Journal of Business Finance & Accounting, 24(7) & (8), September 1997, 0306-686X

AN EXAMINATION OF THE LONG RUN PERFORMANCE OF UK ACQUIRING FIRMS

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INTRODUCTION

Recent work in the US has shown that acquisitions are wealth-reducing events for the acquirer (Agrawal, Jaffe and Mandelker, 1992). The research of Agrawal et al. (1992) shows that the earlier findings of Franks, Harris and Titman (1991), which claimed that post-merger returns were not significantly negative, are specific to the time period of that study (1975–84). In the UK, Limmack (1991) and Higson and Elliot (1993) show that long-run bidder returns are significantly negative. Perhaps the greatest puzzle of all from these papers, given that one might expect markets to 'learn' or become more efficient through time, is the suggestion that takeovers in more recent years have produced significant negative returns for acquiring company shareholders. These results make the study of later acquisitions especially interesting. Fundamentally, the research question is whether these acquisitions really led to negative abnormal returns, or whether these results are the result of some type of specification error. To address this question, a comprehensive data set of large UK takeovers over the period 1984–1992 is analysed, with the focus being upon the long run returns to bidding companies.

The obvious difficulty with an 'event study' research model is that of benchmarking. First, there is a dispute about the appropriate asset pricing model to be used. The market model, market-adjusted returns (or 'zero-one' model), and size-decile adjusted returns have all been used in UK studies, whilst in the US the Dimson-Marsh (1986) and multi-factor models have been employed in more recent studies. There is considerable evidence that the choice of benchmark can have an important impact on the scale of abnormal returns from event studies (e.g. Dimson and Marsh, 1986; Agrawal et al., 1992; Gregory, Matatko, Tonks and Purkis, 1994; Kennedy and Limmack, 1996; and Fama and French, 1996). Given the unresolved dispute over the

* The author is from the School of Business & Economics, University of Exeter. He would like to thank INQUIRE for their support of this project, and Liz Webber of the School of Business & Economics, University of Exeter, for her excellent research assistance. This paper has benefited considerably from the comments of participants at the *JBFA* Capital Markets Conference at Manchester Business School in September 1996 and the INQUIRE Conference at Bath in September 1996, and especially from the comments of Robin Limmack who acted as discussant at the *JBFA* Conference. The usual disclaimer applies concerning errors and omissions.

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appropriate form of asset pricing model it seems important to examine the sensitivity of the conclusions of research into acquiring company returns to different specifications of 'normal' returns.

This study addresses this issue by employing models of 'normal' return including the basic capital asset pricing model (CAPM), the Dimson and Marsh (1986) risk and size adjusted model, two CAPM-type models extended for 'size effects', the simple size-adjusted model of Dimson and Marsh (1986), and the Fama and French (1996) three-factor model. All of these models have appeared in the recent literature. Whilst five of these benchmarks assume that some form of CAPM applies, a sixth assumes that returns are driven only by firm size. Although the main focus is upon abnormal returns in event time, returns accumulated on portfolios formed in calendar time (as in Loughran and Ritter, 1995) are also reported. Additionally, this study looks at two alternative mechanisms for assessing the impact of an acquisition. First, as in Agrawal et al. (1992), the Ibbotson (1975) Returns Across Time and Securities (RATS) method adjusted for firm size is employed. Second, a more general cross-sectional regression is estimated for all listed companies in each year, with appropriate dummy variables included for firms which experience takeover events. This should be a very general model, free from contamination by any mis-specification of the 'true' asset pricing process, unless there is some priced risk factor which differs systematically between acquiring and non-acquiring firms. The conclusion from *all* of these models is unambiguous; takeovers were, on average, wealth reducing events for acquiring companies.

Related to the choice of benchmark is the issue of whether to use pre or post event data to estimate the necessary model parameters. Evidence is presented in this paper which shows that there are fundamental shifts in model parameters between pre and post takeover periods. Whilst there is no significant change in beta, alpha estimates move from being significantly positive in the period before a takeover, to being significantly negative post takeover. Furthermore, as may be expected given the implied change in firm size which results from a takeover, exposure to 'size effects' can be significantly less in the post takeover period. In the Fama-French three-factor model, there is an increase in exposure to 'book to market' effects post takeover. This evidence on parameter value changes suggests that the use of pre-take-over estimates may bias the results from models which use the standard 'cumulative average abnormal return' (CAAR) method.[1] Thus this study follows Agrawal et al. (1992) in using post-event data to estimate the appropriate model parameters.

One matter of serious concern for any study of the long term effects of a particular event is the recent evidence that long-horizon abnormal security returns can be seriously mis-specified (Kothari and Warner, 1997). However, the focus on 'buy and hold' (abnormal performance indicator) abnormal returns adopted here is likely, according to the results of the Kothari and Warner (1997) simulations, to under-state the significance of longer term

negative abnormal performance and to over-state the significance of longer term positive abnormal performance (Kothari and Warner, 1997, Table 3). Returns accumulated using the CAR approach, although less skewed, can over-reject the null hypothesis over long horizon windows for both negative and positive abnormal performance (Kothari and Warner, 1997, Table 4 and Table 1). Given that the central finding of this study is that acquiring firms under-perform, and that this conclusion is obtained using 'buy and hold' returns, the consequence is that the significance of the results is likely to be under-stated rather than over-stated. Given the simulation results reported in Kothari and Warner (1997), it is unlikely that the magnitude of the results obtained from a CAR approach, also reported in this paper, can be explained away by specification errors.

The rest of this paper is organised as follows. The next section summarises the more recent empirical evidence, with an emphasis on the UK. The third section explains the data and the research method used. The fourth section presents the results and the final section draws conclusions from the study.

PREVIOUS EVIDENCE

The first major study of UK takeovers by Firth (1980) used a market model with parameters estimated using pre-event data and revealed significant negative CAARs of −6.3% for the announcement month. Using a similar model and a larger sample of 1048 acquirers covering the period 1955–85 Franks and Harris (1989, p.239) showed significant negative returns of −12.6% in the 24 months following completion when the market model is used, but a significant +4.5% when the CAPM is used as the benchmark. Franks and Harris (1989, p.245) attribute this difference to the average bidder alpha (estimated pre-takeover) of 0.95% per month. Limmack (1991) uses three benchmarks to compute abnormal returns: a conventional market model (where Limmack notes the possible problems resulting from significantly positive alpha values in the pre-bid period for acquirers); a market model using London Business School (LBS) beta and alpha values; and a 'zero-one' model. Results in the paper relating to the announcement period are reported for the second of these models only, Limmack (1991, p.247) noting that the choice of models has little impact on the results. For successfully completed bids, the announcement period returns are an insignificant −0.2%. The longer-term returns are a significant −4.67% for the bid month to 24 months following completion using the LBS beta model, with the figures for the market and 'zero-one' models being a significant −14.96% and −7.43% respectively.

Sudarsanam, Holl and Salami (1996) study announcement period returns associated with 429 UK bidders over the period 1980–90. Overall, they find

significant negative CARs of −4.04% over the period −20 to +40 days around the bid announcement date. A particularly interesting aspect of this paper is that it seeks to explain these returns by regressing the CARs on variables proxying for synergy, ownership and control. They find that the intercept is then a significant −0.54%, and that relative efficiency, and the proportion of large equity holders are significantly negatively associated with returns, whilst complimentary fits between liquid resources and growth opportunities, relatively smaller targets, hostile bids and cash bids all exhibit significantly positive coefficients.

An alternative approach to the estimation of the wealth-effects of UK takeovers is provided by Manson, Stark and Thomas (1994). They analyse the operating cash flow gains following takeover, and compare this to the sum of the acquirer and target companies' operating cash flow pre-takeover. The interesting feature of this research is that the cash flow comparison is made net of estimated industry average cash flows, and is based on the method used in Healy, Palepu and Ruback (1992). Whilst this method has the disadvantage of insuring the performance benchmark against a poor choice of target industry (Manson et al., p. 9, fn. 2), the advantage is an estimate of performance against that which it would have been in the absence of a takeover (Manson et al., p. 6). Based on a sample of 38 takeovers completed between January 1985 and July 1987, they conclude that operational cash flow gains result from takeovers, and that these gains are related to the market's assessment, at the time of takeover, of these gains.

Indirect evidence on the wealth effects of takeovers paid for by shares is provided by Cooke, Gregory and Pearson (1994). This paper investigates the Larson-Gonedes (1969) exchange ratio model in the UK. This model is specified in terms of post-takeover PE ratios, and essentially involves testing the change in wealth of acquirer and acquired company shareholders resulting from the actual exchange ratio used in the merger. Results were presented on the basis of unadjusted, market adjusted and beta adjusted prices; with the latter categorisation, out of a sample of 90 takeovers, 43 resulted in wealth gains for both shareholder groups one month after completion of the merger (44 at time of completion), with a further five resulting in wealth gains for acquiring company shareholders (with acquired company shareholders experiencing wealth losses). In 31 cases acquired company shareholders gain but acquiring shareholders lose out, whilst in 11 cases both shareholder groups lose out; thus acquiring shareholders experience wealth gains in just over 50% of cases, whilst target company shareholders gain in 74 out of 90 cases.

One weakness with the above studies of UK takeovers is that they do not control for size effects. Although this is unlikely to be of major concern in the bid period announcement study of Sudarsanam et al. (1996), size effects will be of importance in any study of long term returns. Whilst it appears that beta does have a role in explaining returns, so does firm size (Kothari, Shanken and

Sloan, 1995; and Kim, 1995). Furthermore, this size effect is time-varying, so that the normal market-model estimate of alpha cannot be relied upon to adequately capture size effects (Dimson and Marsh, 1986, p.137). Size effects have been taken into account in empirical studies in a variety of ways. Most simply, Dimson and Marsh (1986) used size decile control portfolios, where each company is assigned a decile membership based upon its market capitalisation at the beginning of each year. A second method used by Dimson and Marsh (1986) is the risk and size control model, which takes into account both firm and decile betas. Finally, more recent studies (Fama, Booth and Sinquefield, 1993; Loughran and Ritter, 1995; and Fama and French, 1996) have used a multi-factor benchmark approach. In particular, Fama and French (1996) suggest that many apparent anomalies in 'efficient markets' studies can be explained by the use of a three-factor model, where the factors are the excess returns on the market, the difference in returns between companies with high book-to-market (BMV) and low BMV ratios, and the difference in returns between large and small companies. Fama and French (1992 and 1996) claim that the 'book-to-market' effect proxies for financial distress risk. However, differences in returns between high and low BMV firms can also reflect differences in growth opportunities (Fama and French, 1995).

Higson and Elliott (1993) and Kennedy and Limmack (1996) are the only UK studies to consider size effects, and in both cases this is done by using the simple Dimson and Marsh (1986) size-decile control method; the results obtained by a 'zero-one' (or 'market adjusted) model are also reported in both cases. To date, no event studies have been undertaken in the UK using the three-factor benchmark. The Higson and Elliott study covers 726 takeovers between 1974 and 1990 and aggregates returns as the product of monthly returns (PAR). On the size-controlled basis, announcement period returns are a significant −4.4% (−3.90% on a 'zero-one' basis), whilst for the 24 months following completion the returns are an insignificant −10.6% (−10.1% on a 'zero-one' basis). Although the abnormal returns are insignificant, the percentages of positive PARs are significantly less than 50% for both models. Higson and Elliott also show that friendly bids produce much lower returns than hostile bids, and that returns are lower for the period 1983–90 than for the period 1974–82. Kennedy and Limmack (1996, Table 3) show that overall size adjusted returns are negative on a size-adjusted basis (using APIs, which do not imply portfolio re-balancing), with bidder excess returns being a significant −4.92% for the period 12 to 24 months post-bid. Returns pre-bid are significantly positive, whilst those for the eleven months post-bid are not significantly different from zero.

In the US, recent studies which employ size controls are Franks, Harris and Titman (1991) and Agrawal, Jaffe and Mandelker (1992). Franks et al. (1991) report results using a number of benchmarks, but emphasise the results from an eight factor benchmark.[2] The performance measure for acquiring firms is then an adjusted Jensen benchmark (intercept term) of excess returns

regressed on the benchmark for up to 36 months post takeover (18 months being the minimum cut-off for inclusion). Overall, the results of a sample of 346 acquisitions over the period 1975–84 are that the average monthly excess return is −0.11%, which is not significantly different from zero. Although all equity bidders perform less well than cash bidders, unopposed bidders perform less well than opposed bidders, and uncontested bidders perform less well than contested bidders, none of these results is significant at the 5% level. However, Franks et al. (1991, p.95) note that 'the differences remain large enough in economic terms (4–5% per year) to warrant further attention'.

The conclusions of Franks et al. (1991) are contested by Agrawal et al. (1992), who show, using both CAARs derived from the Dimson and Marsh (1986) risk and size return method and the Ibbotson (1975) RATS method, that firms acquiring through mergers (typically financed by equity) have significant negative post acquisition returns, whilst those acquiring by the tender method (typically financed by cash) are not significantly different from zero. Furthermore, Agrawal et al. show that the results of Franks et al. (1991) are explained by the particular time period studied and the aggregation of tender offers and mergers. Agrawal et al. show that for the merger sample as a whole, the CAARs are −4.94% after 24 months, −7.38% after 36 months and increasing to −10.26% after 60 months. The sub-periods 1955–9 and 1980–7 show the worst long run returns of any of the sub-periods investigated, whilst non-conglomerate mergers exhibit worse performance than conglomerate mergers.

DATA AND METHOD

The data set used in this study is a comprehensive list of all successful UK domestic takeovers of listed companies by UK plcs with a bid value greater than £10 million for the period 1984–1992. This cut-off, whilst somewhat arbitrary, was chosen to avoid the problem of 'noise' which could occur when firms acquire very small companies; it can also be argued that larger takeovers are of greater economic significance and therefore more worthy of attention. The source of data of takeovers and dates is the *Acquisitions Monthly AMDATA Database*. The *Stock Exchange Official Yearbooks* were used to identify acquirers which had changed names.[3] The only additional requirement is that LBS Share Price Database returns are available for the acquirer for at least 12 months post-acquisition[4] (except in the case of the size-control benchmark described below), and that the market capitalisation is available for the beginning of the year in which a takeover is either announced or completed and for each subsequent year in which returns are accumulated if the benchmark used is either the Dimson-Marsh (1986) size and risk control model or the basic size control model described below. The maximum number of acquisitions is 452 (for models which do not require market capitalisations)

Table 1

Summary Statistics on Takeovers, 1984–92

Decile of Acquirer	*Number of Acquisitions*	*1984*		*1992*	
		Maximum, £m	*Minimum, £m*	*Maximum, £m*	*Minimum, £m*
Decile 1	158	7421.0	145.9	25638.1	481.5
Decile 2	77	141.2	59.7	462.0	160.6
Decile 3	58	57.8	29.0	159.3	75.9
Decile 4	40	28.8	15.4	75.9	38.4
Decile 5	31	15.4	9.8	38.1	23.1
Decile 6	15	9.8	6.6	23.1	15.4
Decile 7	14	6.6	4.2	15.3	9.4
Decile 8	11	4.2	3.0	9.4	5.7
Decile 9	10	3.0	1.3	5.6	2.9
Decile 10	6	1.3	0.1	2.9	0.1
Total	420				
Overall Summary Figures:	**£ million**				
Maximum size	3700.0				
Minimum size	10.1				
Mean size	137.6				
Median size	33.4				

Notes:
Column 2 shows the number of acquisitions by decile of acquiring company at time of bid announcement, for all those companies which have market capitalisations available on the LBS database for 1 January (420 bids out of the full sample of 452 bids). Columns 3 and 4, and 5 and 6, show the cut-off market values for decline groupings for *all* companies on the LBS data-tapes for 1984 (the first year of the study) and 1992 (the last year of the study). The overall summary figures show the maximum, minimum, mean and median sizes by bid value of the takeovers included in the sample.

and a minimum of 403 (for the Dimson-Marsh, 1986, size and risk control model). The aim of this research is to study abnormal returns in the two years following takeover. As some companies fail to survive for the full 24 months post-takeover,[5] the number of acquiring firms remaining in the sample falls slightly over time. Summary statistics for the takeovers included in the sample are presented in Table 1. As this table shows, the majority of acquirers are concentrated in the top two deciles by market capitalisation, with 157 out of the 398 acquirers for which market capitalisation are available being in the top decile. By contrast, some 75% to 80% of the market value of all UK companies is concentrated in the top decile,[6] thus the portfolio of acquiring firms is biased towards smaller companies. Table 1 also shows that the distribution of the value of the acquisitions themselves is biased towards smaller deals; whilst the average successful bid has a value of £140.3m, the

median value is only £33.6m. The sample maximum and minimum values are £3.7bn and £10.1m respectively. The table also shows the market value cut-offs used to establish the decile portfolios, described below, of all companies.

The first method used to investigate the sample is the standard event-study method. For each benchmark, the abnormal return is measured by estimating the parameters for up to 36 months post takeover completion (excluding the month of completion);[7] the minimum criterion for inclusion was 12 monthly observations post-completion. However, in order to minimise problems caused by survivorship bias, a further analysis was undertaken which, in the case of all models except (3) below, involved substituting both mean values of parameters across sample companies and also substituting beta values of zero;[8] in both cases, the effect on the overall results was minimal. To avoid any downward bias in returns caused by Jensen's inequality when averaging returns across portfolios, raw returns are used throughout this paper (as in Franks et al., 1991 and Agrawal et al., 1992).[9]

The following abnormal returns in respect of each of the six benchmarks is estimated:

Model 1: CAPM:

$$\varepsilon_{it}^{c} = R_{it} - [R_{ft} + \hat{\beta}_i^{c}(R_{mt} - R_{ft})] \tag{1}$$

where:

R_{it} = return on company i in event month t;
R_{mt} = return on market in event month t;
R_{ft} = risk-free (treasury bill) return in event month t;
$\hat{\beta}_i^{c}$ = CAPM beta of company i, estimated by OLS regression on up to 36 months post completion of takeover.

Model 2: Dimson-Marsh risk and size adjusted model (DM):

$$\varepsilon_{it}^{dm} = R_{it} - [R_{st} + (\hat{\beta}_i^{c} - \hat{\beta}_s)(R_{mt} - R_{ft})], \tag{2}$$

where:

R_{st} = return on the size control portfolio in event month t;
$\hat{\beta}_s$ = CAPM beta of size control portfolio.

In this model, the control portfolios are equally weighted average returns on a portfolio of all firms in the decile to which firm i belongs. Decile betas are estimated by an OLS regression of decile excess returns on market excess returns for the 36 months following completion of the takeover.[10]

Model 3: Simple size control portfolio (SS):

$$\varepsilon_{it}^{ss} = R_{it} - R_{st} \tag{3}$$

Model 4: Multi-index model using equally-weighted smaller decile minus large decile returns (SML):

$$\varepsilon_{it}^{sml} = R_{it} - [R_{ft} + \hat{\beta}_i^{sml}(R_{mt} - R_{ft}) + \hat{\gamma}_i^{sml}(R_{610t} - R_{1t})], \tag{4}$$

where:

R_{610t} = the return on an equally weighted portfolio of smaller (deciles 6 to 10) companies in event month t;
R_{1t} = the return on an equally weighted portfolio of largest (decile 1) companies in event month t.

Model 5: Value weighted multi-index model using the Hoare-Govett Index as the measure of smaller company performance:

$$\varepsilon_{it}^{hg} = R_{it} - [R_{ft} + \hat{\beta}_i^{hg}(R_{mt} - R_{ft}) + \hat{\gamma}_i^{hg}(R_{ht} - R_{mt})], \tag{5}$$

where:

R_{ht} = the return on the Hoare-Govett Smaller Companies index in event month t.

The motivation for using the Hoare-Govett Smaller Companies Index (HGSCI) is that this is a value-weighted index of the bottom 80% of companies by market capitalisation; this model has also been used in the analysis of UK unit trust performance (Gregory, Matatko and Luther, 1997). By contrast, Model 4 uses an equally weighted index of the bottom 50% of firms by market capitalisation. The 'size effect' captured by Model 4 is larger, and hence the average 'gamma' coefficients are smaller. It would be possible to estimate a model that used the differences between the extreme decile portfolios (decile 1 being the largest companies and decile 10 the smallest). Note, however, that decile 10 companies in the UK are somewhat idiosyncratic; the modal market capitalisation of these companies is only £1m (to the nearest £1m) in 1992, for example, and these firms are likely to be affected by thin trading problems. Furthermore, because of the takeover size cut-off used in this study, few decile 10 companies are included in the sample.[11]

Finally, Fama and French (1996) report that many efficient markets 'anomalies' can be explained by taking into account size and book-to-market effects through the use of a three factor benchmark. Under this model, abnormal returns are calculated as follows:

Model 6: Fama and French (1996) Value-weighted three factor model:

$$\varepsilon_{it}^{ff} = R_{it} - [R_{ft} + \hat{\beta}_i^{ff}(R_{mt} - R_{ft}) + \hat{\gamma}_i^{ff}(\text{SMB}) + \hat{\delta}_i^{ff}(\text{HML})] \tag{6}$$

where:

SMB = the value-weighted return on small firms minus the value-weighted return on large firms;

HML = the value-weighted return on high BMV firms minus the value-weighted return on low BMV firms.

A difficulty encountered when trying to estimate book to market value (BMV) portfolio returns in the UK is the lack of availability of BMV figures for many firms on Datastream. For every company which has returns and market capitalisation data available on the LBS tape, the SEDOL number was extracted and used to search for BMV ratios on Datastream in each year from 1980 to 1994. Unfortunately, of the firms on the LBS tape for which market capitalisations were available for January 1990, no fewer than 1,026 firms did not have BMV ratios available on Datastream, leaving a final sample of only 960 firms with which to calculate portfolio returns for 1980. This situation improves through time, so that by 1990, for example, 505 firms do not have BMV ratios and 1,650 firms are used to calculate portfolio returns. This suggests that survivorship bias may be a problem when the three factor model is applied to UK data. Whilst the full Fama and French (1996) three factor model has not been tested in the UK, Strong and Xu (1997) show that BMV effects are important in explaining the cross-section of UK share returns. Given the prominence of the three-factor model in the recent US literature, it used to estimate abnormal returns here, although given the *caveats* above, perhaps the results should be treated with some caution.[12]

The SMB and HML portfolios in (6) above are formed, as in Fama and French (1996) by sorting all companies in each year by BMV and market capitalisations; only companies for which both figures are available are included in the portfolios. Again as in Fama and French (1996), value weighted returns are calculated for the bottom 30% of companies by market capitalisation and the top 30% of companies by market capitalisation, and the top 50% of companies by BMV and the bottom 50% of companies by BMV. The differences between these value-weighted returns form the 'small minus big' (SMB) and 'high minus low BMV' (HML) returns.

For each of the benchmarks described by (1) to (6) above, the CAAR, as used in the Agrawal et al. (1992) study, and the Abnormal Performance Index, as used in Limmack (1991) are reported. The *t*-test statistics are the Brown and Warner (1980, p.251–2) Crude Dependence Adjustment test for the CAARs, and a simple cross-sectional *t*-test (Strong, 1992, p.545) for the API. In the case of those models where the parameters are directly estimated from a single regression (Models 1, 4 ,5 and 6) a direct estimate of the abnormal performance post-takeover is the estimated alpha from the regression, as in Franks et al. (1991). Given the time period over which the regression is estimated, this necessarily precludes any estimate of the returns during the announcement period and during the month of completion. The *t*-test statistics in this case are formed from the standard error of the alpha coefficient from the regression. Since time periods used in the estimation can vary, and specific variances differ across firms, it is appropriate to use the

estimated standard errors of the regressions estimated to obtain the parameters in (1, 4, 5 and 6) above. Under the null hypothesis that $\hat{\alpha}_i = 0$, $i = 1, \mathcal{N}$:

$$t = [\sum(\hat{\alpha}_i / SE(\hat{\alpha}_i))]/\sqrt{N} \tag{7}$$

will have an approximate (asymptotic) normal distribution (where N is the number of acquirers) under the usual assumptions about the error terms. Similar test statistics can be used to test the cross-sectional significance of the β_i and γ_i coefficients.

For the purposes of comparison, these regressions can be run also on pre-announcement returns. Conn and Connell (1990) provide evidence that CARs can be highly sensitive to whether pre or post-event periods are used to estimate market model parameters. Pre-announcement regressions for Models (1), (4), (5) and (6) are estimated using data for up to the 36 months ending six months before announcement; as before, the minimum number of observations used is 12 months. This also allows the pre-announcement and post-completion parameters to be compared. Similar to the t-test described for the mean coefficients above, assuming no correlation between pre-announcement and post-completion performance, a t-test for differences in the alpha coefficients is given by:

$$t = \sum_{1=1}^{i=\mathcal{N}} ([\hat{\alpha}_i^{post} - \hat{\alpha}_i^{pre}]/[\mathrm{var}_{\hat{\alpha}_i^{post}} + \mathrm{var}_{\hat{\alpha}_i^{pre}}]^{\frac{1}{2}})/\sqrt{N} \tag{8}$$

where:

α_i^{pre} = the estimated alpha pre-announcement;
α_i^{post} = the estimated alpha post-completion.

Similar tests can be formed for differences in the other coefficients.

None of these models are problem free, but the least problematic would appear to be Model 5. Model 1 suffers from the fact that no allowance is made for firm size effects although an advantage of a CAPM framework is that it avoids contamination caused by the tendency noted elsewhere in the literature (e.g. Franks and Harris, 1989; and Limmack, 1991) of acquiring companies to out-perform (have positive alphas) in the pre-acquisition period.[13] Thus using the market model can cause downward bias in the post-acquisition CARs. Model 2 and Model 3 have one weakness in common, in that they both (of necessity) define decile membership on the basis of beginning of year market capitalisation. Takeovers of over £10 million which involve new equity financing cause an increase in capitalisation and for companies in lower deciles the change can be sufficient to cause a company to be 'promoted' by more than one decile, especially in the case of smaller companies. Summary figures on decile membership cut-offs are given in Table 1. However, decile changes only take place at the beginning of every year so that if takeovers are

assumed to occur uniformly throughout calendar years on average, the decile membership group can potentially be incorrect for the six months following a takeover. Model 3 has the additional disadvantage that decile betas and firm betas are assumed to be identical.[14] As both Models 4 and 5 use a form of post-event extended CAPM incorporating a small-companies effect, they have the advantage of being compatible with recent papers (Kothari, Shanken and Sloan, 1995; and Kim, 1995) which show that beta risk and firm size are significant in explaining the cross-section of expected returns.

The importance and time-varying nature of size effects in the UK during the study period can be estimated by comparing the returns on the *Hoare-Govett Index* relative to the returns on the *FT All Share Index*. For the period 1984–92, the compound return on the *HGSCI* was 258.6% as opposed to 227.4% on the *FTASI*. However, the out-performance of the *HGSCI* varied from a maximum of +15.6% in 1993 to a minimum of −24.8% in 1989, being positive for the years 1984–88 and 1993–94, but negative from 1989–92. On the choice between Models 4 and 5, it is debatable as to whether the idiosyncrasies of the decile 10 companies referred to earlier have undue weight in Model 4. Fama and French (1996) also prefer a value-weighted model. In considering Model 6, it may be that book-to-market is an important issue in takeovers, either as a measure of growth opportunities or distress risk, or as an indicator of over or under valuation. In the context of target firms, Powell (1997) shows that targets with higher market-to-book ratios have a higher likelihood of being subject to a hostile takeover. However, as has already been noted, the lack of any evidence on the empirical validity of the three-factor model in the UK, together with the possible survivorship bias introduced by the availability of BMV figures on Datastream means that some caution is necessary when interpreting the results from Model 6. For this reason the main emphasis in the results section and the conclusions tends to be placed upon Model 5 rather than Models 4 or 6. Nonetheless, in general it is worth noting that all the models generally yield similar conclusions.

Given evidence that beta can be subject to changes over time (e.g. Coutts et al., 1995) it is appropriate to investigate the results which arise from a model which allows beta to vary during the returns window. To allow for this, the second approach used is the RATS method of Ibbotson (1975) and is estimated in the modified form used in Agrawal et al. (1992). The modification allows for size effects by subtracting the decile return from the realised return in each case:

$$R_{it} - R_{st} = \hat{\alpha}_t + \hat{\beta}_t(R_{mt} - R_{ft}) + e_{it}. \tag{9}$$

Whereas CARs derived from (1,2, 4, 5 and 6) above assume beta (and other coefficients) remain constant through time for each firm, the RATS model in the form of (9) implicitly assumes that the difference between decile and firm betas captured by β_t is constant across firms in any time period. The mean

RATS coefficients can be tested for significance by using a t-test similar to that described in (7).

A possible problem with a violation of the assumption of cross-sectional independence of abnormal returns occurs with models based on event time if firms have contemporaneous events in calendar time (Strong, 1992). This can either be avoided by using the method employed in Jaffe (1974) and Mandelker (1974), or by that used in Loughran and Ritter (1995); the latter approach is adopted here. Excess returns on a portfolio of all firms making an acquisition in month τ, or in the months $\tau-1$ to $\tau-24$ are regressed on the benchmarks used in (1), (4), (5) and (6) respectively. Thus, for example, the time-series regression for Model 1 is:

$$R_{p\tau} - R_{f\tau} = a + b(R_{m\tau} - R_{f\tau}) + e_\tau, \tag{10}$$

where:

$R_{p\tau}$ = the return on the portfolio of acquiring firms in calendar month τ.

If acquiring firms under-perform, then the intercepts in the time-series regressions should be significantly less than zero. Note that the time-series regressions will weight acquirers in low-transactions periods more heavily than acquirers in high-transactions periods, since (10) weights each of the 131 months[15] equally, irrespective of the number of firms in the portfolio. By contrast, the event-time approach weights each acquisition equally.

The CAAR, RATS and calendar time approaches assume that the asset pricing process is known. A more general model is simply to regress the excess returns on every security on the LBS database on dummy variables for acquiring firms and firm size, for each year of interest (1984–94):

$$R'_{it} = a + bA_{i\tau} + cY^c_{i\tau} + dY^p_{i\tau} + eS_i + fR'_{m\tau} + u_{i\tau}, \tag{11}$$

where:

R'_{it} = excess return on company i in calendar month τ;
R'_{mt} = excess return on the market in calendar month τ;
A_{it} = dummy variable equal to 1 if firm i announces an acquisition in calendar month τ;
Y^c_{it} = dummy variable equal to 1 if firm i announces or completes an acquisition in the 12 months prior to calendar month τ;
Y^p_{it} = dummy variable equal to 1 if firm i announces or completes an acquisition in the 13–24 months prior to calendar month τ;
S_i = log of market capitalisation at beginning of the year.

This regression is run separately for each of 11 years (note that for comparison to be made with the results from (1) to (6) and (8) above, no estimate of y^p_{it} can be made for 1984, no estimate of y^c_{it} can be made for 1994, and no estimate of A_{it} can be made for either 1993 or 1994). Results are shown for each year separately, but the average coefficients are also calculated and the overall

significance of the result can be assessed by a procedure similar to that described in (7) above.

RESULTS

Overall

The results from Models 1 to 6 above are shown in Table 2; CAARs and APIs are presented, along with the appropriate *t*-test statistics described above. CAARs and APIs are calculated for periods t_0 to t_{c+24} where t_0 is the month of announcement and t_c is the month of completion of the takeover. CAARs through event time are shown in Figure 1, whilst APIs are shown in Figure 2.[16] Overall, the results from all six models are consistent, with none showing significant abnormal returns for the month of announcement. For the announcement period (defined as from the month of announcement up to and including month of completion),[17] the multi-index SML Model 4 and the Fama-French three-factor Model 6 produce significant negative APIs and CAARs. All models show significant negative APIs and CAARs for the 24 months following completion of takeover. Thus the message from all these event-study models is clear; the long-run shareholder wealth effects of recent acquisitions in the UK have been, on average, significantly negative. However, the various models do lead to economically significant differences in the magnitude of the negative abnormal returns.

The absolute magnitudes of the APIs are generally less than those of CAARs for accumulation periods greater than six months. Although APIs may be more meaningful for assessing abnormal returns (Strong 1992, p.540) and do not involve portfolio re-balancing assumptions, CAARs are frequently used in the literature when non-logarithmic returns are employed (e.g. Agrawal et al.). CAARs also exhibit less skewness than APIs (Kothari and Warner, 1996). On a CAAR basis, significant negative returns at the 5% level are found for all models after six months. In all cases, t_0 to t_{c+24}, t_{c+1} to t_{c+24}, t_{c+1} to t_{c+12} and t_{c+13} to t_{c+24} month CAARs are significantly negative. However, the level of the CAARs differs between models, with Model 6 (three-factor) showing the lowest (i.e. biggest negative) CAARs after 12 months and 24 months post completion. Under this model, the CAARs to t_{c+24} are -18.01%, with -10.63% occurring in the 12 months following completion. The closest result to that of the three-factor model is produced by the CAPM (-9.21% after 12 months and -17.73% after 24 months). By contrast, the HG model (Model 5) produces a result reasonably close to (although with larger negative CAARs than) the SML model, with CAARs to t_{c+24} of -12.03%, the equivalent SML CAAR being -14.29%. The DM model (2) and the SS model (3) produce results which are similar (-12.52% compared to -11.82%) with the DM model giving consistently lower CAARs.[18] Thus

Table 2

Abnormal Performance Index and Cumulative Average Abnormal Returns for Alternative Benchmark Models

Time	*CAPM Model 1*		*Dimson-Marsh Model 2*		*Size Decile Model 3*		*Multi-index SML Model 4*		*Multi-index HG Model 5*		*Multi-index FF Model 6*	
	API %	*CAR* %	*API* %	*CAR* %	*API* %	*CAR* %	*API* %	*CAR* %	*API* %	*CAR* %	*API* %	*CAR* %
A_0	−0.30	−0.30	−0.50	−0.50	−0.37	−0.37	−0.71	−0.71*	−0.54	−0.54	−0.59	−0.59
	(−0.57)	(−0.60)	(−0.96)	(−1.32)	(−0.72)	(−0.91)	(−1.33)	(−1.69)	(−1.04)	(−1.35)	(−1.08)	(−1.34)
A_0 to C_0	−0.84	−0.67	−0.63	−0.46	−0.46	−0.30	−1.36†	1.23†	−0.99	−0.85	−1.48†	−1.35†
	(−1.29)	(−0.94)	(−0.98)	(−0.86)	(−0.72)	(−0.52)	(−2.08)	(2.08)	(−1.54)	(−1.50)	(−2.25)	(−2.16)
A_0 to C_6	−3.91‡	−3.78†	−3.77‡	−3.71‡	−3.39‡	−3.32‡	−4.18‡	−5.98‡	−2.98‡	−3.27‡	−4.82‡	−5.12‡
	(−3.56)	(−2.67)	(−3.45)	(−3.50)	(−3.12)	(−2.75)	(−3.67)	(−3.84)	(−2.65)	(−2.87)	(−4.38)	(−4.08)
A_0 to C_{12}	−7.46‡	−9.21‡	−6.16‡	−6.57‡	−5.83‡	−6.10‡	−5.98‡	−8.48‡	−4.78‡	−6.80‡	−8.29‡	−10.63‡
	(−5.01)	(−4.91)	(−4.46)	(−4.68)	(−4.22)	(−3.97)	(−3.61)	(−5.43)	(−3.08)	(−4.51)	(−5.42)	(−6.42)
A_0 to C_{24}	−12.44‡	−17.73‡	−11.25‡	−12.52‡	−11.03‡	−11.82‡	−9.18‡	−14.29‡	−8.15‡	−12.03‡	−12.22‡	−18.01‡
	(−6.37)	(−6.93)	(−6.16)	(−6.54)	(−6.06)	(−5.64)	(−4.29)	(−6.71)	(−4.02)	(−5.86)	(−5.99)	(−7.97)
Median and Nos. Positive and Negative, A_0 to C_{24}												
Median	−14.80		−13.82		−14.12		−12.95		−12.17		−14.14	
No. positive	166		137		142		170		175		161	
No. negative	286		266		276		282		277		291	
Z-test	−5.64‡		−6.43‡		−6.55‡		−5.27‡		−4.80‡		−6.11‡	
Sub-periods												
C_1 to C_{12}	−6.68‡	−8.54‡	−5.57‡	−6.61‡	−5.39‡	−5.80‡	−4.68‡	−7.26‡	−3.83‡	−5.95‡	−6.91‡	−9.28‡
	(−5.10)	(−4.92)	(−4.67)	(−4.70)	(−4.44)	(−4.08)	(−3.32)	(−5.01)	(−2.81)	(−4.26)	(−5.25)	(−6.05)
C_{13} to C_{24}	−5.38‡	−8.52‡	−5.42‡	−5.95‡	−5.53‡	−5.71‡	−3.40†	−5.81‡	−3.54‡	−5.23‡	−4.28‡	−7.38‡
	(−3.84)	(−4.90)	(−3.60)	(−4.58)	(−3.69)	(−4.02)	(−2.59)	(−4.02)	(−2.72)	(−3.75)	(−3.32)	(−4.81)

Notes:
Figures are percentages; models are: 1. the Capital Asset Pricing Model, 2. the Dimson-Marsh (1986) risk and size adjustment model, 3. the Dimson-Marsh (1986) simple size adjustment model, 4. the Fama et al. (1993) type multi-index model using the average of the smallest five deciles minus largest decile returns, 5. a multi-index model using the return on the *HGSCI* minus the return on the *FTASI*, and 6. the Fama and French (1996) three factor model. A_0 is the calendar month of announcement, C_0 the calendar month of completion of the merger, and C_1 the calendar month following completion of the merger, etc. Figures in brackets are *t*-statistics; * = significant at the 10% level, † = significant at the 5% level, and ‡ = significant at the 1% level,in two-tailed tests. The Z-test is the sign test result.

Figure 1

CARs from Alternative Benchmarks

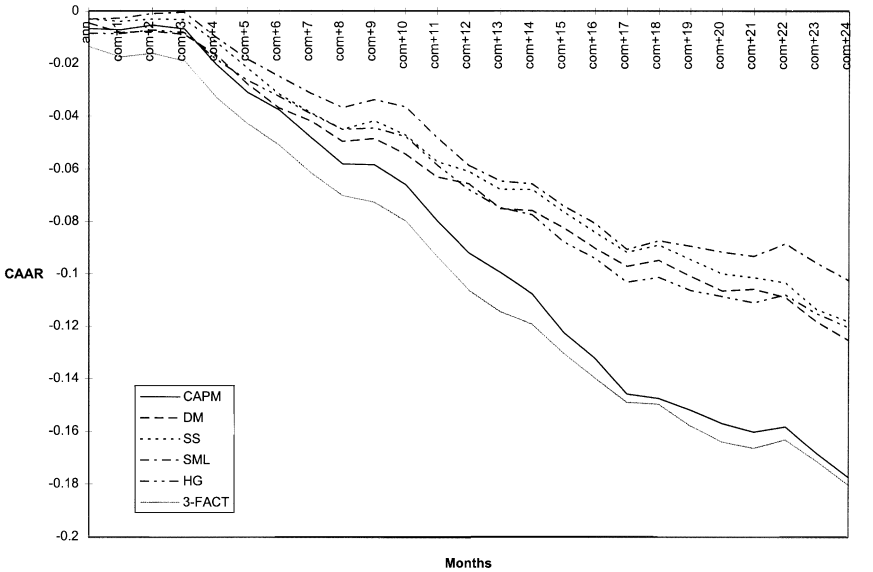
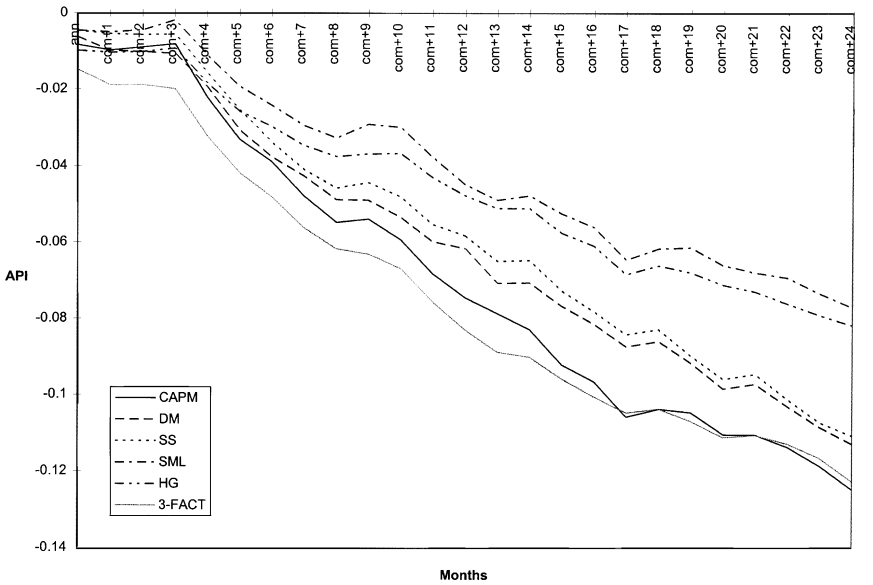


Figure 2

APIs from Alternative Benchmarks



all the size-adjusted models yield 24-month CAARs within around 2.5% of one another, with the CAPM and the three-factor model producing considerably larger negative CAARs.

Using APIs and hence avoiding portfolio re-balancing assumptions, the HG model gives the smallest negative returns (significant at the 1% level) of −8.15% after 24 months, with −4.78% occurring during the first 12 months after completion. The SML Model 4 again gives a result (−9.18%) close to the SML model. Using APIs the largest negative returns (−12.44%) are found under the CAPM, close to the −12.22% produced by the three-factor model. The Dimson-Marsh Model 2 produces an API of −11.25%, whilst the returns from the SS Model (3) are −11.03%. On an API basis, the sub-period returns for t_{c+1} to t_{c+12} and t_{c+13} to t_{c+24} are significantly negative at the 1% level for all models, except (4), which is significant at the 5% level. Although the stress in this paper is placed upon the returns from Model 5, it is nonetheless comforting to find that all six models (and, in particular, the four size-adjusted models) confirm the central conclusion that long term post-acquisition returns are significantly negative. Finally, to check whether these mean APIs are being driven by outliers, median APIs and the number of positive and negative APIs are also shown in Table 2. In all cases, the median APIs are lower (more negative) than the mean APIs, and a simple sign test shows the number of negative APIs to be significantly greater than the number of positive APIs at the 1% level.[19]

The large negative results associated with the three-factor Model 6 raise the inevitable question of whether these are driven by survivorship bias. If *Datastream* has a bias towards the inclusion of surviving companies, the SMB and HML portfolios will have upwardly biased returns. One way of investigating this is to compare the value-weighted return on the portfolio of all companies from which the SMB and HML sets were derived (i.e. the firms for which both market capitalisation and BMV figures are available — the total portfolio) with the value-weighted *FTASI* return. For the period January 1984 to December 1994, the mean monthly difference between the total portfolio return and the *FTASI* return was −0.019%; taking annual averages of the monthly differences, the maximum is 0.268% per month (1987) and the minimum difference is −0.267% per month. For seven of the years the difference is negative, whilst for four of the years the difference is positive. However, the years 1986 and 1987 were important years for takeover activity, and positive differences occurred during both these years. Nonetheless, taking an extreme position and assuming that all of this difference is due to bias,[20] for a takeover occurring in January 1986 the maximum additional (negative) API which could be attributed to such bias is 4.9%. Thus survivorship bias in the benchmark portfolios cannot explain all of the negative APIs under the three-factor model, but it may have a role in explaining why the three-factor abnormal returns are markedly lower than those from any of the size-adjusted benchmarks.

An alternative method of looking at the overall results is by direct examination of the regression model alphas used to estimate the parameters for (1), (4), (5) and (6) above. The averages of the post-completion individual regression alpha parameters are shown in column 9 of Table 3; averages of the beta, and gamma and delta coefficients (where applicable) are also shown. For all four models the post-completion alpha values are significantly negative at the 1% level. The FF Model 6 gives the lowest value estimate of an average of −0.796% per month (unrounded), with the CAPM estimate being −0.769% per month, the SML giving −0.675% per month and the HG Model 5 producing −0.544% per month. These figures are based on up to 36 months observations, with a minimum of 12 months,[21] and therefore care needs to be exercised in drawing inferences about the magnitude of the abnormal returns. Nonetheless, a compounding of the HG estimate of 0.544% by 24 months produces a figure of 13.5%, close to the 24 month CAAR of −12.03%. The implication of these coefficients is that the adverse post-completion performance may continue at a similar rate for up to 36 months after a takeover. It is also worth noting that although the CAPM and SML betas show a significant increase post-acquisition, the HG and three-factor FF model betas do not; as expected with the HG model, the exposure to size effects is significantly reduced by a takeover.[22] Intriguingly, the three-factor FF model shows that exposure to BMV effects increases significantly post-takeover. Pre-takeover, exposure to BMV effects is significantly negative.[23] Post takeover, acquirers show a positive but insignificant exposure to BMV effects. This provides an alternative explanation for the large difference between the three-factor model results and the results from the other size control models. If BMV is a proxy for growth opportunities, it would seem that firms which become acquirers are initially perceived by the market as being 'growth' companies; post takeover, they no longer appear to be seen as such.

Confirmation of the scale of the under-performance from the size-control models can also be found from the RATS coefficient estimate. The mean monthly alpha estimate for the RATS model is -0.0061 (t=−6.91), which implies a 24 month average under performance of 13.54%. Given that the RATS model is effectively specified as the Dimson-Marsh model, it is not surprising that this result is reasonably close to the DM CAAR which is −11.25%. What the RATS model shows is that time-varying betas are not an explanation of poor post-takeover performance. It is also interesting to note that the beta coefficient from the RATS model, which as noted above is the difference between the security and decile beta in period t, is significantly positive at 0.131 (t=8.85). The value of this coefficient is also very close to that suggested by the difference in the average betas given in note 18.

Looking at the times-series regressions of returns shows that the results in calendar time are indicative of larger negative abnormal returns. The coefficients and t-ratios for the four regression models estimated (the time

Table 3

Pre-Announcement and Post-Completion Regression Results for Alternative Benchmark Models

Model	*Pre-announcement Paramaters (436)*				*Post-completion Paramaters (452)*				*Differences in Paramaters (Post - Pre) (426)*			
	Beta	*Gamma*	*Delta*	*Alpha*	*Beta*	*Gamma*	*Delta*	*Alpha*	*Beta*	*Gamma*	*Delta*	*Alpha*
CAPM (1)	1.015‡	n.a.	n.a.	0.017‡	1.105‡	n.a.	n.a.	−0.008‡	0.065‡	n.a.	n.a	−0.020‡
	(73.45)			(10.40)	(106.86)			(−6.70)	(4.47)			(−12.96)
SML(4)	1.187‡	0.472‡	n.a.	0.004‡	1.25‡	0.407‡	n.a.	−0.007‡	0.077‡	−0.057	n.a.	−0.011‡
	(70.69)	(12.69)		(5.15)	(104.67)	(16.594)		(−5.71)	(4.129)	(−1.16)		(−7.87)
HG(5)	1.192‡	1.032‡	n.a.	0.007‡	1.186‡	0.692‡	n.a.	−0.005‡	−0.003	−0.347‡	n.a.	−0.013‡
	(73.46)	(18.69)		(7.22)	(110.29)	(23.52)		(−4.01)	(0.84)	(−2.68)		(−8.51)
FF(6)	1.140‡	0.429‡	−0.151‡	0.006‡	1.130‡	0.349‡	0.123	−0.008‡	−0.002	−0.085	0.269‡	−0.014‡
	(69.43)	(9.53)	(−4.00)	(6.36)	(99.66)	(11.84)	(1.57)	(−7.16)	(1.26)	(0.46)	(4.37)	(−9.72)

Notes:
Figures are the coefficient averages and differences in averages. Models are: 1. the Capital Asset Pricing Model, 2. the Dimson-Marsh (1986) risk and size adjustment model, 3. the Dimson-Marsh (1986) simple size adjustment model, 4. the Fama et al. (1993) type multi-index model using the average of the smallest five deciles minus largest decile returns, 5. a multi-index model using the return on the *HGSCI* minus the return on the *FTASI*, and 6. the Fama and French (1996) three factor model. Figures in brackets in the header row are the number of observations; figures in brackets in the body of the table are *t*-statistics; * = significant at the 10% level, † = significant at the 5% level, and ‡ = significant at the 1% level,in two-tailed tests. The regressions in each case are estimated using up to 36 months data and are given by expressions (1), (4), (5) and (6) in the text. Beta is the sensitivity of the excess returns on the acquiring company to the excess returns on the market (*FTASI*); gamma is the sensitivity of the excess returns on the acquiring company to a 'small firms premium' whilst delta is the sensitivity to BMV effects. Alpha is the intercept term. Pre-announcement parameters are the results of running the regressions on the 36 months ending six months prior to bid announcement, whilst post-announcement parameters are the results of running the regressions on the 36 months post completion of the takeover.

Table 4

Time-Series Regression Results for Alternative Benchmark Models

	Coefficients:				*Adjusted R-squared*
	Alpha	*Beta*	*Gamma*	*Delta*	
CAPM	−0.0066	1.1098	n.a.	n.a.	0.796
	(−2.561)†	(25.14)‡			
SML	−0.0068	1.2177	0.8922	n.a.	0.915
	(−4.108)‡	(36.03)‡	(10.44)‡		
HG	−0.0098	1.2578	0.4549	n.a.	0.865
	(−4.693)‡	(26.31)‡	(5.510)‡		
FF three-factor	−0.0099	1.1659	0.4081	0.134	0.852
	(−4.748)‡	(30.90)‡	(5.495)‡	(1.120)	

Notes:
Time-series models are the Capital Asset Pricing Model , the Fama et al. (1993) type multi-index model using the average of the smallest five deciles minus largest decile returns, a multi-index model using the return on the *HGSCI* minus the return on the *FTASI*, and the Fama and French (1996) three factor model. Figures in brackets are the *t*-statistics; * = significant at the 10% level, † = significant at the 5% level, and ‡ = significant at the 1% level,in two-tailed tests. The regressions in each case are estimated using 131 monthly observations with the dependent variable being the return on a portfolio of acquiring firms minus the risk-free rate, and the independent variables being the benchmark factors implied by expressions (1), (4), (5) and (6) in the text. Beta is the sensitivity of the excess returns on the acquiring company to the excess returns on the market (*FTASI*); gamma is the sensitivity of the excess returns on the acquiring company to the 'small firms premium', and delta is the sensitivity to the BMV factor. Alpha is the intercept term.

series models using the benchmarks employed in (1), (4), (5) and (6)) are reported in Table 4. As there is some weak evidence of heteroscedasticity in some of the regressions, all *t*-statistics are corrected using the White (1980) heteroscedasticity-consistent method. In all cases, the intercept term is significantly negative, with abnormal returns of −0.66% per month under the CAPM, −0.68% under the SML model, −0.98% per month under the HG model, and −0.99% per month under the three-factor FF model. Consistent with the event time results, size effects are significant in both size control and three-factor models, and BMV effects are not significantly different from zero.

The final test on the validity of the results is the cross-sectional test described by (11). Results from this model are provided in Table 5, with each year being shown separately. As individual years' regressions can show evidence of heteroscedasticity in the residuals, all standard errors are reported after correction by the White (1980) method. The dummy variables for acquirers confirm that the return in the month of acquisition is negative (−0.5%) but not significantly so, whereas the returns in the 12 months following completion are a highly significant −0.6% (t=−2.909) per month. The results for the 13 to 24 months following completion have a value of −0.4% per month, although the coefficient just fails to be significant at the 5% level using a conventional

Table 5

Cross-Sectional Regression of Company Returns

Year	A_{it}	Y^r_{it}	Y^p_{it}	S_i	R'_{mt}	*Constant (a)*
1984	−0.011	0.004	n.a.	0.000	0.777‡	−0.005
	(−0.70)	(0.57)		(0.37)	(54.62)	(−0.73)
1985	0.016	0.001	0.010	−0.002‡	0.662‡	0.028‡
	(0.79)	(0.36)	(1.46)	(−3.74)	(33.13)	(3.97)
1986	0.002	−0.000	0.001	−0.003‡	0.507‡	0.067‡
	(0.20)	(0.14)	(0.12)	(−7.13)	(29.98)	(8.57)
1987	0.031*	−0.009†	0.003	−0.004‡	1.049‡	0.075‡
	(1.77)	(−2.53)	(0.80)	(−7.08)	(90.75)	(7.03)
1988	−0.007	−0.004	−0.004	−0.000	0.851‡	0.009
	(−0.50)	(−1.49)	(−1.32)	(−1.32)	(42.10)	(1.41)
1989	−0.002	−0.005	−0.004	0.003‡	0.794‡	−0.075‡
	(−0.16)	(−1.26)	(−1.11)	(9.24)	(56.34)	(−11.13)
1990	−0.012	−0.015†	−0.025‡	0.005‡	0.666‡	−0.114‡
	(−0.60)	(−2.64)	(−3.39)	(11.80)	(41.13)	(−14.82)
1991	−0.003	0.002	−0.016	0.001‡	0.866‡	−0.036‡
	(−0.53)	(0.34)	(−1.60)	(2.99)	(35.97)	(−3.98)
1992	−0.060	−0.018*	−0.017	0.003‡	0.926‡	−0.077‡
	(−1.23)	(−1.68)	(−1.64)	(6.00)	(46.30)	(−7.36)
1993	n.a.	−0.011	−0.009	−0.003‡	0.542‡	0.078‡
		(−0.73)	(−0.93)	(−6.94)	(16.74)	(8.69)
1994	n.a.	n.a.	0.025	−0.002‡	0.769‡	0.035‡
			(1.53)	(−4.07)	(39.65)	(3.88)
Average	−0.005	−0.006‡	−0.004*	−0.000	0.765‡	−0.001
t-test	(−0.322)	(−2.909)	(−1.920)	(0.032)	(140.504)	(−1.210)

Notes:
Regressions are estimated for acquisitions each year 1984–94. Explanatory variables are A_{it} = dummy variable of 1 if a company announces a takeover in the month, Y^r_{it} = dummy variable of 1 if a company has announced or completed a takeover in the previous 12 months, Y^p_{it} = dummy variable of 1 if a company has announced or completed a takeover in the previous 13–24 months, S_i = log of market capitalisation at the beginning of the year, and R'_{mt} = excess return on the market during the month. Figures in brackets are *t*-statistics; * = significant at the 10% level, † = significant at the 5% level, and ‡ = significant at the 1% level, in two-tailed tests.

two-tailed *t*-test (t = 1.92). If, however, the cross-sectional regression is regarded as confirming the results of the 'event study' type models, then a one-tailed test is appropriate since the hypothesis being tested is that acquirers returns are significantly negative, in which case the 13–24 month returns are significant at the 5% level. The comforting conclusion from these cross-sectional regressions is that they seem to confirm the average returns suggested by the size-adjusted models' APIs. Compounding up the mean coefficients for the announcement, 1 to 12 month and 13–24 month periods suggests a total abnormal return of −10.89% after two years, between the APIs shown by the HG and SS models. One other result of interest from the cross-sectional

regressions is that there is considerable year-to-year variability in the value of the dummy coefficients; in particular, it seems that acquisitions completed between 1986 and 1992 yield particularly poor results (takeovers completed during 1986 and 1989 appearing to be the worst cases[24,25]). It seems, therefore, that models which control for size effects, the Ibbotson RATS approach, and a very general cross-sectional model yield broadly similar conclusions.

Partitioned Results

Having confirmed the reliability of the overall results, which seem to suggest that the size adjusted models' results correlate with the conclusions drawn from a general cross-sectional model, the break-down of the results is now investigated. The results are partitioned by form of payment, whether or not the take-over was a conglomerate merger, whether the bid hostile, whether the bid is from a multiple bidding (or 'regular bidder') company (during the period of the study), and whether there was a competing bid for the target. Payment form is partitioned according to whether the bid is for cash (note that cash offers with a loan note alternative[26] were classified as cash offers), for shares (including share offers with a cash alternative)[27] or mixed (for example, shares plus cash, shares plus loan notes, cash plus loan notes).

Conglomerate bids are those where the two-digit SIC codes differ between acquiring and acquired company; if the first two-digits of the SIC code are the same, the bid is classified as a 'horizontal' takeover. Opposition to bids is partitioned according to whether the bid was hostile, whether the bid was agreed, and whether or not the bidder was a 'white knight' (i.e. a second bidder favoured by the target company's board). 'Regular acquirers' were defined as those companies which completed more than one acquisition of value greater than £10m during the period 1984–92. In all cases, partitions are based upon information disclosed in the *AMDATA* database. Given the arguments in the literature in favour of using APIs rather than CAARs, only the former are reported in the results for this section, which are presented in Table 6a; Table 6b shows median APIs and the number of positive and negative APIs for each partitioning. In the context of choosing between APIs and CAARs, it should be borne in mind that Kothari and Warner (1996) show that when abnormal performance is negative, APIs are likely to lead to an under-rejection of the null hypothesis.

In the case of partitioned data, the model chosen can have an impact on the conclusions drawn. In addition, the small number of observations in some of the categories make the drawing of inferences difficult. Using the preferred HG model (5), the 24 month API is not significantly different from zero for cash offers, but significantly negative (−11.57%) for share offers; mixed offers record a positive API of 4.12%, but the number of these is small (35) allowing no statistical inferences to be drawn. A two-sample *t*-test for differences between

Table 6a

Partitioned 24-month Abnormal Performance Index for Alternative Benchmark Models

Partitioned on:	*CAPM Model 1*		*Dimson-Marsh Model 2*		*Size Decile Model 3*		*Multi-index SML Model 4*		*Multi-index HG Model 5*		*Multi-indexFF Model 6*	
	API %	*No. Bids*	*API %*	*No. Bids*	*API %*	*No. Bids*	*API %*	*No. Bids*	*API %*	*No. Bids*	*API %*	*No. Bids*
Payment												
Cash	−9.50†	84	−6.12	80	−5.83	81	−0.60	84	0.31	84	−4.63	84
	(−2.36)		(−1.53)		(−1.46)		(−0.12)		(0.07)		(−1.02)	
Shares	−12.84‡	333	−14.62‡	292	−13.85‡	304	−12.62‡	333	−11.57‡	333	−15.07‡	333
	(−5.90)		(−6.73)		(−6.37)		(−4.99)		(−4.81)		(−6.14)	
Mixed	−6.13	35	7.28	31	2.12	33	0.05	35	4.12	35	−3.32	35
	(−0.27)		(1.89)		(0.36)		(0.01)		(0.67)		(−0.59)	
Conglomerate												
Yes	−14.14‡	269	−11.70‡	241	−11.18‡	248	−11.99‡	269	−11.33‡	269	−13.68‡	269
	(−4.68)		(−4.99)		−(4.86)		(−4.44)		(−4.39)		(−5.21)	
No	−9.92‡	183	−10.56‡	162	−10.81‡	170	−5.05	183	−3.48	183	−10.08‡	183
	(−3.19)		(−3.57)		(−3.65)		(−1.45)		(−1.06)		(−3.10)	
Hostility												
Hostile	−8.59	84	−8.51	80	−8.67	81	−2.74	84	−1.68	84	−4.83	84
	(−1.75)		(−1.75)		(−1.73)		(−0.49)		(−0.31)		(−0.85)	
Recommended												
	−13.70‡	349	−12.04‡	305	−11.79‡	318	−11.30‡	349	−10.35‡	349	−14.57‡	349
	(−6.25)		(−5.97)		(−5.95)		−(4.78)		(−4.64)		(−6.56)	
White Knight												
	−6.27	19	−10.41†	18	−8.49	19	1.26	19	3.76	19	−1.83	19
	(−0.82)		(−1.64)		(−1.22)		(0.03)		(0.43)		(−0.22)	
Regular Acquirer?												
Yes	−7.62†	228	−8.57‡	204	−7.07‡	209	−5.48*	228	−5.18*	228	−7.30†	228
	(−2.85)		(−3.19)		−(2.61)		(−1.89)		(−1.87)		(−2.53)	
No	−17.34‡	224	−13.99‡	199	−15.00‡	209	−12.95‡	224	−11.17‡	224	−17.23‡	224
	(−6.09)		(−5.66)		(−6.15)		(−4.12)		(−3.76)		(−5.96)	
Competing Bid?												
Yes	−9.12	42	−12.15*	38	−10.39	40	−4.10	42	−2.23	42	−6.68	42
	(−1.44)		(−1.72)		(−1.48)		(−0.60)		(−0.33)		(−1.08)	
No	−12.78‡	410	−11.15‡	365	−11.10‡	378	−9.70‡	410	−8.75‡	410	−7.45‡	410
	(−5.31)		(−5.95)		(−5.93)		(−4.31)		(−4.13)		(−3.35)	

Notes:
Figures are percentages; models are: 1. the Capital Asset Pricing Model, 2. the Dimson-Marsh (1986) risk and size adjustment model, 3. the Dimson-Marsh (1986) simple size adjustment model, 4. the Fama et al. (1993) type multi-index model using the average of the smallest five deciles minus largest decile returns, 5. a multi-index model using the return on the *HGSCI* minus the return on the *FTASI*, and 6. the Fama and French (1996) three factor model. Figures in brackets are *t*-statistics; * = significant at the 10 % level, † = significant at the 5% level, and ‡ = significant at the 1% level, in two-tailed tests.

Table 6b

Partitioned 24-month Positive and Negative Abnormal Performance Indices For Alternative Benchmark Models

Partitioned on:	*CAPM Model 1*		*Dimson-Marsh Model 2*		*Size Decile Model 3*		*Multi-index SML Model 4*		*Multi-index HG Model 5*		*Multi-index FF Model 6*	
	Median API: +ve −ve	*No. Bids, z-test*	*Median API: +ve −ve*	*No. Bids, z-test*	*Median API: +ve −ve*	*No. Bids, z-test*	*Median API: +ve −ve*	*No. Bids, z-test*	*Median API: +ve −ve*	*No. Bids, z-test*	*Median API: +ve −ve*	*No. Bids, z-test*
Payment												
Cash	−10.22	84	−5.70	80	−6.92	81	−3.33	84	−2.89	84	−9.88	84
	37	(−1.09)	35	(−1.12)	35	(−1.22)	38	(−0.87)	38	(−0.87)	33	(−1.96)
	47		45		46		46		46		51	†
Shares	−16.82	333	−18.23	292	−17.28	304	−15.92	333	−15.14	333	−19.23	333
	111	(−6.08)	83	(−7.37)	87	(−7.46)	114	(−5.45)	118	(−5.32)	109	(−6.30)
	222	‡	209	‡	217	‡	219	‡	215	‡	224	‡
Mixed	0.62	35	10.15	31	2.12	33	−2.06	35	4.22	35	4.17	35
	18	(0.17)	19	(1.26)	20	(1.21)	18	0.17	19	(0.51)	19	0.51
	17		12		13		17		16		16	
Conglomerate												
Yes	−16.34	269	−13.81	241	−14.24	248	−14.92	269	−13.19	269	−15.69	269
	97	(−4.57)	84	(−4.70)	84	(−5.08)	98	(−4.45)	101	(−4.08)	94	(−4.94)
	172	‡	157	‡	164	‡	171	‡	168	‡	175	‡
No	−12.34	183	−12.52	162	−12.82	170	−10.18	183	−8.32	183	−12.19	183
	69	(−3.33)	53	(−4.40)	58	(−4.14)	72	(−2.88)	74	(−2.59)	67	(−3.62)
	114	‡	109	‡	112	‡	111	‡	109	‡	116	†
Hostility												
Hostile	−15.52	84	−16.60	80	−14.3	81	−6.92	84	−11.94	84	−8.09	84
	27	(3.27)	29	(−2.46)	29	(−2.56)	36	(−1.31)	35	(−1.53)	29	(−2.84)
	57	‡	51	†	52	†	48		49		55	‡
Recommended												
	−14.78	349	−12.91	305	−14.19	318	−14.09	349	−12.46	349	−17.42	349
	132	(−4.55)	101	(−5.90)	107	(−5.83)	126	(−5.19)	132	(−4.55)	124	(−5.41)
	217	‡	204	‡	211	‡	223	‡	217	‡	225	‡
White Knight												
	−9.46	19	−5.91	18	−4.37	19	−2.77	19	−3.72	19	−4.87	19
	7	(−1.15)	7	(−0.94)	6	(−1.61)	8	(−0.69)	8	(−0.68)	8	(−0.69)
	12		11		13		11		11		11	
Regular Acquirer?												
Yes												
	−9.60	228	−11.65	204	−10.14	209	−7.42	228	−8.37	228	−10.15	228
	94	(−2.64)	72	(−4.20)	75	(−5.19)	90	(−3.18)	94	(−2.64)	92	(−2.92)
	134	‡	132	‡	134	‡	138	†	134	‡	136	‡
No	−19.09	224	−18.59	199	−18.78	209	−16.91	224	−14.81	224	−19.02	224
	72	(−5.34)	65	(−4.89)	67	(−5.18)	80	(−4.28)	81	(−4.14)	69	(−5.75)
	152	‡	134	‡	142	‡	144	‡	143	‡	155	‡
Competing Bid?												
Yes	−16.19	42	−26.24	38	−23.05	40	−13.59	42	−9.89	42	−13.44	42
	14	(−2.16)	13	(−1.95)	13	(−2.21)	16	(−1.54)	17	(−1.23)	15	(−1.85)
	18	†	25	*	27	†	26		25		27	*
No	−14.47	410	−12.91	365	−13.93	378	−12.94	410	−12.17	410	−14.54	410
	152	(−5.23)	124	(−6.12)	129	(−6.17)	154	(−5.03)	158	(−4.64)	146	(−5.83)
	258	‡	241	‡	249	‡	256	‡	252	‡	264	

Table 6 (Continued)

Notes:
For each classification, the first figure is the median API, the second figure is the number of positive APIs, and the third the number of negative APIs. Models are: 1. the Capital Asset Pricing Model, 2. the Dimson-Marsh (1986) risk and size adjustment model, 3. the Dimson-Marsh (1986) simple size adjustment model, 4. the Fama et al. (1993) type multi-index model using the average of the smallest five deciles minus largest decile returns, 5. a multi-index model using the return on the *HGSCI* minus the return on the *FTASI*, and 6. the Fama and French (1996) three factor model. The figures in parentheses are the z-statistics from the sign test.

APIs shows cash significantly higher than shares at the 5% level ($t = 2.28$) and shares significantly lower than mixed offers ($t = -2.37$). These findings concerning share and cash offers confirm the conclusions of Franks et al. (1991) and Agrawal et al. (1992). Broadly similar conclusions are drawn from Models 2, 3, 4 and 6, although all of these show negative (but insignificant) returns associated with cash offers, compared to a positive (but insignificant) return of 0.31% from the HG model. Only the CAPM (Model 1) shows cash offers having a significant negative post-acquisition performance. However, the ordering of the returns is the same across all models, with mixed offers showing the strongest results, followed by cash offers with share offers exhibiting the poorest returns. Apart from the HG model and the three-factor model (which has a t-test value of 2.02), the two-sample t-tests for differences between cash and equity offers fail to be significant at the 5% level, although the DM and SS models show significant differences at the 10% level. The sign tests (based upon the number of postive and negative APIs in the sample) reported in Table 6b confirm the significant adverse performance of equity bidders. They also confirm that neither cash nor mixed offers appear to exhibit returns significantly different from zero, with the exception that the three-factor model shows cash offers as being just significantly negative at the 5% level ($z = -1.96$). An alternative test for differences between the partitioned samples is the simple χ^2 test, which avoids any conclusion being driven by possible outliers in the data.[28] Whilst Models 1 to 5 show the difference between cash and equity to be significant at least at the 10% level, the three-factor model shows the difference to be significant at only the 25.8% level.

This finding that equity financed acquisitions perform badly suggests a possible link to the literature on seasoned equity offerings (SEOs). Loughran and Ritter (1995, Table III) find that the sample of firms conducting an SEO under-perform a matched sample of non-issuing firms by an average of 6.7% p.a. over five years. This is reasonably close to the size-adjusted APIs associated with equity-financed acquisisitions (Table 6a) and raises an interesting question for future research. Is the equity-financed acquisition another example of firms with over-valued equity taking advantage of the over-valuation to sell this equity to the market through the medium of a share-for-share offer? Loughran and Ritter note that their results are

> consistent with a market in which companies announce stock issues when their stock is grossly over-valued, the market does not revalue the stock appropriately, and the stock is still substantially over-valued when the issue occurs (1995, p. 47).

Loughran and Ritter also find that issuing firms tend to have had recent improvements in their operating performance; this is compatible with the evidence presented in Table 3 which shows significant out-performance by acquirers in the pre-announcement period.

Turning to the issue of conglomerate versus non-conglomerate acquisitions, under Model (5), conglomerate mergers show significant average negative returns of −11.33% over 24 months, compared to an insignificant −3.48% for non-conglomerate mergers. The two-sample *t*-test for differences is significant at the 10% level ($t = 1.62$). However, the CAPM, DM, SS and three-factor models show both conglomerate and non-conglomerate acquirers exhibiting significant under-performance. Furthermore, the sign tests for both groups (Table 6b) are significantly negative for all models. None of the models show significant χ^2 test results for differences between conglomerate and non-conglomerate acquirers, based on the number of positive and negative abnormal returns. To shed further light on the matter of the relative performance of conglomerate and non-conglomerate acquisitions, the sample was split into 1984–88 and 1989–92 time periods, to capture the effect (as in Powell, 1997) of the 'merger wave' of the mid-1980s. The HG and SML models confirm that in both periods conglomerate acquirers have lower mean and median APIs than non-conglomerate acquirers. Both models show that the difference in performance widens between the two periods; the HG model has conglomerate APIs of −9.46% versus non-conglomerate returns of −5.55% for the first period (both significant at the 5% level), compared to a significant −16.5% and an insignificant +2.14% for the second period. The difference in the second period between conglomerates and non-conglomerates is almost significant at the 5% level ($t = 1.95$).[29] Agrawal et al. (1992) find non-conglomerates have poorer post-acquisition performance; the difference in relative performance in the second time period may be compatible with the idea that conglomerate acquirers have become 'unfashionable' in later years. Whilst it is clear that conglomerate mergers perform poorly in the longer term, to some degree it remains an open issue as to whether or not non-conglomerate mergers produce significant negative abnormal returns. One problem is that the lower number of acquisitions post 1988 makes it difficult to reach any strong conclusions on this matter.[30]

Partitioning on hostility shows that recommended offers have the worst performance over the 24 months following a takeover under all models. This may be because over-payment is more likely in an agreed bid; it may also be because agreed bids are unlikely to be 'disciplinary', which Kennedy and Limmack (1996) show to have marginally less poor post-bid performance than 'non-disciplinary' bids. This result is in spite of the fact that bid resistance leads to higher wealth gains for target firms (Holl and Kyriazis, 1997).

Recommended bids produce average APIs ranging from a significant −10.35% under the HG model to −14.57% (both significant at the 1% level) under the three-factor model. Returns in hostile bids are negative but not significantly so in the case of all models. The SML model shows returns as −2.74%, whilst the HG model shows −1.68%. Sign tests for both the HG and SML models show insignificant negative returns for hostile acquirers, but these are significantly negative in the case of all other models. However, two-sample *t*-tests for differences between hostile and recommended offers are insignificant at the 10% level for all models and this insignificance in differences is confirmed by the χ^2 test results based on the number of positive and negative abnormal returns. So-called 'White Knight' bidders generate small but insignificant positive average returns according to HG and SML models but not according to the other models; the small number of these deals (19) means in general that statistical inferences cannot be drawn.

One other result of significance is that under all of the models regular acquirers do better than one-off acquirers, perhaps suggesting that some 'learning' may take place as a result of the acquisition process. Under the HG model, regular acquirers show a 24 month return of −5.18% (significant at the 10% level), as opposed to −11.17% (significant at the 1% level) for single acquirers. All but the SML and HG models show regular acquirers having significant negative performance at least at the 5% level, whilst all models show single acquirers generating significant negative returns at the 1% level. However, sign tests show regular acquirers exhibiting negative abnormal performance in all cases. The two-sample *t*-test suggests that regular acquirers do significantly less badly under SML, SS and CAPM models (t = 2.02, 2.17 and 2.48 respectively), but the HG, DM and three-factor models show the difference to be significant only at the 10% level. The χ^2 test indicates these differences to be significant only for the three-factor and CAPM benchmarks.

Finally, although the relatively small number of cases where there are competing bids makes the drawing of inferences difficult, there is little to suggest that the arrival of a competing bid causes worse post-acquisition performance on average by the successful acquirer. Under the HG model, acquirers in competitive bids have a post acquisition performance of an insignificant −2.23%, as opposed to −8.75% (significant at the 1% level) in the case of uncontested bids. This general conclusion, in line with that of Franks et al. (1991), is borne out by all models except the DM model. Median APIs and the number of positive and negative APIs do not reveal any significant differences between competing and non-competing bid situations, and neither do two-sample *t*-tests for differences.

CONCLUSIONS

The contribution of this paper has been to show that the post-takeover performance of UK companies undertaking large domestic acquisitions is

unambiguously negative, on average, in the longer term. Under all benchmarks used, this conclusion is unaltered. However, the argument has been that a 'good' benchmark must control for the size effect. Using four different models which control for size and one that controls for both size and BMV effects fails to alter the general conclusion that the two year post takeover performance for these UK acquirers is significantly negative.

This conclusion is strengthened by confirming evidence from the Ibbotson 'RATS' method, by calendar time (as opposed to event time) returns, and by a general cross-sectional regression of all companies with the inclusion of dummy variables for acquiring firms. The only impact that model choice and return accumulation methods have is in terms of the size of the abnormal returns and sometimes on the conclusions drawn from the partitioning of the sample. After size has been controlled for, the scale of abnormal returns (as measured by API) ranges from a minimum of −8.15% (using a multi-factor benchmark with the return on the *Hoare-Govett Index* minus the return on the *FTASI* as a measure of the 'small firms' effect) to −11.25% (using the size and risk control model of Dimson and Marsh, 1986). It is also clear that allowing for 'book-to-market' effects by using the Fama and French (1996) three-factor benchmark does not explain the negative abnormal returns found. Indeed, abnormal returns are particularly striking under this model, although it should be borne in mind that there is, as yet, no conclusive evidence on whether a three-factor model is appropriate in explaining the cross-section of UK share returns.

Thus the evidence in this paper is not compatible with shareholder wealth maximising behaviour on the part of acquiring firms' management. Rather, the evidence is consistent with either Roll's (1986) 'Hubris' hypothesis of takeovers, or with 'managerialist' theories of behaviour.

Partitioning the sample confirms the results found elsewhere that cash offers are associated with post-merger performance which is not significantly different from zero, whilst equity offers are associated with significant negative post-outcome performance. As was noted, this evidence is compatible with acquiring firms using over-valued equity to buy target firms, and this 'over-valued' equity hypothesis is one which has been advanced in explaining the poor performance of firms making seasoned equity offerings (Loughran and Ritter, 1995). Looked at in this way, a share-for-share acquisition is simply one way to issue over-valued equity to the market.

Multi-index models of performance also suggest that conglomerate mergers may be less successful than conglomerate mergers, agreed bids are less successful than hostile bids, and that acquisitions by multiple acquirers are less unsuccessful than those by single acquirers, although these conclusions are sensitive to the benchmark for performance chosen.

Legitimate concerns can be raised about the appropriate benchmark to use when assessing the performance of acquirers, and some of the conclusions with regard to the partitioned samples are dependent on the model chosen.

Nonetheless, unless acquiring firms differ from non-acquirers in some systematic manner in their exposure to risk factors priced by the market, the main conclusion from this analysis is that the long-run post acquisition performance of UK acquiring firms is significantly negative, and that this result is robust to alternative benchmark specifications.

NOTES

1 If size effects were constant, it would be possible to argue that the decrease in alpha and the decrease in exposure to size effects are entirely consistent with a market model in which alpha captures size effects. However, the evidence that 'size effects' are time varying (e.g. Reinganaum, 1992; and Dimson and Marsh, 1995) suggests that it would be dangerous to rely on such a model.

2 Note that Franks and Harris (1989) do not use this same benchmark in their analysis of UK takeovers.

3 *AMDATA* records the name of bidding and acquiring companies at the time of the takeover, whereas the LBS Share Price Database generally records companies by their most recent name.

4 This raises the issue of survivorship bias. As is shown below, dropping this requirement has no material impact on the conclusions drawn.

5 Of the sample of 452 acquirers, 418 have returns recorded to the end of the 24 month period following completion of the acquisition. Of the 34 which do not have recorded returns for the full period, 20 acquirers themselves become subject to successful bids, whilst eight are recorded by the *Stock Exchange Official Yearbooks* as being in receivership, liquidation or administration. The remaining six have missing returns for some of the second year. Omitting these firms from the sample has no material impact on any of the results reported below.

6 The figures vary over time; for example, in January 1986 the proportion was 79.2%.

7 For the small number of takeovers (13) which occurred in 1992, the maximum number of observations is necessarily limited to less than 36 months, the absolute minimum being 24 months if a takeover was completed in December 1992.

8 This has the effect of setting the benchmark return equal to the risk-free rate.

9 However, CARs were also estimated using log returns. The results from this are consistently lower than the results obtained from the analysis using raw returns.

10 Except in the case of takeovers occurring in 1992; see note 4.

11 An SML model based on the differences between these extreme portfolios was estimated, and although the results are not reported here, the abnormal returns were generally more negative than those found under Models 4 and 5, but less negative than those found from Models 2 and 3.

12 It is also questionable whether BMV has any real role in explaining the cross-section of US stock returns (Kothari et al., 1995; and Jaganathan and Wang, 1996).

13 In the following section, it is shown that this effect also holds in the current sample.

14 In fact, as is shown later, the average acquirer has a beta greater than the average decile beta.

15 The portfolios are formed for each month from January 1984 to December 1994, although only acquisitions to the end of December 1992 are considered. Thus the total number of months is 132; however, there were no acquisitions in January 1984 and therefore the number of observations used in the regressions is 131.

16 For the purposes of these charts, the announcement period returns are shown as one month.

17 Whilst the modal difference between announcement and completion months is one month, both events sometimes occur in the same month. The maximum difference (one case only, referred to the Monopolies and Mergers Commission) is nine months. Although under the *City Code on Takeovers and Mergers* the maximum period over which a bid can remain open is 60 days, this timetable can be extended in the case of referrals to the MMC (for a discussion, see Sudarsanam, 1995, chapter 6). In general, this paper treats as one bid any situation where a bidder subsequently increases or amends an offer, in order to capture the full announcement period effects on the bidder.

18 This result is to be expected given that the SS model effectively assumes that the decile beta and acquirer betas are both equal to 1.0. In fact, under the DM method, the average of the betas of the deciles to which acquirers belong is 0.953 whereas the average acquirer beta is 1.101.
19 The reported z-statistic is based upon a null hypothesis of an equal number of positive and negative APIs.
20 It should be noted that the total portfolio includes a wider universe of companies than the *FTASI*, and will therefore have greater exposure to 'size effects'; as may be expected, the difference between total portfolio returns and *FTASI* returns is positively correlated with the size effect, as measured by SMB (the Pearson correlation coefficient is 0.19).
21 See notes 4 and 7.
22 Although this result is not found using either the SML or three-factor models, where the decreases are not statistically significant.
23 Loughran and Ritter (1995) find similar negative exposure to BMV effects for their group of large firms issuing equity between 1973 and 1992.
24 Note that these references are to the year of takeover, not to the year in which the coefficient appears.
25 An alternative model, excluding the market return was run as a cross-check on these results. Although the R^2 from these regressions is far lower, the overall significance levels of the dummy variables of interest are very similar, and the conclusions on overall significance are identical.
26 An important reason for this being offered is that it allows the postponement of UK capital gains tax.
27 The UK City Code on Takeovers and Mergers required that a cash alternative must be made available to target firm shareholders if either (a) 15% or more of the voting rights had been acquired in the previous 12 months, or (b) the bid was triggered by the 30% toehold limit being exceeded. Cooke (1986) discusses the City Code requirements in more detail.
28 I am grateful to Robin Limmack for suggesting this test.
29 The figures for the SML model were −7.54% versus −4.77% for the first period (both significant at the 5% level), compared to −17.95% (significance) and −1.32% (insignificance) for the second period. The second period two-sample t-test for differences is 1.83.
30 There were 332 acquisitions in the period 1984–88, of which 134 were non-conglomerate and 198 were conglomerate. By contrast, 1989–92 figures were 120 in total, of which 49 were non-conglomerate and 71 were conglomerate.

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