Asset Allocation Dynamics and Pension Fund Performance

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Asset Allocation Dynamics and Pension Fund Performance*

Using a data set on more than 300 U.K. pension funds' asset holdings, this article provides a systematic investigation of the performance of managed portfolios across multiple asset classes. We find evidence of slow mean reversion in the funds' portfolio weights toward a common, time-varying strategic asset allocation. We also find surprisingly little cross-sectional variation in the average ex post returns arising from the strategic-asset-allocation, market-timing, and security-selection decisions of the fund managers. Strategic asset allocation accounts for most of the time-series variation in portfolio returns, while market timing and asset selection appear to have been far less important.

I. Introduction

> The most fundamental decision of investing is the allocation of your assets. How much should you own in stocks? How much should you own in bonds? How much should you own in cash reserves? According to a recent study, that decision has accounted for an astonishing 94% of the differences in total returns achieved by institutionally managed pension funds. (Bogle [1994], p. 235)

This quote from the chairman and founder of the Vanguard Group of mutual funds might lead one to think that the domination of managed portfolio returns by the component attributed to the strategic asset allocation decision is an established scientific verity. While many academics doubtless believe in the comparative importance of the strategic asset allocation decision, the fact is that

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the recent study to which Bogle refers is one of only two published studies on this question: Brinson, Hood, and Beebower (1986) and the follow-up study, Brinson, Singer, and Beebower (1991), both in the *Financial Analysts Journal*. Stated differently, remarkably little is known empirically about the investment performance of multiple-asset-class portfolios.[1] In addition, many of the methodological choices made in these studies have not been subject to sensitivity analysis, an exercise that might change their central conclusions.[2]

To the best of our knowledge, this article provides the first systematic academic investigation of the performance of such managed portfolios. We analyze a data set provided by the WM Company containing 9 years of monthly information on the holdings in eight classes of assets by 306 U.K. pension funds. Hence, we have a sample that is well suited to a detailed examination of fund performance in terms of market timing (variations over time in the 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 2 years longer than the average duration of an investment management contract in the United Kingdom. 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 the importance of the strategic asset allocation (long-run allocation of funds across asset classes) to portfolio performance.

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

The industrial organization of the U.K. pension fund industry offers

1. Most of the studies on U.S. 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).

2. For example, Jahnke (1997) has criticized the Brinson et al. studies on a number of grounds, some of which are, at best, opaque to us and have been ably addressed in Singer (1997). Nevertheless, several of his criticisms are potentially important, including the interpretation of the comparative economic and statistical importance of and presumption of a fixed long-term asset allocation and the limited number of asset classes and time-series observations used in their analysis. The first potential problem is particularly relevant in our application.

an interesting case study. Over the period under investigation, U.K. pension fund managers faced arguably the smallest set of externally imposed restrictions and regulations on their investment behavior of any group of institutional investors anywhere in the world. They were, by and large, unconstrained by their liabilities: U.K. 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 important, in their choice of investments. Unlike many of their counterparts in continental Europe and elsewhere, U.K. 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 U.S. counterparts, U.K. pension fund managers faced no substantive regulatory controls on or real threat of litigation over imprudent investment behavior during this period.

This relative freedom together with the presence of large actuarial surpluses accounts for several important differences between the portfolio holdings of U.S. and U.K. pension funds. U.S. pension funds are far more heavily invested in lower volatility domestic bonds than their U.K. counterparts, while, conversely, U.K. pension funds have a far larger weighting in higher volatility equities. The general absence of constraints on investment behavior 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, we should not be surprised if there is comparatively little cross-sectional variation in performance compared with the striking differences observed in U.S. data. U.K. fund managers are explicitly evaluated in relative terms, and the U.K. fund management industry is highly concentrated, suggesting that firms risk losing substantial market share in the event of bad relative performance. Our data permit us to see whether these incentive effects or the efforts to translate the absence of constraints into active management dominate actual portfolio behavior.

The structure of this article is as follows. We begin with a brief review of pension funding arrangements in the United Kingdom (Sec. II) and a description of our data set (Sec. III). We then analyze the asset allocation decisions of fund managers. We decompose changes in portfolio weights over time into return and cash-flow components (Sec. IV) and performance into security-selection and market-timing components (Sec. V). Section VI concludes.

II. Pension Funding in the United Kingdom

Pension trust law is very flexible in the United Kingdom, enabling the trust deed to be drawn up in virtually any way that suits 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 to preserve the trust capital and to apply the capital and its income according to the trust deed, and members can sue for compensation if they suffer loss as a result of negligence by trustees. In addition, pension fund managers were, over the sample period, 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 the possibility of recourse only through the courts.[3]

The U.S. and U.K. pension fund industries differ significantly in terms of their concentration. Lakonishok, Shleifer, and Vishny (1992) report that none of the independent investment counselors in the defined benefit group they considered for the United States had a market share above 3.7%. In contrast, the top five U.K. asset management groups (Mercury Asset Management, Phillips and Drew Fund Management, Gartmore Pension Fund Managers, Morgan Grenfell Asset Management, and Schroder Investment Management) managed 1,154 funds among them as of year-end 1993, accounting for about 80% of the market, according to Lambert (1998).

Another unusual feature of the U.K. experiment concerns overfunding. Huge pension fund surpluses, equivalent to half the value of pension fund assets at the time, built up during the early 1980s. This may have lowered the pressure on fund managers to earn high levels of return in the short term.[4] Furthermore, most fund managers wished to be seen as offering a ‘‘balanced’’ service, in part because U.K. fund managers tended in the past not to want to be typecast. In contrast, U.S. fund managers are usually characterized by an investment style and cannot subsequently change their style if selected by a client with the aid of a consultant.

3. The changes introduced by the 1995 Pensions Act bring the U.K. pensions regulatory framework closer to the prudent-man principle established by the U.S. Employee Retirement Income Security Act of 1974. However, substantial differences remain, e.g., the compensation scheme established by the 1995 Act explicitly sought to avoid the problems with deliberate underfunding. Similarly, the trustees must now conduct an asset-liability modeling exercise that obliges them to establish a strategic or long-run asset allocation (the ‘‘statement of investment principles’’).

4. For example, Hart (1992), in his discussion of Lakonishok et al. (1992), hypothesized that overfunded plans’ fund managers have only relatively weak incentives to pursue high investment returns on pension assets. In contrast, the incentives to perform are likely to be much stronger in the case of underfunded schemes where the sponsoring company is responsible for making up any shortfall.

All of the managers in our data set were in place throughout the sample period 1986–94, and the average length of tenure of a pension fund manager in the U.K. is 7.25 years (Prosser 1995). The largest fund management groups have the most secure reputations and, according to Kay, Laslett, and Duffy (1994), use their track records to retain existing clients or to attract new clients, rather than to extract higher fees. In addition, U.K. pension fund trustees 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 investment mandates. These considerations all point to substantial disincentives to manage portfolios actively in ways that risk large differences in relative performance.[5]

The fees charged by a fund management group are related, to some extent, to managerial performance, 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 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.

These institutional arrangements reveal important features of the U.K. experiment:

1. U.K. pension fund managers have a weak incentive to add value and 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 resulting limits are so flexible as to be effectively unenforced 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. Over the course of a mandate, most U.K. pension fund managers earn fees related solely to the value of assets under management and not to their relative performance against either a predetermined bench-

5. To the extent that U.K. managers provide services beyond adding value, they are more akin to financial institutions such as bank trust departments and insurance companies that produce commodity financial services.

mark or their peer group (i.e., there is generally no specific penalty for underperforming and no specific reward for outperforming an agreed upon benchmark).

4. The heavy concentration in the U.K. industry is likely to lead to portfolios being dominated by a small number of ‘‘house positions’’ in respect of asset allocations, with each fund management house’s preferred position similar to the others to reduce the risk of relative underperformance.[6]

III. Data Description

Our data consist of monthly observations on 306 U.K. pension funds from 1986 to 1994 provided to us by the WM Company. The sample is complete in the sense that it contains all of the funds that maintained the same single, externally appointed fund management group throughout the period and that submitted continuous return records to WM. For each fund, we have data on the overall portfolio and eight constituents: U.K. equities, international equities, U.K. bonds, international bonds, U.K. index-linked bonds, cash, U.K. property, and international property. For each asset class, each fund reported initial market value and net investment, the mean (time-weighted) asset value, income received, and return over the month. Compared with Brinson et al. (1986), we have more time-series observations (108 to 40), more funds (306 to 91), and data on more asset classes (8 to 3), including holdings of international equities and bonds. For each group of assets, every fund in the sample reported initial market value, net investment in the asset over the month, the mean (time-weighted) asset value over the month, income received over the month, and return on the asset. All assets were denominated in pounds sterling.

WM uses a range of value-weighted asset-class benchmarks to assess the performance of the funds in its stable. The set of external indices that it used includes:[7]

1. U.K. Equities: Financial Times Actuaries (FTA) All-Share Index;
2. International Equities: FTA World (excluding U.K.) Index;
3. U.K. Bonds: British Government Stocks All-Stocks Index;
4. International Bonds: J.P. Morgan Global (excluding U.K.) Bonds Index;

6. For a comparative analysis of the incentives operating in the U.S. pension fund industry, see Lakonishok et al. (1992).

7. There is one exception to the use of these indices. For 1986 only, the WM PUT Property Index was used to measure returns on U.K. property.

5. U.K. Index Linked: British Government Stocks Index-Linked All-Stocks Index;
6. Cash: LIBID (London Inter-Bank Bid Rate) 7-day deposit rate; and
7. U.K. Property: Investment Property Databank (IPD) All-Property Index.

All of these indices assume that income is reinvested.[8] The WM Pension Fund Index for total assets is 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 contains less than 0.5% of the total portfolio value in our sample.

These benchmarks have the virtue of being independently calculated indices that are immediately publicly available and widely used for short-term performance measurement in the United Kingdom. However, several of them, most notably international equities and cash, have weightings that can differ substantially from those of the pension funds. Accordingly, we also use the WM2000 peer-group indices that contain all funds ranked below the largest 50 funds tracked by WM. Their weightings are more typical of those achieved by single externally appointed fund managers.[9]

Pension funds of very different size populate our sample. As of December 1994, the smallest fund funds had assets just above £1 million, and 28 funds had assets below £10 million. At the other end of the scale, two funds had assets between £10 billion and £20 billion. The vast majority of funds in our sample had assets between £10 million and £1 billion, and the median fund size was £54.4 million.[10]

An important component of our experiment is the examination of the persistence of investment performance over time. Accordingly, we

8. Property returns are particularly subject to measurement problems so we briefly explain how these were computed. Returns on the All-Property Index are designed to approximate daily continuous compounding by assuming that rental income is received in mid-month. They are computed as capital value at the end of the month plus capital value at the beginning of the month plus 1/12 times the annual rental income, all divided by the capital value at the beginning of the month plus 1/2 times the net investment during the month minus 1/2 times the average monthly rental income.

9. We were unable to obtain information on the exact transactions costs (spreads and commissions) or running costs (management and custody fees, property security, and insurance costs, etc.) incurred by the various funds. Hence, returns are gross of all these costs, except dealing spreads, which are automatically included. 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 with index returns, an appropriate penalty when funds could have been passively managed at extremely low cost in the external benchmarks.

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

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 restriction to managers who are in place over the whole sample introduces another potential problem that has recently received substantial attention in the literature, namely, survivor bias.[11] Funds were excluded from our data set because there was a change in manager or in management structure or because a fund left or joined part of the way through the sample period, not necessarily because of poor performance. Nevertheless, there is a trade-off between greater precision induced by larger samples and the potential bias induced by sample selection in our performance measures.

Fortunately, we are in a position to assess directly some of the facts regarding survivor bias in our sample in two ways that are reported in tables 1 and 2. Table 1 presents the annual portfolio allocation across eight categories of assets for all funds in our sample along with the aggregate portfolio weights for the entire population of U.K. pension funds tracked by WM (1,034 at the end of 1994). Reassuringly, the differences between the aggregate asset allocation of the pension funds in our sample and the overall asset allocation of the WM universe seem numerically and economically trivial year by year: we would expect to observe large differences if managers systematically lost their mandates by making bad market timing bets. Table 2 reports the corresponding aggregate returns in each asset class and for the aggregate portfolio for both the entire WM universe and our subset of it. If survivor bias infected the funds included in our subsample, they should be more successful ex post than those in the overall universe monitored by WM, both on average and increasingly over time, peaking toward the end of the sample as poorly performing funds systematically dropped out. As is readily apparent, neither tendency arises on average or over time across asset classes and for the overall portfolio.[12] In short, the cost in terms of inducing potential survivor bias seems to be

11. For recent examinations of survivor bias, see Grinblatt and Titman (1989, 1992); Brown, Goetzmann, Ibbotson, and Ross (1992); Brown and Goetzmann (1995); and Malkiel (1995).

12. For example, the WM2000 return actually exceeded that of our universe by an economically trivial 6 basis points over the whole sample. Similarly, the peer-group index underperformed the value-weighted portfolio by 28 basis points per year during the first half of the sample but outperformed the latter portfolio by 39 basis points 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. Finally, the correlation between the return on the external and peer-group indices and the value- and equal-weighted portfolios constructed from our sample of funds all exceed 0.995.

TABLE 1 Aggregate Portfolio Weights for U.K. and U.S. Pension Funds, 1986–94

	1986	1987	1988	1989	1990	1991	1992	1993	1994
WM sample:									
Domestic equities	50.4	52.9	54.0	53.2	53.6	54.9	55.6	54.8	53.6
International equities	19.6	15.7	15.8	19.6	16.8	20.4	21.4	23.8	22.5
Domestic bonds	12.2	12.0	10.5	7.5	6.5	5.2	4.7	4.6	5.3
International bonds	.9	.8	.8	1.3	2.5	3.5	3.9	3.3	2.8
Index bonds	3.1	2.9	2.8	2.3	2.4	2.0	2.2	2.8	3.6
Cash/other investments	3.6	5.2	4.9	5.1	6.3	4.0	3.8	3.7	4.2
Domestic property	9.2	9.1	10.0	9.8	10.9	9.1	7.6	6.5	7.6
International property	1.0	1.4	1.2	1.2	1.0	.9	.8	.5	.4
WM universe:									
Domestic equities	51	54	53	53	54	56	56	56	54
International equities:	20	14	16	21	18	21	22	24	22
North America	N.A.	N.A.	N.A.	7	5	6	6	5	4
Continental Europe	N.A.	N.A.	N.A.	7	7	8	8	9	8
Japan	N.A.	N.A.	N.A.	5	3	4	4	4	5
Total Pacific (except Japan)	N.A.	N.A.	N.A.	1	2	2	3	5	4
Others	N.A.	N.A.	N.A.	1	1	1	1	1	1
Domestic bonds	13	13	10	6	6	5	4	4	6
International bonds	0	1	1	2	3	4	4	4	4
Domestic index-linked bonds	3	3	3	2	3	2	3	3	4
Cash/other investments	4	5	6	6	7	4	4	4	4
Domestic property	8	9	10	9	8	7	6	5	6
International property	1	1	1	1	1	1	1	0	0

	1986	1990	1994
U.S. pension funds:			
Domestic equities	45.6	42.1	44.8
International equities	2.6	4.5	8.3
Domestic bonds	37.8	38.9	34.2
International bonds	.0	.0	2.0
Index bonds	.0	.0	.0
Cash/other investments	7.8	9.8	7.5
Domestic property	6.2	4.7	3.2
International property	.0	.0	.0

NOTE.—For each year the table shows the percentage of the total portfolio invested in the assets held by U.K. and U.S. pension funds. The ‘‘WM sample’’ figures are based on the U.K. 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 U.S. pension funds were supplied by Greenwich Associates. N.A. = not available.

TABLE 2 U.K. Pension Fund Investment Performance with Respect to Different Benchmarks, 1986–94 (Annual Percentages)

Year	External Index (1)	WM2000 Return (2)	Value Weighted Return (3)	Equal Weighted Return (4)	% of Out-Performers (5)
U.K. equities:					
1986	24.24	22.76	23.35	23.64	36.4
1987	7.65	7.07	6.30	8.08	48.5
1988	10.90	9.60	10.48	9.76	31.4
1989	30.80	30.36	30.83	30.59	47.5
1990	−10.20	−10.43	−10.20	−9.58	58.2
1991	18.84	18.06	18.22	18.14	40.4
1992	18.59	19.13	18.71	19.36	66.7
1993	24.90	24.23	24.78	24.97	48.7
1994	−6.01	−5.83	−5.78	−6.08	48.2
1986–94	13.30	12.77	12.97	13.21	44.8
International equities:					
1986	31.62	31.29	31.56	30.98	45.4
1987	−9.12	−22.60	−20.16	−21.78	.6
1988	27.03	19.92	21.65	19.16	2.8
1989	27.38	34.19	33.67	34.50	97.5
1990	−39.77	−31.35	−32.27	−31.39	98.0
1991	21.30	18.26	19.41	18.00	12.8
1992	15.51	18.02	17.88	17.46	79.2
1993	22.55	34.21	32.91	34.36	97.5
1994	.85	−4.57	−3.60	−4.96	2.2
1986–94	11.11	10.82	11.23	10.70	39.8
U.K. bonds:					
1986	10.85	11.39	11.35	11.22	63.3
1987	14.17	14.91	15.63	14.96	76.0
1988	6.52	7.94	8.35	7.81	80.5
1989	7.85	7.16	6.73	7.61	35.9
1990	9.15	6.31	7.53	8.85	28.1
1991	14.95	16.46	16.78	18.19	84.1
1992	17.05	17.35	17.08	17.13	63.5
1993	19.01	22.92	22.14	22.15	76.7
1994	−6.43	−9.64	−8.74	−9.37	14.1
1986–94	10.35	10.53	10.76	10.95	77.3
International bonds:					
1986	16.92	26.18	18.95	24.85	69.8
1987	−12.24	−3.09	−1.01	−5.60	74.5
1988	9.06	7.27	3.01	10.11	37.0
1989	18.58	15.41	16.16	16.80	24.7
1990	−7.88	−.14	−1.94	−1.08	75.9
1991	17.91	18.85	18.24	18.58	58.2
1992	26.25	25.24	25.79	23.99	22.8
1993	13.29	15.64	16.88	15.32	59.0
1994	−4.11	−5.31	−5.78	−6.04	38.1
1986–94	8.64	11.11	10.03	10.77	68.8
U.K. index-linked bonds:					
1986	6.60	5.06	4.97	5.46	11.2
1987	6.37	4.94	5.97	5.74	43.0
1988	11.37	12.37	12.27	12.15	84.6
1989	13.53	13.79	13.54	13.49	53.8
1990	5.61	3.69	4.11	4.21	13.4
1991	5.22	4.52	4.72	4.91	24.1
1992	15.24	16.17	16.69	17.63	84.8
1993	17.33	20.04	19.11	19.91	89.7
1994	−7.26	−9.14	−8.26	−8.54	13.6
1986–94	8.22	7.94	8.12	8.33	51.7

TABLE 2 (***Continued***)

Year	External Index (1)	WM2000 Return (2)	Value Weighted Return (3)	Equal Weighted Return (4)	% of Out-Performers (5)
Cash/other investments:					
1986	10.71	12.13	8.75	12.58	60.5
1987	9.63	9.02	8.06	11.12	48.8
1988	9.48	8.51	8.67	9.23	44.8
1989	13.12	12.73	13.32	13.38	55.2
1990	14.65	12.71	9.90	13.45	34.9
1991	11.53	11.47	9.57	11.67	50.4
1992	9.68	11.64	8.80	12.54	64.3
1993	5.46	7.49	9.38	6.34	68.8
1994	4.87	3.45	4.66	4.08	42.2
1986–94	9.90	9.90	9.01	10.49	59.5
U.K. property:					
1986	4.21	5.50	6.59	3.80	44.5
1987	13.97	16.62	18.56	16.76	65.7
1988	26.55	26.63	27.58	26.46	41.4
1989	17.64	16.99	16.57	16.05	23.0
1990	−5.60	−5.61	−9.77	−9.28	30.5
1991	−1.84	−.83	−2.28	1.24	77.8
1992	.27	−.32	−1.97	−.57	56.8
1993	11.57	14.31	18.60	15.53	82.9
1994	14.20	13.27	11.84	11.94	22.0
1986–94	9.00	9.62	9.52	9.10	39.1
International property:					
1986	N.A.	8.52	13.76	.70	N.A.
1987	N.A.	−15.90	−11.62	−36.32	N.A.
1988	N.A.	12.01	11.09	−9.83	N.A.
1989	N.A.	26.88	22.13	15.40	N.A.
1990	N.A.	−7.60	−16.75	−16.65	N.A.
1991	N.A.	−4.42	−2.45	−13.46	N.A.
1992	N.A.	−9.83	−3.16	−9.22	N.A.
1993	N.A.	−2.48	−4.94	5.26	N.A.
1994	N.A.	−14.08	−9.28	−9.04	N.A.
1986–94	N.A.	−.13	−8.13	.00	N.A.
Total:					
1986	20.94	21.33	20.47	22.01	66.3
1987	5.92	2.15	3.33	2.85	16.5
1988	14.32	11.86	13.57	11.75	14.4
1989	25.86	27.16	26.35	27.40	77.1
1990	−11.89	−11.53	−11.19	−10.91	72.1
1991	16.58	16.40	15.28	16.45	49.0
1992	17.04	18.06	16.60	18.03	72.6
1993	24.42	25.31	24.95	25.75	73.8
1994	−3.54	−4.67	−3.75	−4.97	18.4
1986–94	12.18	11.79	11.73	12.03	42.8

NOTE.—For each asset class, column 1 gives the compound return on the value-weighted external indices described in Section III. Returns in column 2 are based on the peer-group WM2000 index, which is a value-weighted index constructed by WM. Columns 3 and 4 present the value- and equal-weighted returns on the U.K. pension funds that report, in a given year, their returns on a given asset class. Column 5 reports the percentage of the pension funds whose returns in a given year and in a given asset class exceeded the return on the external index listed in the first column. N.A. = not available, or could not be calculated.

small relative to the gains in precision from lengthening the sample.[13]

Before proceeding, it is worth describing two regularities that pose the greatest empirical challenge to the interpretation of U.K. pension fund performance. The first concerns the behavior of the overall asset allocation of U.K. pension funds, namely, the substantial trend toward domestic and international equities and away from domestic bonds with more modest movements 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.

By 1993, domestic and international equities constituted more than 78% of the aggregate portfolio value of U.K. 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% prevailing in 1986. The allocation of more than 20% to international equities is even more striking.[14] In contrast, U.K. pension funds decreased their holdings of U.K. bonds from 12% to 5%, while international bonds experienced a modest increase, rising from 1% to 3%. The proportion invested in U.K. 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 indicate that the pension funds included in our sample had not reached stable long-term asset allocations.

For comparison, the final columns in table 1 give the average portfolio holdings for U.S. pension funds at the end of 1986, 1990, and 1994, respectively. These figures confirm the striking differences between the holdings of U.K. and U.S. pension funds. U.K. 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 3 percentage points or so less in cash compared with their U.S. counterparts.

The second striking regularity is the remarkably low cross-sectional variation in average total return across the funds in our sample. We found that the semi-interquartile range runs from 11.47% to 12.59% per year and less than 300 basis points separates the funds in the 5th and

13. While there is no evidence of survivor bias on average, 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 both within and across asset classes.

14. Pension fund assets invested in U.K. equities actually declined between 1975 and 1983 before rising dramatically between 1984 and 1993. A pronounced jump in international equity holdings followed the abolition of U.K. exchange controls in 1979: the average allocation to international equities rose from 6% in 1979 to 20% in 1986, temporarily declining in 1987 and 1988 before surging past this level between 1988 and 1993.

95th percentiles. To be sure, there is somewhat greater cross-sectional variability in particular asset classes. For example, the annualized semi-interquartile range for U.K. 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 even 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 U.S. equity mutual funds.

IV. Pension Fund Asset Allocation Strategies and Performance

We exploit the information on U.K. pension fund asset allocations over time in two steps. We noted earlier that the funds tilted their asset allocation toward 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 (i.e., a change in the strategic asset allocation) or the reward for a market timing bet that turned out well ex post. We need a better understanding of asset allocation dynamics in order to identify any market-timing or security-selection ability among our sample of managers. Accordingly, the next section studies various aspects of aggregate portfolio dynamics and the concomitant cross-sectional variation in asset allocation across individual funds. Armed with the results from this exercise, the next section then 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).

A. The Evolution of Aggregate Portfolio Weights

We employ a simple decomposition to help identify the factors causing portfolio weights to change. Asset classes that enjoy large positive relative returns also experience an increase in their allocations in the total portfolio unless fund managers deliberately rebalance portfolios as this occurs.

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

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

where r_{jt} is the rate of return on U.K. pension funds' holdings of asset class j and NCF_{jt} is the rate of net cash flow into that asset class during month t. Using this relation, the portfolio weight of asset class j (ω_{jt}) can be written as

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

Taking log-differences, it follows that

$$\Delta\log(\omega_{jt}) = \log(1 + r_{jt} + NCF_{jt}) - \log\left[1 + \sum_{k=1}^{M}\omega_{kt}(r_{kt} + NCF_{kt})\right], \quad (3)$$

so that, to a close approximation,

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

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

$$\mathrm{var}(\Delta\log(\omega_{jt})) \approx \mathrm{var}(r_{jt} - r_{pt}) + \mathrm{var}(NCF_{jt} - NCF_{pt}) + 2\mathrm{cov}(r_{jt} - r_{pt}, NCF_{jt} - NCF_{pt}). \quad (5)$$

The decomposition in (4) enables us to measure the extent to which changes in aggregate portfolio weights are caused by differential returns across asset classes, as indicated by $r_{jt} - r_{pt}$, or by shifts in net cash flows across asset classes, as indicated by $NCF_{jt} - NCF_{pt}$. Shifts 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 allocation. In contrast, revisions associated with the second component result from the active strategy of rebalancing the portfolio by redirecting cash flows across asset groups, although rebalancing toward the long-run or strategic asset allocation would generally be viewed as part of a passive, not active, investment strategy. The dramatic increase in the allocation to equities might simply reflect the fact that stocks generated higher mean returns than the other asset categories over the sample.

Panel A of table 3 reports the sample means of $\Delta\log(\omega_{jt})$ and its two components, $r_{jt} - r_{pt}$ and $NCF_{jt} - NCF_{pt}$ (see [4]). The only asset class for which differential returns contributed positively to its asset allocation was U.K. equities, the only asset class whose mean return exceeded that of the total portfolio over the sample. Thus, any increase in the portfolio weights of the remaining asset classes must have been due by definition to net purchases. The large flow of funds out of U.K. bonds was almost entirely due to net sales, while international bonds saw a similar percentage increase due to net purchases. In contrast, the declining weights in index bonds and international property were entirely due to poor relative returns for these asset classes.

Panel B of table 3 reports the percentage of the short-term variation in aggregate asset allocations, as measured by the variance in percentage changes in portfolio weights, accounted for by variations in, respectively, return differentials, net cash flow differentials, and their covariance (see [5]). The results suggest that return differentials (1) largely account for the monthly variation in the weights allocated to U.K. and international equities and to U.K. property, (2) account for much of the monthly variation (of the order of 40%–50%) in the weights allocated to conventional and index-linked U.K. 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.

B. *The Evolution of Individual Funds' Portfolio Weights*

The above observations concern only the dynamics of the aggregate portfolio weights. We are also interested in cross-sectional aspects of the asset allocation dynamics, both for their implications concerning performance measurement and for our understanding of pension fund behavior. Accordingly, consider the fund-specific version of (4):

$$\Delta\log(\omega_{ijt}) \approx r_{ijt} - r_{ipt}r_{ipt} + NCF_{ijt} - NCF_{ipt}, \tag{6}$$

where i indexes pension funds. Subtracting equation (4) from (6) yields

$$\begin{aligned}\Delta\log(\omega_{ijt}) - \Delta\log(\omega_{jt}) \approx {} & [(r_{ijt} - r_{ipt}) - (r_{jt} - r_{pt})] \\ & + [(NCF_{ijt} - NCF_{ipt}) \\ & \quad - (NCF_{jt} - NCF_{pt})] \\ \equiv {} & \psi_{ijt}.\end{aligned} \tag{7}$$

Equation (7) is in the form of a fixed-effects dummy-variable model: $\Delta\log(\omega_{jt})$ is a time effect common across funds, and the composite residual on the right-hand side of (7) is a fund-specific effect with a

TABLE 3 **Identifying the Sources of Changes to Aggregate Portfolio Weights across Asset Classes**

	U.K. Equities	International Equities	U.K. Bonds	International Bonds	U.K. Index Bonds	Cash/Other Investments	U.K. Property	International Property
A. Mean percentage change in portfolio weight (annualized)	.98	3.97	−11.51	12.72	−.99	3.72	−3.91	−14.09
Due to differential returns	1.19	−.54	−.92	−1.10	−3.63	−1.50	−2.19	−10.81
Due to net cash flow differentials	−.21	4.51	−10.58	13.82	2.64	5.23	−1.71	−3.28
B. Percentage of monthly variance in portfolio weights:								
Due to variance of differential returns	91.13	60.31	39.82	16.10	40.06	15.18	76.31	50.91
Due to variance of net cash flow differentials	7.25	12.75	21.54	40.34	22.06	42.53	8.34	17.29
Due to covariance between these	1.62	26.94	38.64	43.55	37.88	42.29	15.34	31.80

NOTE.—Each month a value-weighted aggregate portfolio was formed by aggregating the individual pension funds' investments. The time series of the aggregate portfolio holdings were then used to compute portfolio weights for each asset class. The results in this table are based on the decomposition of percentage changes in aggregate portfolio weights into a return differential component and a net cash flow component using equation (4). For each asset class, panel A reports the mean annual percentage change in the aggregate portfolio weights and how this mean change decomposes into differential return and net cash flow components. Panel B reports the decomposition of the monthly variance of changes in the aggregate portfolio weights due to the variance in differential returns, the variance in net cash flow differentials, and the covariance between these.

nonzero mean. However, the standard model typically postulates that the time- and fund-specific effects are uncorrelated both with each other and cross-sectionally, whereas the absence of such a correlation need not be a feature of our data.[15] Nevertheless, we consider this model to be a useful baseline and can envisage other models in which relative performance evaluation leads managers to follow strategies that make this a natural decomposition.

Panel A of table 4 describes the extent to which individual fund portfolio weights conform to the fixed-effects model. We report the cross-sectional distribution of the variance ratio:

$$[\mathrm{var}(\Delta\ \log(\omega_{jt})) + \mathrm{var}(\psi_{ijt})]/\mathrm{var}(\Delta\ \log(\omega_{ijt})), \tag{8}$$

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 close to unity for all asset classes. Similarly, the changes in most fund asset allocations relative to the value-weighted average 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 (i.e., those with the largest positive correlations between $\Delta\ \log(\omega_{jt})$ and ψ_{ijt}) lie between 0.85 and 0.97, and the corresponding ratios for the 25th percentile lie between 0.94 and unity. There is somewhat greater spread in the variance ratios associated with negative correlations between $\Delta\ \log(\omega_{jt})$ and ψ_{ijt}, with ranges of 1.05 to 1.41 and 1.14 to 1.81 at the 75th and 95th percentiles, respectively. Nevertheless, changes in the asset allocations of most funds appear to largely, although not entirely, involve random variations about a common trend.[16]

Panel B of table 4 reports the fractiles of the percentage changes in the funds' portfolio weights in excess of the corresponding aggregate change, that is, $\Delta\ \ln(\overline{\omega}_{ij}) - \Delta\ \ln(\overline{\omega}_{j})$, where averages are taken over time. 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. This range of variation is generally large relative to the average annual rates of change in the asset allocations themselves: of the order of 51 and 76 basis points for U.K. and international equities, respectively, -106 basis

15. 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 (7).

16. We also examined the coefficient from the regression of $\Delta\ \log(\omega_{ijt}) - \Delta\ \log(\omega_{jt})$ on $\Delta\ \log(\omega_{jt})$, which should be zero in the same circumstances. 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 5% level ranged from 20% to a little more than 40%.

TABLE 4 Determinants of Individual Funds' Portfolio Weight Changes: Cross-Sectional Results

	U.K. Equities	International Equities	U.K. Bonds	International Bonds	U.K. Index Bonds	Cash/Other Investments	U.K. Property	International Property
A. Variance ratios (variance of time effect + variance of composite residual)/total variance:								
5%	.845	.876	.944	.893	.920	.954	.917	.974
10%	.894	.902	.964	.962	.950	.970	.937	.997
25%	.950	.943	.995	.988	.997	.991	.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 the aggregate change (percentage points per year):								
5%	−1.36	−1.42	−1.06	−.62	−.69	−.95	−.69	−.22
10%	−1.00	−1.13	−.72	−.13	−.46	−.60	−.54	−.19
25%	−.18	−.78	−.38	−.09	−.12	−.37	−.13	−.15
50%	.24	−.45	−.12	.07	.04	−.16	.15	−.14
75%	.75	.00	.41	.12	.34	.04	.45	−.02
90%	1.14	.51	.76	.20	.77	.25	.61	.08
95%	1.46	.75	1.17	.50	.98	.38	.77	.20

C. Correlation between the cross section of mean net cash flow rates and mean return differentials	−.331	−.246	−.426	−.399	.004	−.202	−.256	−.349
D. Cross-sectional distribution of correlation between net cash flows and return differentials:								
5%	−.273	−.261	−.316	−.311	−.275	−.174	−.232	−.076
10%	−.228	−.197	−.187	−.195	−.251	−.155	−.164	−.068
25%	−.152	−.106	−.115	−.121	−.078	−.091	−.091	−.067
50%	−.048	.022	−.019	−.048	−.023	−.012	−.005	.032
75%	.033	.057	.044	.056	.005	.088	.067	.057
90%	.107	.139	.147	.061	.107	.149	.159	.129
95%	.142	.157	.212	.079	.186	.211	.187	.142

Note.—These results are based on the decomposition of individual funds' monthly portfolio weights into a return differential and a net cash flow component using equation (7). Panel A reports the cross-sectional distribution of the ratios 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 the individual funds' portfolio weight in each asset class (eq. [8]). Panel B presents the cross-sectional distribution of the individual funds' mean changes in portfolio weights relative to the mean aggregate change in portfolio weights (in percentage points per year). Panel C reports the correlations between the cross-sectional distribution of mean net cash flow rates and mean return differentials, while panel D reports the cross-sectional distribution of the sample time-series correlation between individual funds' net cash flow differentials and their return differentials.

points for U.K. bonds, 35 basis points for U.K. property, and between −10 and 16 basis points for the remaining asset classes. The substantial overall drift toward equities over the sample conceals a wide range of drift rates across the individual funds.

Panel C of table 4 sheds some light on both the size and timing of any rebalancing toward or away from asset classes that experienced good or bad performance relative to the aggregate peer-group benchmark. While the aggregate asset allocation shifted toward asset classes that performed relatively well over the sample, the cross-sectional correlation between average excess net cash flow (i.e., the time-series mean of $[(NCF_{ijt} - NCF_{ipt}) - (NCF_{jt} - NCF_{pt})]$) and the corresponding average excess asset class return (i.e., the time-series mean of $(r_{ijt} - r_{ipt}) - (r_{jt} - r_{pt})$) is negative for all asset classes, except index-linked bonds, with correlations between −0.20 and −0.43. Thus the funds with the highest relative return within a given asset class were also the ones with the smallest net cash flow into that asset class, suggesting that cash flows are used to stabilize the actual asset allocation around a common (and possibly dynamically changing) strategic asset allocation.

Moreover, panel D of table 4 shows 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 $[(NCF_{ijt} - NCF_{ipt}) - (NCF_{jt} - NCF_{pt})]$ and $[(r_{ijt} - r_{ipt}) - (r_{jt} - r_{pt})]$. The median time-series correlation is numerically and economically close to zero, and the 5th percentile (i.e., the funds with correlations smaller than those of 95% of the fund universe) is closer to zero than the corresponding cross-sectional correlation 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, adds weight to our finding that funds exhibited a tendency to rebalance toward their strategic asset allocations when relative asset returns moved out of line.

These statistics measure the average behavior of individual fund asset allocations, but they reveal little about any mean reversion tendencies they may exhibit. Any such mean reversion would have to be quite pronounced to be reliably estimated in a sample as short as ours. Panel A of table 5 reports 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 fairly low probabilities of between one-twentieth and one-third of crossing over the average. The time-series standard errors of the sample transition probabilities are sufficiently small that we may infer that the corresponding population probabilities are far from the null value of 50%, both economically and statistically. Similarly, panel B of table 5 provides the sample probabilities for the transitions from

TABLE 5 Dynamics of Individual Funds' Portfolio Weights

	U.K. Equities	International Equities	U.K. Bonds	International Bonds	U.K. Index Bonds	Cash/Other Investments	U.K. Property	International Property
A. Markov switching probabilities:								
I. Year-on-year stayer probabilities:								
Probability	.778	.738	.794	.771	.893	.672	.869	.946
Standard error	.046	.047	.044	.086	.038	.057	.032	.074
II. Beginning-to-terminal probabilities	.617	.503	.537	.500	.786	.473	.593	.714
B. Cross-sectional distribution of persistence coefficients:								
5%	.739	.785	.821	.822	.837	.509	.862	.770
10%	.779	.832	.856	.846	.883	.568	.907	.910
25%	.842	.886	.893	.899	.926	.682	.935	.939
50%	.904	.930	.939	.942	.948	.775	.954	.962
75%	.946	.958	.967	.981	.990	.851	.973	.982
90%	.968	.974	.985	.986	1.012	.893	.993	.984
95%	.977	.986	.988	.992	1.027	.905	1.001	.989

NOTE.—Using the individual funds' portfolio weights in December of each year, we computed the proportion of funds with an above-average weight in a given asset class that continue to have an above-average weight 12 months later. Time-series averages of these transition probability estimates are reported as year-on-year stayer probabilities in part I of panel A. Their standard errors are also based on the time series of these transition probabilities. Part II of panel A reports the corresponding transition probability estimates for the event that a fund that initially has an above-average portfolio weight in a given asset class continues to have an above-average weight in this asset class at the end of the sample. Panel B reports the cross-sectional distribution of first-order autocorrelation coefficients for the differential between individual funds' portfolio weights in a given asset class and the average portfolio weight in that asset. These coefficients were computed using the time series of portfolio weights.

initial to final relative weight but without standard errors since there is only one time-series data point per fund. The point estimates are also consistent with slow mean reversion, 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 B of table 5 provides further evidence of slow mean reversion by reporting results from a regression of $\omega_{ijt} - \omega_{jt}$ on a constant and the lagged dependent variable. The slope coefficients above the 50th percentile range from 0.90 to unity for all asset classes except cash, which has a median coefficient of 0.78. Similarly, the *t*-statistics (for the null hypothesis that portfolio weights follow a random walk) have rejection rates of around 5% at the 5% critical level, except for domestic and international equities and cash, which had rejection rates of 14%, 11%, and 30%, respectively.

Our analysis so far appears to indicate slow mean reversion by individual funds toward a commonly changing strategic asset allocation but with random and, in the case of some funds, quite substantial short-term deviations from this longer-term process. However, the story remains incomplete because of the absence of information on pension fund liabilities. This 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, say, the elimination of pension fund surpluses required by the 1986 Finance Act or the increasing indexation of liabilities prompted in large measure by the 1985 Social Security Act (see Blake 1995).

V. Active and Passive Management Return Decompositions

We use the simple decomposition proposed by Brinson et al. (1986) to separate portfolio returns into components arising from active and passive management. Suppose there are M asset classes, and let ω_{njt} be the "normal" or strategic asset allocation of a fund in the jth asset class at time t, ω_{ajt} the actual portfolio weight, r_{njt} the "normal" portfolio return, and r_{ajt} the actual portfolio return. Then, as an arithmetic identity,

$$\sum_{j=1}^{M} \omega_{ajt} r_{ajt} \equiv \sum_{j=1}^{M} \omega_{njt} r_{njt} + \sum_{j=1}^{M} \omega_{njt}(r_{ajt} - r_{njt}) + \sum_{j=1}^{M} (\omega_{ajt} - \omega_{njt}) r_{njt} + \sum_{j=1}^{M} (\omega_{ajt} - \omega_{njt})(r_{ajt} - r_{njt}), \quad (9)$$

or total return $\equiv$ normal return + return from security selection + return from market timing + residual return. This is a useful decomposition if both the residual term is small compared with the other components (since it represents the component of returns that is not attributable to either timing or selectivity)[17] and we have good measures of ‘‘normal’’ portfolio returns and weights. In fact, the residual return in our sample proved to be small relative to the normal return but of the same order of magnitude as the selectivity return. Natural measures of normal portfolio returns are the various external or peer-group benchmark indices.

One reasonable concern about the interpretation of the security selection components is that it represents only performance evaluation relative to a benchmark with an implicit beta of unity. To be sure, although relative performance evaluation is the norm in the United Kingdom, this practice might conceal more substantial cross-sectional variations in risk-adjusted returns relative to alternative benchmarks. However, it turns out that this is not the case: we found that risk adjustment using single or multiple indices with both time-invariant and time-varying betas across asset classes changes the location of the cross-sectional distribution of mean raw returns but leaves its shape virtually unchanged.[18] Stated differently, there was near perfect correlation between average total returns and a variety of unconditional and conditional Jensen (1972) measures across asset classes and for the overall portfolio of each fund.[19]

The choice of normal portfolio weights is more problematic. Genuine performance measures should reflect investors’ ex ante information on future asset returns. However, we only observe actual portfolio weights, and these reflect realized returns. So information on ex post returns and portfolio weights will permit only noisy performance measurement. In the absence of any information on the funds’ asset-liability modeling exercises that might enable us to draw inferences about their associated strategic asset allocations, we were reduced to experimenting with a few simple, empirically plausible models. Accordingly, we take care to note the possible biases in performance measures engendered in samples such as ours that possess a relatively small time-series dimension.

17. The ambiguity can be eliminated by allocating the residual return to one of the other components. For example, Bodie, Kane, and Marcus (1995) add the residual return to the return from security selection.

18. For more details, see Blake, Lehmann, and Timmermann (1998).

19. This can be explained by the tendency of betas to cluster around unity. For example, the semi-interquartile ranges of the beta estimates from single-index Jensen regressions applied to the most important asset classes were 0.99–1.01 (U.K. equity), 0.80–0.92 (international equity), 1.02–1.15 (U.K. bonds), 0.92–1.03 (U.K. property), and 0.98–1.08 (total portfolio).

The first model, proposed by Brinson et al. (1986), takes the average portfolio allocation over the sample as the normal portfolio weights:

$$\omega_{njt} = \sum_{t=1}^{T} \omega_{ajt}/T \tag{10}$$

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 investment opportunities are stationary. However, this is an unattractive assumption in our case, since U.K. pension funds were not apparently 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 (9).

The systematic increase in equity exposure over the period is the most obvious nonstationarity in our data set. 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

$$\omega_{njt} = \omega_{aj1} + (t/T)(\omega_{ajT} - \omega_{aj1}). \tag{11}$$

Since $\sum_{j=1}^{M} (\omega_{ajT} - \omega_{aj1}) = 0$, this measure has the important property that the normal portfolio weights are confined to lie in the interval [0, 1] at each point in time. Benchmark portfolio weights increase (or decrease) linearly in time between the initial and terminal weights.[20]

Table 6 summarizes the aggregate evidence produced by these different normal portfolio weight models, while table 7 displays key fractiles of the cross-sectional distribution of the average returns to the normal, market-timing, and security-selection components of performance for

20. However, both sets of normal portfolio weights are sample dependent, inducing potential biases in this otherwise straightforward 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, thereby inducing some of this postulated selection ability to be attributed to the strategic asset allocation decision. Similarly, a good market timer need not confront an equal number of positive and negative signals over the sample, thereby 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 a market timing or security selection ability when they, in fact, have no such abilities. Again the tilt toward managers with the presumed ability is incorrectly classified as part of the strategic asset allocation, while the effect on the measured normal return depends on whether these managers happened to be lucky or unlucky over the sample period. In particular, funds that tilted toward U.K. equities based on an erroneous belief that their managers possessed superior performance ability experienced higher measured normal returns owing to the good performance of U.K. equities over our sample period.

TABLE 6 **Decomposition of U.K. Pension Funds' Returns from International Equity (Average Annual Percentages)**

	Normal Return	Security Selection	Market Timing	Residual Return	Total Return
A. Constant benchmark for normal portfolio weights, external benchmarks for normal returns:					
Mean return	12.305	.010	−.342	.061	12.034
t-value	2.29	.02	−2.16	1.00	2.29
B. Trended benchmark for normal portfolio weights, external benchmarks for normal returns:					
Mean return	12.262	.031	−.299	.041	12.304
t-value	2.30	.05	−1.64	.69	2.29
C. Constant benchmark for normal portfolio weights, peer-group benchmarks for normal returns:					
Mean return	11.989	.326	−.215	−.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:					
Mean return	11.970	.322	−.197	−.061	12.034
t-value	2.28	2.66	−1.11	−2.23	2.29

NOTE.—For each fund, the monthly returns were decomposed into returns from normal asset allocation, selection, timing, and a residual (eq. [9]). 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). External benchmark returns (panels A and B) are based on external indices described in Section III, while the peer-group benchmarks use the returns on the WM2000 indices (panels C and D). The results that assume a constant benchmark for the normal portfolio weights compute this as the simple sample average of the individual funds' time series of portfolio weights (eq. [10]). The results that assume a trended benchmark for the normal portfolio weights adjust these for a linear trend using the initial and terminal portfolio weights (eq. [11]).

each asset class, as well as the maximum and minimum values and their associated Bonferroni *p*-values.[21] The most noteworthy feature is the robustness of the results across models with very different dynamics and drifts. The constant mean and linear trend models each yield normal portfolio returns that are numerically close both on average (table 6) and fractile by fractile (table 7), despite both the substantial shift toward equities over the sample period and the considerable cross-sec-

21. The Bonferroni *p*-value bounds the marginal significance level of the largest *t*-statistic in absolute value with p_0 when its *p*-value is p_0/N, where N is the number of *t*-statistics examined simultaneously.

TABLE 7 Cross-Sectional Distribution of Return Components (Annual Percentages)

	Normal Return		Security Selection		Market Timing		R^2 of Normal Return		Portfolio Change Measure
	No Trend	Trended	No Trend	Trended	No Trend	Trended	No Trend	Trended	
A. Sorted on individual components:									
Minimum	10.71	10.86	−4.76	−4.76	−1.89	−1.12	.530	.512	−6.34
5%	11.74	11.62	−1.52	−1.60	−.78	−.73	.904	.899	−.34
10%	11.98	11.81	−1.07	−1.04	−.70	−.64	.938	.932	−.16
25%	12.19	12.11	−.56	−.49	−.51	−.49	.958	.954	−.03
50%	12.33	12.30	.00	.03	−.35	−.31	.976	.974	.06
75%	12.44	12.48	.56	.59	−.16	−.14	.982	.981	.13
90%	12.56	12.59	1.03	1.03	−.01	.04	.987	.986	.25
95%	12.65	12.65	1.56	1.43	.13	.13	.990	.988	.32
Maximum	13.13	13.35	4.73	4.87	.94	1.37	.995	.994	6.13
Bonferroni bounds:									
Minimum t-value	. . .	. . .	−2.64	−2.69	−4.69	−4.44	. . .	. . .	−2.16
p-value	. . .	. . .	1.000	1.000	.0004	.0014	. . .	. . .	1.000
Maximum t-value	. . .	. . .	4.20	3.72	2.39	2.05	. . .	. . .	3.11
p-value	. . .	. . .	.004	.031	1.000	1.000	. . .	. . .	.293

	Normal Return		Security Selection		Market Timing		Total Return	Portfolio Change Measure
	No Trend	Trended	No Trend	Trended	No Trend	Trended		
B. Sorted on total returns:								
Minimum	12.40	12.32	−4.76	−4.76	−.29	−.21	7.22	.08
5%	12.37	12.44	−1.49	−1.47	−.73	−.79	10.59	.09
10%	12.14	12.07	−.26	−.21	−1.15	−1.09	10.95	.21
25%	12.32	12.17	−.45	−.40	−.31	−.16	11.47	.14
50%	12.50	12.42	.13	−.02	−.70	−.62	12.05	−.17
75%	12.26	12.16	.65	.68	−.14	−.04	12.61	.02
90%	11.94	12.03	.74	.82	−.16	−.25	13.14	.30
95%	12.38	12.30	1.12	1.13	−.40	−.32	13.41	−.08
Maximum	12.20	12.05	4.73	4.87	−.04	.11	15.03	6.13

NOTE.—The cross-sectional distributions are based on external benchmarks and use the decompositions of returns into normal asset allocation, security selection, market timing, and residual components from equation (9). The columns labeled ‘‘no trend’’ compute the normal portfolio weights as the simple sample average of the time series of portfolio weights (eq. [10]), while the columns labeled ‘‘trended’’ allow for a linear trend in normal portfolio weights, computed from the initial and terminal portfolio weights using equation (11). R^2 of normal returns is computed from a regression of actual returns on normal returns that imposes a coefficient of unity on the normal return component. 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 1 month earlier. Panel A performs sorts on individual return components while the sorts in panel B are performed according to the funds’ total returns and hence are comparable across rows. Ellipses indicate that information could not be calculated for a given cell.

tional 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 find this consistency reassuring in the absence of a single compelling model for normal portfolio weights.

The cross-sectional variation in the ex post performance measures from these decompositions is also remarkably narrow (see table 7). The semi-interquartile ranges are only 25–40 basis points for the mean annualized normal and market-timing components of portfolio returns and a still 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. Clearly, there is very similar behavior among the bulk of these funds in these three dimensions of average performance.

The results reveal something about the abilities of the managers in question. Panels A and B of table 6 report the decomposition when the normal returns are set equal to the external benchmarks. The average normal return of about 12.31% per year exceeds the mean aggregate annual portfolio return of 12.03%. In contrast, U.K. pension funds earned an economically small negative return from active portfolio management on average, although there is some variation in the security-selection component. The mean annualized return from security selection at 1 basis point is insignificant at conventional levels, while that from market timing at −34 basis points is statistically significant. In addition, around half of the funds had negative selectivity estimates, and more than 80% had negative, albeit economically small, timing estimates.[22] Our aggregate findings are similar to those of Brinson et al. (1991), but differ from those of Brinson et al. (1986), who found a small negative return from selection on average.

Table 7 also reports the portfolio change measure suggested by Grinblatt and Titman (1989). This is calculated as $r_{jt}\,(\omega_{ijt} - \omega_{ijt-1})$, where ω_{ijt-1} is the strategic asset allocation prevailing one month earlier. It therefore measures the return to changing portfolio weights, so that any correlation between weight changes and returns over the previous month can be treated as arising from abnormal performance. Again,

22. The coefficients on squared excess benchmark returns from Treynor and Mazuy (1966)–style regressions 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). The cross section of these coefficients had a semi-interquartile range of −0.66–0.045. In common with similar regressions involving U.S. 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 positive market timing ability for all but perhaps a few managers.

the table shows the narrowness of the cross-sectional distribution of this performance measure.[23]

The results also demonstrate the importance of the strategic asset allocation decision. For our first two definitions of ‘‘normal’’ weights, we found that 96% of the total variation in monthly portfolio returns could be explained by the normal asset-class holdings across funds on average. In fact, normal asset-class holdings explained more than half of the variability in portfolio returns for the fund with the smallest contribution to return variability from this component. Brinson et al. (1986, p. 39) put the aggregate fraction of total variation attributable to the strategic asset allocation at 93.6%[24] and concluded that ‘‘investment policy [i.e., the strategic asset allocation] dominates investment strategy [market timing and security selection],’’ a finding that has led others, such as Bogle (1994, p. 235), to conclude that the ‘‘94% figure suggests that long-term fund investors might profit by concentrating more on the allocation of their investments between stock and bond funds and less on the question of which particular stock and bond fund to hold.’’ In other words, the practitioner literature has come to view the comparative statistical importance of strategic asset allocation performance as direct evidence of the central economic role of this decision.

This view is false, however. Ignoring any error in identifying actual strategic asset allocations, the domination of pension fund returns by the returns to passive management actually reflects the absence of extensive attempts at active management by U.K. fund managers. That is, the large coefficient of variation that we find describes the behavior of portfolio managers, not the economic role of asset allocation decisions. Similarly, we would be unable to conclude that active management decisions were economically more important just because we found a market in which the active management component dominated the time-series and cross-sectional variations in average portfolio returns. Rather, we should ask whether active management earned positive expected excess risk-adjusted returns, which is a somewhat different question.

Now the evidence in panel A of table 7 and in Blake et al. (1998) suggests the absence of abnormal performance by all but perhaps a few of the funds. Nevertheless, it is interesting to ascertain how much of the cross-sectional variation in average raw returns is attributable to the various components. Panel B of table 7 provides one simple answer to that question by displaying the average returns to the normal, market-timing, and security-selection components at each given fractile of average total return, with the funds having been sorted on the basis of average total returns over the sample. There appears to be no relation

23. These findings are very robust to using a horizon longer than a single month.
24. This figure is a little higher than the 91.5% reported by Brinson et al. (1991).

between average total return and the portfolio change measure except for the most extreme performers. There is an apparent, if modest, inverse relation between average total return and that of the normal asset allocation and a weak positive one between average total return and the market-timing component. However, there is a strong relation between average total return and the security selection component: the unconditional cross-sectional distribution of the average reward to security selection, reported in panel A, is numerically close to the comparable distribution conditioned on the average total return, reported in panel B. That is, cross-sectional variation in average total return is dominated by the ex post average reward to security selection, a component of active management to which, according to theory, there is little, if any, ex ante abnormal reward.

Panels C and D of table 6 report the changes to the decomposition when the peer-group indices replace the external benchmarks in the definition of ''normal'' returns. The mean return from security selection, at an economically modest 0.32% per year, is now positive and significant, while the mean 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. For reasons discussed in Blake et al. (1998), this improvement in measured performance arising from the shift from external to peer-group benchmarks suggests that relative performance evaluation, which is standard in the U.K. pension fund industry, plays an important role in the maintenance of money manager reputations and, indeed, in the retention of investment mandates (our sample of fund managers had retained their mandates for much longer than the average U.K. fund manager).[25]

In any event, our main finding is that the strategic asset allocation, however measured, accounts for most of the ex post variation of U.K. pension funds' returns, while the security-selection component dominates the cross-sectional variation in their average total returns. Even so, 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 of funds have negative market-timing estimates, however measured. A randomly selected pension fund would have been better served by applying its strategic asset allocation

25. Lakonishok et al. (1992), using a procedure related to the portfolio change measure, found that active fund management impaired performance for pension fund managers aggregated by investment style. Coggin et al. (1993) found positive and significant stock selection skills and negative timing ability among their sample of U.S. equity pension fund managers.

to passively managed index funds.[26] Finally, our sample of fund managers have retained the loyalty of their clients for much longer than the average manager; any survivor bias would shift the distribution to the left.

VI. Conclusion

From the outset, several aspects of the experimental design implicit in our U.K. pension fund data struck us as critical for understanding performance evaluation in this universe. Chief among these are the legal and economic environments in which the funds operate. In our view, the empirical regularities we observe in these data are a consequence of the incentives arising from the industrial organization and regulatory environment facing the U.K. pension fund industry.

The structure of the industry is similar to that associated with producers of a commodity product for reasons noted by Lakonishok et al. (1992). The industry is dominated by five large money management firms concerned with maintaining their reputations for service and reliable, if similar and unspectacular, performance, the structure one would expect if there were no ex ante differences in performance ability. In contrast, one would expect substantial dispersion in market shares and performance if there were active managers with differing degrees of management skill, as is observed in the United States. Similarly, as industrial organization reasoning suggests, these large firms use their reputations to acquire new clients and retain old ones, as opposed to increasing their fees, and are systematically successful at doing so.

These observations about underlying incentives appear to account for many of the robust regularities we report. Managers had the incentive to produce similar results, and the empirical evidence suggests they did so, despite the reasonably wide cross-sectional variation in asset allocation dynamics. That is, we found surprisingly little cross-sectional variation in average ex post returns to strategic asset allocation, market timing, and security selection. Long-run asset allocations, however modeled, account for the bulk of the time-series variation in returns, providing more robust empirical support for the quote at the beginning of this article. However, we believe that this finding reflects more on managerial behavior (i.e., the absence of extensive attempts at active management) than on the economic role of asset allocation decisions. What cross-sectional variation we found is dominated by the security-selection component and appears to reflect random ex post returns to a zero expected-excess-return activity.

26. By the end of the sample, 16% by value of U.K. equity holdings were invested passively in index funds.

Our results are compatible with the notion that the rule of the game (i.e., that pension plan sponsors are buying what is essentially a commodity product) is imperfectly understood or acted upon by trustees in at least one dimension. Most funds would have been better off with their strategic asset allocations placed in passive index funds, and yet they purchased active management services that resulted in the uneven, if still modest, level of cross-sectional variation in security selection and the more uniformly poor market-timing performance. Perhaps there is an agency problem of the sort discussed by Lakonishok et al. (1992) in that plan sponsors or corporate treasury departments can justify their empires only if they engage in active management to some extent. In any event, some such agency problems seem to be important for understanding the industrial organization of the U.K. pension fund management industry.

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