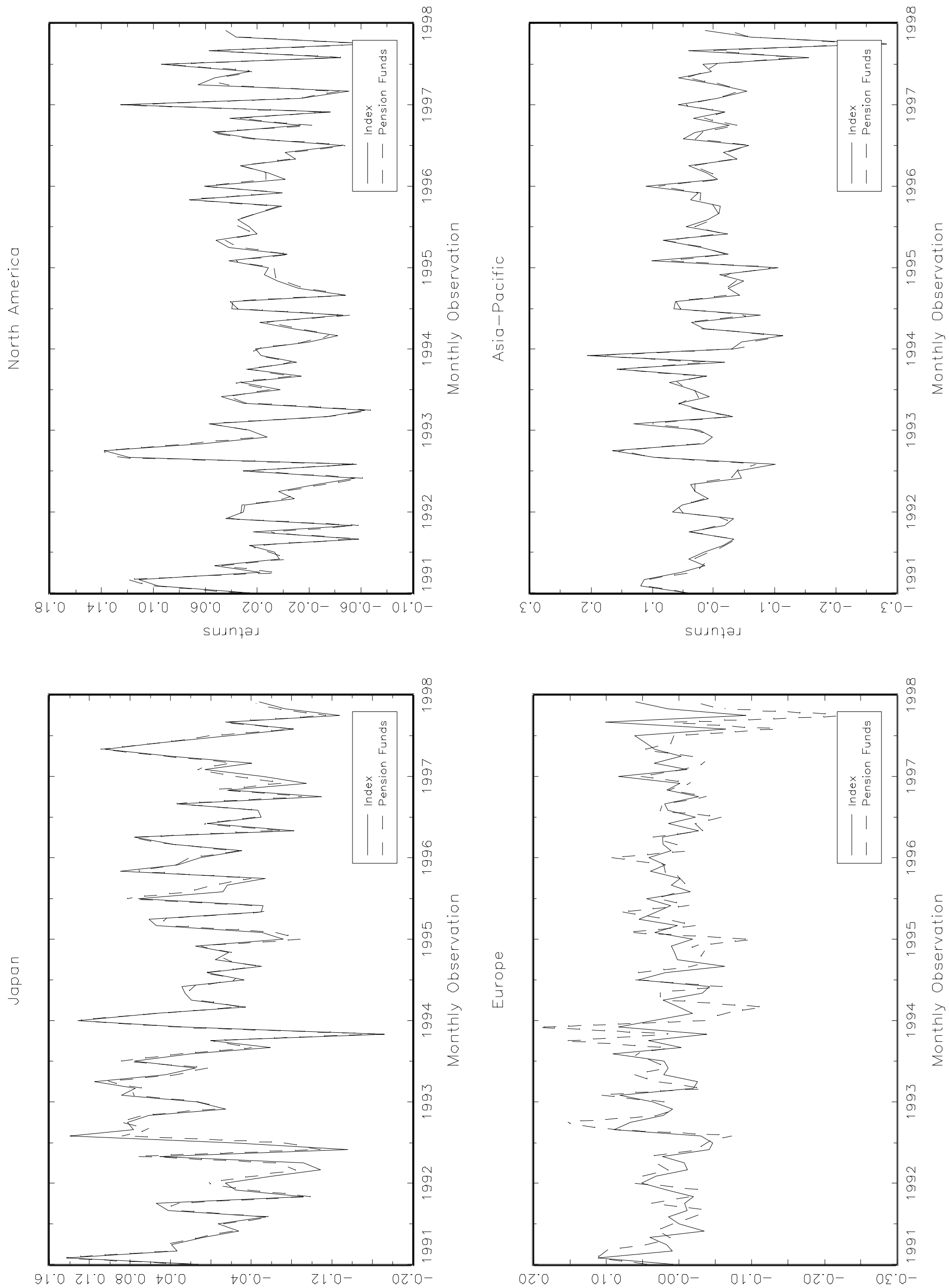

Figure 4: Time-series of returns on FT/SP indices and on UK pension funds

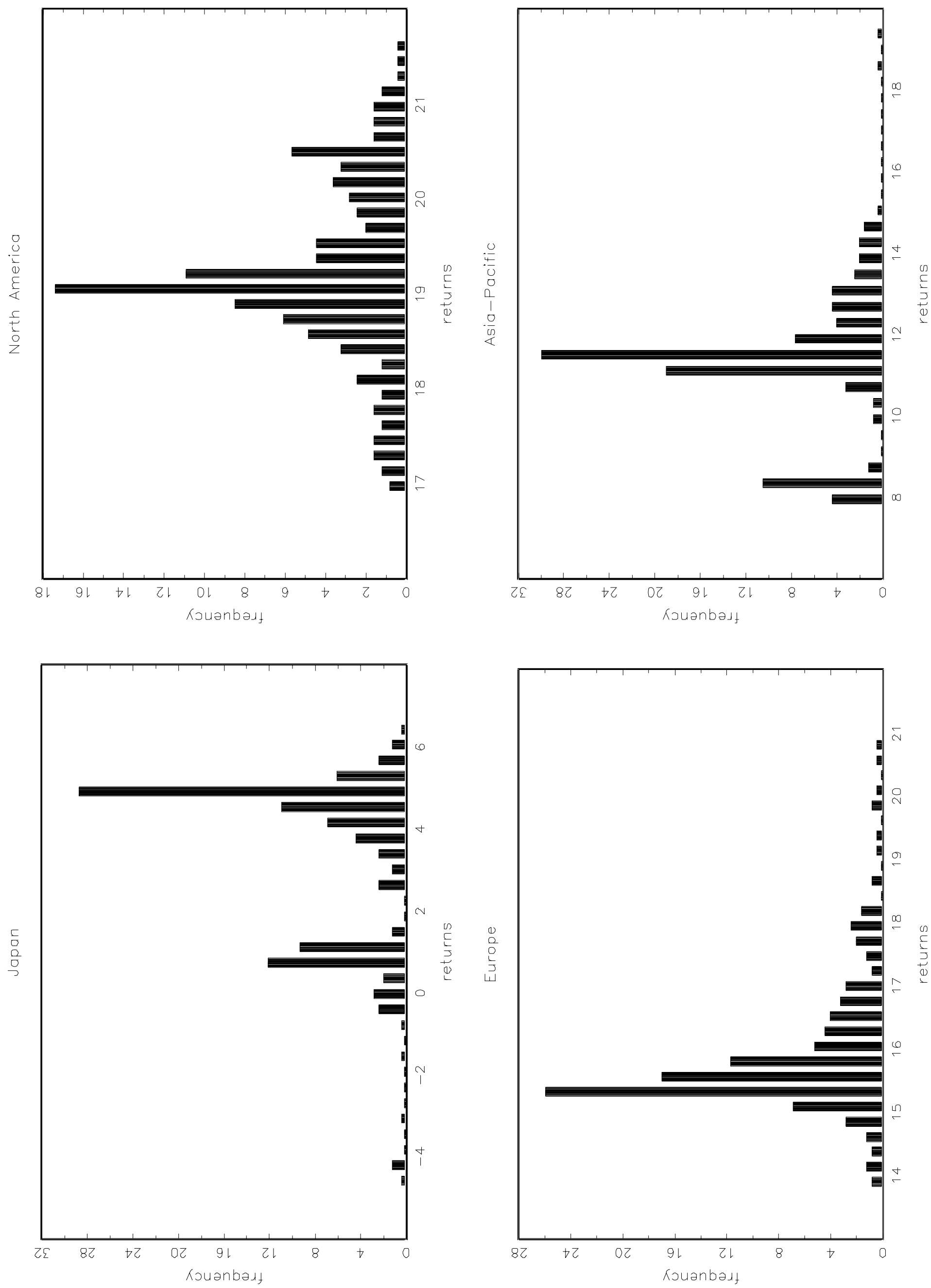

Figure 5: Histogram of individual funds' mean returns in four regional markets

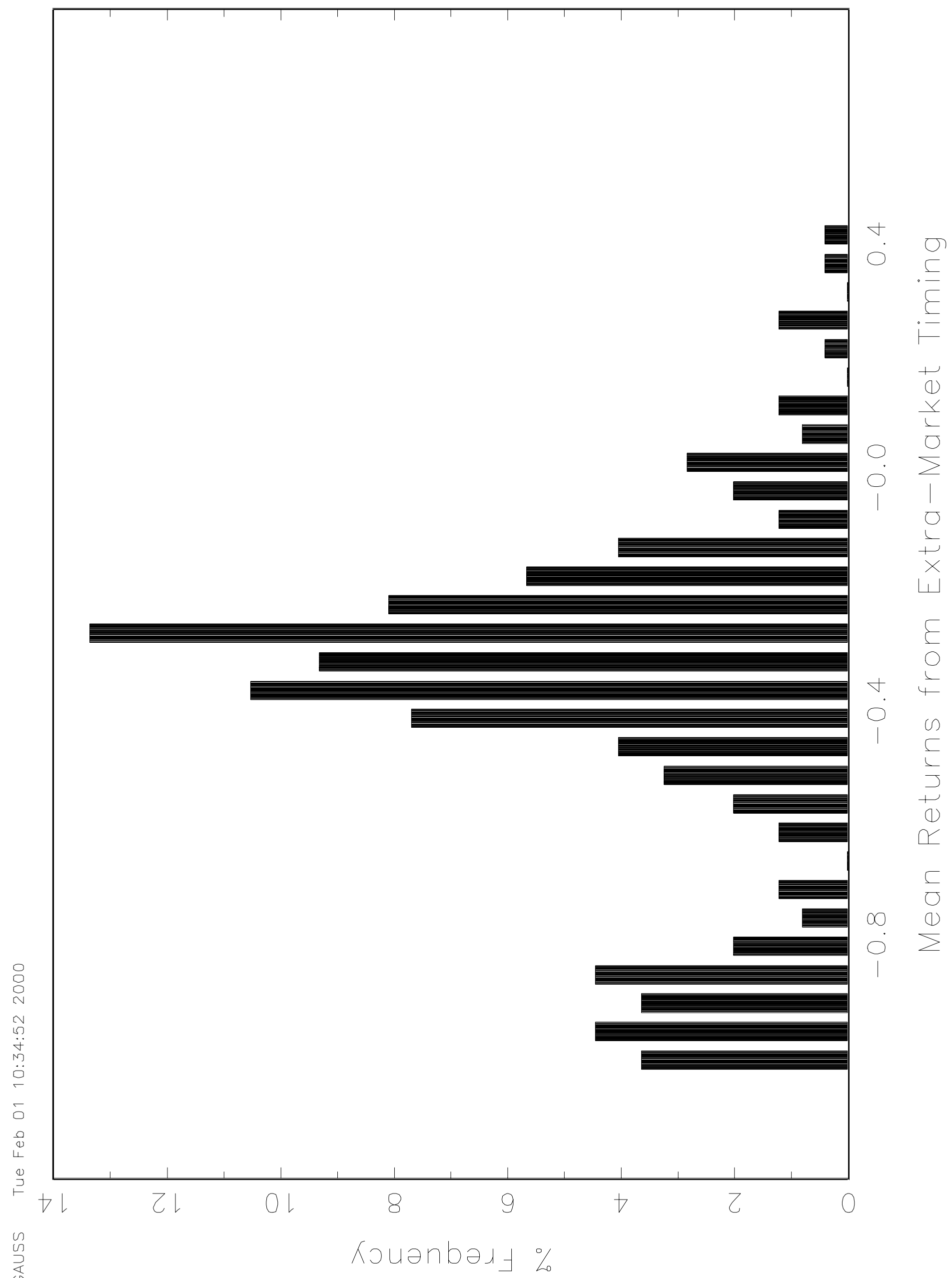

Figure 6: Cross-sectional distribution of mean returns from extra-market timing

Table 1. Identifying the Sources of Changes to Aggregate Portfolio Weights Across Regions

	Japan	North America	Europe	Asia-Pacific
A. Average annual basis point change in portfolio weight	-14	-202	179	37
B. Mean annual percentage change in portfolio weight	-0.88	-10.07	3.58	3.31
- due to differential returns	-8.96	6.90	3.13	-0.57
- due to net cash flow differentials	8.08	-16.97	0.45	3.88
C. Variance decomposition: Percentage of portfolio weight change due to differential returns	31.51	24.46	29.05	39.02

Note: Each month a value-weighted aggregate portfolio was formed by aggregating the individual pension funds' investments. The time-series of the aggregate portfolio holdings were then used to compute portfolio weights for each region. The results in this table are based on the decomposition of percentage changes in aggregate portfolio weights into a return differential component and a net cash flow component (see equation (4) in the text). The variance decomposition is based on equation (5) in the text. The sample covers 247 UK pension funds over the period 1991:1 - 1997:12.

Table 2. Individual Funds' Portfolio Weights: Cross-sectional Dynamics

	Japan	North America	Europe	Asia-Pacific
A. Year-on-year Markov switching (stayer) probabilities				
Probability	0.738	0.650	0.780	0.679
(standard error)	(0.132)	(0.121)	(0.083)	(0.045)
B. Beginning-to-terminal probabilities				
Probability	0.439	0.520	0.699	0.675

Note: Using the individual funds' portfolio weights in December each year, we computed the proportion of funds with an above-average weight in a given asset class that continued to have an above-average weight 12 months later. Time series averages of these transition probability estimates are reported as year-on-year stayer probabilities in Panel A.Their standard errors are also based on the time series of these transition probabilities. Panel B reports the corresponding transition probability estimates for the event that a fund which initially has an above-average portfolio weight in a given asset at the end of the sample. The sample covers 247 UK pension funds over the period 1991:1 - 1997:12.

Table 3. Dynamics of Individual Pension Funds' Net Cash Flows

I. 1-month return horizon

Percentage of significant regressors

	Regressors:									
	Intercept		Return in Japan		Return in North Am.		Return in Europe		Return in Asia-Pacific	
	positive	negative	positive	negative	positive	negative	positive	negative	positive	negative
A. Multivariate Regressions (equation (8))										
Japan	0.0	0.0	12.2	0.0	0.0	1.6	0.0	4.9	4.9	4.5
North America	0.0	38.1	1.2	0.0	1.6	0.0	0.8	0.0	4.5	2.8
Europe	0.0	0.0	0.0	3.6	0.0	14.6	12.6	0.8	0.0	0.0
Asia-Pacific	0.0	0.0	0.8	0.4	0.4	0.0	0.0	3.6	6.5	0.8
B. Univariate Regressions (equation (9))										
Japan	0.0	0.0	7.3	0.0						
North America	0.0	51.8			2.0	0.8				
Europe	0.0	0.0					9.3	0.0		
Asia-Pacific	0.4	0.0							8.1	0.8

Median coefficient estimates

	Intercept	Return in Japan	Return in North Am.	Return in Europe	Return in Asia-Pacific
C. Multivariate Regressions (equation (8))					
Japan	0.010	0.142	-0.072	-0.056	-0.221
North America	-0.014	0.036	0.064	-0.050	0.000
Europe	-0.001	-0.040	0.004	0.132	-0.010
Asia-Pacific	0.004	-0.019	0.044	-0.185	0.176
D. Univariate Regressions (equation (9))					
Japan	0.007	0.232			
North America	-0.013		-0.025		
Europe	-0.0001			0.128	
Asia-Pacific	0.0044				0.162

Table 3 (continued) Dynamics of Individual Pension Funds' Net Cash Flows.

II. 6-month return horizon

Percentage of significant regressors	Intercept		Return in Japan		Return in North Am.		Return in Europe		Return in Asia-Pacific	
	positive	negative	positive	negative	positive	negative	positive	negative	positive	negative
E. Multivariate Regressions (equation (8))										
Japan	1.2	0.0	0.0	1.2	0.0	17.4	0.8	0.8	7.3	0.0
North America	0.0	20.6	0.8	0.0	0.0	1.2	0.8	0.0	0.0	2.0
Europe	0.0	0.0	0.0	0.4	4.9	0.0	0.0	0.8	0.8	0.4
Asia-Pacific	0.0	0.0	1.6	0.4	7.3	0.0	0.0	14.2	4.0	0.8
F. Univariate Regressions (equation (9))										
Japan	0.4	0.0	0.0	0.0						
North America	0.0	18.6			0.0	2.8				
Europe	0.0	0.0					1.2	0.0		
Asia-Pacific	0.4	0.0							1.6	4.9

Median coefficient estimates	Intercept	Return in Japan	Return in North Am.	Return in Europe	Return in Asia-Pacific
G. Multivariate Regressions (equation (8))					
Japan	0.003	-0.161	-0.399	1.179	0.093
North America	-0.006	0.142	-0.347	0.180	-0.116
Europe	0.000	-0.036	-0.006	0.157	-0.098
Asia-Pacific	0.005	0.009	0.421	-0.934	0.197
H. Univariate Regressions (equation (9))					
Japan	0.003	-0.165			
North America	-0.011		-0.598		
Europe	0.000			0.067	
Asia-Pacific	0.002				0.087

Note: For each fund we regressed the net cash flow on an intercept and the returns in the four regions and also on the excess return in that region over the return on the world market portfolio. Results for a 1-month return horizon are reported in part I, while part II reports results for a 6-month horizon. The table reports the percentage of coefficients that were statistically significant at the 5 percent critical level (Panels A, B, E and F) and the median of the estimated coefficients (Panels C, D, G and H). The sample covers 247 UK pension funds over the period 1991:1 - 1997:12.

Table 4. Estimates from Bivariate Garch(1,1) Models

The following bivariate GARCH model was estimated:

$$\begin{aligned} r_{jt+1} &= \gamma_{0j} + \gamma_{1j} Yield_{jt} + \gamma_{2j} Def_t + \gamma_{3j} I_t^{us} + \gamma_{4j}(I_t^{us} - I_t^{uk}) + \eta_{jt+1} \\ \eta_{jt+1} &\sim N(0, \sigma_{jj,t}^2) \\ \sigma_{kk,t} &= \alpha_{kk} + \beta_{k0}\eta_{kt}^2 + \beta_{k1}\sigma_{kk,t-1} \quad k,l = j, uk \\ \sigma_{kl,t} &= \rho_{kl}\sqrt{\sigma_{kk,t}\sigma_{ll,t}}, \;\; k,l = j, uk \end{aligned}$$

	Japan	North America	Europe	Asia-Pacific	UK
Conditional mean					
Constant	0.0092	0.0057	0.0186	-0.0315	0.0071
	(0.0137)	(0.0104)	(0.0089)	(0.0199)	(0.0097)
$Yield_{jt}$	-0.0001	0.0029	-0.0006	0.0178	-0.0007
	(0.0030)	(0.0042)	(0.0030)	(0.0059)	(0.0028)
Def_t	0.0401	0.0257	0.0152	0.0229	0.0210
	(0.0084)	(0.0077)	(0.0072)	(0.0108)	(0.0077)
I_t^{us}	-0.0048	-0.0042	-0.0027	-0.0061	-0.0032
	(0.0015)	(0.0016)	(0.0012)	(0.0019)	(0.0013)
$I_t^{us} - I_t^{uk}$	0.0029	0.0016	0.0008	0.0027	-0.0016
	(0.0015)	(0.0012)	(0.0010)	(0.0019)	(0.0013)
Conditional volatility					
Constant	0.0008	0.0019	0.0016	0.0014	0.0005
	(0.0004)	(0.0016)	(0.0002)	(0.0006)	(0.0002)
η_{kt}^2	0.1050	0.0988	0.1326	0.1908	0.2450
	(0.0858)	(0.496)	(0.0753)	(0.1094)	(.1076)
$\sigma_{kk,t-1}$	0.7032	0.1836	0.0000	0.5407	0.6399
	(0.155)	(0.583)	(0.075)	(0.1415)	(0.1498)
ρ_{kl}	0.2608	0.4860	0.5162	0.5040	0.2607
	(0.0511)	(0.0418)	(0.0410)	(0.0426)	(0.0511)

Note: The table presents maximum likelihood estimates from bivariate GARCH models estimated pairwise on returns in the four regional markets and the UK stock market. Standard errors are provided in brackets beneath the estimates. The estimations are based on the constant conditional correlation specification proposed by Bollerslev (1990). r_{jt+1} is the return in region $j$in period $t+1$, $Yield_{jt}$ is the dividend yield in region j, Def_t is the default premium on Baa over Aaa rated bonds, I_t^{us} is the 1-month UST-bill rate and I_t^{UK} is the 1-month UK T-bill rate. The estimation period is 1970:1 - 1997:12.

Table 5. Projections of Portfolio Weights on Conditional Moments

	A: Expected Returns				B: Expected Returns, Volatility			
	Japan	North Am.	Europe	Asia-Pac	Japan	North Am.	Europe	Asia-Pac
Expected return								
Median $\widehat{\beta}_{1j}$	-12.60	50.52	8.711	-4.24	-13.74	43.67	10.74	-4.36
% of regressions with $\widehat{\beta}_{1j} > 0$	29.15	97.98	91.09	0.00	28.74	96.76	93.52	0.00
% of regressions with $t_{\widehat{\beta}_{1j}} > 2$	3.64	93.52	80.16	0.00	3.24	90.69	82.19	0.00
Conditional volatility								
Median $\widehat{\beta}_{2j}$					-2.49	-2.76	-2.21	-0.46
% of regressions with $\widehat{\beta}_{2j} < 0$					98.38	98.79	95.55	92.31
% of regressions with $t_{\widehat{\beta}_{2j}} < -2$					82.19	53.44	6.07	17.00
Median R^2	0.295	0.420	0.122	0.620	0.355	0.456	0.145	0.630

	C: Expected Returns, Covariance				D:Exp. Returns, Volatility, Covariance			
	Japan	North Am.	Europe	Asia-Pac	Japan	North Am.	Europe	Asia-Pac
Expected return								
Median $\widehat{\beta}_{1j}$	-12.60	49.96	6.16	-4.29	-13.57	20.17	-0.313	-3.56
% of regressions with $\widehat{\beta}_{1j} > 0$	29.55	97.98	88.26	0.00	29.15	82.59	45.75	0.00
% of regressions with $t_{\widehat{\beta}_{1j}} > 2$	3.64	93.52	61.94	0.00	3.24	59.92	3.24	0.00
Conditional volatility								
Median $\widehat{\beta}_{2j}$					-6.92	-11.20	4.96	3.81
% of regressions with $\widehat{\beta}_{2j} < 0$					96.36	100.00	2.83	4.86
% of regressions with $t_{\widehat{\beta}_{2j}} < -2$					79.35	98.79	0.00	0.81
Conditional covariance								
Median $\widehat{\beta}_{3j}$	0.74	4.43	-16.20	-2.60	20.65	45.04	-25.75	-13.67
% of regressions with $\widehat{\beta}_{3j} < 0$	42.51	29.96	99.60	93.12	27.13	0.40	98.39	94.74
% of regressions with $t_{\widehat{\beta}_{3j}} < -2$	27.53	0.00	94.33	75.30	17.81	0.00	95.95	87.85
Median R^2	0.299	0.433	0.252	0.652	0.485	0.588	0.292	0.701

Note: This table reports statistics characterizing the cross-sectional distribution of regression coefficients from linear projections of individual funds' portfolio weights (ω_{ijt}) on expected excess returns ($\widehat{\rho}_{jt}$), conditional volatility ($\widehat{\sigma}_{jj,t}$) and conditional covariances ($\widehat{\sigma}_{juk,t}$) with UK stock returns:

$$\omega_{ijt} = \alpha_{ij} + \beta_{1ij}\widehat{\rho}_{jt} + \beta_{2ij}\sqrt{\widehat{\sigma}_{jj,t}} + \beta_{3ij}\sqrt{\widehat{\sigma}_{juk,t}} + \varepsilon_{ijt}$$

Table 6. Summary Statistics for International Equity Returns

	Japan	North America	Europe	Asia-Pacific	World ex UK
A: Mean return (% per annum)					
FT/S&P index	-0.73	20.02	16.50	13.46	13.28
Sample (value-weighted)	2.85	19.60	16.00	11.40	12.58
Sample (equal-weighted)	3.23	19.21	15.93	11.31	12.51
B: Proportion of outperformers relative to FT/S&P index (%)	97.2	20.2	20.7	8.9	13.27
C: Correlation (FT/S&P index, sample)	0.977	0.993	0.989	0.989	0.924

Note: For each of the four regions under consideration this table reports the mean return (annual percentage) for the Financial Times/Standard & Poor index, and the value- and equal-weighted portfolios comprising the funds in our sample. We also report the proportion of outperformers relative to the index and the correlation between the time-series of monthly returns on the indices and on the value-weighted portfolios for each of the regions. The sample covers 247 UK pension funds over the period 1991:1 - 1997:12.

Table 7. Market Timing and Public Information

$A: \ \rho_{jt+1} = c_j + \beta_{1j}\Delta\omega_{jt} + \beta_j'\mathbf{Z}_t + \varepsilon_{jt+1}$

	Japan	North America	Europe	Asia-Pacific
Median $\widehat{\beta}_{1j}$	0.226	-0.047	0.020	0.257
% of regressions with $\widehat{\beta}_1 > 0$	91.09	29.96	60.73	88.66
% of regressions with $t_{\widehat{\beta}_1} > 2$	2.43	0.00	5.26	5.67

$B: \frac{\rho_{jt+1}}{\widehat{\sigma}_{jj,t}^{1/2}} = c_j + \beta_{1j}\Delta\omega_{jt} + \beta_j'\mathbf{Z}_t + \varepsilon_{jt+1}$

	Japan	North America	Europe	Asia-Pacific
Median $\widehat{\beta}_{1j}$	3.47	-0.94	0.47	3.33
% of regressions with $\widehat{\beta}_{1j} > 0$	89.07	28.74	66.80	85.02
% of regressions with $t_{\widehat{\beta}_{1j}} > 2$	2.43	0.00	5.67	4.05

$C: \frac{\rho_{jt+1}}{\widehat{\sigma}_{juk,t}^{1/2}} = c_j + \beta_{1j}\Delta\omega_{jt} + \beta_j'\mathbf{Z}_t + \varepsilon_{jt+1}$

	Japan	North America	Europe	Asia-Pacific
Median $\widehat{\beta}_{1j}$	8.46	-1.28	0.47	5.65
% of regressions with $\widehat{\beta}_{1j} > 0$	91.90	31.58	62.75	88.26
% of regressions with $t_{\widehat{\beta}_{1j}} > 2$	2.02	0.00	5.26	5.26

Note: This table tests whether funds anticipated future excess returns (relative to average world ex-UK returns), ρ_{jt+1}, by adjusting their portfolio weights ($\Delta\omega_{jt}$) prior to the return movement. The regression controls for the effect of public information, $\mathbf{Z}_t$. A positive and significant estimate of β_{1j} indicates market timing skills. Panel A uses excess returns as the dependent variable. Panels B and C adjust excess returns for the own-market conditional volatility, $\widehat{\sigma}_{jj,t}^{1/2}$, and the conditional covariance with UK stock returns, $\widehat{\sigma}_{juk,t}^{1/2}$, both obtained from the bivariate GARCH estimations reported in Table 4.

Table 8: Market Timing Skills in Up and Down Markets

A: $\Delta\omega_{jt} = \beta_{1j} I_{\{\rho_{jt+1} \geq 0\}} + \beta_{2j} I_{\{\rho_{jt+1} < 0\}} + \varepsilon_{jt}$

	Japan	North America	Europe	Asia-Pacific
Median $\widehat{\beta}_{1j}$	0.0029	-0.0011	0.0031	0.0010
% of regressions with $\widehat{\beta}_{1j} > 0$	99.60	7.69	98.38	91.50
% of regressions with $t_{\widehat{\beta}_{1j}} > 2$	9.31	0.00	2.83	0.81
Median $\widehat{\beta}_{2j}$	-0.0023	-0.0022	-0.0011	-0.0005
% of regressions with $\widehat{\beta}_{2j} < 0$	98.79	93.93	80.57	64.77
% of regressions with $t_{\widehat{\beta}_{2j}} < -2$	2.02	9.31	0.40	0.00

B: Henriksson-Merton tests of market timing: excess returns

	Japan	North America	Europe	Asia-Pacific
% of funds with positive market timing test	95.1	63.6	82.2	79.3
% of funds with positive and significant test	3.2	0.0	9.3	4.9

C: Henriksson-Merton tests of market timing: unexpected excess returns

	Japan	North America	Europe	Asia-Pacific
% of funds with positive market timing test	96.4	63.6	71.3	88.3
% of funds with positive and significant test	2.0	0.0	4.0	2.0

Note: Panel A tests whether funds increased their portfolio weights, $\Delta\omega_{jt}$, in anticipation of a positive sign for next period's excess return in a given region, j, relative to the global average (ρ_{jt+1}). The panel also tests whether the funds had market timing skills in down markets. $I_{\{\rho_{jt+1} \geq 0\}}$ is an indicator function that takes a value of unity whenever excess returns in period $t+1$ are non-negative and otherwise is zero. $I_{\{\rho_{jt+1} < 0\}}$ takes a value of unity when ρ_{jt+1} is negative. Market timing skills should show up as a positive value of β_{1j} and a negative value of β_{2j}. The Henriksson-Merton tests (Panels B and C) consider the null hypothesis that the sign of $\Delta\omega_{jt}$ and ρ_{jt+1} are independently distributed. A positive and significant value of this test again indicates market timing skills. A 5 percent critical value was assumed throughout the table to assess statistical significance.

Table 9: Tests for Extra-Market Timing Skills

$$\Delta\omega_{jt} = \beta_{1j} I_{\{\rho^u_{jt+1} \geq 0\}} + \beta_{2j} I_{\{\rho^u_{jt+1} < 0\}} + \beta_{3j} I_{\{\rho^e_{jt+1} \geq 0\}} + \varepsilon_{jt}$$

	Japan	North America	Europe	Asia-Pacific
Median $\widehat{\beta}_{1j}$	0.0003	-0.0022	-0.0046	0.0007
% of regressions with $\widehat{\beta}_{1j} > 0$	79.76	0.00	2.43	91.90
% of regressions with $t_{\widehat{\beta}_{1j}} > 2$	0.00	0.00	0.00	0.00
Median $\widehat{\beta}_{2j}$	1.132	-1.512	1.587	0.113
% of regressions with $\widehat{\beta}_{2j} < 0$	14.98	94.74	1.62	5.26
% of regressions with $t_{\widehat{\beta}_{2j}} < -2$	0.00	2.83	0.00	0.00
Median $\widehat{\beta}_{3j}$	0.045	-0.024	0.0062	0.031
% of regressions with $\widehat{\beta}_{3j} > 0$	94.74	22.67	59.11	94.74
% of regressions with $t_{\widehat{\beta}_{3j}} > 2$	1.62	0.00	5.67	9.72

Note: This table tests whether the funds changed their portfolio weights, $\Delta\omega_{jt}$, in correct anticipation of the sign of that part of next period's excess return that is unpredictable through public information. $I_{\{\rho^u_{jt+1} \geq 0\}}$ is an indicator function for the event that the unexpected excess return is non-negative. Conversely, $I_{\{\rho^u_{jt+1} < 0\}}$ takes a value of unity whenever unexpected excess returns were negative in period $t+1$. $I_{\{\rho^e_{jt+1} \geq 0\}}$ is an indicator for the sign of the expected excess return. A positive sign of β_{1j} indicates that funds correctly anticipated returns above what was expected given public information, while a negative sign of β_{2j} suggests that the funds correctly anticipated negative excess returns below what was expected given public information. A positive sign of β_{3j} suggests that public information - as reflected in expected returns - influenced the portfolio weights.