International Asset Allocation with Time-varying Investment Opportunities*

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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 model portfolio weight dynamics as a function of time-varying conditional moments. 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. Consequently, controlling for the effect of state variables that track time-variations in investment opportunities significantly affects estimates of returns from international market timing. Our estimates suggest that the portfolio movements that were orthogonal to such state variables accounted for a net loss of 0.2 per cent per annum for the average fund.

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I. Introduction

"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."

Barry Riley, *Financial Times,* 16 December 1997.

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. The advantage of working with data on UK pension funds is that they face very few restrictions on their investment behavior. 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 among 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 and 98 percent of all funds generated a positive and significant coefficient on expected returns for North America and Europe, respectively. These regions account for more than 75 percent of the funds' international equity holdings. 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 coefficient on either own-market volatility or conditional covariance with global stock returns varied from 48 percent (Europe) to 98 percent (Japan). These results indicate that time-varying conditional moments are essential for explaining and evaluating institutional investors' international asset allocations.

The withdrawal of funds from North America together with the fact that the US stock market paid substantially higher returns than 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 a massive 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 may be 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

[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).

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. While the importance of time-varying investment opportunities is widely acknowledged in studies of domestic fund performance, to our knowledge no study has previously investigated these effects in the context of international asset allocation. Once we control for the effects of public information, there is no evidence of what Graham & Harvey (1996) refer to as extra-market-timing ability, i.e. anticipating return movements beyond that which could have been predicted using public information. Our estimates indicate that the median fund earned a negative return from international market-timing of around -0.2 percent per annum.

The plan of the paper is as follows. Section II provides a description and initial characterization of our data set. Section III analyzes the extent to which the funds' investment strategies in foreign markets can be explained by time variations in the investment opportunity set. Section IV examines evidence on returns from the part of international market-timing that is not explained by time-varying investment opportunities 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.

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.

Interest in analyzing international portfolio flows has grown recently. 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 countries, Tesar and Werner (1994, 1995) analyze the evolution in aggregate holdings of foreign assets in five major economies. They find, among other things, that investors' turnover rate in foreign equity investments is high relative to their home market turnover rate. 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.

Our data set is unique relative to those analyzed in previous studies in that it is organized by individual pension funds' asset holdings. Froot, O'Connell & Seasholes (1999) use daily data on international transactions over the period 1994 to 1998 to shed light on the relationship between foreign asset trades and stock returns. Their data consists of detailed records on aggregate holdings of pension, endowment and mutual funds and of governments. Choe, Kho & Stulz (1999) examine transactions of foreign investors on the Korea Stock Exchange over the period November 1996 to December 1997. While our data set is not well suited for studying the price impact of foreign investors in a particular domestic market, it

[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.

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.

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

[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.

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 and this might be pure coincidence).

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 monthly 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.

III. Portfolio Weights and Time-varying Investment Opportunities

This section addresses 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 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.[4] 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:

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

To capture possible time variations in conditional volatilities and covariances, we model returns in the context of a bivariate generalized ARCH model. The

[4]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.

contribution of foreign equity holdings to a pension fund's total volatility from foreign equity holdings is determined in part by their own volatility and in part by their covariance with global returns. Let $\mathbf{r}_{t+1} = (r_{jt+1}\ r_{wt+1})'$, where r_{jt+1} and r_{wt+1} are region j and global equity returns in month $t+1$, respectively. We follow Bollerslev (1990) and model returns as follows:

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

where $\boldsymbol{\eta}_{t+1} = (\eta_{jt+1}, \eta_{wt+1})'$ is the set of heteroskedastic return innovations defined as $\boldsymbol{\eta}_{jt+1} = \sigma_{jj,t}\varepsilon_{kt+1}$, where $\boldsymbol{\varepsilon}_{t+1} = (\varepsilon_{jt+1}, \varepsilon_{wt+1})'$ are normal, independent and identically distributed residuals so that $\boldsymbol{\eta}_{t+1} \sim N(0, \Sigma_t)$. $\Sigma_t = [\sigma^2_{kl,t}]$ is the conditional covariance matrix and ψ_{kl} is the conditional correlation coefficient which is assumed to be constant and was always found to be non-negative. Finally $\mathbf{Z}_t = \{Yield_{jt}, Def_t,\ I^{us}_t, I^{us}_t - I^{uk}_u\}$ 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 global equity returns.

Table 1 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 globally.

Armed with this specification of the time-varying opportunity set 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.[5] 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.[6] Consistent with theoretical models of intertemporal asset allocation (e.g. Brennan et. al (1997)), we use portfolio weights, ω_{ijt}, as the dependent variables. For each fund, i, and each region, j, we estimate a set of time-series regressions:

$$\omega_{ijt} = \alpha_{ij} + \sum_{k=1}^{4} \beta_{1ik}\widehat{\rho}_{kt} + \sum_{k=1}^{4} \beta_{2ik}\widehat{\sigma}_{kk,t} + \sum_{k=1}^{4} \beta_{3ik}\widehat{\sigma}_{kw,t} + \varepsilon_{ijt}, \tag{3}$$

where $\widehat{\rho}_{kt}$ is the expected excess return in region k, while $\widehat{\sigma}_{kk,t}$ is the conditional return volatility and $\widehat{\sigma}_{kw,t}$ is the conditional covariance with global equity returns, all estimated from (2). These moments are based on information at time $t-1$ and hence are known by the time the fund decides on ω_{ijt}. As we shall see below, it is important that the same regressors appear in each equation to ensure consistency of the parameter estimates.

For each fund, i, the weights sum to one across the regions, j:[7]

$$\sum_{k=1}^{4} \omega_{ikt} = 1. \tag{4}$$

This means that the coefficients in equation (3) are subject to the adding-up constraints

$$\sum_{k=1}^{4} \alpha_{ik} \quad = \quad 1$$

[5]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 $\rho^e_{jt} = (\mathbf{e}'_j - \boldsymbol{\omega}'_t)\boldsymbol{\mu}_t$.

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

[7]In the context of allocation or demand systems, the equivalent condition is that the sum of the individual components equals a predetermined aggregate. This is known as the adding-up criterion, c.f. Bewley (1986).

$$\begin{aligned}\sum_{k=1}^{4}\beta_{1ik} &= 0.\\ \sum_{k=1}^{4}\beta_{2ik} &= 0 \qquad (5)\\ \sum_{k=1}^{4}\beta_{3ik} &= 0.\end{aligned}$$

To see the consequences of these constraints on parameter estimation, we write the full set of constraints on a given fund's portfolio weights as follows:

$$W_i \equiv \begin{pmatrix} \omega_{i11} & \omega_{i21} & \omega_{i31} & \omega_{i41} \\ \omega_{i12} & \omega_{i22} & \omega_{i32} & \omega_{i42} \\ \vdots & \vdots & \vdots & \vdots \\ \omega_{i1T} & \omega_{i2T} & \omega_{i3T} & \omega_{i4T} \end{pmatrix} = X\beta_i + U_i, \qquad (6)$$

where X is a $T \times p$ matrix of predetermined regressors, β_i is a $p \times 4$ matrix of coefficients and U_i is a $T \times 4$ matrix of innovation terms; p is the number of regressors which in our case equals one plus the number of moments the portfolio weights are projected on.

The system of portfolio weight equations can also be written in the following convenient way:

$$\begin{aligned} vec(W_i) &= (I \otimes X)vec(\beta_i) + vec(U_i) \qquad (7)\\ E[vec(U_i)vec(U_i)'] &= \Omega_i \otimes I, \end{aligned}$$

where $vec(.)$ is the vector stacking operator, $\otimes$ is the Kronecker product, Ω_i is a 4×4 symmetric covariance matrix and I is the identity matrix of suitable dimension. Using this notation, the adding-up constraint can be written as follows

$$W_i \boldsymbol{\iota}_4 = X\beta_i \boldsymbol{\iota}_4 + U_i \boldsymbol{\iota}_4 = \boldsymbol{\iota}_T, \qquad (8)$$

where $\boldsymbol{\iota}_4$ and $\boldsymbol{\iota}_T$ are 4×1 and $T \times 1$ vectors of ones, respectively. These constraints have important implications for estimation of the covariance matrix for the system of portfolio weights. This will be singular since

$$\Omega_i \boldsymbol{\iota}_4 = T^{-1}E[U_i'U_i]\boldsymbol{\iota}_4 = T^{-1}E[U_i'U_i\boldsymbol{\iota}_4] = \mathbf{0}. \qquad (9)$$

More intuitively, this is an implication of the constraint that, at each point in time, t, and for each fund, i, the innovations must add up to zero:

$$\sum_{j=1}^{4} \varepsilon_{ijt} = 0. \tag{10}$$

Standard generalized least squares (GLS) and maximum likelihood (ML) methods can therefore not be used to estimate coefficients in the full system of portfolio weights (6). Instead it is necessary to delete one column from (6) and constrain the estimators.[8] Letting Ω_{11i} be the 3×3 covariance matrix of full rank for the first three sets of portfolio weights, while W_{1i} and β_{1i} are the first three columns of W_i and β_i, the log-likelihood function of the subsystem comprising the first three weights is

$$\ln(L(vec(W_{1i})) = \frac{-3T}{2}\ln(2\pi) - \frac{T}{2}\ln(|\Omega_{11i}|) - \frac{1}{2}tr(\Omega_{11i}^{-1}(W_{1i} - X\beta_{1i})'(W_{1i} - X\beta_{1i})). \tag{11}$$

Taking derivatives now yields the following maximum likelihood estimators for β_{1i} and Ω_{11i} :

$$\begin{aligned} vec(\widehat{\beta}_{1i}) &= (I \otimes (X'X)^{-1}X')vec(W_{1i}) \\ \widehat{\Omega}_{11i} &= (W_{1i} - X\widehat{\beta}_{1i})'(W_{1i} - X\widehat{\beta}_{1i})/T, \end{aligned} \tag{12}$$

c.f. Bewley (1986). The estimated coefficients of a particular column in (6) scaled by their standard errors follow a t-distribution, although a Wald test of coefficient restrictions across columns will not have a standard distribution.

Table 2 summarizes the empirical results based on the above estimators. Panel A projects portfolio weights onto expected returns alone. More than 95 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 five percent for Asia-Pacific. The result for Asia-Pacific may initially seem puzzling, but is related to the importance of the dividend yield 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

[8] The estimation results are invariant to which column is deleted from (6). The estimated coefficients of the deleted equation can be derived from the adding-up constraints (5).

that 94 and 98 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 volatility as a regressor in the portfolio weight equation. Panel B shows that the majority of funds - in excess of 70 percent in Japan, North America and Asia Pacific - generated negative coefficients on own-market volatility. This indicates that the funds decreased their allocation towards regions whose volatility was expected to go up. Panel C reports the outcome from using expected returns and conditional covariance with global stock returns as regressors. Conditional covariances are not quite as important as own-market volatility, but they predominantly have the right sign for the three main regions, namely Japan, North America and Europe.[9]

Panel D shows the results from regressions that include all three explanatory variables. Own-market expected returns continue to have the right sign and be statistically significant for almost all funds' weights in North America and Europe. Likewise, own-market conditional volatility generates a negative coefficient estimate for the majority of funds, while conditional covariance with global returns produces a negative coefficient estimate for around 60 percent of the funds in the two largest markets, North America and Europe. However, the sign of the coefficients of some of the volatility and covariance regressors is now difficult to interpret since the two series are driven by a common component and hence strongly correlated. While the individual coefficient estimates are difficult to interpret, one would expect that the total effect of the own-market variance and covariance with the global market return is negative. We therefore tested whether the sum of the coefficients on own-market variance and covariance with the global market return, $\beta_{2ij} + \beta_{3ij}$, are negative. The results, reported at the bottom of Panel D, show that the total effect of these variables is predominantly negative for the three main regions, namely

[9]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 indeed influence portfolio holdings. The difference between these findings may be explained by our use of time-varying conditional moments.

Japan, North America and Europe.

So far the results capture covariance effects by modelling the portfolio weights as a function of the conditional covariances between regional and global market returns. Such covariances represent the regional equity holdings' contribution to systematic risk in the context of a single-factor international CAPM. These covariances are likely to capture a large fraction of the regional returns' contribution to total portfolio risk. However, since the pension funds hold significant parts of their portfolios in foreign equity, the inter-regional return covariances will also contribute to the total portfolio risk. To investigate the effect of the inter-regional covariances, we obtained conditional covariance estimates that were used in the following linear regressions:

$$\omega_{ijt} = \alpha_{ij} + \sum_{k=1}^{4} \beta_{1ik}\widehat{\rho}_{kt} + \sum_{j=1}^{4}\sum_{k=1}^{4} \beta_{2ijk}\widehat{\sigma}_{jk,t} + \varepsilon_{ijt}. \tag{3'}$$

Here $\widehat{\sigma}_{jk,t}$ is the conditional covariance between returns in regions j and k. The results from these regressions are shown in Panel E of Table 2. They are similar to those obtained in Panel D. For North America and Europe the vast majority of funds continue to have a positive coefficient on the own-market expected return and a negative coefficient on the own-market volatility, while Japan and Asia-Pacific produce weaker results. The effect of the inter-regional covariances, computed as $\sum_{k \neq j} \widehat{\beta}_{2ijk}$, is negative for the vast majority of funds' investments in Japan. A negative effect is also observed for close to half of the funds in Asia Pacific. In contrast, there is not much of a negative effect for North America and Europe. Again these results are related to the findings for the conditional variances since Japan and Asia Pacific are the regions where own-market volatility does not have a strong effect on the portfolio weights (but conditional covariances do), while for North America and Europe own-market volatility has a strong negative effect on portfolio weights but the conditional covariances do not.

To further investigate whether information on variances and covariances help to predict variation in asset weights, we undertook the predictive information test proposed by Diebold & Mariano (1995). To do so, we first computed the squared forecast error differential, $dif_t = e_{t*}^2 - e_t^2$, where e_{t*} is the forecast error (i.e., the difference between the actual and predicted weight) based on the full model (3) that includes time-varying first and second moments while e_t is the forecast error from a simpler model that only projects portfolio weights on a constant and expected returns. Based on these forecast errors we then computed the statistic $\sqrt{T}\overline{dif}/sd(dif_t)$, where $\overline{dif} = \sum_{t=1}^{T} dif_t/T$ and $sd(dif_t)$ are estimates of the mean and standard deviation of dif_t, respectively. This gives a test statistic that is asymptotically normally distributed. The results showed that adding second-moment information led to a significantly better forecast for 22%, 19%, 72% and 2% of the funds' weights in Japan, North America, Europe and Asia Pacific, respectively.

To demonstrate graphically the importance to the evolution in portfolio weights of the time variation in conditionally expected returns, volatilities and covariances, Figure 2 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.[10]

We finally computed the optimal portfolio weights based on the first and second moment estimates and compared these to the observed portfolio weights. For a mean-variance optimizing investor the optimal portfolio weights on the j'th security is given by (c.f. Bohn & Tesar (1996))

$$\omega_{jt}^{*} = \alpha \mathbf{e}_j' \Sigma_t^{-1} \boldsymbol{\mu}_t + \eta_{jt}, \tag{13}$$

where α is the investor's coefficient of relative risk aversion, $\mathbf{e}_j$ is a zero-one vector selecting the jth regional return, Σ_t is the conditional covariance matrix between the regional returns, $\boldsymbol{\mu}_t$ is the vector of expected returns and η_{jt} represents a 'hedge factor' that captures risks beyond those captured by the regional return processes. These are likely to be important here and may represent the (unmodeled) effect of asset-liability matching. Since we do not observe this hedge factor and also do not know the true value of α we simply test the broad implications of the model that there should be a positive correlation between the observed weights (ω_{jt}) and the optimal portfolio weights (ω_{jt}^{*}). As it turns out, our results closely match those found in Table 2. For the two largest markets, North America and Europe, we

[10]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. 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.

found a positive correlation between the observed weights and the optimal weights for 91 and 56 percent of the funds, respectively. For Japan and Asia Pacific we did not find a positive correlation between the actual and optimal portfolio weights for many funds. This is unsurprising in light of the zero or negative coefficient on the mean return for most of the funds' investments in these two regions, c.f. Panel D in Table 2.

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. 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 even 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, it is far from obvious which external index provides the most suitable representation of benchmark returns: Kang & Stulz (1997), for example, show that foreign investors' holdings of Japanese equities are concentrated in the largest firms.

Figure 3 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 series are clearly strongly correlated. This impression is confirmed by the sample correlations reported in the last row of Table 3. 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 3 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.[11] However, this region is the only one in which a typical UK pension 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.[12] 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.

B. Conditional Market-Timing Tests

To test whether UK pension funds possess market-timing skills after controlling for public information, we ran a range 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

[11] 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 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, we found once again that only a small portion of the funds generated statistically significant outperformance.

[12] 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.

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 expected-return-maximizing funds ought to increase allocations to regions with above-average expected returns:

$$\rho_{jt+1} = c_j + \beta_{1j}\Delta\omega_{jt} + \beta_j' \mathbf{Z}_t + \varepsilon_{jt+1}. \tag{14}$$

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, $\mathbf{Z}_t$, considered in equation (2). Market-timing skills should show up in the form of a positive coefficient estimate, $\widehat{\beta}_{1j}$.

Panel A of Table 4 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 expected 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 ($\widehat{\sigma}_{jj,t+1}$) or by the conditional covariance with returns on UK stocks ($\widehat{\sigma}_{juk,t+1}$), both obtained from the bivariate GARCH model (2). The results, reported in Panels B and C of Table 4, do not change very much, suggesting that the evidence on market-timing is robust in the presence of time-varying risk.

C. Directional Tests

We next conducted a Merton-style market-timing regression based on indicator functions (taking the values unity or zero) 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:

$$\Delta\omega_{jt} = \beta_{1j} I_{\{\rho_{jt+1}\geq 0\}} + \beta_{2j} I_{\{\rho_{jt+1}<0\}} + \varepsilon_{jt}. \tag{15}$$

An unconditional 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 5 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, these unconditional regressions suggest 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 considerable 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 5. When applied to the four regions, we find only 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 perform conditional tests by regressing 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} (based on the earlier regression of regional returns on the lagged instruments, $\mathbf{Z}_t$)[13]

$$\Delta\omega_{jt} = \beta_{1j} I_{\{\rho^u_{jt+1} \geq 0\}} + \beta_{2j} I_{\{\rho^u_{jt+1} < 0\}} + \beta_{3j} I_{\{\rho^e_{jt+1} \geq 0\}} + \varepsilon_{jt}, \tag{16}$$

If funds can predict the part of future differential returns unaccounted for by current public information (Graham and Harvey call this extra-market-timing ability) β_{1j} should be positive and β_{2j} should be negative.

Table 6 shows very little evidence of extra-market-timing skills. While 80 and 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 5 appears to reflect publicly available information as evidenced by the many positive estimates of β_{3j} for Japan and Europe.

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 outcome using total future excess returns, the results, as shown in Panel C of Table 5, 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.

D. Overall Measures of Market-Timing

As a means of providing an overall summary measure of market-timing skills, 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

[13]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 induce perfect collinearity.

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 any 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.

Our evidence so far suggests that genuine market-timing skills are very weak. However, it also raises the possibility 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 extra-market-timing activities, we compute for each region the return from that part of the portfolio weight that is orthogonal to time-varying moments, $\widehat{\omega}^u_{ijt} = \omega_{ijt} - \widehat{\omega}_{ijt}$, where $\widehat{\omega}_{ijt}$ is the projection of ω_{ijt} on the conditional mean, variance and covariance from equation (2), rescaled to sum to unity. For each fund (i) the $\widehat{\omega}^u_{ijt}$ sum to zero (across region j), so these weights 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:

$$\sum_{j=1}^{4} \omega^u_{ijt} r_{jt}. \tag{17}$$

The mean of the time-series average of this measure is -0.16 percent per annum when the portfolio weights are projected on expected returns, variances and covariances.[14] Figure 4 provides a histogram of the statistic, demonstrating that there are two clusters of funds. The vast majority of funds belong to the cluster with a mean return from extra-market-timing of around -0.25 percent per year. A smaller cluster of funds is centered around a mean returns of 0.25 percent per year. Only 29 out of 247 or eleven percent of the funds generated positive mean returns from extra-market-timing. None of these time-series means was individually statistically significant, however.

[14]When the portfolio weights were projected on expected returns and variances, leaving covariances with global returns out, the average of the market timing statistic was -0.20.

V. Conclusion

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 has previously been known about the factors influencing investors' market-timing and strategic asset allocation decisions in international equity markets.

Several new insights into institutional investors' behavior and performance in foreign equity markets have resulted from this study, chief of which is our finding that portfolio weights are highly correlated with time-varying expected returns, volatilities and conditional covariances with global 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 partially 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.

While we find weak evidence in support of international market-timing skills based on standard, unconditional performance regressions, this evidence becomes much weaker in tests that account for a time-varying global investment opportunity set. Our estimates suggest that, when we orthogonalize portfolio weight movements with respect to predictable time-varying moments, the average extra-market-timing performance was -0.2 per cent per annum.

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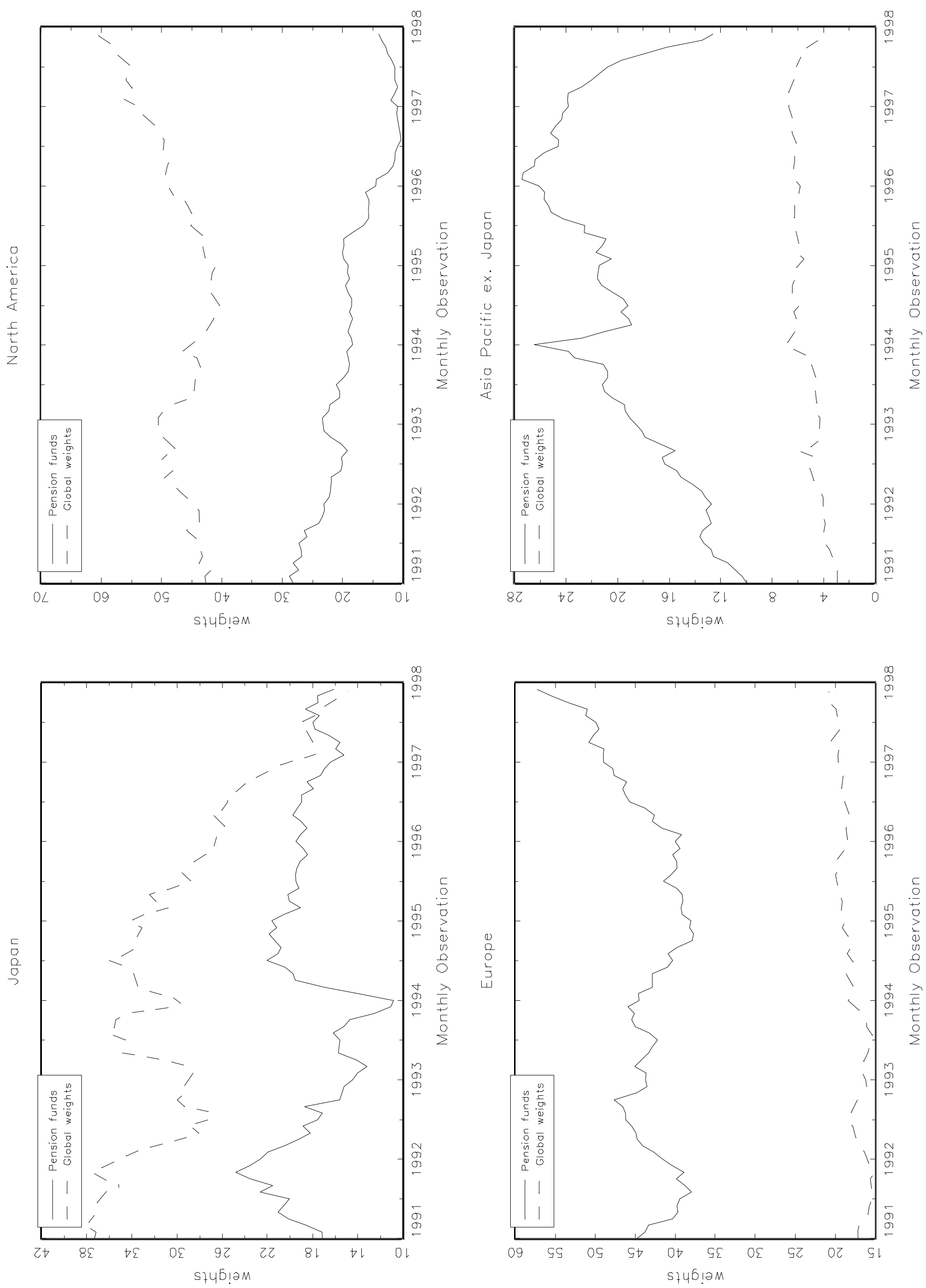

Figure 1: UK pension funds' portfolio weights and global market weights

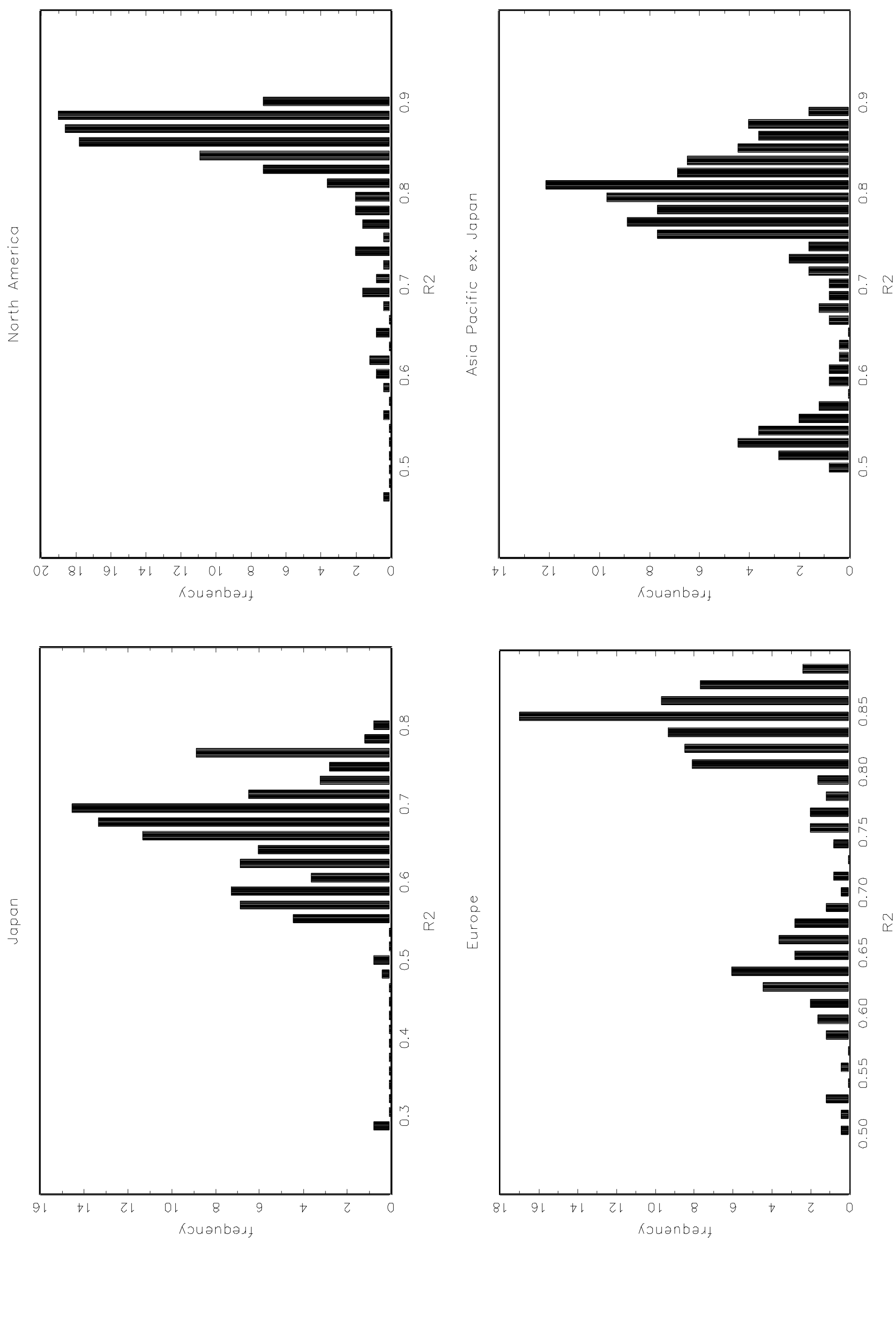

Figure 2: Histogram of R^2 from regressions of portfolio weights on conditional moments

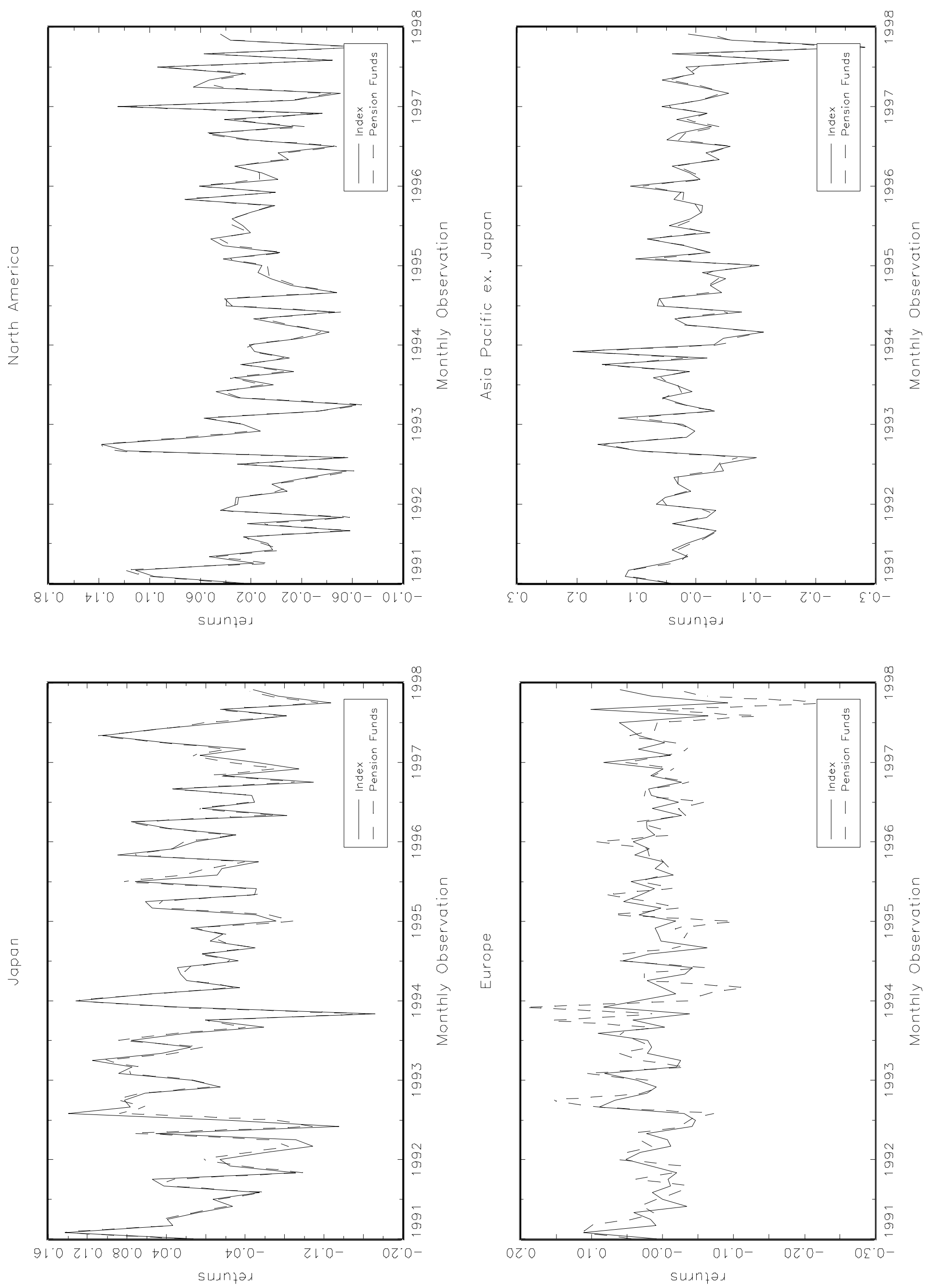

Figure 3: Time-series of returns on FT/SP indices and on UK pension funds

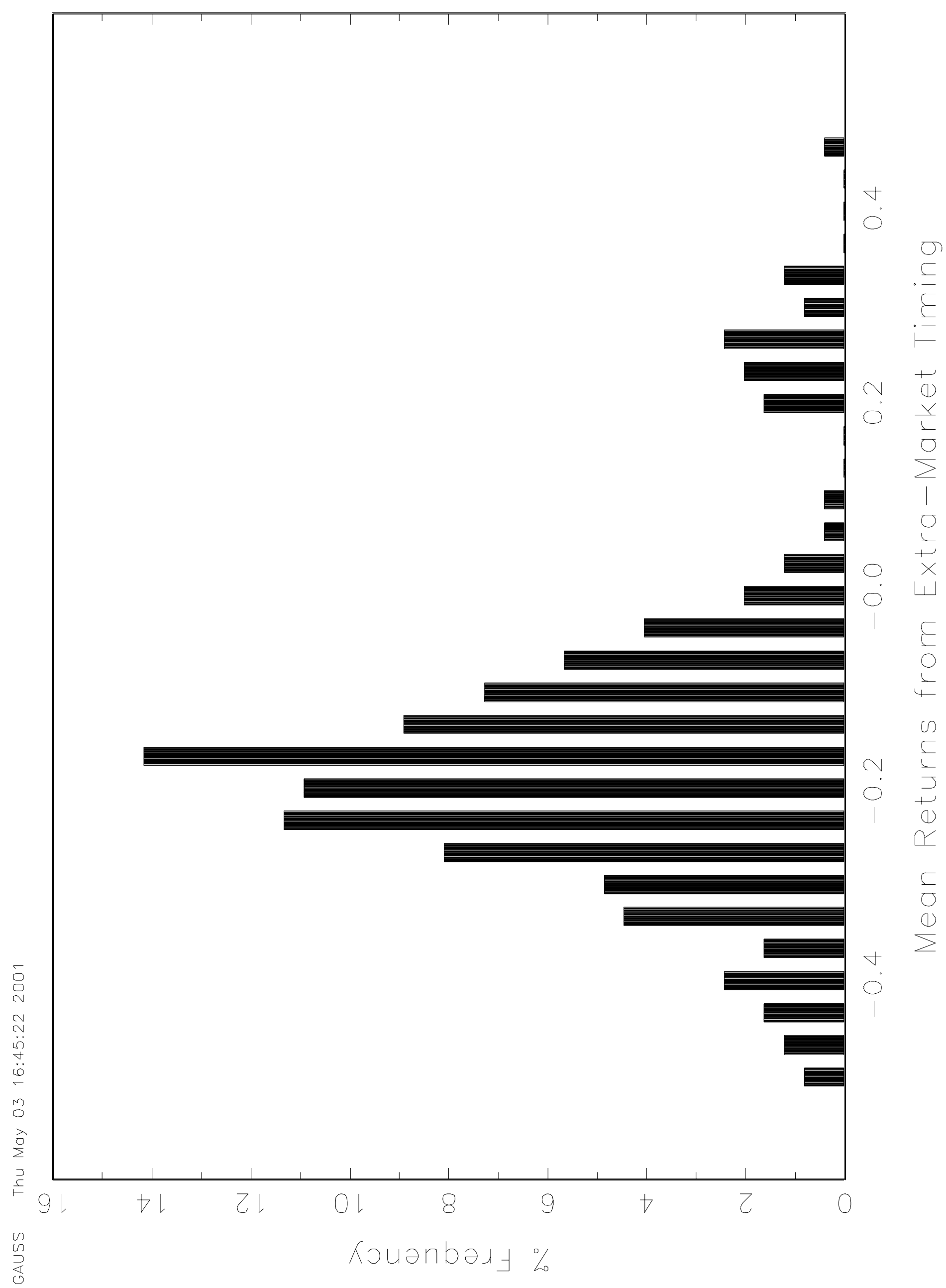

Figure 4: Cross-sectional distribution of mean returns from extra-market timing

Table 1. Estimates of conditional means and volatilities from bivariate GARCH(1,1) model

	Japan	North America	Europe	Asia-Pacific	UK
Conditional mean					
Constant	.0074	-.0013	.00144	-.0406*	-.0008
	(0.0132)	(0.0105)	(0.0088)	(0.0181)	(0.0090)
$Yield_{jt}$	-.0003	.0003	-.0004	.0173*	.0008
	(0.0031)	(0.0046)	(0.0029)	(0.0051)	(0.0037)
Def_t	.0416*	.0290*	.0187*	.0245*	.0308*
	(0.0085)	(0.0074)	(0.0070)	(0.0104)	(0.0062)
I_t^{us}	-.0046*	-.0027	-.0028*	-.0049*	-0.0034*
	(0.0015)	(0.0016)	(0.0012)	(0.0019)	(0.0014)
$I_t^{us} - I_t^{uk}$	0.0029	.0004	.0010	.0032	.0007
	(0.0015)	(0.0012)	(0.0011)	(0.0018)	(0.0011)
Conditional volatility					
Constant	.0011*	.0004	.0004	.0010*	.00004*
	(0.0005)	(0.0002)	(0.0005)	(0.0004)	(0.0002)
η_{kt-1}^2	.1121	.0997	.0559	.2111*	.0860
	(0.1026)	(0.0630)	(0.1978)	(0.0851)	(.0774)
$\sigma_{kk,t-1}$	.6111*	.7667*	0.7300*	.6130*	.7071*
	(0.1555)	(0.1450)	(0.2869)	(0.1280)	(0.1227)
ρ_{kl}	.6447*	.8940*	.7263*	.6260*	.8940*
	(0.0320)	(0.0112)	(0.0258)	(0.0347)	(0.0112)

Note: The table presents maximum likelihood estimates from the following bivariate GARCH model:

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

$\eta_{jt+1} = \sigma_{jj,t}\varepsilon_{jt+1},\ \varepsilon_{jt+1} \sim N(0,1)$

$\sigma_{kk,t}^2 = \alpha_{kk} + \beta_{k0}\eta_{kt}^2 + \beta_{k1}\sigma_{kk,t-1}^2 \quad k,l = j,w$

$\sigma_{kl,t} = \psi_{kl}\sigma_{kk,t}\sigma_{ll,t},\ k,l = j,w$

The model was estimated pairwise on returns in the four regional markets and on the global stock market index. Standard errors are provided in brackets beneath the estimates. The estimations are based on the constant conditional correlation specification proposed by Bollerslev (1990). The regressors are defined as follows: 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 US T-bill rate and I_t^{UK} is the 1-month UK T-bill rate. An asterisk implies significance at the 5% level. The sample covers 247 UK pension funds over the period 1991:1 - 1997:12.

Table 2. Projections of portfolio weights on conditional moments

	Japan	North America	Europe	Asia Pacific
A. Expected returns				
Expected returns				
Median $\widehat{\beta}_{1j}$	-5.11	24.82	57.30	-3.61
% of regressions with $\widehat{\beta}_{1j} > 0$	29.55	96.36	98.79	5.26
% of regressions with $t_{\widehat{\beta}_{1j}} > 2$	6.07	93.93	98.79	0.00
Median R^2	.510	.805	.923	.724
B. Expected returns and volatility				
Expected returns				
Median $\widehat{\beta}_{1j}$	-5.72	24.98	51.14	-3.61
% of regressions with $\widehat{\beta}_{1j} > 0$	29.14	96.35	99.60	5.26
% of regressions with $t_{\widehat{\beta}_{1j}} > 2$	5.67	94.33	93.12	0.00
Conditional volatility				
Median $\widehat{\beta}_{2j}$	-1.74	-0.79	0.30	-0.14
% of regressions with $\widehat{\beta}_{2j} < 0$	97.57	88.66	33.20	71.66
% of regressions with $t_{\widehat{\beta}_{2j}} < -2$	74.08	9.71	0.00	11.74
Median R^2	.590	.830	.794	.630
C. Expected returns, Covariance				
Expected returns				
Median $\widehat{\beta}_{1j}$	-5.95	25.78	54.69	-3.24
% of regressions with $\widehat{\beta}_{1j} > 0$	28.74	97.17	98.79	4.86
% of regressions with $t_{\widehat{\beta}_{1j}} > 2$	5.67	94.33	93.12	0.00
Conditional covariance				
Median $\widehat{\beta}_{3j}$	-3.92	-0.39	0.58	1.55
% of regressions with $\widehat{\beta}_{3j} < 0$	89.47	63.97	48.18	14.17
% of regressions with $t_{\widehat{\beta}_{3j}} < -2$	80.16	16.60	7.29	0.00
Median R^2	.614	.846	.782	.781

D. Expected returns, volatility, and covariance				
Expected returns				
Median $\widehat{\beta}_{1ij}$	-5.99	25.81	48.67	-3.07
% of regressions with $\widehat{\beta}_{1ij} > 0$	29.15	97.57	98.79	0.00
% of regressions with $t_{\widehat{\beta}_{1ij}} > 2$	5.67	94.74	92.31	0.00
Conditional volatility				
Median $\widehat{\beta}_{2ij}$	-2.56	-2.06	1.46	-0.72
% of regressions with $\widehat{\beta}_{2ij} < 0$	97.53	89.07	34.82	78.95
% of regressions with $t_{\widehat{\beta}_{2ij}} < -2$	63.97	11.34	5.67	0.40
Conditional covariance				
Median $\widehat{\beta}_{3ij}$	0.96	-0.38	-3.06	4.00
% of regressions with $\widehat{\beta}_{3ij} < 0$	28.74	62.13	57.09	17.00
% of regressions with $t_{\widehat{\beta}_{3ij}} < -2$	6.48	9.31	12.96	3.24
% of regressions with $\widehat{\beta}_{2ij} + \widehat{\beta}_{3ij} < 0$	85.83	98.38	53.44	14.98
% of regressions with $t_{\widehat{\beta}_{2ij}+\widehat{\beta}_{3ij}} < -2$	1.21	26.32	11.74	2.83
Median R^2	.669	.860	.820	.701

E. Expected returns, volatility, and covariance				
Expected returns				
Median $\widehat{\beta}_{1ij}$	-10.72	30.74	32.18	-2.47
% of regressions with $\widehat{\beta}_{1ij} > 0$	17.00	96.36	98.79	10.53
% of regressions with $t_{\widehat{\beta}_{1ij}} > 2$	5.26	91.90	95.14	4.05
Conditional volatility				
Median $\widehat{\beta}_{2ij}$	3.00	-19.97	-9.13	-1.16
% of regressions with $\widehat{\beta}_{2ij} < 0$	23.89	70.04	94.33	74.90
% of regressions with $t_{\widehat{\beta}_{2ij}} < -2$	6.89	47.78	0.40	14.17
Conditional covariance				
% of regressions with $\widehat{\beta}_{\sum_{k\neq j}\widehat{\beta}_{2ijk}} < 0$	85.02	19.43	15.38	41.30
% of regressions with $t_{\sum_{k\neq j}\widehat{\beta}_{2ijk}} < -2$	22.27	0.00	0.00	6.47
Median R^2	.766	.863	.774	.843

Panels A-D of 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 ($\widehat{\rho}_{kt}$), conditional volatility ($\widehat{\sigma}_{kk,t}$) and conditional covariances ($\widehat{\sigma}_{kw,t}$) with global stock returns:
$\omega_{ijt} = \alpha_{ij} + \sum_{k=1}^{4} \beta_{1ik}\widehat{\rho}_{kt} + \sum_{k=1}^{4} \beta_{2ik}\widehat{\sigma}_{kk,t} + \sum_{k=1}^{4} \beta_{3ik}\widehat{\sigma}_{kw,t} + \varepsilon_{ijt}$.
Panel E is based on the specification
$\omega_{ijt} = \alpha_{ij} + \sum_{k=1}^{4} \beta_{1ik}\widehat{\rho}_{kt} + \sum_{j=1}^{4}\sum_{k=1}^{4} \beta_{2ijk}\widehat{\sigma}_{jk,t} + \varepsilon_{ijt}$,
where $\widehat{\sigma}_{jk,t}$ is the conditional covariance between returns in regions j and k.
The sample covers 247 UK pension funds over the period 1991:1 - 1997:12.

Table 3. Summary statistics for international equity returns

	Japan	North America	Europe	Asia-Pacific	World ex UK
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
Proportion of outperformers relative to FT/S&P index (%)	97.2	20.2	20.7	8.9	13.27
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 4. Market timing and public information

	Japan	North America	Europe	Asia-Pacific
$A: \ \rho_{jt+1} = c_j + \beta_{1j}\Delta\omega_{jt} + \beta_j'\mathbf{Z}_t + \varepsilon_{jt+1}$				
Median $\widehat{\beta}_{1j}$	.226	-.047	.020	.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
$B: \frac{\rho_{jt+1}}{\widehat{\sigma}_{jj,t+1}} = c_j + \beta_{1j}\Delta\omega_{jt} + \beta_j'\mathbf{Z}_t + \varepsilon_{jt+1}$				
Median $\widehat{\beta}_{1j}$	3.47	-0.94	.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
$C: \frac{\rho_{jt+1}}{\widehat{\sigma}_{jw,t+1}} = c_j + \beta_{1j}\Delta\omega_{jt} + \beta_j'\mathbf{Z}_t + \varepsilon_{jt+1}$				
Median $\widehat{\beta}_{1j}$	8.46	-1.28	.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 correctly 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 for β_{1j} indicates market timing skills. Panel A uses excess returns as the dependent variable. Panels B and C respectively adjust excess returns for the own-market conditional volatility ($\widehat{\sigma}_{jj,t+1}$) and the conditional covariance with UK stock returns ($\widehat{\sigma}_{juk,t+1}$), both obtained from the bivariate GARCH estimations reported in Table 1. The sample covers 247 UK pensions funds over the period 1991:1 - 1997:12.

Table 5: Market timing skills in up and down markets

	Japan	North America	Europe	Asia-Pacific
$A : \Delta\omega_{jt} = \beta_{1j} I_{\{\rho_{jt+1} \geq 0\}} + \beta_{2j} I_{\{\rho_{jt+1} < 0\}} + \varepsilon_{jt}$				
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	.81
Median $\widehat{\beta}_{2j}$	-.0023	-.0022	-.0011	-.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				
% 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				
% 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

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 the excess return in period $t+1$ is 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 (1981) 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. The sample covers 247 UK pension funds over the period 1991:1 - 1997:12.

Table 6: Tests for extra-market timing skills

	Japan	North America	Europe	Asia-Pacific
Median $\widehat{\beta}_{1j}$	0.0003	-0.0022	-0.00046	0.0007
% of regressions with $\widehat{\beta}_{1j} > 0$	79.76	0.00	2.43	91.96
% of regressions with $t_{\widehat{\beta}_{1j}} > 2$	0.00	0.00	0.00	0.00
Median $\widehat{\beta}_{2j}$	1.130	-1.512	1.587	.1130
% 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 $\widehat{\beta}_{3j}$	.045	-0.024	.0062	.031
% of regressions with $\widehat{\beta}_{3j} > 0$	94.74	22.67	59.11	4.86
% 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. It is based on the regression equation:

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

$I_{\{\rho^u_{jt+1} \geq 0\}}$ is an indicator function for the event that the unexpected excess return in period $t+1$ is non-negative. Conversely, $I_{\{\rho^u_{jt+1} < 0\}}$ takes a value of unity whenever unexpected excess returns are 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 for β_{1j} indicates that funds correctly anticipated returns above that expected given public information, while a negative sign for β_{2j} suggests that the funds correctly anticipated negative excess returns below that expected given public information. A positive sign for β_{3j} suggests that public information - as reflected in expected returns - influenced the portfolio weights. The sample covers 247 UK pension funds over the period 1991:1 - 1997:12.