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ABSTRACT

Performance Measurement using Multiple Asset Class Portfolio Data: A Study of UK Pension Funds*

Using a data set containing 364 UK pension funds' asset holdings, this paper provides a systematic investigation of the performance of managed portfolios across multiple asset classes. We find surprisingly little cross-sectional variation in the *ex-post* average performance across the UK pension fund portfolios as a whole as well as within asset classes. This finding we ascribe to the strong incentive the fund managers had not to underperform relative to their peer group. For domestic equities, by far the most important component of the portfolios, we find that fund size is the only variable that appears to account for an important fraction of the cross-sectional variation in measured performance.

JEL Classification: G12
Keywords: pension funds, asset allocation, performance measurement

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NON-TECHNICAL SUMMARY

This paper investigates the investment performance of the UK pension fund industry using data provided by the WM Company on 364 pension funds over the period 1986–94. The data comprises the pension funds' monthly holdings in eight classes of assets. Hence, we have a sample that is well-suited to a detailed examination of fund performance in terms of market timing (allocation of funds across asset classes) and security selection (allocation of funds within asset classes).

The industrial organization of the UK pension fund industry provides for an interesting case study. Over the period under investigation, UK pension fund managers faced arguably the smallest set of externally-imposed restrictions and regulations on their investment behaviour of any group of institutional investors anywhere in the world. They were, by and large, unconstrained by their liabilities: UK pension funds were running large actuarial surpluses until almost the end of the period under investigation. In addition, trustee (i.e. pension plan) sponsors interferred very little (if at all) in their day-to-day operations and, more importantly, in their choice of investments. Unlike many of their counterparts in continental Europe and elsewhere, UK pension fund managers were free to invest in almost any security in any asset class in any currency denomination and in any amount (although they did face trustee resistance to the use of derivatives, at least in the early part of the period, and there are statutory limits on self-investment in the sponsoring company). Finally, in contrast with their US counterparts, UK pension fund managers faced no substantive regulatory controls on or real threat of litigation over imprudent investment behaviour during this period.

This relative freedom accounts for several important differences between the portfolio holdings of US and UK pension funds. US pension funds are far more heavily invested in lower volatility domestic bonds than their UK counterparts, while, conversely, UK pension funds have a far larger portfolio weighting in higher volatility equities. As a consequence of the general absence of constraints on their investment behaviour, the WM data set should enable us to identify the genuine investment skills of a group of fund managers in a way that is not possible with other data sets on investment performance generated under more restrictive conditions. On the other hand, UK pension plan sponsors had relatively weak incentives to be concerned about overall fund performance during our sample period and, hence, this relative freedom might have resulted in comparatively little cross-sectional variation in measured performance in contrast with the striking differences observed in US data.

Moreover, fund manager performance was explicitly evaluated in relative terms in the long run, possibly exacerbating the incentive to produce similar results. Our data permit us to see if these incentive effects dominate actual portfolio behaviour or whether fund managers translated the absence of constraints into actively managed portfolios that sought to outperform their peer group.

The twin objectives of adding absolute value and performing well relative to the peer group provide fund managers with a mixed set of incentives. On the one hand, genuine *ex-ante* ability that translates into superior *ex-post* performance increases the assets under management and, thus, the base on which the management fee is calculated. On the other hand, this incentive is not particularly strong and active management subjects the manager to non-trivial risks. The incentive is weak because the prospective increase in the fee is second order, being the product of the *ex-post* return from active management and the management fee and thus around two full orders of magnitude smaller than the base fee itself. Moreover, the *ex-post* return from active management of a truly superior fund manager will often be negative and occasionally large as well, resulting in poor performance relative to managers who did not attempt to profit from active management irrespective of their ability. The probability of relative underperformance large enough to lose the mandate is likely to be at least an order of magnitude larger than the management fee. Hence, on balance, it would appear that the risk of underperforming due to poor luck would outweigh the prospective benefits from active management for all but the most certain security selection or market-timing opportunities.

These observations about underlying incentives appear to account for many of the robust regularities we report. We found surprisingly little cross-sectional variation in *ex-post* average performance across pension fund portfolios as a whole as well as within asset classes either from Jensen-style regressions, which ignore the information on asset allocations, or from decompositions of returns into components due to normal or strategic asset allocation, market timing, security selection, and a residual. While there are robust differences in *average* performance in the Jensen-style regressions from sources such as risk adjustment and the use of alternative peer-group and external benchmarks, the same kinds of modifications left the cross-sectional variation in measured abnormal performance virtually identical to that of the raw return data. Similarly, normal or strategic asset allocations, however modelled, account for the bulk of the cross-sectional variation in overall fund returns and exhibit remarkably little cross-sectional variation. To be sure, we found some evidence compatible with the hypothesis that some funds experienced either persistently good or persistently bad performance and some evidence of a non-trivial cross-sectional variation in the measured returns attributable to security selection. Nevertheless, these regularities are second order compared with the central observations on cross-sectional variation in measured abnormal performance and the central role of the strategic asset allocation.

Fund size is the one variable that apparently accounts for an important fraction of cross-sectional variation in measured performance in equity returns. This regularity makes it that much easier to understand the role of relative performance evaluation in the UK pension fund industry. Due to the considerable cross-sectional variation in fund size, it happens that two-thirds of the funds outperformed the peer-group benchmark in UK equities. Not surprisingly, large funds were over-represented among the relative underperformers.

Hence, the vast majority of managers could point to either their good performance or to a size handicap as a way of justifying the retention of their mandates. Since fund managers had reasonably strong incentives to produce similar, though not identical, results, the noise in asset returns suggests that funds also had a fair chance of producing performance good enough to warrant being offered new mandates as well. The industrial organization of the UK pension fund industry appears to be a clear case of economic actors following their incentives.

1 Introduction

A large literature in finance has studied the investment performance of US mutual funds, predominantly those invested in US equities. Some work has also been devoted to the performance of US pension funds.[1] However, investment performance both in asset categories apart from equities and by institutions outside the US has been much less intensively researched.[2] This omission is potentially important since differences in asset characteristics, institutional and legal frameworks and, indeed, differing investment cultures and compensation schemes make it far from clear that the same findings would obtain for non-US fund managers' investment performance.

To our knowledge, this paper provides the first systematic academic investigation of the performance of managed portfolios across multiple asset classes. We analyze a data set provided by the WM Company containing nine years of monthly information on the holdings in eight classes of assets by 364 UK pension funds. Hence, we have a sample that is well-suited to a detailed examination of fund performance in terms of market timing (allocation of funds across asset classes) and security selection (allocation of funds within asset classes). While this is a relatively short period, it is still nearly two years longer than the average duration of an investment management contract in the UK. Moreover, as it happens, a number of robust empirical regularities emerge from these data, suggesting that we have a sufficiently long sample to provide a fair assessment of some aspects of the performance of UK pension fund managers.

The industrial organization of the UK pension fund industry provides for an interesting case study. Over the period under investigation, UK pension fund managers faced arguably the smallest set of externally-imposed restrictions and regulations on their investment behavior

[1]Recent examples of studies in the latter category include Lakonishok, Shleifer, and Vishny (1992), Coggin, Fabozzi, and Rahman (1993), Christopherson, Ferson and Glassman (1995).

[2]Most of the studies on US mutual fund performance have not analyzed data on holdings of different types of assets, although there are some exceptions, e.g., Elton, Gruber, Das, and Hlavka (1993).

of any group of institutional investors anywhere in the world. They were, by and large, unconstrained by their liabilities: UK pension funds were running large actuarial surpluses until almost the end of the period under investigation. In addition, trustee (i.e., pension plan) sponsors interfered very little (if at all) in their day-to-day operations and, more importantly, in their choice of investments. Unlike many of their counterparts in continental Europe and elsewhere, UK pension fund managers were free to invest in almost any security in any asset class in any currency denomination and in any amount (although they did face trustee resistance to the use of derivatives, at least in the early part of the period, and there are statutory limits on self-investment in the sponsoring company). Finally, in contrast to their US counterparts, UK pension fund managers faced no substantive regulatory controls on or real threat of litigation over imprudent investment behavior during this period.

This relative freedom accounts for several important differences between the portfolio holdings of US and UK pension funds. US pension funds are far more heavily invested in lower volatility domestic bonds than their UK counterparts, while, conversely, UK pension funds have a far larger portfolio weighting in higher volatility equities. As a consequence of the general absence of constraints on their investment behavior, the WM data set should enable us to identify the genuine investment skills of a group of fund managers in a way that is not possible with other data sets on investment performance generated under more restrictive conditions. On the other hand, UK pension plan sponsors had relatively weak incentives to be concerned about overall fund performance during our sample period and, hence, this relative freedom might have resulted in comparatively little cross-sectional variation in measured performance in contrast with the striking differences observed in US data. Moreover, fund manager performance was explicitly evaluated in relative terms in the long run, possibly exacerbating the incentive to produce similar results. Our data permit us to see if these incentive effects dominate actual portfolio behavior or whether fund managers translated the absence of constraints into actively managed portfolios that sought to outperform their peer group.

Another virtue of the UK experiment concerns the nature of the benchmarks used to correct for systematic risk. Benchmark inefficiency is a central theme of both the theoretical and empirical literatures on performance evaluation, because of the difficulty in distinguishing benchmark inefficiency from abnormal performance. As an empirical matter, Lehmann and Modest (1987), Grinblatt and Titman (1989, 1994), and Elton, Gruber, Das, and Hlavka (1993) have found that measured US equity mutual fund performance can depend critically on the benchmark used in the analysis. Elton, Gruber, Das, and Hlavka (1993) and Ferson and Schadt (1996) highlight some of the misspecification problems associated with performance measurement that arise when the funds under consideration hold assets, such as international equities and bonds, that are excluded from the benchmark index.[3]

Our data permit us to deal with some of these issues. Since we know the structure of the asset allocation of these pension funds, we can use benchmarks that do not suffer from defects of asset coverage. That is, we can compare asset class returns with asset-specific benchmarks in both unconditional and conditional single-index models and with appropriate multiple-index benchmarks that, in both cases, represent all of the different asset categories actually held by the pension funds. Moreover, an institutional feature of the UK pension fund industry is that managers are explicitly compared against peer-group benchmarks and data on these have been provided to us by WM as well. Hence we can assess the importance of relative performance evaluation, a topic which has received some attention in recent papers such as Brown, Harlow, and Starks (1996) and Chevalier and Ellison (1995).

Multiple asset class data provides an opportunity to search over different dimensions of abnormal performance ability. Any similarities in overall performance might conceal nontrivial

[3]Elton, Gruber, Das, and Hlavka (1993) found that the inclusion of a bond index and a small cap equity index substantially altered the estimated alphas in their analysis of the performance of a large universe of US mutual funds.

differences when disaggregated across asset classes and over time. Similarly, any such differences might show up as abnormal performance in related asset classes, an outcome that would arise as spillover effects in performance across asset classes. Moreover, these asset allocation data permit us to generate more precise measures of any market timing ability by these fund managers along the lines, for example, of Brinson, Hood, and Beebower (1986) and Brinson, Singer, and Beebower (1991).

The opportunities afforded by multiple asset class portfolio data engender new problems as well. Chief among them is the problem of distinguishing between short term tactical asset allocation and long term strategic asset allocation decisions. There was a substantial and systematic increase in the allocation to both domestic and international equities over the sample which complicates the interpretation of the short-term dynamics in portfolio weights. Accordingly, we propose new decompositions of portfolio weight changes which seek to measure the relative importance of passive and active fund management, both in the short and long run.

The structure of the paper is as follows. We begin with a brief review of pension funding arrangements in the UK. We then examine the overall investment performance of UK pension funds relative to two sets of benchmarks: a set of industry-standard and a set of peer-group benchmarks (Section 3). We discuss the potential for survivorship bias in Section 4 and how the choice of benchmarks affects the relative performance of fund managers (Section 5); this section also examines whether there is any persistence in relative performance over time,[4] whether there are spillovers in relative performance between asset categories, and whether relative performance is influenced by the size of the fund under management. Finally, we

[4]See, for example, Hendricks, Patel, and Zeckhauser (1993). Malkiel (1995) found evidence of persistence in portfolio performance in a sample of US equity fund managers during the 1970's but not for the 1980's. Lehmann and Modest (1987), Grinblatt and Titman (1992), and Christopherson, Ferson, and Glassman (1995) find persistence mainly among poorly performing managers.

analyze the asset allocation decisions of fund managers (Section 6). Changes in portfolio weights over time are decomposed into return and cash flow components. We also decompose performance into security selection and market timing components. Section 7 concludes the paper.

2 Pension Funding in the UK

There are four main types of pension scheme operating in the UK. First, the basic state scheme pays a flat-rate fully indexed-linked pension on a pay-as-you-go basis. Membership in this scheme is mandatory for all of the UK's employed and self-employed population with earnings above the small earnings exemption limit. Second, the state earnings-related pension scheme (SERPS) pays a defined benefit pension (currently 25 per cent of average revalued earnings within an earnings band), again fully index-linked on a pay-as-you-go basis. Membership in this scheme is automatic for employees unless the employee's pension has been specifically contracted-out into an eligible private sector scheme; the self-employed are not eligible for membership. Third, occupational pension schemes (which are established under trust law) are funded and pay a defined benefit pension (typically one-sixtieth of final salary for each year of membership up to a maximum of forty years); the pension in payment is indexed to retail price inflation up to a maximum of 5% p.a.. There are about 150,000 occupational schemes in the UK, covering 10.7 million active members and 7 million pensioners. Most such schemes are very small, having only a few members each, and generally have their assets managed on a pooled basis by insurance companies. Finally, personal pension schemes are funded on a defined contribution basis. There are about 5.5 million personal pension schemes in the UK and roughly 4 million people are drawing personal pensions which are index-linked up to a maximum of 5% p.a.. Blake and Orszag (1997) estimated that the values of the pension rights associated with each of these four schemes in 1993 were as follows: L686 billion for the basic state scheme, L178 billion for SERPS, L711 billion for occupational schemes and L73 billion for personal schemes. In comparison, personal net financial wealth totalled L702 billion and housing wealth was about L929 billion in the same year. Pensions rights therefore comprise a very large fraction of aggregate wealth.

Trust law is very flexible in the UK, enabling the trust deed to be drawn up in virtually any way that suited the sponsor, so that the sponsor can ensure effective control of the fund through the appointment of the trustees. To be sure, the trustees have a fiduciary duty under the Trustee Investments Act of 1961 to preserve the trust capital and to apply the capital and its income according to the trust deed and members can sue for compensation under the 1925 Trustee Act if they suffer loss as a result of negligence by trustees. In addition, pension fund managers are authorized by the Investment Managers Regulatory Organization, a self-regulatory organization established under the Financial Services Act of 1986. Nevertheless, there was no external regulatory oversight of pension funds during our sample period, leaving pensioners with only the opportunity for remedy through the courts.[5]

Another unusual aspect of the UK experiment concerns overfunding. Huge pension fund surpluses built up during the early 1980's: Keating and Smyth (1985) estimated these at L77 billion in 1985, equivalent to half the value of pension fund assets at the time. The main causes of overfunding were the UK equity market boom between 1974 and 1987, which led to huge

[5]The changes introduced by the 1995 Pensions Act bring the UK pensions regulatory framework closer to the prudent-man principle established by the US Employee Retirement Income Security Act of 1974. However, substantial differences remain: for example, the compensation scheme established by the 1995 Pensions Act explicitly sought to avoid the problems with deliberate underfunding faced by the Pension Benefit Guaranty Corporation. Similarly, the trustees must now conduct an asset-liability modeling exercise that obliges them to establish a strategic or long-run asset allocation. As in the US, the trustees are usually advised in this process by pension consultants and/or actuaries.

increases in pension asset values, and the 1980-81 recession, which caused substantial layoffs among pension scheme members leading to a freezing of their deferred pension liabilities. The 1986 Finance Act required that pension fund surpluses be reduced to no more than 5% of liabilities within 5 years of 1987 and this could be achieved by any of the following means: employee contribution holidays, employer contribution holidays, enhanced payments to pensioners and deferred pensioners, or a return of contributions to the employer (subject to a 40% tax rate without any offset). Not surprisingly, the most common choice among firms was the employer contribution holiday. By the mid-1990's, pension fund surpluses had largely vanished[6] but, for most of our sample, firms generally could not use their pension funds as a tax haven and any superior performance increasing asset values would have resulted in further reductions in employer contributions.

Some general information on the structure of the UK fund management industry, albeit toward the end of our sample, can be gleaned from a survey conducted by Prosser (1995). Many of the fund managers in our sample will have come from the same management groups. The top four fund management groups in the UK–Mercury Asset Management (MAM), Phillips and Drew Fund Management, Gartmore Pension Fund Managers, and Schroder Investment

[6] Apart from the decline in contributions into schemes, there is another important explanation for the reduction in surpluses over this period. Prompted by the high inflation of the 1970s, the government introduced a series of Acts (beginning with the 1973 Social Security Act and ending with the 1995 Pensions Act) that led to pension scheme liabilities becoming increasingly indexed against inflation. For example, pensions in payment now have limited price indexation up to 5% p.a.. A key piece of legislation for our sample period was the 1985 Social Security Act which gave so-called 'early leavers' the right to a preserved pension in the scheme that they were leaving; part of the preserved pension (the 'guaranteed minimum pension') had to be fully indexed to increases in national average earnings and the rest was subject to limited price indexation up to a maximum of 5% p.a. compound. Because, over the sample period, pension scheme liabilities were increasingly becoming 'real' liabilities, pension fund managers had to begin switching from assets fixed in nominal terms to assets that provided greater long term hedges against inflation, particularly equities.

Management—collectively managed 1154 funds between them at the end of 1993. Most fund managers wished to be seen as offering a 'balanced' service, in part because UK fund managers tended not to want to be typecast in the past. In contrast, US fund managers are usually characterized by an investment style and, if selected by a client with the aid of a consultant, cannot subsequently change their style. At least 70 per cent of funds with assets exceeding L100 million use more than one fund manager. Of the 95 per cent of funds that were not indexed, 40 per cent were managed on a fully discretionary basis with the trustees exerting no influence on tactical asset allocation or security selection, 19 per cent were managed on a multi-asset mandate against a specific benchmark, often customized to reflect any idiosyncrasies in the strategic asset allocation, 17 per cent were managed on a specialist, single-asset mandate with different specialists and benchmarks in each asset category, and the remaining 19 per cent were managed by other (unspecified) methods. Recently fund sponsors have begun to monitor their managers much more closely: the proportion of fully discretionary mandates fell from 58 per cent to 40 per cent of the total between 1993 and 1994 alone.

Performance measurement services are provided by WM or CAPS (Combined Actuarial Performance Services). Interviews with industry practitioners suggest that managers tend to be assessed on their relative performance against other fund managers: good managers are those who appear consistently in the upper quartile, even in years in which all fund managers performed badly. Potential new managers are invited to join the 'beauty contest' for a mandate—an investment management contract—only if they have enjoyed a good relative performance record over the previous three years.

One perhaps surprising feature of pension fund management in the UK, in contrast with the US, is that there is rarely a change of fund manager, although the terms of mandates, such as the benchmark to be used, sometimes change. To be sure, some of this stability arises from the expense associated with shifting management. However, the principal cause appears to be the efforts of fund managers to maintain the confidence of their clients through consistent track records based on relative performance. Indeed, one fund manager informed us that managers do not get fired for past bad performance but rather for lack of confidence in future performance which, for example, might be signaled by major changes in personnel or systems or because important clients begin to leave.

In fact, all of the 364 managers in our data set were in place throughout the sample period 1986-1993 and the average length of tenure of a pension fund manager in the UK is 7.25 years (Prosser (1995)). The largest fund management groups have the most secure reputations: 'Just as "no one ever got fired for buying IBM", so "no one ever got sued for hiring MAM"' (Kay, Laslett, and Duffy (1994)). According to Kay, Laslett, and Duffy (1994), the largest fund management groups appear to use their track records to retain existing clients or to attract new clients, rather than to extract higher charges. In addition, UK pension fund trustees tend to place a high value on the service provided by the fund manager. Good service and good personal relationships between fund managers and trustees can compensate for periods of poor investment performance and so also help to retain mandates in these circumstances. These considerations all point to substantial disincentives to actively manage portfolios in ways that risk large differences in relative performance.

The management fees charged by a fund management group are related to some extent to the performance of the manager, either directly or indirectly. In the case of balanced management, the fee is proportional to the value of the fund and therefore rises if the fund manager adds value or if the fund does well by chance. However, specialist mandates tend to be more directly performance-related than balanced mandates. The fee in this case involves a value-related component designed to cover the fund manager's costs plus a component related to the fund's outperformance of an agreed upon benchmark. In most cases, performance is measured relative to the peer group, not to external benchmarks, and relative performance benchmarks can give managers the incentive to place bets that do not deviate too much from industry norms.

To get some notion of the size of the fees charged by UK pension fund managers, we obtained the current fee structure of three major UK fund managers. MAM's management fees for balanced, segregated funds is as follows (reported in Kay, Laslett, and Duffy (1994)). For funds up to L50 million in value: 0.75% on the first L1 million, 0.5% on the next L4 million, 0.3% on the next L5 million, and 0.15% on the next L40 million. For funds between L50m and L100m, 0.175% on the first L50m and 0.15% on the next L50m. The rate for funds between L100 million and L200 million is 0.15% and the fee is negotiable for funds greater than L200 million. Gartmore's management fee (for 1996) is as follows: 0.5% on the first L25 million, 0.3% on the next L50 million, and 0.2% thereafter, except for a negotiable fee above L150 million. The fees of another large fund management group (which asked not to be identified, but also for 1996) are 0.5% on the first L20 million, 0.3% on the next L30 million, 0.25% on the next L50 million, and 0.175% thereafter. In each case, the marginal fee is falling but the total fee is indirectly performance-related because the fee increases with the value of the fund (in practice, the fee is paid quarterly depending on the value of the fund at the end of each quarter). Further details of the investment environment faced by UK pension funds during the late 80's and early 90's are contained in Stevenson (1993) and Blake (1995).

The institutional arrangements that we have just reviewed reveal important features of the UK experiment. They help us to place identifying restrictions on the behavior of our group of fund managers over the period under investigation and, hence, to improve the power of our tests. The most important lessons that we learn are:

(1) UK pension fund managers have a weak incentive to add value and they are largely unconstrained in the way in which they attempt to do so. While the strategic asset allocation may be set by the trustees in principle, any such limits are so flexible as to be effectively

unenforced both because of wide tolerance in allowable deviations of short-run from long-run asset allocations and because the strategic asset allocation itself can be renegotiated in most cases.

(2) Fund managers know that their relative performance against their peer-group rather than their absolute performance determines their long-term survival in the industry.

(3) In the short term (during the course of the current mandate), most UK pension fund managers earn fees related solely to the value of assets under management and not to their relative performance against either a predetermined benchmark or their peer-group (i.e., there is, in general, no specific penalty for underperforming and no specific reward for outperforming an agreed upon benchmark).[7]

3 Data Description

Our data consists of monthly observations on 364 UK pension funds covering the period 1986 - 1994. The sample is complete in the sense that it contains all of the pension funds that maintained the same single, externally-appointed fund management group throughout the period 1986-1993. Later we added data for 1994 causing a small decline in the number of funds still with the same manager to around 330 for reasons discussed in Section 4. For each fund, data was available on 8 constituents of the portfolio: UK equities, international equities, UK bonds, international bonds, UK index-linked bonds, cash, UK property, and international property. In addition, we had data on the total portfolio of each fund. For each group of assets, every fund in the sample reported initial market value (IMV), net investment in the asset over the month (NI), the mean (time-weighted) asset value over the month ($MeanAsset$), income received over the month ($Income$) and return on the asset. Returns were calculated as follows:

[7]For a comparative analysis of the incentives operating in the US pension fund industry, see Lakonishok et. al. (1992).

$$r_{i,j,t} = \frac{IMV_{i,j,t+1} - IMV_{i,j,t} - NI_{i,j,t} + Income_{i,j,t}}{(MeanAsset)_{i,j,t}} \tag{1}$$

where $r_{i,j,t}$ is the return on the i'th fund's j'th asset holding in period t. All assets were denominated in pounds sterling. WM uses a range of asset-class benchmarks to assess the performance of the funds in its stable.

The original set of external indices that it used is listed below:

UK Equities: Financial Times Actuaries (FTA) All-Share Index.

International Equities: Morgan Stanley Capital International (MSCI) World Index (excluding UK) for 1986, followed by FTA World (excluding UK) Index for 1987-1994.

UK Bonds: British Government Stocks All-Stocks Index.

International Bonds: JP Morgan Global (excluding UK) Bonds Index.

UK Index Linked: British Government Stocks Index-Linked All-Stocks Index.

Cash: LIBID 7 day deposit rate.

UK Property: WM PUT Property Index for 1986, followed by International Property Database (IPD) All-Property Index for 1987 - 1994.

Total Assets: WM Pension Fund Index.

All of these indices are value-weighted and assume that income is reinvested. Although a bond index, the BGS All-Stocks Index is referred to as a stock index since stocks and bonds are interchangeable terms in the UK. The BGS indices use a mixture of government bonds across all maturities. For the cash assets, the 7-day LIBID rate was used on the grounds that, although not strictly a performance measure, it is widely used as an institutional standard. The property index assumes that gross income from property is reinvested. The WM Pension Fund Index for total assets is a value-weighted measure based on all pension funds monitored by WM. No index was available for international property during the sample period. However, this is not a major problem for our analysis since international property comprises less than

0.5 per cent of the total portfolio value in our sample.

The above benchmarks have the virtue of being independently-calculated indices that are immediately publicly available and they are widely used for short-term performance measurement in the UK. WM selects them to match the actual behavior of UK pension fund managers as closely as possible and periodically changes them to reflect changes in pension fund behavior as well as the introduction of higher quality alternative benchmarks. For example, WM initially used UK bond indices that contained only bonds with remaining maturities in excess of 15 years and subsequently switched to all-stock indices. Similarly, they switched to the more comprehensive FTA World Index and the IPD All-Property Index during the sample period.

Despite such changes, pension fund managers were concerned that these external indices did not and, indeed, could not reflect their actual behavior sufficiently accurately, particularly in the case of international investments where UK funds' holdings could not be expected to match the relevant world indices. While the UK stock market accounts for less than 10 per cent of the global stock market capitalization, no fund held international assets in the same market-valued weights as the world's other stock exchanges for a number of reasons. First, managers might not invest in some markets because market liquidity is too thin or trading is insufficiently free from corruption. Second, restrictions on foreign investment in some markets may preclude a fund from achieving the market-value weighting. Third, funds might simply prefer to concentrate on just a few international markets. Fourth, and most importantly, UK pension funds exhibit the same puzzling home country bias found in all capital markets.

For reasons such as these, WM developed a new set of indices: the WM universe or sub-universe indices which use value-weighted portfolios of the population of funds actually tracked by WM. To capture the effect of these peer-group benchmarks, we also report results based on the WM2000 indices which comprise all funds ranked below the largest 50 funds tracked by WM. These are the indices more typical of the weighting achieved by single externally-appointed fund managers.[8] WM regards these peer-group weighted averages as the most appropriate measures of medium- to long-term performance.

Some of the assets are not very liquid and, hence, it is hard to get their market values. For example, bonds are carried at market values when possible but some of the prices for illiquid bonds are 'indicative' quotes from market makers as opposed to transaction prices, a practice not likely to be a problem for all but the most illiquid bonds. In aggregate, less than 5 per cent of total bond holdings are in illiquid corporate bonds and the rest are in government bonds which have substantial liquidity. Property is the other asset class for which the unavailability of market prices is a major problem. Property is valued: (1) once every three years with one-third of the property revalued each year, (2) over a five-year period with 20 per cent of the portfolio being reappraised annually or (3) annually (in the case of smaller funds). There are also two alternative ways of carrying the value of a property between valuations. Either property is carried at book value with any expenditure incurred on the property added to book value, but with no allowance being made for capital growth, or an estimate of capital growth is added to the book value between valuations. Not surprisingly, we are reluctant to draw strong conclusions regarding the property component of pension fund asset allocations.

Other facets of the data set should be noted. First, we were not able to get information on the exact transaction costs (spreads and commissions) or running costs (custody fees, property

[8]These WM indices are prepared and used in the following way. They are not available instantaneously from the end of each month because the funds do not supply WM immediately with their asset mixes and returns. Suppose we are at the end of one quarter (e.g. end of March). One week later (end of first week of April) WM sends out a flash report of performance based on the latest asset mix and index returns known to them. After another week WM sends out a second flash report based on the real fund data for the funds that have sent them information by this time. This updating continues until the sixth week (middle of May) when the whole universe has submitted returns to WM. Since this exercise is only performed quarterly, the monthly returns for April and May will be based on the full set of March weights. These weights are then multiplied by the corresponding asset returns and summed to derive the WM pension fund index.

security and insurance costs and so on) incurred by the various funds. Therefore the fund returns used in this study are gross of all these costs, except dealing spreads which are automatically included in the data reported to WM.[9] However, the commission schemes quoted in Section 2 are likely to provide a good indication of the fees charged by pension fund managers. Second, we were not able to get meaningful data on the trading activity of funds. Ten or fifteen years ago, a measure such as asset turnover (i.e., (gross cost of asset purchases + net proceeds from asset sales)/mean asset value)) might have provided a reliable indicator of activity. However, with the growth of indexing, hedging, and asset-liability management, any transactions designed to match an index, hedge a part of the underlying portfolio, or respond to the increasing maturity of the liabilities might increase turnover without actually signaling an increase in active management. In fact, the opposite conclusion might be nearer the truth. Prosser (1995) found that 5 per cent of funds were managed on an indexed basis, 68 per cent of fund managers were using currency hedging, 52 per cent were using other derivatives, and 9 per cent were using portfolio insurance in 1994. Blake (1992) reports a survey by Greenwich Associates which found that 30 per cent of pension funds were using asset-liability modeling in one form or another.

Pension funds of very different sizes populate our sample. By the end of the sample, the smallest funds had assets just above L1 million and 28 funds had assets below L10 million. At the other end of the scale, two funds had assets between L10 billion and L20 billion. The vast majority of funds in our sample had assets between L10 million and L1 billion, and the median

[9]Note that, in contrast, the index returns are gross of all costs including dealing spreads. This has the effect of marginally penalizing fund managers when their performance is compared against the index returns. Of course, such a penalty is appropriate when funds could have been passively managed at extremely low cost in the external benchmarks.

fund size in December 1994 was L54.4 million.[10]

Before proceeding with the analysis it is worth first describing two regularities that pose the greatest empirical challenges to any interpretation of the performance of the UK pension fund industry. The first concerns the behavior of the overall asset allocation of UK pension funds: the substantial trend toward domestic and international equities and away from domestic bonds with more modest variations in the allocation to other asset categories. The second involves the cross-sectional variation in returns across pension funds. We briefly describe these regularities in turn.

Table 1 reports the annual portfolio allocation across eight categories of assets for all the funds contained in the sample supplied by WM. For comparison, the aggregate portfolio weights for the entire population of UK pension funds tracked by WM are also reported in the table. Reassuringly, the asset allocation of the pension funds in our sample is very similar to the overall asset allocation of the WM universe.

By 1993, domestic and international equities comprised more than 78 per cent of the aggregate portfolio value of UK pension funds, by far the highest pension fund equity allocation in the world and a substantial increase in equity exposure compared with the already high level of 70 per cent prevailing in 1986. Pension fund assets invested in UK equities declined between 1975 and 1983 and then rose dramatically between 1984 and 1993. International equity investment experienced a more striking increase. Prior to 1979, funds incurred penalties when investing in foreign assets, which had to be purchased using 'investment currency' that effectively taxed such investments. The jump in international equity holdings after the abolition of UK exchange controls in 1979 is discernible in the data: after the abolition of these controls, the average portfolio weight in international equities rose from 6 per cent in 1979 to 20 per

[10] Adjusting for the growth in assets over the sample period, which averaged 8.8 per cent per year, similar size distributions for the funds' total assets were obtained at the beginning and middle of the sample.

cent in 1986, temporarily declining in 1987 and 1988 before again significantly increasing again between 1988 and 1993. Originally, most international investments were in North American, continental European, and Japanese stocks, but the Pacific Rim (apart from Japan) became increasingly important in the later part of the period. Only a relatively small fraction of UK pension fund assets was invested in emerging markets by the end of 1993.

In contrast, UK pension funds decreased their holdings of UK bonds from 12 to 5 per cent while international bonds experienced a modest increase, rising from one to three per cent. The proportion invested in UK index-linked bonds (introduced for the first time in 1982) was quite stable, if low, throughout the sample. The increase in equity exposure and decrease in bond holdings over the period clearly indicates that the pension funds included in our sample had not reached stable long-term asset allocations. The movement away from property was linked to a combination of high operating costs and poor returns for this asset over the period. Cash holdings displayed some volatility over the period, as one might expect from a short-term asset.

The final columns in Table 1 give the average portfolio holdings for US pension funds at the end of 1986, 1990 and 1994, as supplied by Greenwich Associates. These figures confirm the striking differences between the holdings of UK and US pension funds. UK pension funds hold around 10 percentage points more of their portfolio in domestic equities and around 15 percentage points more in international equities. Similarly, they hold around 30 percentage points less in domestic bonds and three percentage points or so less in cash compared with their US counterparts. Historically the equity premium–the mean excess return on stocks relative to T-bills–has been very large both in the US and UK (in excess of six per cent in each country, albeit with considerable variation), suggesting that there should be substantial cross-country differences in pension fund performance, and the much larger fraction of portfolio wealth invested by UK pension funds in equities is further testimony to the unique lack of constraints on their choice of portfolio holdings.[11]

These regularities pose a challenge to any analysis of the performance evaluation of the UK pension funds. Taken together, UK pension funds tilted their asset allocation toward equities, which did well over the sample, and away from domestic bonds, which only slightly outperformed cash and did more poorly than the aggregate pension portfolio.[12]

The final column of Table 2 reports the average cross-sectional variation in total return. As is readily apparent, there is remarkably little cross-sectional variation in annualized total returns across the funds in our sample. The semi-interquartile range runs from 11.47 percent to 12.59 percent per year and less than 300 basis points separates the funds in the 5th and

[11]Almost all funds in the sample reported holdings of UK and international equities and cash/other investments. In addition, between 200 and 350 funds report holdings of UK bonds and around 200 funds report holdings of UK property. Far fewer funds hold the more 'exotic' assets such as index-linked bonds, international bonds, and international property. Interestingly, many funds almost entirely eliminated their holdings of UK bonds and replaced them with international bonds over the period 1986-1993, a particularly surprising finding in light of the behavior of US pension funds. This finding can perhaps be explained by the fact that most UK pension fund liabilities are indexed to inflation and by the corresponding low level of real returns on UK bonds. Over the period 1946-1995, the geometric mean real return on UK equities was 6.65 per cent while it was only -0.11 per cent on gilts and 0.79 per cent on T-bills (BZW (1996)).

[12]Not surprisingly, examination of the plots of monthly portfolio weights reveal the same trends along with some intra-year variability. However, Table 1 gives no indication of the degree of the cross-sectional heterogeneity in asset allocations. Time series plots of the cross-sectional standard deviation in the weights allocated to the individual asset classes show little evidence of trends save for an increase in the dispersion of portfolio weights in cash/other investments, UK property, and international bonds during 1990. Note also that the trends suggests the presence of a common component in UK pension funds' portfolio weights. For example, the cross-sectional dispersion of portfolio weights would have increased if only a fraction of the pension funds increased their holdings of UK equities from 1986 to 1994 while the remaining funds kept their portfolio weights in that asset class constant. Grinblatt, Titman and Wermers (1995), found evidence of a strong common component in mutual funds' holdings of individual equities and have interpreted this as evidence of herd behaviour among mutual funds. We provide evidence on the cross-section of asset allocations in Section 6.

95th percentiles. To be sure, there is somewhat more cross-sectional variability in particular asset classes. For example, the annualized semi-interquartile range for UK equity returns is of the order of 150 basis points and the corresponding 5th-95th percentile range is 400 basis points. The corresponding ranges are larger for international equity returns, with a semi-interquartile range of more than 200 basis points and a 5th-95th percentile range of 450 basis points. Nevertheless, these ranges are small compared with those observed in other performance evaluation settings, such as in the analysis of US equity mutual funds.

This comparatively narrow range of cross-sectional return variability, coupled with the likelihood that a nontrivial fraction of this variability can be attributed to cross-sectional differences in strategic asset allocations, suggests that any differences in performance ability among the managers in our sample must show up conditionally. This statement is nothing more than the observation that an unconditional distribution with low variability can conceal highly variable distributions once nontrivial conditioning information is taken into account. Hence, a diligent search for abnormal performance ability must consider performance conditional on risk exposures (since the unconditional cross-sectional distribution might conceal both low beta funds with good performance and their converse), on size (because there might be diseconomies of scale in asset management resulting from market impact), on prior performance (since funds in the upper tail for the whole period might consistently be in the upper tail over subperiods as well), and on performance in other asset categories (since a manager who performs well in one asset class might perform well in another). Of course, the unconditional regularity will be all the more striking if conditional cross-sectional variation proves to be small as well, as seems likely to us *a priori*.

Finally, Table 3 reports summary statistics describing the volatility and comovements among aggregate asset class returns. It presents the standard deviations of the monthly returns on the (value-weighted) portfolio components (along the diagonal). Returns on UK equities were the most volatile throughout the sample followed by those on international equities. Unsurprisingly, returns on cash/other investments had by far the smallest standard deviation. The off-diagonal elements of Table 3 contain the correlations between the monthly returns. Returns on UK equities were highly correlated with returns on international equities and were also positively correlated with returns on the various bond categories.

4 Survivorship Bias

An important component of our experiment is the examination of the persistence of investment performance over time. Accordingly, we found it essential to use a sample containing performance data on the same fund management groups over an extended period since the power of our tests increases with sample size. However, the selection of managers who are in place over the whole sample introduces another potential problem that has recently received substantial attention in the literature, namely survivorship bias.[13] Hence, there is a tradeoff between greater precision induced by larger samples and the bias induced by sample selection in our performance measures.

Funds were excluded from the sample supplied to us by WM for one of five reasons. First, funds that switched between managers are excluded from the sample, the potentially most pernicious source of survivorship bias. Second, company takeovers mean that funds are merged and merged funds are excluded. Third, funds might withdraw themselves from the WM measurement service with no explanation. Fourth, funds that switched from in-house to external management are excluded from the sample since this constitutes a change in management.

[13] For recent examinations of survivorship bias, see Brown and Goetzmann (1995), Brown, Goetzmann, Ibbotson, and Ross (1992), Grinblatt and Titman (1989, 1992), and Malkiel (1995). There is no consensus regarding the magnitude of survivorship bias. Malkiel finds that survivorship bias can account for mean returns of about 1.5 per cent per annum, while Grinblatt and Titman (1989) estimate the survivorship bias to be somewhat smaller at 0.4 of a per cent per year.

Fifth, some fund management groups permit WM to measure only a proportion of the funds in their stable in order to save costs and occasionally they will rotate these funds, a practice called 'dynamization', and such funds are dropped from our sample.[14] To summarize, funds were excluded from our data set because there was a change in manager, in management structure, or because they left or joined part of the way through the sample period. Funds are, however, not necessarily withdrawn because of poor performance.

Fortunately, we are in a position to directly assess some of the facts regarding survivorship bias in our sample. In addition to the performance and benchmark data discussed so far, WM supplied us with the returns on an index that includes the entire population of funds (1034 at the end of 1994) that they track for each asset class and for the aggregate portfolio as well. The annual value-weighted return of this index is given in Table 4 as the WM2000 return, while the annual value-weighted return of the funds in our sample is shown in the third column. If survivorship bias infected the funds included in our subsample–that is, making them more successful *ex post* than the average fund manager in the total sample of funds monitored by WM–we should observe higher returns in the value-weighted portfolio constructed from our sample of funds compared with those from the total population of pension funds tracked by WM. Moreover, the return differentials should be increasing over time, peaking toward the end of the sample as we systematically excluded more funds that performed poorly.

As is readily apparent, neither tendency arises on average or over time across asset classes and for the overall portfolio in Table 4. For example, the WM2000 return actually exceeded that of our universe by an economically trivial six basis points over the whole sample. It would also be hard to argue that survivorship bias infects the last half of the sample period since the peer-group index underperformed the value-weighted portfolio by -0.28 of a per cent per year

[14]The decrease in the number of fund managers included in our sample between 1993 and 1994 described in the previous section was mainly caused by this dynamization practice.

during the first half of the sample, but outperformed the latter portfolio by 0.39 of a per cent per year during the second half. Moreover, the time path of the signs in the return differential is the perfectly symmetric +, -, -, +, -, +, -, -, +. In addition, the differences are generally economically small in each year across asset classes, well within the range of variation that would arise from modest differences in the underlying portfolios. Hence, our sample of funds does not appear to be obviously affected by survivorship bias, if at all.[15]

A final reason why survivorship bias does not appear to be an important issue in our sample is the similarity between the evolution of portfolio weights in our sample and those of the population of pension funds in the WM universe. The year-by-year differences in the overall asset allocations seem numerically and economically trivial, yet we might have expected to observe large differences if fund managers systematically lost their mandates by making bad market timing bets. In Section 6, we show that the 'normal' portfolio weights, i.e., the average or trend-adjusted portfolio weights across asset classes, account for around 95 per cent of the variation in portfolio returns. In view of the similarity between the two sets of portfolio weights reported in the two panels in Table 1, we would not expect any substantial average differences between the returns on the funds in our sample and those on the population tracked by WM.

In short, the cost in terms of inducing potential survivorship bias seems to be small relative to the gains in precision from lengthening the sample. There is no evidence of survivorship bias on average. Of course, our calculations shed little light on any potential bias in the most extreme performers in the sample since the far left tail of the distribution has only a marginal effect on average performance. Hence, we should be cautious in drawing inferences about the left tail of the cross-sectional return distribution within and across asset classes.

[15] As a further check of the relationship between the returns on the external and peer-group indices and the value- and equal-weighted portfolios constructed from our sample of funds, we computed the correlation matrix for these four (total) return series. All elements of this correlation matrix exceeded 0.995.

5 The Performance of Pension Fund Asset Class Returns

In this section, we examine the performance of UK pension funds within asset classes. We begin by reporting some broad regularities in return performance followed by a more systematic evaluation of benchmark-adjusted returns for both the UK equity component in isolation and the total fund. We conclude with an attempt to identify funds that display some consistency in benchmark-adjusted performance in relation to fund size, prior performance, and performance in other asset classes.

5.1 Regularities in the Performance of Asset Class Returns

For each year 1986 through 1994 and for each asset group, Table 4 reports the mean return on the (value- and equal-weighted) portfolios of the funds in the sample, the mean returns on two benchmark indices, and the proportion of pension fund managers outperforming the relevant external index listed in Section 3 above.[16] In the case of domestic equities, the median manager outperformed the FTA All-Share index in only two of the nine years in our sample (1990 and 1992) and only 44.8 per cent of the funds outperformed the index over the entire sample. Similarly, the median international equity manager outperformed the index in only three out of nine years. However, the fraction of outperformers varies over time much more for international equities than for UK equities: the proportions of outperformers lie between 31 and 66 per cent for UK equities but ranges from 0 to 98 per cent for international equities. This merely confirms our suspicion outlined in Section 3 above, namely that general indices for international equities may not provide appropriate benchmarks for tracking the performance of UK pension funds in this asset category.[17]

[16]A similar performance comparison relative to a single benchmark was conducted for US equities by Lakonishok, Shleifer, and Vishny (1992).

[17]For example, sizable country bets by international equity managers appear to be made more frequently than market sector bets by domestic equity managers.

Performance in UK equities proves to be negatively correlated with the size of a fund's holdings of this asset class. At the beginning of each year, we sorted portfolios into quartiles based on the size of the pension funds' holdings in UK equities. This procedure was repeated for all nine years in the sample. The funds in the smallest-size quartile outperformed the funds in the largest-size quartile in eight of nine years, and for the complete sample the mean return on the smallest-size quartile portfolio was 13.47 per cent per annum against 12.80 per cent per year for the largest-size quartile portfolio. Further statistics on the relationship between fund size and portfolio performance are reported in Section 5.4 below.

Conventional bond fund managers have the highest proportion of outperformers of all asset categories (77 per cent with UK bonds and 69 per cent with international bonds). The proportion of outperformers in the case of index-linked bonds is also very volatile from year to year, ranging from 11 per cent to 90 per cent, although the average (value-weighted) return is very close to the external index average. Cash/other investments exhibited lower volatility than in the case of index-linked bonds, although the average (value-weighted) return is 89 basis points lower than that of the external benchmark. The average return on UK property generated by the funds compares favourably with the benchmark returns. International property delivered a disastrous performance over the period which accounts for its systematically declining weighting in pension fund portfolios.

The average value-weighted return exceeds the average WM2000 return in the case of both UK and international equities and both UK conventional and index-linked bonds. In other words, the average pension fund manager in our sample has outperformed the *peer-group* average on more than 80 per cent of the overall portfolio. At the same time, however, our average fund manager has failed to outperform the *external* benchmark in the case of UK equities along with index-linked bonds and cash, that is, on 60 per cent of the overall portfolio.

The key to these findings, of course, is the performance of UK equities which accounts for no less than half the value of the aggregate portfolio over the period. The average UK equity fund manager in our sample beat the peer-group average by 20 basis points on a value-weighted basis but underperformed the external index by 33 basis points before management fees and such are taken into account. For the total portfolio, the average value-weighted return is broadly comparable with the peer-group average but underperformed relative to the external total returns benchmark by 45 basis points, with only 43 per cent of funds beating this benchmark over the whole period.[18] Hence, the average fund manager in our sample, despite its longer relationship with its client than other pension fund managers monitored by WM and its outperformance of the peer-group average in UK equities, nevertheless had the same total performance as the peer-group and underperformed the market as a whole, although some of that underperformance is due to the fact that the external benchmarks do not account for dealing spreads and other costs.

5.2 The Performance in UK Equities against Single-Index Benchmarks

In this subsection, we conduct a more detailed examination of the performance of the domestic equity portfolios of the managers in our sample. In part, we do so to draw comparisons with the existing academic literature which mainly covers equity mutual funds. We also do so because of the central role in pension fund performance of UK equities which account for over half of fund value and, thus, well over half of fund risk exposure, however measured.

We investigated several variations on the basic Jensen regression in the single-index model:

[18]The fact that only 43 per cent of the funds outperformed the total returns index over the whole sample period while the median fund manager outperformed this index in five out of nine years arises because of the very poor performance of the funds during years in which the median fund manager underperformed the index.

$$r_{i,j,t} - r_{f,t} = \alpha_{i,j,t} + \beta_{i,j,t}(r_{m,j,t} - r_{f,t}) + \epsilon_{i,j,t} \tag{2}$$

where $r_{i,j,t}$ is the return on the i'th fund manager's j'th asset class (in this case UK equities) in period t, $r_{f,t}$ is the return on a 1-month T-bill, and $r_{m,j,t}$ is the return in period t on the j'th external index, in this case the UK equity index. For multiple asset class portfolios such as ours, the usual interpretation of the Jensen alpha as an indicator of the funds' security selection skills is only strictly valid if equity managers have been instructed to maximize risk-adjusted returns separately for the equity component of the overall portfolio. Even in this case, the Jensen alpha only provides a consistent estimator of selection ability if the fund managers do not also possess market timing skills unless we know or can model the fund $\beta_{i,j,t}$ (Jensen (1972), Admati et. al. (1986), Lehmann and Modest (1987), Grinblatt and Titman (1989)), an exercise we perform to some extent below.

Table 5 reports five versions of the standard Jensen regression. The first is the original Jensen regression with time-invariant alphas and betas which provides performance measures conditional only on differences in unconditional betas. Second, we follow Ferson and Schadt (1996) by permitting the beta in equation (2) to vary over time, allowing for predictable variation in risk exposures and, implicitly, in benchmark returns, on the grounds that fund managers should not be credited for changing their portfolio weights in the light of public information.[19] Third, we allow for predictable variation in alphas as well in a third regression, after the fashion of Christopherson, Ferson, and Glassman (1995).[20] Fourth, given the possibility that the value-weighted nature of the UK equity index might bias the alphas, we added the monthly returns

[19]Under the null hypothesis of no market timing ability, this procedure only affects the precision of the selectivity estimates since the time variation in betas is uncorrelated with the realizations of the index return, making it a component of the residual.

[20]As is common in the literature, we assume that $\beta_{i,j,t}$ in the second set of regressions and both $\alpha_{i,j,t}$ and $\beta_{i,j,t}$ in the third set are linear functions of predetermined variables, $\underline{z}_{t-1}$, which include the instruments commonly used in asset pricing applications: the lagged values of the dividend yield, the T-bill yield and the

on the Hoare-Govett small cap index to the unconditional Jensen regression. Finally, since all of these procedures are suspect if managers possess market timing ability, we also added the squared excess benchmark return to the unconditional Jensen regression, following Treynor and Mazuy (1966). If fund managers possess market timing ability, market timers should earn positive excess returns when benchmark returns are high while selection skills should show up as positive alphas in the absence of benchmark error under plausible assumptions (see Jensen (1972), Admati, Bhattacharya, Pfleiderer, and Ross (1986), Lehmann and Modest (1987), and Grinblatt and Titman (1989)).

The behavior of the Jensen alphas from these models should differ depending on the nature of the underlying economic environment and the hypothesized market timing ability of the fund managers. If the investment opportunity set is unchanging—that is, if benchmark returns and their first few moments are time invariant—and the fund managers have no market timing ability, all models with the same benchmark will produce alphas and betas with the same expected values. In particular, the cross-sectional distribution of the alphas should be identical across models, holding the benchmark constant. The interpretation is more problematic if the investment opportunity set is time-varying—that is, if the mean, volatility, and, perhaps, higher moments of the benchmark return experience predictable variation. The Jensen alphas and betas would be biased estimates of their unconditional means in this case, even if managers possess no market timing ability, so long as the covariance between the fund beta and conditional benchmark return volatility is nonzero. Hence, conditioning on public information as in the second and third models and on the squared excess market return as in the fifth model can materially alter the distribution of the alphas to the extent that conditional variation in betas is negatively correlated with population alphas. Finally, conditioning on public information might eliminate some of the cross-sectional variation in measured alphas to the extent that

long-term gilt yield. See Pesaran and Timmermann (1995) for a recent evaluation and references.

fund betas are correlated with conditional market risk premiums and volatilities.

Table 5 reports a number of summary statistics describing the cross-sectional distribution of the alphas from these models. We provide several fractiles of their distribution—5%, 10%, 25%, 50%, 75%, 90% and 95%—as well as their maximum and minimum values and their associated Bonferroni p-values.[21] We also supply the mean alpha and its t-statistic.[22]

Several regularities emerge from these models. Excluding the last column in Table 5, average performance is economically and statistically negligible, the largest alpha being measured in the conditional Jensen regression of Christopherson, Ferson and Glassman (1995) at only 33 basis points annualized. Similarly, the fraction of funds with positive alphas is less than 50% for all models except the conditional alpha model, where 58% of the estimates were positive, with 8% significant at the 5% level. In addition, the most extreme outperformer and underperformer had one-sided t-statistics with Bonferroni p-values well below the 0.0001 level, except for the case of the largest outperformer identified by the Ferson-Schadt regression, which had a marginal significance level of 0.015. Taken together, and, ignoring any concern about benchmark error and survivorship bias, these statistics suggest that there is little evidence of abnormal performance on average in this industry or indeed much evidence of extreme out- or underperformance that is significant at any reasonable level.

[21]Since the t-statistics of these alphas are interdependent and the number of alphas exceed the number of time series observations, we cannot construct a joint test of their significance. Moreover, the joint test generally has very low power if a small subset of the alphas differ from zero in the population, as we would generally expect *a priori* in the case of abnormal performance on the hypothesis that it is not pervasive. For both reasons, we report p-values based on the Bonferroni inequality, which in this case states that the marginal significance of the largest t-statistic in absolute value is less than π when its p-value is π/N, where N is the number of t-statistics examined simultaneously.

[22]Following the procedure established by Fama and MacBeth (1973), the standard error of this average alpha was computed from the time series of returns on the equal-weighted portfolio, although this procedure ignores the small downward bias associated with the omission of the sample squared Sharpe ratio of the index, c.f. Shanken (1992).

However, the main regularity concerns the shape of the cross-sectional distribution. The annualized semi-interquartile range in each of these models is about 150 basis points, similar to that of the raw UK equity returns. Furthermore, the 5th-95th percentile range is within 15 basis points of that of the raw UK equity returns. This implies that conditioning on alternative models for betas changes the location of the cross-sectional distribution of raw returns but leaves its shape virtually unchanged. Pension funds with similar performance by any of these measures also face similar risk exposures. Moreover, any shifts in the betas of these funds had sufficiently low correlations with benchmark returns or publicly available conditioning information so as to leave the cross-sectional distribution of the *ex post* alphas unchanged.[23] This suggests that market timing switches between asset classes do not seem to be an important contributor to cross-sectional variation in average equity returns within the UK pension fund industry.

Of course, we have an alternative to this risk-adjusted performance evaluation approach that is more appropriate if managers are evaluated relative to peer-group benchmarks. That is, we can replicate WM's performance evaluation by comparing fund performance in UK equities with that of the WM2000 UK Equity Index ($r^*_{m,j,t}$) as in:

$$\alpha_{i,j,t} = r_{i,j,t} - r^*_{m,j,t}. \tag{3}$$

In contrast with the previous methods, the peer-group approach requires no estimation of risk exposures, since relative performance evaluation implicitly sets $\beta_{i,j,t}$ to one. Recent empirical evidence, e.g. Brown, Harlow, and Starks (1996) and Chevalier and Ellison (1995), suggests the importance of relative performance evaluation for US equity managers as well.

[23] An even stronger result was that the location of the individual alpha estimates within the cross-sectional alpha distribution was found to be very robust with respect to the choice of risk-adjustment procedure. For example, using the unconditional and conditional Jensen procedures, the cross-sectional rank-correlation between the funds' mean (raw) excess returns and their alpha estimates was 0.99 and 0.88, respectively.

The results reported in the last column of Table 5 indicate that the UK pension fund industry practice of relative performance (or peer-group) evaluation provides an important identifying restriction on the assessment of managerial performance. Nearly two thirds of the funds outperformed the relative equity benchmark with 48 funds (16% of the total) having relative performance alphas that are significant at the 5% level. Many fewer funds earned negative alphas and fewer than 15 of these were significant at the 5% level. Average performance was positive: the mean alpha estimate was 0.459% per year with a t-value of 4.04. Of course, the shape of the cross-sectional distribution of these relative performance alphas is identical to that of the corresponding raw returns since relative evaluation only changes the location of the distribution.

Our sample of fund managers comprises those who have maintained the longest continuous relationships with their clients from the universe of fund managers monitored by WM. The relative performance benchmark suggests an obvious reason as to why this is the case. Nearly two-thirds of these fund managers beat the UK equities peer-group index, although only 16 per cent of the funds beat this index by a statistically significant margin. Of course, this observation begs several questions. For example, it does not explain why trustees retain poorly performing managers or why trustees appear to disregard the fact that performance is much poorer when the comparison is made with external benchmarks. Nevertheless, relative performance evaluation, which permitted two-thirds of the funds that held their mandates for longer than average to be in the top half of the class, is probably an important part of any explanation of the industrial organization of the UK pension fund industry.

5.3 The Performance of the Overall Portfolios

We conducted a comparable analysis for the total returns of the pension funds in our sample. Since we have data on asset-class-specific benchmarks, we implement the Jensen regression using a multi-index benchmark. That is, we compare the excess total return of the i'th fund with the excess returns on the entire set of indices:

$$r_{i,p,t} - r_{f,t} = \alpha_{i,p} + \sum_{j=1}^{M} \beta_{i,j}(r_{m,j,t} - r_{f,t}) + \epsilon_{i,p,t} \tag{4}$$

where M is the number of asset groups in the sample (seven in our case since we do not have an index for international property). Hence, $\alpha_{i,p}$ is the multifactor analogue of the standard Jensen measure, the only difference being that the excess return on the portfolio is calculated relative to a set of asset-class-specific indices and not to a single market index, and the potential pitfalls in its interpretation in the presence of market timing ability parallels that in the single index case.[24]

UK pension fund managers as a group tended to slightly underperform over the period in terms of their total return: 138 funds had positive alphas with only 9 (3% of the total) of these being significant at the 5% level, while 168 alphas were negative, of which 6 (2%) were significant as well. Their semi-interquartile range ran from -0.71% to 0.55%, an annualized range (115 basis points) that differed from that obtained from raw returns by only one basis point, while the alpha estimate for the equal-weighted portfolio was a minuscule -0.11% with a t-value of -0.17. The left tail of the cross-sectional distribution is neither long nor dense and the Bonferroni p-value for the most underperforming fund has a marginal significance level of only 0.62. Only the Bonferroni test statistic for the most successful fund is suggestive of abnormal performance, with a p-value less than 0.00001 indicating sharp rejection of the null of no outperforming funds at any conventional level. Of course, this rejection could still reflect benchmark error and survivorship bias as well.[25]

[24]We did not employ analogues of the time-varying beta models or of the Treynor-Mazuy regressions in this case since either approach would greatly increase the number of parameters, straining an already modest-sized sample.

[25]As we have noted previously, the asset class indices might still have biases associated with value weighting

Hence, once again, the cross-sectional distribution of these aggregate portfolio alphas suggests an absence of systematic abnormal performance and differs little from the cross-sectional distribution of average raw total portfolio returns. If anything, the cross-sectional variation in average total returns is more similar to its risk-adjusted counterpart than was the case with equity returns. The remarkably narrow range of realized average total returns remains a striking regularity even after conventional risk adjustment.

Relative performance evaluation is also conducted for the overall portfolio. By construction, adjusting for the peer-group benchmark return changes the location of the distribution relative to the raw total returns without changing its shape. We found that 197 of the funds (64% of the total) outperformed the peer group benchmark return over the whole sample, 41 (13% of the total) significantly so at the 5% level. Average fund performance was quite close to that of the peer-group benchmark, being an economically and statistically negligible 6 basis points below the benchmark return, but underperformed the external benchmark by a more substantial 45 basis points, c.f. Table 4. However, one justification for downplaying the external benchmark in performance evaluation lies perhaps in the extreme year-to-year variability in the fraction of funds that outperformed the external benchmarks, with proportions that were below 20% or above 66% in different years.

5.4 Portfolio Performance and Fund Size

The results above reveal two features concerning abnormal performance in the UK pension fund industry. First, a variety of benchmark corrections suggest that there are a few funds with robustly measured extreme abnormal performance, with the evidence for outperformance being stronger than that for underperformance (although this result may be an artifact of

and asset coverage, country coverage in the case of international securities and sector coverage in the case of domestic securities.

survivorship bias). The second fact strikes us as of greater economic significance: the shape of the cross-sectional distribution of average raw total and UK equity returns is broadly unaffected by risk adjustment, with even extreme ranges such as the 5th-95th percentile spread virtually unchanged. This suggests that cross-sectional variation in risk exposures does not conceal cross-sectional variation in abnormal performance.

It is possible that performance depends on other attributes of the fund. One such attribute is fund size. If it tends to be the larger funds that underperform the peer-group, then this would add credence to the often-made claim that size is the anchor of performance. A recognized fund size bias in measured performance would facilitate a better understanding of the role of performance evaluation in the UK pension fund industry.

Accordingly, we formed equal-weighted portfolios, based on quartiles sorted according to the value of assets at the beginning of each year, starting with the smallest funds. This procedure generated four time series on the portfolio returns for each asset class, the abnormal performance of which is presented in Table 6. Panel A reports the results for multi-index Jensen regressions for these portfolios. A size effect is observed most clearly for UK equities where the smallest-fund quartile has a positive alpha and the largest a negative alpha, neither of which is significantly different from zero at conventional levels but the difference between them (0.79%) has a t-statistic of 3.33 and an associated marginal significance level of less than 0.001. Panel B of Table 6 confirms these results when using the relative performance measurement procedure. Each of the fund-size-sorted portfolios has positive mean excess returns relative to the peer-group benchmark, increasing from an economically and statistically insignificant 2.7 basis points per year for the large-fund portfolio to 72 basis points per year for the small-fund portfolio. The remaining asset classes reveal no clear pattern save for international bonds and equities, which indicate a direct, rather than an inverse, relationship between fund size and

Jensen alpha.[26] Perhaps most importantly, there is no systematic relationship between fund size and abnormal performance for the overall pension fund portfolios.

Nevertheless, the finding of an inverse relationship between fund performance and fund size in UK equities could be an important part of the explanation for mandate retention in the UK. UK pension funds hold a very substantial proportion of UK equities and large UK funds hold large fractions of their portfolios in UK equities as well. These funds can surely argue that an annualized performance differential of the order of 70 basis points per year reflects the impact of the trading of large funds in a market in which they are important players.

5.5 The Relationship Between Current and Prior Portfolio Performance

Lehmann and Modest (1987), Grinblatt and Titman (1992), Hendricks, Patel, and Zeckhauser (1993), and Brown and Goetzmann (1995) find evidence of persistence in the performance of (mainly) the worst-performing mutual funds in the US. However, this finding is incompatible with the good measured performance of small UK pension funds.[27] Nevertheless, it is obviously of some interest to ascertain whether it is the size component of fund value or the prior performance component that is driving the negative relationship observed in UK equities between fund size and performance.

In this section, we consider whether the current performance of a given UK pension fund is correlated with its future performance. We adopted two approaches to maintain comparability with both the literature and the evidence in the previous subsections. The first simply examines the relationship between current and future rankings of relative portfolio returns without

[26] One possible explanation (which we are not able to test) lies in the economies of scale to information gathering in global asset markets.

[27] Funds can be small because they have just been established or because they were previously large but have generated large negative returns.

adjusting for their correlation with one or more indices. This approach is appropriate for investors with the bulk of their wealth invested in a single pension scheme and corresponds with our previous analysis of the persistence of raw portfolio returns. In the second approach, we investigate the persistence of Jensen measures obtained from average asset class returns after correcting for the correlation with the multiple-index benchmark. This benchmark correction is more appropriate for investors with only a fraction of their wealth invested in a particular pension scheme. In essence, this distinction reflects the difference between the Sharpe and Jensen-Treynor-Black approaches to the measurement of performance.

A useful indicator of persistence in performance is provided by the matrix of transition probabilities P_{ij} which measures the probability that a fund will belong to the j'th quartile in month t+12 given that it belonged to the i'th quartile in month t. The matrix of transition probabilities for the quartile-sorted portfolios is presented in the form of four-by-four contingency tables collected in Table 7.[28] The reported transition probabilities were calculated as an average of the estimated annual transition probabilities. If the future ranking of a portfolio is independent of its current ranking then the transition probabilities (which sum to 100 per cent across the columns) will take the value 25 per cent, while, for the opposite case where fund rankings never change, the transition probabilities are represented by a diagonal matrix with 100s along the main diagonal.

Four factors—two creating problems of interpretation and two creating potential statistical biases—might affect these transition probability estimates, each with empirically distinct implications. First, the funds managed by systematically better or worse informed managers would tend to remain in the top and bottom cells along the main diagonal, respectively, because we know that such managers retained their mandates, ignoring the reasons why systematically

[28]The persistence of performance of UK pension fund managers has also been examined by Brown, Draper, and McKenzie (1993), and Lakonishok, Shleifer, and Vishny (1992) analyze the returns of US pension funds in this fashion.

poor performers were able to do so. Second, the distribution of risk profiles across funds might also be important. Nontrivial exposure to a risk factor with persistently high (low) returns over the sample period would lead to consistently good (bad) performance, while low exposure would lead to correspondingly poor (good) performance *ceteris paribus*, thereby tending to increase the size of the transition probabilities along the main diagonal in Table 7. The potential importance of these first two factors suggests that we cannot uncritically associate persistence in measured probabilities with persistence of abnormal performance. In essence, this confusion arises from the standard practice of using raw rather than risk-adjusted returns.

There are also two factors that can create statistical biases. First, funds with substantial exposure to a volatile risk factor about which their managers possess no superior information should oscillate between the top and bottom quartiles more frequently than those with little exposure to it. We found little evidence of oscillation between extreme quartiles, suggesting either that systematic differences in risk factor exposures across funds are not important for explaining the transition probability estimates or that the risk factor returns are persistent over time.[29] Finally, the estimated transition probabilities are derived from sample returns that are noisy estimates of the corresponding mean returns. If the cross-sectional variation in sample returns (in contrast with the population mean returns) is dominated by noise (as seems reasonable *a priori*), we would expect some reversion from the extreme quartiles to the middle ones, since extremely positive (negative) prior returns reflect, on average, positive (negative) sampling error. This consideration suggests that the P_{11} and P_{44} estimates are, if anything, biased downward, strengthening any evidence of persistence.

The estimated transition probabilities in Table 7 suggest that funds are more likely to remain in the top and bottom quartiles than random chance would suggest, although the

[29]As an empirical check on the importance of persistence in risk factor returns for persistence in fund performance, we sorted the funds into quartiles based on their beta-estimates and examined their subsequent abnormal returns. There was no systematic difference in the performance of low- and high-beta funds.

transition probabilities alone do not give any indication of the economic magnitude of these effects.[30]

With one exception, *every* diagonal element of the estimated transition matrices for each asset class exceeds 25%, albeit often by numerically and statistically unimpressive amounts. The estimated probability that a fund that is currently in the top or bottom quartile will continue to be in the same quartile next year is above 29 per cent for UK equities, international equities and bonds, and cash/other investments. Much lower persistence is observed for UK bonds, while there is some persistence for the top, but not the bottom, quartile in the case of UK index-linked bonds and UK property. However, the short sample precludes precise estimation of these transition probabilities and, hence, only a few of the estimates (marked by asterisks) proved to be significantly different from 25 per cent at the 5% level, according to standard errors based on the time series of the individual transition probabilities. For the overall portfolio, the transition probabilities for the top and bottom quartiles lie around 29 per cent.[31]

Persistence might be easier to detect over longer holding periods than a year as in Lakonishok, Shleifer, and Vishny (1992), who found some evidence of persistence at horizons of 2 to 3 years but not at the one-year horizon. For our sample, the transition probability estimates fell dramatically when based on two non-overlapping three-year holding periods. In the case of UK equities, the estimates of P_{11} and P_{44} fell to 25 per cent and 27 per cent, respectively (both figures being insignificantly different from 25%). The apparent peak in persistence at the one-year horizon parallels the results of Hendricks, Patel, and Zeckhauser (1993). In contrast, Lehmann and Modest (1987) and Grinblatt and Titman (1992) find evidence of persistence

[30]In view of the small number of funds that hold international property, this asset class was left out as a separate category in the estimated transition probabilities reported in Table 7.

[31]The estimated transition probabilities based on the other benchmark-adjusted returns were very similar to the ones reported here for raw returns.

in mutual fund performance across five-year intervals, a test we cannot undertake with any precision due to our relatively short sample.

We also tested for persistence in performance using a variant of the approach employed by Hendricks, Patel, and Zeckhauser (1993). For December of each year, we calculated the abnormal performance measures (3) and (4), using only data from the beginning of the sample up to December of the year in question. We then sorted the funds into four equal-weighted portfolios based on the rank of their abnormal performance measure over the most recent 12-month period. The performance of the portfolios over the subsequent year was noted and the procedure repeated after 12 months. Since we need one year of data to get an initial estimate of abnormal performance, this procedure leaves us with returns on four portfolios measured over 96 months.[32]

Panels A and B of Table 8 reveal that UK equities and cash/other investments are the only asset classes for which there is any strong evidence of persistence in performance and then only in respect of peer-group comparisons. For the multi-index benchmark case, the individual alphas from the quartile-sorted Jensen regressions are insignificant at conventional levels for all asset classes, although the difference between the annualized alphas of the highest and lowest prior performance portfolios are 126 and 62 basis points for the UK equity and cash/other investments portfolios, respectively. These regularities are also reflected in the peer-group benchmark-adjusted returns, where the annualized average raw return differentials are 146 and 292 basis points for these two asset classes, respectively. In both cases, the sample means are also ordered from largest to smallest across the four quartiles. In contrast, there is no indication of persistence of either above or below normal performance for the remaining asset categories or for the total fund irrespective of the benchmark used. To be sure, many of

[32] On average each of the four portfolios contained around 80 funds for UK equities, international equities, cash/other investments and total holdings and somewhat fewer funds for the other asset classes.

the other coefficients in the table are larger but the volatility of the quartile-sorted portfolio returns is so large as to render those estimates insignificant at conventional levels.[33]

We provide an alternative characterization of the persistence of abnormal performance in panel C of Table 8. We formed zero net investment portfolios each December by taking a long position in those funds that had positive alphas over the preceding year and a short position in those that had negative alphas and tracked the performance of these constructed portfolios over the subsequent twelve months in a manner similar to Brown, Goetzmann, Ibbotson, and Ross (1992) and Hendricks, Patel, and Zeckhauser (1993). The results remain consistent with the hypothesis that there is measured persistence in UK equity returns and no measurable persistence for any other asset group or for the total returns of the funds. Once again, the magnitude of the effect in UK equities is modest, of the order of just 0.5 of a per cent annualized.

Of course, fund size reflects in part cumulative past performance while our prior return measure reflects recent performance, making it somewhat difficult to disentangle the two effects. That these measured effects are related shows up in portfolio composition: only 15 per cent of the quartile containing the smallest funds were also in the quartile of worst-performing funds, whereas 32 per cent of the largest funds were contained in this quartile. Evidence such as this makes it hard to tell if size is the anchor of current performance or the result of prior performance.

To disentangle the two effects, we ran single-index Jensen regressions for each UK equity portfolio with the portfolio's own (size-adjusted or prior-performance-adjusted) quartile return included as an additional regressor. This procedure can be justified on the hypothesis that the single index regressions omitted some important risk factor and that the betas on the size- or prior- performance- quartile portfolios are constant.[34] The results covering the period

[33]These results are not sensitive to the number of categories used in the sorting procedure. We also experimented with eight groups and obtained very similar results to those reported in the case of four groups.

[34]We prefer to view this as more of an exercise in data description than as an exploration of alternative

1987-1994 (96 months) are presented in Table 9. One year of data is lost due to the initial sort. The 5%-95% range for the alpha estimates based on the standard benchmark regressions is 400 basis points from −1.86 to 2.11 per cent. When the funds' size-sorted-quartile portfolio returns were included in the regression, the range fell by a large amount to 319 basis points, from -1.54 to 1.65 per cent. However, when the benchmark and the excess returns on the prior-performance-sorted portfolios are included as regressors, the range fell by a much smaller amount to 374 basis points, from -1.82 to 1.92 percent. This suggests that fund size accounts for a non-trivial proportion of the cross-sectional variation in abnormal performance, while prior performance does not.

5.6 Spillover Effects in Performance

As noted earlier, it is common practice in the UK for performance in each asset category to be compared with the associated peer-group return despite the fact that around 60 per cent of UK pension funds are managed on a balanced mandate, suggesting that most sponsors should be more interested in the 'bottom line' total return on their fund than how the total return is achieved. Nevertheless, such comparisons enable sponsors to identify the relative strengths and weaknesses of their fund manager in different activities. For example, the fund manager might be good at UK stock selection but poor at picking bonds. Increasingly sponsors, particularly those of the larger funds, are using evidence on persistent strengths and weaknesses to alter their fund management arrangements. For instance, in the example above, the sponsor might retain their existing fund manager to pick equities, but choose a different bond manager.[35]

risk/return relations.

[35]In interviews with industry practitioners, we discovered that fund managers are beginning to specialize in different areas, the main divisions being between security selection and market timing and between different geographical regions. The main reason for this appears to be more thorough evaluation of balanced managers by plan sponsors. For example, the trustees and their advisers might not find the presentation of a balanced manager entirely convincing, perhaps becoming concerned that the lack of a coherent investment philosophy

Most of the fund managers in our sample are balanced fund managers and, hence, sell themselves as being good at all aspects of the business. If they are, there should be spillover effects between different asset categories—that is, good performance in one asset category would be associated with good performance in other asset categories. Moreover, the grouping of categories for which there were positive spillovers would also be revealing - strong spillover effects across asset classes would validate the claim that the fund manager was offering a genuinely value-increasing balanced service, while spillover effects between similar asset classes might suggest specialist ability in a particular type of asset as opposed to special expertise across asset classes. For example, a value-oriented manager might be good at identifying undervalued stocks in the same industries in different markets (which would show up as spillovers between domestic and international equities) but not across the other asset classes. In contrast, signs of spillover effects in funds managed on the basis of specialist mandates in each asset category would provide evidence of the sponsor's ability in selecting good managers in different asset categories.[36]

We undertook the following simple nonparametric analysis to measure the relationship between the relative performance of a fund in different asset categories. For each asset category and for each year, we ranked the fund managers on the basis of their relative returns. The process was repeated for all years in the sample and the fund managers were again ranked according to their average rank throughout the sample. We adopted this procedure (which emphasizes the short-term correlation between the returns of various managers) rather than one that ranks funds on the basis of mean returns over the entire sample because of the

in a particular specialty might reflect the absence of a strong team in that field. When pressed, such fund managers might admit this comparative weakness and the plan sponsor might solicit mandates from other managers in the specialty area in question.

[36]The issue of multi-managed funds and the coordination problems involved was first raised by Rudd and Clasing (1982, Chapter 6).

evidence in the literature that 'hot hands' appears, if anything, to be a short-term phenomenon (c.f., Hendricks, Patel, and Zeckhauser (1993)). Using this average of the rankings over the sample, we next computed Spearman coefficients measuring the correlation between any pairs of vectors of rankings of the fund managers. These statistics, which are asymptotically normally distributed, are reported in Table 10.

Overall the evidence on positive spillover effects between asset groups is, at best, very weak. The sole exception is the consistent, if modest, correlation between domestic and international equity returns, however benchmark-adjusted. That is, there is some evidence that funds which earned high (low) mean returns in UK equities in a given year also tended to earn high (low) mean returns in international equities during the same year. One possible explanation is that this indicates an industry bias in the equity selection procedures of our group of fund managers.

6 The Performance of Pension Fund Asset Allocations

The performance of UK pension funds within asset classes is interesting in its own right, both because of the implications for the industrial organization of the UK pension fund industry and because of the possibility of direct comparisons with US pension and mutual funds. The preceding analysis, however, disregards a unique feature of our data: the information on the asset allocations of UK pension funds over time. This section corrects for this by conducting a systematic analysis of the performance of the short-term asset allocations of UK pension funds.

We proceed in two steps. We noted earlier that the funds tilted their asset allocation towards equities and away from domestic bonds over the sample period and it is difficult to determine whether this reflected a change in desired *ex ante* risk exposure - that is, a change in the strategic asset allocation - or the reward for a market timing bet - that is, a tactical asset allocation - that turned out well *ex post*. We need a better understanding of the asset allocation dynamics in order to identify any market timing or security selection ability among our sample of managers. Accordingly, the next subsection studies various aspects of aggregate portfolio dynamics and the concomitant cross-sectional variation in asset allocation across funds. Armed with the information we extract from this exercise, the subsequent subsection provides a variety of decompositions of the market timing and security selection skills of fund managers along the lines of Brinson, Hood, and Beebower (1986).

6.1 The Evolution of Portfolio Weights

To help identify the factors causing portfolio weights to change over time, we employ a simple decomposition of the changes in these weights into cash-flow-related and returns-related components. Asset classes that enjoyed large positive relative returns would also experience an increase in their weights in the total portfolio, unless fund managers deliberately rebalance portfolios as this occurs.

We first apply this decomposition to the aggregate portfolio. Let $W_{j,t}$ be the total holding in asset class j at the end of month t, computed across all the pension funds in the sample, and let W_t be the total holding across all asset classes. These weights must satisfy the accounting identity:

$$W_{j,t} = W_{j,t-1}(1 + r_{j,t} + NCF_{j,t}),$$

where $r_{j,t}$ is the rate of return on UK pension funds' holdings of asset class j during month t and $NCF_{j,t}$ is the rate of net cash flows into asset class j during month t. Using this relation, the portfolio weight of asset class j can be written as:

$$\omega_{j,t} = \frac{W_{j,t}}{W_t} = \frac{\frac{W_{j,t-1}}{W_{t-1}}\left(\frac{W_{j,t}}{W_{j,t-1}}\right)}{W_t/W_{t-1}} = \omega_{j,t-1}\frac{1 + r_{j,t} + NCF_{j,t}}{1 + \sum_{k=1}^{M}\omega_{k,t}(r_{k,t} + NCF_{k,t})}. \tag{5}$$

Taking log-differences, it follows that:

$$\Delta \log(\omega_{j,t}) = \log(1 + r_{j,t} + NCF_{j,t}) - \log(1 + \sum_{k=1}^{M} \omega_{k,t}(r_{k,t} + NCF_{k,t})),$$

so that, to a close approximation:

$$\Delta \log(\omega_{j,t}) \approx r_{j,t} - r_{p,t} + NCF_{j,t} - NCF_{p,t}, \tag{6}$$

where $r_{p,t}$ is the value-weighted total return on the portfolio during month t, while $NCF_{p,t}$ is the value-weighted average rate of net cash flow for the total portfolio during month t.

This decomposition is particularly informative since it allows us to measure the extent to which changes in portfolio weights are caused by differential returns across asset classes, as indicated by the component $r_{j,t}$-$r_{p,t}$, or by shifts in the net cash flows across asset classes, as indicated by $NCF_{j,t}$-$NCF_{p,t}$. Changes in the portfolio weights due to the first component arise from the passive investment strategy of 'buy-and-hold', reinvesting asset income in the same asset categories, and distributing any net inflows into the pension fund according to the *ex post* asset-class weights. In contrast, the changes due to the second component result from the more active strategy of rebalancing the portfolio by redirecting cash flows across asset groups, although we recognize that rebalancing toward the long-run or strategic asset allocation would generally be viewed as part of a passive, not active, investment strategy. Insofar as equities generated higher mean returns during the sample than the other asset groups, this might account for at least a fraction of the increase in the weight allocated to equities from 1986 to 1994.

As a simple measure of how much of the short-term changes in the aggregate value-weighted portfolio weights of the various asset classes can be explained by the two components, we calculated the R^2 from a least-squares regression of $\Delta \log(\omega_{j,t})$ on a constant and $r_{j,t}$-$r_{p,t}$. The outcome of this exercise is reported in panel A of Table 11. The results suggest that return differentials: (1) largely account for the monthly variation in the weights allocated to UK and international equities and UK property; (2) account for much of the monthly variation (of the order of 70 per cent) in the weights allocated to conventional and index-linked UK bonds and international property; and (3) seem to explain a much smaller proportion of the monthly variation in the allocations to international bonds and cash/other investments.

It is also interesting to decompose the overall percentage change in portfolio weights over the sample period into return and net cash flow components using (6). For this purpose, we report in panel B of Table 11 the sample means of $\Delta\log(\omega_{j,t})$ and its two components, $r_{j,t}$ - $r_{p,t}$ and $NCF_{j,t}$ - $NCF_{p,t}$. The only asset class for which differential returns contributed positively to its asset allocation was UK equities. This occurs because it is the only asset class whose mean return exceeded the mean return on the total portfolio during the sample period. Thus, by definition, any increase in the portfolio weights of the remaining asset classes must have been due to positive net cash flows for those assets.[37]

These observations concern only the dynamics of the aggregate portfolio weights. We are also interested in the cross-sectional aspects of asset allocation dynamics, both for their implications for performance measurement and for our understanding of pension fund behavior. Accordingly, consider the fund-specific version of (6):

$$\Delta\log(\omega_{i,j,t}) \approx r_{i,j,t} - r_{i,p,t} + NCF_{i,j,t} - NCF_{i,p,t} \tag{7}$$

[37]Industry practitioners confirm that there is no automatic rebalancing of portfolio weights, especially by the majority of balanced managers: portfolio weights are generally allowed to drift in the direction of return differentials. However, rebalancing is much more common with specialist managers or with the small (but growing) proportion of balanced managers using ALM. In such cases, funds either employ specialist asset allocation managers to switch between asset classes as part of an active management strategy or pursue a passive rebalancing strategy, either every quarter or every year. With both approaches, managers might use a futures overlay in the short run to achieve the desired rebalancing but will physically rebalance across asset categories in the long run.

where i indexes pension funds. Subtracting equation (6) from (7) yields:

$$\begin{aligned} \Delta \log(\omega_{i,j,t}) - \Delta \log(\omega_{j,t}) \;\approx\; & [(r_{i,j,t} - r_{i,p,t}) - (r_{j,t} - r_{p,t})] + \\ & [(NCF_{i,j,t} - NCF_{i,p,t}) - (NCF_{j,t} - NCF_{p,t})] \end{aligned} \tag{8}$$

Note that (8) is in the form of a standard fixed-effects dummy-variable model. That is, $\Delta \log(\omega_{j,t})$ is a time effect and the composite residual on the RHS of (8) is a fund specific effect with nonzero mean. However, such a model typically postulates that the time- and fund-specific effects are uncorrelated both with each other and cross-sectionally, whereas such an absence of correlation need not be a property of the measured components in these data.[38] Nevertheless, we consider this model to be a useful baseline and can envisage models in which relative performance evaluation leads managers to follow strategies that make this a natural decomposition.

Panel A of Table 12 provides information on the extent to which the individual fund portfolio weights conform to the fixed-effects model. We report the cross-sectional distribution of the variance ratio:

$$\begin{aligned} & [Var\{\Delta \log(\omega_{j,t})\} + Var\{(r_{i,j,t} - r_{i,p,t}) - (r_{j,t} - r_{p,t}) \\ & +(NCF_{i,j,t} \; - NCF_{i,p,t}) - (NCF_{j,t} - NCF_{p,t})\}]/Var\{\Delta \log(\omega_{i,j,t})\} \end{aligned} \tag{9}$$

which should be unity if the data satisfy the correlation structure of the fixed-effects model. The model clearly fits well on average: the median variance ratio is numerically and economically close to unity for all asset classes. Similarly, the changes in most of the funds' asset

[38]This formulation also differs from the standard model in that the time effect is a value-weighted average of the individual asset class weights as opposed to the usual least squares or weighted least squares estimator of the intercept in a regression based on (8).

allocations relative to the value-weighted average appear to have only modest and typically negative correlations with the aggregate allocation in its asset class. For example, the variance ratios for the 5th percentile of funds—that is, those funds with nearly the smallest positive correlations—lie between 0.85 and 0.97 and the corresponding ratios for the 25th percentile lie between 0.95 and unity. There is a bit more spread in the variance ratios associated with negative correlations between $\Delta\log(\omega_{i,j,t})$ and $\Delta\log(\omega_{j,t})$ with ratios in the ranges of 1.05 to 1.41 and 1.14 to 1.81 at the 75th and 95th percentiles, respectively. Hence, changes in the asset allocations of most funds relative to the corresponding changes in aggregate allocations are largely, but not entirely, idiosyncratic to the individual funds.[39]

Panel B of Table 12 reports the fractiles of the changes in the funds' portfolio weights (measured in percentage points per year) in excess of the corresponding aggregate change. More than 140 basis points (and more than 200 basis points for the more important asset classes) separated the funds in the 5th and 95th percentiles for all asset classes except for international property, which had a much tighter spread of 42 basis points. Note also that this range of variation is generally large relative to the average rates of change in the asset allocations themselves: of the order of 51 and 76 basis points for UK and international equities, respectively, -106 basis points for UK bonds, 35 basis points for UK property and between -10 and 16 basis points for the remaining asset classes. In particular, we note that the substantial overall drift towards equities over the period itself conceals a wide range of drift rates across the individual funds.

It is also interesting to examine the typical behavior of the two components of these drifts:

[39]We also examined the coefficient from the regression of $\Delta\log(\omega_{i,j,t}) - \Delta\log(\omega_{j,t})$ on $\Delta\log(\omega_{j,t})$ which should be zero in the same circumstance. We chose to report the variance ratio because the dummy variable model is a variance decomposition. Since both measures reflect the same correlations, it is unsurprising that they produced similar results. For example, the number of regression coefficients significant at the five per cent level ranged from 20% to a bit more than 40%.

the average excess net cash flow into a particular asset class (i.e., the mean over i of [($NCF_{i,j,t}$-$NCF_{i,p,t}$) - ($NCF_{j,t}$-$NCF_{p,t}$)]) and the corresponding average relative return for the asset class (i.e, the mean over i of [($r_{i,j,t}$-$r_{i,p,t}$)-($r_{j,t}$-$r_{p,t}$)]). Panel C of Table 12 sheds some light on both how much and when funds rebalance toward or away from asset classes that have experienced good or bad performance relative to the aggregate peer-group benchmark. In aggregate, funds increased the asset allocation towards asset classes that performed relatively well over the sample. However, the cross-sectional correlation between the average excess net cash flows and the corresponding average excess asset class returns is negative for all asset classes (except index linked bonds) with correlations between -0.20 and -0.45. Thus the funds with the relatively highest returns within a given asset class were also the ones with the smallest net cash flows into that asset class.

Panel C also reveals that this average behavior does not show up as substantial rebalancing year-by-year by reporting the cross-sectional distribution of the sample time series correlations between [($NCF_{i,j,t}$-$NCF_{i,p,t}$)-($NCF_{j,t}$-$NCF_{p,t}$)] and [($r_{i,j,t}$-$r_{i,p,t}$)-($r_{j,t}$-$r_{p,t}$)]. The median time series correlation is numerically and economically close to zero and the 5th percentile time series correlation (that is, the funds with the correlations smaller than 95% of the fund universe) is closer to zero than the corresponding cross-sectional correlation for each asset class for all asset classes except international equities and index-linked bonds. The substantial average cross-sectional correlation coupled with the weak correlations in the year-on-year figures suggest that funds did exhibit a tendency to rebalance towards their *strategic* asset allocations.

These statistics measure the average behavior in the drifts of the asset allocations of individual funds but reveal little about their short run dynamics and, in particular, any mean reversion tendencies that they may exhibit. Any such mean reversion would have to be quite pronounced to be reliably estimated in a short sample such as ours. Panel D of Table 12 provides two indicators of the nature of any such mean reversion. First, we present Markov chain estimates for the probability of individual fund asset allocations remaining above or below the industry average weight each year: these range from 67% to 95% for all asset classes, implying probabilities of between one twentieth to one third of crossing over the average. The time series standard errors of the sample transition probabilities are sufficiently small that it is clear that the corresponding population probabilities are far from the null value of 50%, both economically and statistically. Second, Panel D provides similar information for the sample probabilities for the transitions from initial to final relative weights (but without standard errors since there is only one time series data point per fund). The point estimates are consistent with slow mean reversion as well with stayer probabilities between 47% and 79%. Taken together, the Markov chain evidence suggests that any mean reversion tendencies in the relative portfolio weights are quite slow.

Panel E of Table 12 provides supporting evidence for the proposition that mean reversion in portfolio weights toward the aggregate weight is quite slow. It presents the results from a regression of $\omega_{i,j,t}$-$\omega_{j,t}$ on a constant, its lagged value, and the change in its lagged value—that is, a standard Dickey-Fuller regression. The distribution of the coefficients on $\omega_{i,j,t-1}$-$\omega_{j,t-1}$ above the 50th percentile ranges from 0.90 to unity for all asset classes except cash which has a median coefficient of 0.78. Similarly, the t-statistics have rejection rates for the null hypothesis at the 5% significance level of the order of 5%, except for domestic and international equities with rejection rates at 14% and 11%, respectively, and cash with one of 30%.

From a modeling perspective, these observations leave us with a conundrum. Clearly, the overall asset allocation of these funds was in flux over the sample period. This fact makes it difficult to distinguish between short-term attempts to profit from supposed superior information and any long-run shifts in desired risk exposure as might have arisen from the elimination of excess overfunding as required by the 1986 Finance Act and the increasing indexation of deferred liabilities as required by (in particular) the 1985 Social Security Act. Nevertheless, the evidence we have gathered appears to indicate slow mean reversion towards a changing strategic asset allocation.

6.2 Decomposing Overall Returns into Returns from Stock Selection and Market Timing

We use a simple decomposition proposed by Brinson, Hood, and Beebower (1986) to separate portfolio returns into components due to active and passive management. Suppose there are M asset classes and let $\omega_{n,j,t}$ be the 'normal' or strategic asset allocation of a fund in the j'th asset class at time t, $\omega_{a,j,t}$ be the actual portfolio weight, $r_{n,j,t}$ the 'normal' portfolio return, and $r_{a,j,t}$ the actual portfolio return. Then, as an arithmetic identity:

$$\begin{aligned}\sum_{j=1}^{M}\omega_{a,j,t}r_{a,j,t} \quad = \quad & \sum_{j=1}^{M}\omega_{n,j,t}r_{n,j,t}+\sum_{j=1}^{M}\omega_{n,j,t}(r_{a,j,t}-r_{n,j,t})+ \\ & \sum_{j=1}^{M}(\omega_{a,j,t}-\omega_{n,j,t})r_{n,j,t}+\sum_{j=1}^{M}(\omega_{a,j,t}-\omega_{n,j,t})(r_{a,j,t}-r_{n,j,t,})\end{aligned} \tag{10}$$

or Total Return = Normal Return + Return from Stock Selection + Return from Market Timing + Residual Return. For this to be a useful decomposition of asset class returns, the residual term should be small compared with the other components, since it represents the component of returns that is not attributable to either timing or selectivity, and we also need good measures of 'normal' portfolio returns and weights. In fact, the average residual for our sample proves to be small relative to the normal asset allocation component but is of the same order as the selectivity component, reducing the attractiveness of the decomposition. Natural measures of normal portfolio returns are the various external or peer-group benchmark indices.

However, the choice for the normal portfolio weights is more problematic. In the absence of any information on the funds' asset-liability modeling exercises which might enable us to draw inferences about appropriate strategic asset allocations, we were reduced to experimenting with a few simple models with an eye towards spanning the range of empirically relevant ones. The main alternatives are a straightforward static model and one in which the weights of each fund trend over the sample. We also explore a few strategic asset allocation models that allow for different short run dynamics.

The first model, proposed by Brinson, Hood, and Beebower (1986), takes the average portfolio allocation over the sample as the normal portfolio weights: $\omega_{n,j,t} = \sum_{t=1}^{T} \omega_{a,j,t}/T$ for all t. This definition seems reasonable if the funds are in a steady state in the sense that they have achieved their target portfolio composition across major asset groups and that long-run risk/return tradeoffs are stationary. It implicitly assumes that the mean values of the portfolio weights do not change over time, implying that they are effectively drawn from a stationary distribution. This is an unattractive assumption in our case, since UK pension funds were not in a state of equilibrium over the sample period. Nevertheless, it provides a useful benchmark, and any similarity between the decompositions generated under this palpably false model and those produced using more dynamic models will indicate a robustness in the decomposition given in (10).

The systematic increase in pension fund equity exposure over the period is the most obvious nonstationarity. A particularly simple way of accounting for nonstationary portfolio weights is to include a trend in these weights, letting the normal portfolio weights increase (or decrease) linearly in time between the initial and terminal weights. Hence, our second measure of the 'normal' portfolio weights is: $\omega_{n,j,t} = \omega_{a,j,1} + (t/T)(\omega_{a,j,T} - \omega_{a,j,1})$. Since $\sum_{j=1}^{M} \omega_{a,j,T} - \sum_{j=1}^{M} \omega_{a,j,1} = 0$, this measure has the important property that the normal portfolio weights sum to unity at each point in time, t, and are confined to lie in the interval [0,1]. Benchmark portfolio weights increase (or decrease) linearly in time between the initial and terminal weights.

In both cases, however, we are using sample-dependent estimates of the normal portfolio weights, thereby inducing potential biases in this otherwise simple decomposition. For example, a fund's asset allocation manager, knowing that a particular asset class manager has good selection skills, might increase the allocation to that manager, thus attributing some of this postulated selection ability to the strategic asset allocation. Similarly, a good market timer need not confront an equal number of positive and negative signals in the sample, biasing the measured long-term asset allocation in the direction of the more frequently observed signal. In both cases, these biases affect the magnitude but not the sign of the timing and selectivity components. These effects are reversed when the asset allocation manager believes that the portfolio managers possess market timing or security selection ability when they, in fact, have no such ability. In this case, the tilt toward managers with the presumed ability is incorrectly classified as part of the strategic asset allocation, while the effect on the measured return to the normal asset allocation depends on whether these managers happended to be lucky or unlucky over the sample period. In particular, funds that tilted toward UK equities based on an erroneous belief that their managers possessed superior performance ability experienced higher measured normal asset allocation returns due to the good performance of UK equities over the period.

Both normal weight models are deterministic save for the randomness induced by averaging over the sample or using the sample endpoints in the calculations. Our remaining choices reflect as much a search for robustness as an exploration of economically or statistically relevant alternatives. The first follows from the observation that the change in actual portfolio weights arising from net cash inflows or outflows could represent either passive or active management (or some convex combination of the two), depending on whether one viewed these flows as passive rebalancing or as active reallocation of assets. Multiplying ($\omega_{i,j,t}$-$\omega_{i,p,t}$), the change in portfolio weight due to either active or passive management, by $r_{j,t}$, we get an alternative measure of the returns to active and passive fund management. In addition, we are concerned that our attempts to decompose performance in this way might actually obscure genuine ability in some fashion. Accordingly, we also set the strategic asset allocation equal to the allocation prevailing six months earlier, treating any correlation between subsequent portfolio weights and returns as due to abnormal performance.[40]

Table 13 summarizes the evidence produced by these different models for normal portfolio weights. We also present several fractiles of the cross-sectional distribution of the average returns to the normal, market timing, and security selection components of performance for each asset class as well as the maximum and minimum values and their associated Bonferroni p-values.

What is the most noteworthy observation is the robustness of the results across models with very different dynamics and drifts. The constant mean and simple trend models yield normal portfolio returns that are numerically close fractile by fractile, despite both the substantial shift toward equities over the sample period and the considerable cross-sectional variation in the drifts of individual fund asset allocations. Similarly, the fractiles relating to the average market timing and selectivity components agree numerically up to the tens of basis points. We obviously find this consistency reassuring in the absence of a single compelling model for normal portfolio weights.

The cross-sectional variation in the performance measures from these decompositions is also remarkably small (Panel E). The semi-interquartile ranges are of the order of only 25 to 40 basis points for the mean annualized normal and market timing components of portfolio

[40]The choice of six months is fairly arbitrary, although it was chosen because it generated greater abnormal performance than using one to five lags. In a related but different context, Grinblatt and Titman termed this the portfolio change performance measure. However, j indexed securities in their analysis, not asset classes as in our setting. Nevertheless, the underlying intuition is the same—funds with superior market timing ability should, on average, shift out of asset classes prior to negative returns and into them prior to any appreciation and those with superior selection skills should shift into sectors when their picks seem particularly good *ex ante*.

returns and a modest 110 basis points for the security selection component, while the annualized differences between the 5th and 95th percentiles are roughly three times the corresponding semi-interquartile ranges. It is difficult to imagine drawing any conclusion other than this: there is not much to choose between the average performances of the bulk of these funds in these three dimensions of portfolio returns.

The results also indicate something about the abilities of the managers in question. Panels A and B of Table 13 report the decomposition when the normal returns equal the external benchmarks. The average normal return at about 12.35 per cent per year, constitutes by far the largest component of the aggregate annual portfolio return whose mean was 12.03 per cent. The mean annualized return from security selection at -5 basis points is insignificant at conventional levels, while that from market timing at -31 basis points is statistically significant. In addition, better than half of the funds had negative selectivity estimates and more than 80 per cent had negative, albeit economically small, timing estimates.[41] Hence, UK pension funds, when compared with all other participants in the UK financial markets, earned an economically small negative return from active portfolio management on average, although there is some variation in the security selection component. Our finding that active fund management contributes negatively to fund value is similar to that of Brinson, Hood, and Beebower (1986), the main difference being that both active components are negative on average in our sample, while

[41] The coefficients on squared excess benchmark returns from the Treynor and Mazuy (1966) regressions of the previous section provide an alternative measure of the market timing ability of managers within asset classes under plausible assumptions (see Jensen (1972), Admati, Bhattacharya, Pfleiderer, and Ross (1986), Lehmann and Modest (1987), and Grinblatt and Titman (1989)) when positive. The cross-section of these coefficients had a semi-interquartile range of -0.66 to 0.045. As is common in such regressions involving US mutual fund data, there are more negative than positive coefficients with the distribution of both the coefficients and their t statistics skewed to the left, suggesting that we are measuring something other than market timing ability. In any event, the results are incompatible with the presence of nontrivial market timing ability for all but perhaps a few managers.

they find negative returns from market timing and positive returns from security selection on average.

For both definitions of 'normal' weights, we found that 96 per cent of the total variation in monthly portfolio returns could be explained by the normal asset class holdings across funds on average.[42] In fact, normal asset class holdings explained 52 per cent of the variability in portfolio returns for the fund with the *smallest* contribution to return variability from this component.

Panels C and D of Table 13 report the changes to the decomposition when the peer-group indices replace the external benchmarks in the definition of 'normal' returns. The return from security selection, at an economically modest 0.32 per cent per year, is now positive and significant, while the return from market timing remains negative after this change of benchmarks. In this case, the semi-interquartile range of the security selection component ran from -0.26 to 0.88, while that of the market timing component ran from -0.37 to -0.07. These results are consistent with our earlier finding that our sample of fund managers earned a higher mean return on their holdings of UK equities when compared with their peer-group, but earned a lower mean return when compared with the external index for UK equities.

Coggin et al (1993), in their study of stock picking and market timing skills of US pension fund managers, found that US equity pension fund managers had positive and significant stock selection skills and negative timing skills. Our results also differ from those of Grinblatt and Titman (1989), one of the few studies in the performance literature that uses data on portfolio weights. Their analysis assumed that the long-run asset allocation is equal to the initial one, treating all portfolio weight changes save for those that rebalance back to the initial allocation as the returns from active portfolio management. But such rebalancing constitutes active *trading* when there are large return differentials between asset classes as managers rebalance to

[42]This figure is a little higher than the 91.5 per cent reported by Brinson, Singer, and Beebower (1991).

restore the initial asset allocation. Otherwise, asset classes with relatively high mean returns will experience portfolio weight increases as we have seen in the UK case. Comparing the returns on portfolios whose weights are updated every quarter with portfolios with annual or no updating from the initial weights, Grinblatt and Titman (1989) find evidence that active fund management contributes positively to fund value. In contrast, Lakonishok, Shleifer, and Vishny (1992), based on a procedure where actual portfolio returns are compared with returns on a portfolio without active transactions, obtained results similar to ours, namely that active fund management does not contribute positively to fund value. Even using Grinblatt and Titman's portfolio change measure, it is clear from the figures reported in the last column of Panel E in Table 13 that we find little evidence of economically or statistically significant abnormal performance.

In any event, the main point remains the comparative importance of the long run or strategic asset allocation, however measured, and the correspondingly modest gains over a passive investment strategy from even the best performing managers. The bulk of the selectivity measures are both economically and statistically small in absolute value with more negative than positive estimates. Moreover, the vast majority funds have negative market timing measures, irrespective of the method of construction. A randomly selected pension fund would have been better served by applying their asset allocation to passively managed index funds.

7 Conclusion

From the outset, several aspects of the experimental design implicit in our UK pension fund data struck us as critical for understanding performance evaluation in this universe. Chief among these are the legal and economic environment in which the funds operated. In our view, the empirical regularities we observe in these data reflect the industrial organization and regulatory environment of the UK pension fund industry.

UK pension funds were essentially unregulated during our sample period and they ran large surpluses until almost the end of the period. While trustees gave fund managers a great deal of freedom in their choice of asset allocation and investment strategy, both the fee structure and performance evaluation framework in the industry affected the incentives of the fund managers to attempt to outperform the market. UK pension fund managers were set the objective of adding value, in most cases on the basis of a multi-asset balanced mandate, but their fee was related to the year-end value of the funds they manage, rather than explicitly to the value-added they achieved. Moreover, their long-term success in the fund management industry depends on their relative performance against their peer-group.

These twin objectives provide fund managers with a mixed set of incentives. On the one hand, genuine *ex ante* ability that translates into superior *ex post* performance increases the assets under management and, thus, the base on which the management fee is calculated. On the other hand, this incentive is not particularly strong and active management subjects the manager to nontrivial risks. The incentive is weak because the prospective increase in the fee is second order, being the product of the *ex post* return from active management and the management fee and thus around two full orders of magnitude smaller than the base fee itself. Moreover, the *ex post* return from active management of a truly superior fund manager will often be negative and occasionally large as well, resulting in poor performance relative to managers who did not attempt to profit from active management irrespective of their ability. The probability of relative underperformance large enough to lose the mandate is likely to be at least an order of magnitude larger than the management fee. Hence, on balance, it would appear that the risk of underperforming due to poor luck would outweigh the prospective benefits from active management for all but the most certain security selection or market timing opportunities.

These observations about underlying incentives appear to account for many of the robust regularities we report. We found surprisingly little cross-sectional variation in *ex post* average performance across pension fund portfolios as a whole as well as within asset classes either from Jensen-style regressions, which ignore the information on asset allocations, or from decompositions of returns into components due to normal or strategic asset allocation, market timing, security selection, and a residual. While there are robust differences in average performance in the Jensen-style regressions from sources such as risk adjustment and the use of alternative peer group and external benchmarks, the same kinds of modifications left the cross-sectional variation in measured abnormal performance virtually identical to that of the raw return data. Similarly, normal or strategic asset allocations, however modeled, account for the bulk of the cross-sectional variation in overall fund returns and exhibit remarkably little cross-sectional variation. To be sure, we found some evidence compatible with the hypothesis that some funds experienced either persistently good or persistently bad performance and some evidence of a nontrivial cross-sectional variation in the measured returns attributable to security selection. Nevertheless, these regularities are second order compared with the central observations on cross-sectional variation in measured abnormal performance and the central role of the strategic asset allocation.

Fund size is the one variable that apparently accounts for an important fraction of cross-sectional variation in measured performance in equity returns. This regularity makes it that much easier to understand the role of relative performance evaluation in the UK pension fund industry. Due to the considerable cross-sectional variation in fund size, it happens that two thirds of the funds outperformed the peer-group benchmark in UK equities. Not surprisingly, large funds were overrepresented among the relative underperformers.

Hence, the vast majority of managers could point to either their good performance or to a size handicap as a way of justifying the retention of their mandates. Since fund managers had reasonably strong incentives to produce similar, though not identical, results, the noise in asset returns suggests that funds also had a fair chance of producing performance good enough to warrant being offered new mandates as well. The industrial organization of the UK pension fund industry appears to be a clear case of economic actors following their incentives.

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Table 1. Aggregate Portfolio Weights for UK and US Pension Funds, WM Sample and Universe, 1986 - 1994.

	1986	1987	1988	1989	1990	1991	1992	1993	1994	1986	1990	1994
				WM Sample						US Pension Funds		
Domestic Equities	50.4	52.9	54.0	53.2	53.6	54.9	55.6	54.8	53.6	45.6	42.1	44.8
International Equities	19.6	15.7	15.8	19.6	16.8	20.4	21.4	23.8	22.5	2.6	4.5	8.3
Domestic Bonds	12.2	12.0	10.5	7.5	6.5	5.2	4.7	4.6	5.3	37.8	38.9	34.2
International Bonds	0.9	0.8	0.8	1.3	2.5	3.5	3.9	3.3	2.8	0.0	0.0	2.0
Index Bonds	3.1	2.9	2.8	2.3	2.4	2.0	2.2	2.8	3.6	NA	NA	NA
Cash/Other Investments	3.6	5.2	4.9	5.1	6.3	4.0	3.8	3.7	4.2	7.8	9.8	7.5
Domestic Property	9.2	9.1	10.0	9.8	10.9	9.1	7.6	6.5	7.6	6.2	4.7	3.2
International Property	1.0	1.4	1.2	1.2	1.0	0.9	0.8	0.5	0.4	NA	NA	NA
				WM Universe								
Domestic Equities	51	54	53	53	54	56	56	56	54	NA	NA	NA
International Equities:	20	14	16	21	18	21	22	24	22	NA	NA	NA
North America	NA	NA	NA	7	5	6	6	5	4	NA	NA	NA
Continental Europe	NA	NA	NA	7	7	8	8	9	8	NA	NA	NA
Japan	NA	NA	NA	5	3	4	4	4	5	NA	NA	NA
Total Pacific (ex. Japan)	NA	NA	NA	1	2	2	3	5	4	NA	NA	NA
Others	NA	NA	NA	1	1	1	1	1	1	NA	NA	NA
Domestic Bonds	13	13	10	6	6	5	4	4	6	NA	NA	NA
International Bonds	0	1	1	2	3	4	4	4	4	NA	NA	NA
Domestic Index-Linked Bonds	3	3	3	2	3	2	3	3	4	NA	NA	NA
Cash/Other Investments	4	5	6	6	7	4	4	4	4	NA	NA	NA
Domestic Property	8	9	10	9	8	7	6	5	6	NA	NA	NA
International Property	1	1	1	1	1	1	1	0	0	NA	NA	NA

Note: For each year the table shows the percentage of the total portfolio wealth invested in the assets held by UK pension funds. The WM sample figures are based on the UK pension funds included in our sample while the WM universe includes all fund managers tracked by WM. All figures are based on asset values measured at the end of the year. The figures for the US pension funds were supplied by Greenwich Associates.

Table 2. Fractiles of Total Returns (Annualized percentages) by Asset Class.

	UK Equities	Intl. Equities	UK Bonds	Intl. Bonds	UK Index Bonds	Cash/Other Investments	UK Property	Total
Minimum	8.59	4.42	6.59	-0.64	5.59	2.67	3.05	7.22
5%	11.43	8.59	9.44	2.18	7.20	5.46	5.07	10.60
10%	11.85	9.03	9.95	7.56	7.81	7.60	6.58	10.96
25%	12.44	9.64	10.43	8.30	7.91	8.97	8.03	11.47
50%	13.13	10.65	10.79	11.37	8.22	10.25	8.75	12.06
75%	13.93	11.76	11.22	13.37	8.45	11.72	9.99	12.59
90%	14.81	12.52	11.70	14.55	8.80	14.20	10.84	13.13
95%	15.46	13.14	12.05	18.15	8.89	16.13	11.36	13.39
Maximum	17.39	14.68	17.23	26.34	10.07	19.73	13.53	15.03

Note: The table shows the fractiles of the cross-sectional distribution of returns on individual asset classes as well as on the total portfolio.

Table 3. Standard Deviations of and Correlations Between Monthly Returns on Assets Held by UK Pension Funds, 1986-1994.

	UK Equities	Intl. Equities	UK Bonds	Intl. Bonds	UK Index Bonds	Cash/Other Investments	UK Property	Total
UK Equities	5.41	0.769	0.382	0.200	0.347	0.122	-0.053	0.986
Intl. Equities		4.81	0.187	0.425	0.217	0.223	-0.047	0.853
UK Bonds			2.27	0.305	0.620	0.172	-0.052	0.398
Intl. Bonds				1.73	0.357	0.115	-0.178	0.282
UK Index Bonds					2.32	-0.024	0.059	0.377
Cash/Other Inv.						0.65	0.042	0.169
UK Property							1.48	0.185
Total								3.80

Note: For each month in the sample, value weighted portfolios of the individual assets held by UK pension funds were formed. The diagonal of the table reports the monthly standard deviations of these returns (as percentages) while the off-diagonal cells present the time-series correlations between the returns (as proportions).

Table 4. UK Pension Fund Investment Performance with Respect to Different Benchmarks, 1986-1994 (Annual Percentages).

Year	External Index	WM2000 return	Value wgt. return	Eq. wgt. return	% of out-performers	External Index	WM2000 return	Value wgt. return	Eq. wgt. return	% of out-performers	External Index	WM2000 return	Value wgt. return	Eq. wgt. return	% of out-performers
	UK Equities					International Equities					UK Bonds				
1986	24.24	22.76	23.35	23.64	36.4	34.22	31.29	31.56	30.98	24.0	10.85	11.39	11.35	11.22	63.3
1987	7.65	7.07	6.30	8.08	48.5	-9.12	-22.60	-20.16	-21.78	0.6	14.17	14.91	15.63	14.96	76.0
1988	10.90	9.60	10.48	9.76	31.4	27.03	19.92	21.65	19.16	2.8	6.52	7.94	8.35	7.81	80.5
1989	30.80	30.36	30.83	30.59	47.5	27.38	34.19	33.67	34.50	97.5	7.85	7.16	6.73	7.61	35.9
1990	-10.20	-10.43	-10.20	-9.58	58.2	-39.77	-31.35	-32.27	-31.39	98.0	9.15	6.31	7.53	8.85	28.1
1991	18.84	18.06	18.22	18.14	40.4	21.30	18.26	19.41	18.00	12.8	14.95	16.46	16.78	18.19	84.1
1992	18.59	19.13	18.71	19.36	66.7	15.51	18.02	17.88	17.46	79.2	17.05	17.35	17.08	17.13	63.5
1993	24.90	24.23	24.78	24.97	48.7	22.55	34.21	32.91	34.36	97.5	19.01	22.92	22.14	22.15	76.7
1994	-6.01	-5.83	-5.78	-6.08	48.2	0.85	-4.57	-3.60	-4.96	2.2	-6.43	-9.64	-8.74	-9.37	14.1
1986-1994	13.30	12.77	12.97	13.21	44.8	11.11	10.82	11.23	10.70	39.8	10.35	10.53	10.76	10.95	77.3
	International Bonds					UK Index-Linked Bonds					Cash/Other Investments				
1986	16.92	26.18	18.95	24.85	69.8	6.60	5.06	4.97	5.46	11.2	10.71	12.13	8.75	12.58	60.5
1987	-12.24	-3.09	-1.01	-5.60	74.5	6.37	4.94	5.97	5.74	43.0	9.63	9.02	8.06	11.12	48.8
1988	9.06	7.27	3.01	10.11	37.0	11.37	12.37	12.27	12.15	84.6	9.48	8.51	8.67	9.23	44.8
1989	18.58	15.41	16.16	16.80	24.7	13.53	13.79	13.54	13.49	53.8	13.12	12.73	13.32	13.38	55.2
1990	-7.88	-0.14	-1.94	-1.08	75.9	5.61	3.69	4.11	4.21	13.4	14.65	12.71	9.90	13.45	34.9
1991	17.91	18.85	18.24	18.58	58.2	5.22	4.52	4.72	4.91	24.1	11.53	11.47	9.57	11.67	50.4
1992	26.25	25.24	25.79	23.99	22.8	15.24	16.17	16.69	17.63	84.8	9.68	11.64	8.80	12.54	64.3
1993	13.29	15.64	16.88	15.32	59.0	17.33	20.04	19.11	19.91	89.7	5.46	7.49	9.38	6.34	68.8
1994	-4.11	-5.31	-5.78	-6.04	38.1	-7.26	-9.14	-8.26	-8.54	13.6	4.87	3.45	4.66	4.08	42.2
1986-1994	8.64	11.11	10.03	10.77	68.8	8.22	7.94	8.12	8.33	51.7	9.90	9.90	9.01	10.49	59.5
	UK Property					International Property					Total				
1986	4.21	5.50	6.59	3.80	44.5	N/A	8.52	13.76	0.70	N/A	20.94	21.33	20.47	22.01	66.3
1987	13.97	16.62	18.56	16.76	65.7	N/A	-15.90	-11.62	-36.32	N/A	5.92	2.15	3.33	2.85	16.5
1988	26.55	26.63	27.58	26.46	41.4	N/A	12.01	11.09	-9.83	N/A	14.32	11.86	13.57	11.75	14.4
1989	17.64	16.99	16.57	16.05	23.0	N/A	26.88	22.13	15.40	N/A	25.86	27.16	26.35	27.40	77.1
1990	-5.60	-5.61	-9.77	-9.28	30.5	N/A	-7.60	-16.75	-16.65	N/A	-11.89	-11.53	-11.19	-10.91	72.1
1991	-1.84	-0.83	-2.28	1.24	77.8	N/A	-4.42	-2.45	-13.46	N/A	16.58	16.40	15.28	16.45	49.0
1992	0.27	-0.32	-1.97	-0.57	56.8	N/A	-9.83	-3.16	-9.22	N/A	17.04	18.06	16.60	18.03	72.6
1993	11.57	14.31	18.60	15.53	82.9	N/A	-2.48	-4.94	5.26	N/A	24.42	25.31	24.95	25.75	73.8
1994	14.20	13.27	11.84	11.94	22.0	N/A	-14.08	-9.28	-9.04	N/A	-3.54	-4.67	-3.75	-4.97	18.4
1986-1994	9.00	9.62	9.52	9.10	39.1	N/A	-0.13	-8.13	0.00	N/A	12.18	11.79	11.73	12.03	42.8

Note: For each asset class column one gives the compound return on the external indices described in Section 3. Returns in column two are based on the peer-group WM2000 index, which is a value-weighted index tracked by WM. Columns three and four present the value- and equal-weighted returns on the UK pension funds that report, in a given year, their returns on a given asset class. The fifth column, marked "percentage of out-performers", reports the percentage of the pension funds whose returns on a given asset class exceeded the return on the external index listed in the first column.

Table 5. Performance in UK Equities against a Variety of Benchmarks

	Unconditional Alpha	Conditional Alpha (Ferson-Schadt)	(Christopherson et al)	Small Cap-Adjusted	Treynor-Mazuy	Peer-group Adjusted
Minimum	-4.59	-3.85	-6.54	-4.70	-5.07	-4.19
5%	-1.90	-1.95	-1.61	-1.87	-1.79	-1.35
10%	-1.49	-1.58	-1.18	-1.44	-1.51	-0.92
25%	-0.85	-0.91	-0.44	-0.83	-0.81	-0.33
50%	-0.15	-0.17	0.29	-0.14	-0.07	0.35
75%	0.70	0.58	1.03	0.68	0.74	1.16
90%	1.49	1.36	2.09	1.51	1.60	2.03
95%	2.14	1.90	2.55	2.15	2.06	2.69
Maximum	4.68	3.92	8.13	4.78	4.08	4.62
Range of Alpha Estimates:						
Positive	140	136	177	139	138	194
(of which significant)	(24)	(13)	(24)	(27)	(57)	(48)
Negative	166	170	129	167	168	112
(of which significant)	(29)	(27)	(12)	(27)	(67)	(13)
Bonferroni Bounds						
Minimum t-value	-5.18	-5.93	-5.24	-6.90	-5.08	-3.91
(p-value)	(< 0.0001)	(< 0.0001)	(< 0.0001)	(< 0.0001)	(< 0.0001)	(0.0138)
Maximum t-value	5.11	3.90	4.14	5.65	4.74	6.35
(p-value)	(< 0.0001)	(0.0146)	(0.0052)	(< 0.0001)	(0.0003)	(< 0.0001)
Average Alpha Estimate	-0.047	-0.127	0.332	-0.022	0.001	0.459
(t-value)	(-0.22)	(-0.57)	(0.66)	(-0.10)	(-0.01)	(4.04)

Note: This table reports the cross-sectional distribution of alpha-estimates from Jensen regressions. The unconditional Jensen regression gives the alpha estimate from a regression of the funds' equity excess returns on the excess return on the market index. The conditional Jensen regressions refine the standard equation by allowing the beta (and alpha) to depend linearly on a set of predetermined factors (Ferson-Schadt and Christopherson et al, respectively). In addition to using excess returns on the market index as a regressor, the small cap regression also includes returns on a small cap index. The Treynor-Mazuy model uses as regressors an intercept term and the level and squared value of the excess return on the market index. The peer-group model simply subtracts peer-group returns from the pension funds' equity returns. All alphas are in annualized percentage terms.

Table 6. UK Pension Funds' Alpha Values in Different Asset Categories: Quartile-Sorted According to Fund Size (Average Annual Percentages).

Quartile	UK Equities	Intl. Equities	UK Bonds	Intl. Bonds	Index Bonds	Cash/ Other Inv.	UK Property	Intl. Property	Total
A. Multi-Index Benchmark (Equation (4))									
I (Smallest)	0.352	-3.189	0.676	-3.989	0.106	0.53	-0.999	NA	-0.315
	(0.91)	(-1.28)	(1.00)	(-1.20)	(0.31)	(0.73)	(-0.98)	NA	(-0.47)
II	0.063	-2.492	0.575	-0.805	-0.344	1.545	-0.384	NA	-0.360
	(0.16)	(-1.13)	(0.92)	(-0.35)	(-0.52)	(1.57)	(-0.36)	NA	(-0.59)
III	0.213	-1.464	1.130	-1.886	0.074	0.764	-0.937	NA	0.110
	(0.68)	(-0.76)	(1.69)	(-0.85)	(0.23)	(0.93)	(-1.15)	NA	(0.21)
IV (Largest)	-0.435	-1.041	0.249	2.247	0.137	0.247	-0.334	NA	-0.268
	(-1.36)	(-0.60)	(0.28)	(0.91)	(0.46)	(0.29)	(-0.26)	NA	(-0.53)
B. Peer-Group Benchmark (Equation (3))									
I (Smallest)	0.716	-0.421	0.496	-1.631	0.306	0.733	-1.064	-0.396	0.311
	(4.60)	(-0.33)	(1.20)	(-0.60)	(0.89)	(1.07)	(-1.39)	(-0.10)	(1.23)
II	0.456	-0.396	0.298	-1.245	-0.273	1.056	-0.396	-4.356	0.157
	(2.75)	(-0.58)	(0.92)	(-0.73)	(-0.51)	(1.23)	(-0.48)	(-0.64)	(0.88)
III	0.503	0.103	0.737	-1.161	0.287	0.633	-0.794	-0.804	0.422
	(4.36)	(0.51)	(2.68)	(-0.61)	(1.05)	(0.95)	(-1.30)	(-0.18)	(3.69)
IV (Largest)	0.027	0.439	0.175	1.271	0.283	0.702	-0.668	1.920	0.037
	(0.19)	(1.26)	(0.28)	(0.52)	(1.20)	(1.03)	(-1.00)	(0.62)	(0.15)

Note: Based on their size at the beginning of each year the funds were sorted into quartiles, and four equal-weighted portfolios were formed for the following calendar year corresponding to the smallest group of funds (quartile I), the second smallest group of funds (quartile II), and so forth. This procedure was repeated each calendar year, generating four time series, each with 96 observations. For these portfolios Jensen regressions were used to estimate the mean of the risk-adjusted returns. These alpha-estimates are reported in Panel A in the case of the multi-index benchmark and in Panel B in the case of the peer-group benchmark. The reported numbers measure the alpha estimates of the portfolios, with numbers in brackets showing the t-values based on Newey-West heteroskedasticity- and autocorrelation-consistent standard errors.

Table 7. Transition Probabilities for UK Pension Funds in Different Asset Categories: Quartile-Sorted According to Annual Returns (Percentages).

Quartile at time t	Quartile at time t+12 I	II	III	IV	Quartile at time t+12 I	II	III	IV	Quartile at time t+12 I	II	III	IV	Quartile at time t+12 I	II	III	IV
	UK Equities				International Equities				UK Bonds				International Bonds			
I (Highest)	36.68*	22.86	17.93*	22.53	31.08	22.40	19.79	26.74	27.78	18.51	26.39	27.31	31.25	21.88	25.00	21.88
II	26.81	28.95*	23.19	21.05*	22.57	28.82	25.17	25.44	21.30	29.63	25.46	23.61	25.00	25.00	31.25	18.75
III	21.22	27.80	26.64	24.34	25.35	23.26	31.25	20.14	21.30	28.24	28.24	22.22	15.63	34.38	28.13	21.88
IV (Lowest)	15.30*	20.39	32.24*	32.07*	21.00	25.52	23.78	29.69	29.63	23.61	19.91	26.85	28.13	18.75	15.63	37.50
	UK Index-Linked Bonds				Cash/Other Investments				UK Property				Total			
I (Highest)	33.93	25.00	10.71*	30.36	34.96*	22.27	18.75*	24.02	30.86	21.88	21.48	25.78	28.62	26.64	22.53	22.20
II	17.86	26.79	33.93	21.43	21.88*	30.86*	27.73	19.53*	19.53	29.69	28.52	22.27	23.19	26.81	26.81	23.19
III	23.21	26.79	28.57	21.43	21.68	28.91	28.32*	21.09	22.66	25.39	26.56	25.39	24.67	23.19	26.81	25.33
IV (Lowest)	25.00	21.43	26.79	26.79	21.48	17.97*	25.20	35.35*	26.95	23.05	23.44	26.56	23.52	23.36	23.85	29.28

Note: For each year, mean excess returns, relative to a peer-group benchmark, were computed and the fund managers were sorted into quartiles according to their excess returns. Transition probabilities of moving from a particular quartile in month t to another quartile in month t+12 were computed as the average transition rates across the nine years included in the sample (1986-1994). Asterisks indicate that a transition probability estimate differs from 25 per cent at the five per cent critical level, using the standard errors from the time series of transition probability estimates.

Table 8. UK Pension Funds' Alpha Values in Different Asset Categories: Quartile-Sorted According to Previous-Year Returns (Average Annual Percentages).

Quartile	UK Equities	Intl. Equities	UK Bonds	Intl. Bonds	Index Bonds	Cash/ Other Inv.	UK Property	Intl. Property	Total
A. Multi-Index Benchmark (Equation (4))									
I (Highest)	0.574	-1.938	0.771	-1.345	0.216	1.464	-0.953	NA	0.048
	(1.55)	(-0.91)	(1.08)	(-0.48)	(0.68)	(1.79)	(-0.77)	NA	(0.08)
II	0.243	-2.082	0.585	1.761	-0.296	0.315	-0.162	NA	-0.197
	(0.75)	(-0.89)	(0.89)	(0.94)	(-0.46)	(0.51)	(-0.16)	NA	(-0.37)
III	0.071	-1.695	1.017	-0.698	0.081	0.448	-0.983	NA	-0.334
	(0.21)	(-0.81)	(1.90)	(-0.32)	'(0.20)	(0.39)	(-0.96)	NA	(-0.58)
IV (Lowest)	-0.688	-2.452	0.261	-4.151	-0.008	0.849	-0.556	NA	-0.348
	(-1.74)	(-1.08)	(0.30)	(-1.67)	(-0.03)	(0.96)	(-0.62)	NA	(-0.56)
B. Peer-Group Benchmark (Equation (3))									
I (Highest)	1.145	0.264	0.818	-2.572	-0.194	2.366	-0.295	4.092	0.331
	(5.37)	(0.55)	(2.03)	(-0.94)	(-0.32)	(2.42)	(-0.47)	(0.95)	(1.27)
II	0.604	0.068	0.335	1.325	0.275	1.117	-0.597	2.508	0.315
	(4.75)	(0.10)	(1.12)	(0.68)	(1.44)	(2.25)	(-0.80)	(0.77)	(2.29)
III	0.275	0.086	0.367	-1.291	0.184	0.215	-1.129	-0.120	0.211
	(2.32)	(0.08)	(1.15)	(-0.80)	(1.05)	(0.23)	(-1.92)	(-0.03)	(1.55)
IV (Lowest)	-0.313	-0.677	0.187	-0.229	0.331	-0.555	-0.903	1.092	0.069
	(-1.72)	(-1.18)	(0.30)	(-0.12)	(0.98)	(-0.76)	(-1.12)	(-0.33)	(0.30)
C. Zero Net Investment Portfolios									
Multi-Index	0.458	0.606	-0.105	0.349	0.015	6.334	-0.009	NA	0.208
	(4.18)	(0.69)	(-0.47)	(1.82)	(1.49)	(1.50)	(-0.03)	NA	(1.34)
Peer-Group	0.478	0.257	0.115	-0.219	0.003	2.246	0.094	0.120	0.056
	(5.23)	(0.69)	(1.30)	(-1.24)	(0.33)	(1.10)	(0.30)	(0.82)	(0.49)

Note: 1. At the end of each calendar year, benchmark-adjusted returns were computed for the assets held by the pension funds in the sample. Based on their mean, risk-adjusted returns, the funds were then sorted into quartiles, and four equal-weighted portfolios were formed for the following calendar year corresponding to the best performing quartile (quartile I), the second best performing funds (quartile II), and so forth. This procedure was repeated each calendar year, generating four time series, each with 96 observations. For these portfolios, Jensen regressions were used to estimate the mean of the risk-adjusted returns. These alpha estimates are reported in Panel A in the case of the multi-index benchmark and in Panel B in the case of the peer-group benchmark.
2. The zero net investment portfolios were based on a similar procedure, except that a single portfolio with long positions in funds that historically had a positive estimate of alpha, and short positions in funds with a negative alpha estimate, were formed so that the net cost of the portfolio equals zero. As in panels A and B, the reported figures in panel C measure the alpha estimates of the resulting portfolios, with the figures in brackets showing the t-values based on Newey-West heteroskedasticity- and autocorrelation-consistent standard errors.

Table 9. Fractiles of Alpha Estimates, Correcting for Size- and Performance-sorted Quartile Effects (Average Annual Percentages, 1987-1994).

	Benchmark Only	Benchmark & Size	Benchmark & Prior Performance	Benchmark, Size & Prior Performance
Minimum	-5.24	-5.18	-5.24	-5.23
5%	-1.86	-1.54	-1.82	-1.70
10%	-1.27	-1.02	-1.29	-1.19
25%	-0.75	-0.54	-0.74	-0.59
50%	-0.04	-0.10	-0.02	-0.10
75%	0.57	0.55	0.70	0.57
90%	1.63	1.31	1.56	1.38
95%	2.11	1.65	1.92	1.84
Maximum	4.62	3.86	4.85	4.73
Range of Alpha Estimates:				
Positive	147	134	148	140
(of which significant)	(30)	(17)	(29)	(25)
Negative	159	172	158	166
(of which significant)	(19)	(18)	(27)	(22)
Bonferroni Bounds				
Minimum t-value	-4.74	-3.75	-5.00	-4.67
(p-value)	(0.0003)	(0.0271)	(< 0.0001)	(0.0004)
Maximum t-value	4.40	3.91	4.31	3.77
(p-value)	(0.0017)	(0.0144)	(0.0025)	(0.0248)

Note: At the beginning of each year, the funds were sorted into quartiles based on either current size or (risk-adjusted) performance over the previous year. For the subsequent 12-month period, excess returns on the corresponding equal-weighted-quartile portfolios were computed and the procedure repeated to get monthly time series of excess returns for the period 1987-1994. Alpha-estimates were computed from regressions of a given fund's excess returns on an intercept, excess returns on the market portfolio, and the excess return on the quartile-sorted portfolio to which the fund belonged at a given point in time. The table reports the cross-sectional distribution of these alpha estimates.

Table 10. Spillover Effects Between Performance in Different Asset Classes: Normalized Spearman Coefficients.

	Intl. Equities	UK Bonds	Intl. Bonds	UK Index Bonds	Cash/Other Investments	UK Property
A. Unconditional Jensen Regressions						
UK Equities	5.92**	1.34	0.20	-0.35	0.32	2.14*
International Equities		-0.59	1.66	0.43	2.51*	0.12
UK Bonds			1.18	0.76	-1.83	0.34
International Bonds				1.57	1.79	0.28
UK Index Bonds					0.34	-0.03
Cash/Other Investments						1.8
B. Conditional Jensen Regressions (Ferson-Schadt)						
UK Equities	4.00**	-2.02*	0.64	-0.21	1.88	1.48
International Equities		-1.40	1.04	-1.61	0.43	-1.54
UK Bonds			0.82	1.79	-1.27	0.83
International Bonds				0.44	1.47	0.57
UK Index Bonds					0.53	0.39
Cash/Other Investments						1.55
C. Peer-Group Adjusted Returns						
UK Equities	2.71**	1.30	0.33	-1.12	0.64	1.54
International Equities		-0.53	1.47	0.26	1.73	1.08
UK Bonds			1.45	0.89	-1.43	0.15
International Bonds				1.14	1.30	-0.80
UK Index Bonds					1.35	0.16
Cash/Other Investments						1.33

Note: For each year, pension fund managers were ranked on the basis of their risk-adjusted performance over the preceding 12 months, and their average ranking was computed for the full sample (1986-1994). Then Spearman coefficients were computed for the correlation between the average rankings of fund managers in different assets. The reported test statistic is asymptotically normally distributed. A single asterisk indicates statistical significance at the five percent level, while a double asterisk indicates statistical significance at the one percent level.

Table 11. Identifying the Sources of Changes to Portfolio Weights in Different Asset Categories.

	UK Equities	Intl. Equities	UK Bonds	Intl. Bonds	UK Index Bonds	Cash/Other Investments	UK Property	Intl. Property
A. Percentage of Monthly Variation in Portfolio Weights due to Differential Returns	92.1	83.9	69.4	32.3	67.3	13.9	90.1	75.8
B. Mean Percentage Change in Portfolio Weight (Annualized)	0.98	3.97	-11.51	12.72	-0.99	3.72	-3.91	-14.09
- due to differential returns	1.19	-0.54	-0.92	-1.10	-3.63	-1.50	-2.19	-10.81
- due to net cash flow differentials	-0.21	4.51	-10.58	13.82	2.64	5.23	-1.71	-3.28

Note: 1. For each asset, panel A reports the percentage of the monthly variations in changes to portfolio weights due to differential asset returns. These results are based on the decomposition of changes in portfolio weights suggested in Section 6.
2. Using the same decomposition, panel B reports the mean percentage change in the portfolio weights and how this mean change decomposes into returns-related and net cash flow-related parts.

Table 12. Evolution of Portfolio Weights.

	UK Equities	Intl. Equities	UK Bonds	Intl. Bonds	UK Index Bonds	Cash/Other Investments	UK Property	Intl. Property
A. Variance Ratios: (Variance of Time Effect + Variance of Composite Residual)/Total Variance								
5%	0.845	0.876	0.944	0.893	0.920	0.954	0.917	0.974
10%	0.894	0.902	0.964	0.962	0.950	0.970	0.937	0.997
25%	0.950	0.943	0.995	0.988	0.997	0.991	0.988	1.003
50%	1.036	1.024	1.030	1.020	1.132	1.017	1.074	1.137
75%	1.163	1.127	1.243	1.103	1.407	1.045	1.204	1.248
90%	1.289	1.237	1.386	1.122	1.792	1.098	1.355	1.333
95%	1.397	1.329	1.451	1.139	1.814	1.150	1.508	1.600
B. Cross-Sectional Distribution of Changes in Portfolio Weights in Excess of Aggregate Change (Percentage Points Per Year)								
5%	-1.36	-1.42	-1.06	-0.62	-0.69	-0.95	-0.69	-0.22
10%	-1.00	-1.13	-0.72	-0.13	-0.46	-0.60	-0.54	-0.19
25%	-0.18	-0.78	-0.38	-0.09	-0.12	-0.37	-0.13	-0.15
50%	0.24	-0.45	-0.12	0.07	0.04	-0.16	0.15	-0.14
75%	0.75	0.00	0.41	0.12	0.34	0.04	0.45	-0.02
90%	1.14	0.51	0.76	0.20	0.77	0.25	0.61	0.08
95%	1.46	0.75	1.17	0.50	0.98	0.38	0.77	0.20
C.								
I. Correlation Between Cross-Section of Mean Net Cash Flow Rates and Mean Return Differentials								
	-0.331	-0.246	-0.426	-0.399	0.004	-0.202	-0.256	-0.349
II. Cross-Sectional Distribution of Correlation between Net Cash Flows and Return Differentials								
5%	-0.273	-0.261	-0.316	-0.311	-0.275	-0.174	-0.232	-0.076
10%	-0.228	-0.197	-0.187	-0.195	-0.251	-0.155	-0.164	-0.068
25%	-0.152	-0.106	-0.115	-0.121	-0.078	-0.091	-0.091	-0.067
50%	-0.048	0.022	-0.019	-0.048	-0.023	-0.012	-0.005	0.032
75%	0.033	0.057	0.044	0.056	0.005	0.088	0.067	0.057
90%	0.107	0.139	0.147	0.061	0.107	0.149	0.159	0.129
95%	0.142	0.157	0.212	0.079	0.186	0.211	0.187	0.142
D. Markov Transition Probabilities								
I. Year-on-Year Stayer Probabilities								
Probability	0.778	0.738	0.794	0.771	0.893	0.672	0.869	0.946
(S.E.)	(.046)	(.047)	(.044)	(.086)	(.038)	(.057)	(.032)	(.074)
II. Beginning-to-Terminal Probabilities								
Probability	0.617	0.503	0.537	0.500	0.786	0.473	0.593	0.714
E. Cross-Sectional Distribution of Persistence Coefficients								
5%	0.739	0.785	0.821	0.822	0.837	0.509	0.862	0.770
10%	0.779	0.832	0.856	0.846	0.883	0.568	0.907	0.910
25%	0.842	0.886	0.893	0.899	0.926	0.682	0.935	0.939
50%	0.904	0.930	0.939	0.942	0.948	0.775	0.954	0.962
75%	0.946	0.958	0.967	0.981	0.990	0.851	0.973	0.982
90%	0.968	0.974	0.985	0.986	1.012	0.893	0.993	0.984
95%	0.977	0.986	0.988	0.992	1.027	0.905	1.001	0.989

Note: 1. Panel A reports the cross-sectional distribution of the ratio of the variance of the common time effect plus the variance of the fund-specific change in the portfolio weights relative to the total variance of the change in a fund's portfolio weight (equation (9) in the paper).
2. Panel B presents the cross-sectional distribution of the funds' mean changes in portfolio weights relative to the mean aggregate change in portfolio weights (in percentage points per year).
3. Correlations between the cross-sectional distribution of mean net cash flow rates and mean return differentials are given in Panel C.
4. The proportion of funds with above-average weights in a given asset class that continue to have an above-average portfolio weight in that asset class is reported as stayer probabilities in panel D.
5. Panel E reports the first-order autocorrelation coefficients from a standard augmented Dickey-Fuller univatiate regression (with four lags), using the differential between the i'th fund's and the average fund's weight in a given asset class as the dependent variable.

Table 13. Decomposition of UK Pension Funds' Total Portfolio Returns (Average Annual Percentages).

A. Constant Benchmark for Normal Portfolio Weights; External Benchmarks for Normal Returns					
	Normal	Selection	Timing	Residual	Total
Mean return	12.373	-0.058	-0.333	0.052	12.034
(t-value)	(2.32)	(-0.11)	(-2.10)	(0.85)	(2.29)
B. Trended Benchmark for Normal Portfolio Weights; External Benchmarks for Normal Returns					
	Normal	Selection	Timing	Residual	Total
Mean return	12.329	-0.036	-0.288	0.031	12.034
(t-value)	(2.32)	(-0.07)	(-1.58)	(0.52)	(2.29)
C. Trended Benchmark for Normal Portfolio Weights; Peer-Group Benchmarks for Normal Returns					
	Normal	Selection	Timing	Residual	Total
Mean return	11.989	0.326	-0.215	-0.066	12.034
(t-value)	(2.27)	(2.84)	(-1.42)	(-2.20)	(2.29)
D. Trended Benchmark for Normal Portfolio Weights; Peer-Group Benchmarks for Normal Returns					
	Normal	Selection	Timing	Residual	Total
Mean return	11.970	0.322	-0.197	-0.061	12.034
(t-value)	(2.28)	(2.66)	(-1.11)	(-2.23)	(2.29)

E. Cross-sectional Distribution of Return Components							
	Normal Return		Security Selection		Market Timing		Portfolio Change
	No Trend	Trended	No Trend	Trended	No Trend	Trended	Measure
Minimum	10.869	11.003	-4.849	-4.858	-1.922	-1.113	-5.856
5%	11.797	11.679	-1.591	-1.662	-0.785	-0.729	-0.602
10%	12.041	11.869	-1.151	-1.110	-0.700	-0.620	-0.398
25%	12.261	12.167	-0.617	-0.569	-0.501	-0.476	-0.191
50%	12.400	12.377	-0.075	-0.031	-0.343	-0.289	-0.017
75%	12.516	12.558	0.478	0.378	-0.156	-0.135	0.153
90%	12.643	12.666	0.944	0.952	-0.012	0.038	0.400
95%	12.730	12.718	1.494	1.434	0.141	0.130	0.562
Maximum	13.130	13.355	4.625	4.748	0.944	1.372	4.585
Bonferroni Bounds							
Minimum t-value	NA	NA	-2.61	-2.70	-4.72	-4.42	-2.94
(p-value)	NA	NA	(1.000)	(1.000)	(0.0004)	(0.0015)	(0.521)
Maximum t-value	NA	NA	4.07	3.67	2.39	2.05	3.00
(p-value)	NA	NA	(0.007)	(0.038)	(1.000)	(1.000)	(0.417)

Note: 1. For each fund, the monthly stock returns were decomposed into returns from normal asset allocation, selection, timing, and a residual (c.f. Brinson et al (1986)). Then the mean of these components across the funds was calculated. T-values for these means were computed using the time series standard errors of the returns components as in Fama and MacBeth (1973), and are reported in brackets. External benchmark returns (Panels A and B) are based on the external indices described in Section 3 of the paper, while the peer-group benchmarks use the returns on the WM 2000 indices (Panels C and D).
2. The cross-sectional distributions in Panel E are based on external benchmarks. The portfolio change measure, proposed by Grinblatt and Titman (1989), computes the return to changes in portfolio weights relative to the portfolio weights that were in effect six months earlier.