# A Panel-Based Investigation Into the Relationship Between Stock Prices and Dividends

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**Abstract**

This paper investigates the presence of cointegration between stock prices and dividends for a panel of 56 large UK companies. Using new techniques which account for integrated processes in a panel context we demonstrate that stock prices and dividends are cointegrated, with an implied common discount rate of 5.8%.

# 1    Introduction

This paper examines the relationship between share prices and share dividends. Most financial textbooks point out that in a well-functioning capital market these two variables should be related (e.g. Brealey and Myers, 1986); the present value of the share should be equal to the dividend stream discounted by the return earned on securities of comparable risk.[1] A substantive body of empirical research supports the view that share prices are influenced by changes in company dividends. For example, questionnaire surveys of investors indicate that dividend information plays an important part in their assessment of the current value of a share (Arnold and Moizer, 1984; Pike, Meerjanssan and Chadwick, 1993)[2] while interviews with financial managers suggest that companies take a great deal of care when setting their dividend level (Lintner, 1956). In addition, stock market studies demonstrate that share prices respond to dividend news; dividend increases tend to be associated with share price increases while dividend cuts are usually associated with share price falls (Pettit, 1972; Ahrony and Swary, 1980; Abeyrathna et al., 1996).[3, 4]

This paper investigates whether a long-run relationship exists between dividends and share prices using the cointegration methodology. The use of this methodology has become

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[1]     Of course, this simple relationship only holds in a world of certainty where investors have access to perfect information. Nevertheless, Shiller (1981) demonstrates that a similar relationship should hold in a world of uncertainty where investors have rational expectations.

[2]     For example, Arnold and Moizer (1984) reported that 87% of the 202 investment analysts in their survey "almost always" estimated future dividend yields when valuing shares. In Pike et al. (1993) dividend information was ranked third behind price/earnings ratios and net assets per share, in terms of usefulness for share valuation.

[3]     A small minority of authors have suggested that dividend cuts may not be seen as bad news by investors; instead, they may indicate that a company has profitable investments which it wishes to fund from internal cash resources. For example, Woolridge and Ghosh (1984) cite the example of Gould Inc. where the share price increased by 2% on news of a 60% cut in dividends; the authors suggested that Gould's management had convinced the market that the reduction in dividends was to be used to fund innovative investment.

[4]     Researchers such as Watts (1973), Johnson and Jensen (1995) and De Angelo, De Angelo and Skinner (1996) argue that dividends do not act as a signal to investors about the future prospects of the firm. These authors point out that dividend changes tend not to be followed by earnings changes of the same sign. For example, they demonstrate that dividend cuts usually indicate that the firm has already experienced several years of financial pressure and are usually followed by earnings increases.

increasingly popular in explaining the prices of a variety of financial assets. In particular it has been applied to test present value models for stock prices:

$$P_t - \frac{D_t}{R} = \left(\frac{1}{R}\right) E_t \left[\sum_{i=0}^{\infty} \left(\frac{1}{1+R}\right)^i \Delta D_{t+1+i}\right] \qquad \textbf{(1)}$$

where $P_t$ denotes the real stock price at time $t$, $D_t$ is the time $t$ real dividend, $R$ is the expected real return (assumed constant), a $\Delta$ denotes first differences and $E_t$ denotes the rational expectations operator. If both (real) stock prices and dividends are non-stationary then, under a no-bubbles assumption such that the right-hand-side of (1) is $I(0)$, $P_t$ and $D_t$ will be cointegrated with the cointegrating vector equal to $[-1, 1/R]$. For example, Shiller (1981), Campbell and Shiller (1987, 1988 and 1989) and MacDonald (1993) have employed this technique to investigate whether the current stock price is equal to the value of the future stream of dividends discounted geometrically back to the present by a constant real interest rate.[5]

However, the results from these tests have been unimpressive. For example, using annual US price and dividend data for the Standard & Poor's Composite Index from 1871 to 1979, Campbell and Shiller (1987) were unable to reject the hypothesis that no long-run relationship existed between these two variables at the 10% level; their estimate of the real rate of interest from employing the Engle-Granger two-step cointegration test of 3.2% was less than half the average value of this variable (8.2%) over the sample period. The findings of Phillips and Ouliaris (1986) were even less impressive; these researchers were unable to reject the null hypothesis of no cointegration at the 25% level.

MacDonald and Power (1995) attribute these poor results to a failure to incorporate a

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[5]     Campbell and Shiller (1987) and MacDonald and Speight (1991) have used this technique to examine the value of bonds while MacDonald and Taylor (1993) have adopted this approach when analysing spot foreign exchange rates.

measure of the company's growth prospects into the model. Once this deficiency is remedied, they argue that "a unique long-run relationship can be shown to exist". Specifically, they demonstrate that the inclusion of a retention ratio variable[6] into the analysis to proxy for future growth prospects yields a more accurate estimate of the real rate of interest of 7.2%. They also calculate the error correction model implied by the cointegrating relationship which shows how prices adjust to their long-run equilibrium levels and indicate that the predicted values from this model "track the actual data well, capturing most of the important turning points". Their results demonstrate that this error correction model forecasted values of the Standard & Poor's Index accurately in a twelve year out-of-sample period from 1976-1987; with the exception of 1987 when the market crashed, all the forecasted values lay within two standard errors of the actual observations. The authors suggest therefore that the poor findings of the previous studies may be attributable to an incorrect specification of the present value relationship.

A number of criticisms can be levelled at these previous studies. First, they all employ the same US data set in their investigations; price and dividend information for the Standard and Poor's Composite Index are used in all three studies. Therefore the findings may be country specific and may not apply outside the USA. Second, no dividend information are provided by Standard & Poor prior to 1926 for this index; the dividend data for the early years of these investigations are constructed by adjusting a data series in Cowles (1938).[7] Again, the conclusions may be the result of data problems from splicing together two different dividend

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[6]     This retention ratio is defined as the difference between earnings and dividends divided by the reported earnings for that year. MacDonald and Power argue that since most firms finance investment out of retained profits, the higher this ratio the greater the likelihood that firms will invest in new capital expenditure projects which will lead to growth in the future; this prospect of growth, they suggest, is incorporated into share values.

[7]     Indeed, the dividend data employed in Campbell and Shiller (1987) differed from that used in earlier papers such as Shiller (1981); they incorporate corrections to the dividend series in Cowles' original book which were highlighted in the second edition of Cowles' text. See chapter 26 in Shiller (1989) for a full description of the data sets employed in these studies.

series. Third, the three studies have analysed data stretching back to the 1870s to estimate the long-run cointegrating relationship between prices and dividends. Of course, the nature of the equity market in the US has changed dramatically over the 119-year period spanned by these studies; regulations, technology, trading procedures and the number and expertise of market participants have altered over this long time span. Previous results, therefore, may not adequately characterise the advanced capital markets of the 1970s, 1980s and 1990s. Fourth, the studies have employed different specifications of the model when undertaking their tests. For example, Campbell and Shiller (1987) and West (1988) test the basic expected present value model, Campbell *et al*. (1997) study an approximate present value relationship, while MacDonald and Power (1995) examine augmented present value models. Comparison of the results is therefore difficult. Finally, researchers such as Campbell and Shiller (1987) have conceded that the power of their test of the present value model for stocks is "low"; their failure to uncover a long-run relationship between stock prices and dividends may be the fault of their test rather than the absence of such a relationship in practice.

This paper attempts to address these limitations. We test the present value model for 56 UK equities over a relatively short time period using a panel cointegration approach. This novel method enables researchers to test for a long-run relationship over a smaller number of years by exploiting the cross-sectional variation between stocks as an alternative to extending the time series dimension.

The remainder of the paper is organised as follows. Section 2 describes the data used in this study while Section 3 investigates the time series properties of this data. Section 4 outlines the panel cointegration test and analyses the results of this test for our sample of UK companies. Finally, Section 5 offers some preliminary conclusions and highlights avenues for future research.

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## 2    Data sources

All firms on the London International Official Stock Exchange List according to the Datastream database in December 1996 were initially considered for this study. From this list of approximately 1600 companies, only those firms which satisfied a number of criteria were included in the final sample. First, companies had to have a full set of stock price information and a complete set of dividend data over the period 1968 to 1996. Second, companies had to have a December year-end date and to retain this financial year end throughout the 29-year time span. Third, only companies which were not classified as having been taken over were included in the final sample. Some 56 companies satisfied these selection criteria and are analysed in this study. This sample is drawn from a wide group of industries[8] and the companies vary in size from £5,181m to £23m; the findings should therefore not be specific to any one sector or type of firm. This sample suffers from a survivorship bias since only those companies which have existed for the full 29 year period are included in the analysis. However, to study such a long-run relationship between dividends and prices at an individual company level such bias is inevitable. Nevertheless, this bias may have implications for interpreting the econometric results, and they should be interpreted with this caveat in mind.

For each company, annual stock price information was obtained from Datastream from 31st December 1968 to 31st December 1996. Obviously other dates could have been chosen but December has the benefit that a large proportion of UK companies chose to report their annual results in this month (Barron, 1984)[9] and financial statement information for all the firms in the

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[8]     There were 15 industries represented in our sample of 56 companies. However, the sample was not evenly spread among these different sectors; 12 firms were located in the Engineering (General) sector, 4 firms in the Chemical sector and 3 firms in both the Textiles and Electrical sectors. The remaining companies were drawn from the Oil, Health, Leisure, Vehicle Engineering, Transport, Food and Packaging sectors among others. Therefore, as one would expect, some of the newer industries are not represented in the sample.

[9]     Barron (1984) found that some 41.1% of the UK firms in his sample chose December as their financial year-end date. March was the next most popular month (20.4%) and September was ranked third (8.8%).

sample related to the 12-month period ending on that date. Dividend information for each share[10] was also extracted from the Datastream database; the annual dividend paid (Item 434) was obtained for each company. The series were deflated by the consumer price index and those data were then analysed to determine whether there is a long-run relationship between real stock prices and real stock dividends using the cointegration methodology for a panel data set.

## 3    Orders of Integration

### 3.1    *Single equation test of stationarity*

In this section we investigate the orders of integration of the real price and dividend series. Single-equation (augmented) Dickey-Fuller statistics of the price and dividend series are calculated. The lag length in the ADF test is allowed to vary across individual companies so as to mop up any residual serial correlation. To conserve space the results are not reported in full but very few series reject the null hypothesis of non-stationarity. Using small-sample critical values from MacKinnon (1991), only twelve price and five dividend series could be deemed stationary at the 5% level.

However, the power of this test is known to be very low and recently proposed panel approaches to unit root testing can also be applied to analyse our data set. The panel approach gains power from the cross-equation restriction on the first-order partial autocorrelation functions. Specifically, we use the Levin and Lin (1992, 1993) test which we now briefly outline.

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[10]    This information relates to "ordinary and preference dividends paid during the period, often representing the previous years final and current years interim dividend". Therefore, preference dividends are included in the data, however, we do not believe that this inclusion significantly effects our analysis of ordinary equities since (i) 11 of firms in the analysis had no preference shares in issue throughout the period being examined and (ii) those firms that did issue preference shares tended to pay a relatively constant preference dividend from one year to the next; indeed, 5 firms maintained the same preference dividend throughout the 29-years studied. Unfortunately, dividend information relating solely to ordinary shares was not available for the whole time period examined in this study.

## 3.2    *Panel test of stationarity*

Let $x_{it}$ denote the panel series under consideration, where $i$ is the cross section dimension that runs from 1 to $N$. Perform the following regressions

$$\Delta x_{it} = \sum_{L=1}^{p_i} \hat{\pi}_{iL} \Delta x_{it-L} + \hat{\alpha}_{mi} d_{mt} + \hat{\varphi}_{it} \qquad (2)$$

$$x_{it-1} = \sum_{L=1}^{p_i} \hat{\varphi}_{iL} \Delta x_{it-L} + \hat{\gamma}_{mi} d_{mt-1} + \hat{v}_{it-1} \qquad (3)$$

where a circumflex indicates a fitted value, $\varphi_{it}$ and $v_{it-1}$ represent estimated residuals, $p_i$ measures the idiosyncratic lag length for share $i$, and $d_{mt}$ denotes an appropriate set of deterministic variables defined for three alternative models ($m = 1, 2$ or 3). Thus $d_{1t}$ contains no deterministic variables, $d_{2t}$ contains a constant term and $d_{3t}$ contains both a constant and deterministic trend terms. Then, the *t*-ratio on $\delta_i$ in

$$\hat{\varphi}_{it} = \hat{\delta}_i \hat{v}_{it-1} + \hat{\varepsilon}_{it} \qquad (4)$$

is the panel equivalent of a univariate Dickey-Fuller unit root test. To control for heterogeneity across individuals both $\varphi_{it}$ and $v_{it-1}$ are deflated by the standard error of equation (3) and subsequently denoted by a tilde. Levin and Lin demonstrate that the *t*-ratio on $\delta$ in

$$\tilde{\varphi}_{it} = \tilde{\delta} \tilde{v}_{it-1} + \tilde{\varepsilon}_{it} \qquad (5)$$

has a standard normal distribution for Model 1 but diverges to negative infinity for Models 2 and 3. Fortunately the following adjusted *t*-ratio has been constructed to overcome this problem; it has a standard normal distribution and can be used to test the null hypothesis that $\delta_i = 0$ for $i = 1$ to $N$

$$t_\delta^* = \frac{t_\delta - NT_{AVG}\hat{S}_{NT}\hat{\sigma}_\varepsilon^{-2}RSE(\hat{\delta})\mu_{m\tilde{T}}^*}{\sigma_{m\tilde{T}}^*} \qquad (6)$$

where $\hat{S}_{NT} = \frac{1}{N}\sum_{i=1}^{N}\frac{\hat{\sigma}_{xi}}{\hat{\sigma}_{\varphi i}}$ , $\sigma_{\varphi i}$ is the residual standard error from (4), $\sigma_{xi}$ is the long-run standard deviation of $x$, $RSE(\delta)$ is the reported standard error of $\delta$, $\sigma_\varepsilon$ is the standard error of equation (5), $T_{AVG}$ is the average number of observations per individual in the panel (= $T$ - $p_{AVG}$ - 1), and $p_{AVG}$ is the average lag order of the individual ADF statistics, $p_i$. Finally, $\sigma_{m\tilde{T}}^*$ and $\mu_{m\tilde{T}}^*$ denote mean and standard deviation adjustments given in Table 2 of Levin and Lin (1993) for each model $m$.

The results of applying this procedure to our data set are given in Table 1. The Levin and Lin procedure requires that the data be generated independently across individuals in the panel. They suggest that this assumption may be facilitated by the removal of time specific aggregate effects (which is done by subtracting the cross section averages from the observed data before starting the above regressions). We report statistics both with and without this adjustment for Levin and Lin's Models 2 and 3.

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*Table 1*
*Panel Unit Root Tests*

|  | Real Prices | | Real Dividends | |
| --- | --- | --- | --- | --- |
|  | Model 2 | Model 3 | Model 2 | Model 3 |
| With time dummies | 1.42 | 11.20 | -5.91 | -1.76 |
| Without time dummies | 3.61 | 9.77 | 2.14 | -3.64 |

*Notes*:  The tests are distributed as standard normals.  Test statistics below -1.96 indicate rejection of the null of non stationarity.

The tests indicate that the real stock price series are $I(1)$, irrespective of (i) whether the deterministic variables are included in the system or (ii) whether cross sectional time dummies are incorporated into the analysis.  For real dividends the picture is less clear cut; the test results depend on whether time dummies are included in the equations and whether a trend is included. If both time dummies and a trend are included, or if both are excluded, the Levin and Lin test statistics fail to reject non stationarity.  If either time dummies or a trend are included, the data indicate stationary real dividends.  We experimented with an alternative panel unit root test procedure due to Im, Pesaran and Shin (1997).  Unfortunately, this too gave mixed results which were very sensitive to the lag lengths specified in the individual regressions.  Campbell and Shiller (1987) also note that "there is some evidence for trend stationarity of log dividends". This trend may be attributable to managers' reluctance to raise dividends in line with increases in earnings; instead, Lintner (1956) reported that companies moved to a target dividend level (based on a percentage of earnings) over a period of three years. Lintner explained this caution in terms of managers' unwillingness to cut dividends paid to investors.  Fama and Babiak (1968) modelled the time series behaviour of US companies' dividend payments and agreed with Lintner's conclusions that managers did not wish to increase dividend payments which might have to be cut in the near future.

While our integration results for the dividend series are ambiguous we will proceed to the

cointegration tests for two reasons. First, simply eyeballing the data would seem to indicate that the series in question are non stationary. Second, we believe that the panel test of cointegration outlined in section 4.2 provides a better framework for assessing the stationarity of prices, dividends, and combinations of the two. Evidence of cointegration between stock prices (which are unequivocally $I(1)$) and dividends (uncertain order of integration) would imply some degree of support for the present value model *and* that dividends are indeed $I(1)$.[11]

## 4    Cointegration

### *4.1    Single equation cointegration*

Before analyzing the panel cointegration properties of our data set we first examine whether individual companies' stock prices and dividends are cointegrated using the standard Engle-Granger approach. If the 5% critical value of -3.554 from MacKinnon (1991) is employed, just four of our 56 panellists reject the null hypothesis of no cointegration; the findings confirm the evidence from previous US studies that there appears to be no long-run relationship between stock prices and dividends which the present value model suggest ought to exist. However, the power of this single equation approach to cointegration has been shown to be low and, as with our unit root testing, in the following subsection we shall appeal to the power of the panel approach.

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[11]    Cointegration would not imply non-stationary dividends if the coefficient on stock prices is zero, but this is ruled out by the normalisation imposed in the following empirical work.

*4.2    Panel cointegration*

If the coefficients in the cointegrating relationship between stock prices and dividends were known *a priori* rather than having to be estimated, the Levin and Lin procedure could be applied to the constructed variable in a panel framework.[12] However, analogous to the fact that Dickey-Fuller critical values are not applicable to generated residuals from a cointegrating regression, Levin and Lin critical values are (usually) inappropriate in a panel context. Furthermore, Pedroni (1995) notes that while in the single series case the dependency of the residuals on the distributional properties of the estimated coefficients in the spurious regression can be accounted for by simply altering the critical values, the effect can be much more dramatic, and less easy to remove, in a panel because of the cross sectional dimension to the structure of the residuals. The effect of this dependency also hinges crucially on the alternative hypothesis being tested. Consider a homogeneous panel

$$y_{it} = \alpha + \gamma_t + x_{it}\beta + e_{it} \qquad (7)$$

where the alternative hypothesis states that a cointegrating relationship is similar across the different members of the panel. Pedroni shows that a superconsistency-type result means that in this case the asymptotic distributions of unit root tests are invariant to whether the residuals are known or estimated. However, in the more general case of a heterogeneous panel

$$y_{it} = \alpha_i + \gamma_t + x_{it}\beta_i + e_{it} \qquad (8)$$

where the cointegrating relationship can differ between individual panellists, the standard panel unit root test statistic becomes nonconvergent, with serious implications for the test statistic.

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[12]    This method has been used to test for strong-form purchasing power parity (Frankel and Rose, 1996) and for cointegration between spot and forward exchange rates (MacDonald and Marsh, 1997).

Pedroni (1995) presents three tests of the null hypothesis that $e_{it}$ are non stationary: $Z_{\hat{v}_{NT}}$ based on variance ratios, $Z_{\hat{\rho}_{NT}}$ based on residual autoregressions, and $Z_{t_{\hat{\rho}_{NT}}}$ based on a $t$ test; this latter test corresponds to a Dickey-Fuller or Phillips-Perron-type test depending on the method used to estimate cross sectional covariances. Specifically, he defines

$$Z_{\hat{v}_{NT}} = \left( \sum_{i=1}^{N} \sum_{t=1}^{N} \hat{L}_{11i}^{-2} \hat{e}_{it-1}^2 \right)^{-1}$$

$$Z_{\hat{\rho}_{NT}} = \left( \sum_{i=1}^{N} \sum_{t=1}^{N} \hat{L}_{11i}^{-2} \hat{e}_{it-1}^2 \right)^{-1} \sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \left( \hat{e}_{it-1} \hat{e}_{it} - \hat{\lambda}_i \right) \qquad (9)$$

$$Z_{t_{\hat{\rho}_{NT}}} = \left( \tilde{\sigma}_{NT}^2 \sum_{i=1}^{N} \sum_{t=1}^{N} \hat{L}_{11i}^{-2} \hat{e}_{it-1}^2 \right)^{-1/2} \sum_{i=1}^{N} \sum_{t=1}^{T} \hat{L}_{11i}^{-2} \left( \hat{e}_{it-1} \Delta \hat{e}_{it} - \hat{\lambda}_i \right)$$

where $L_i$ is the lower triangular Cholesky decomposition of a consistent estimate of the asymptotic covariance matrix of the residuals from a first order VAR of $x_i$ and $y_i$, and $\tilde{\sigma}_{NT} = \frac{1}{N} \sum_{t=1}^{N} \left( \hat{L}_{11i} \hat{\sigma}_i \right)^2$. The variable $\hat{\lambda}_i$ is based on the estimated residuals from the autoregression $e_{it} = \rho e_{it-1} + \mu_{it}$, such that $\hat{\lambda}_i = \frac{1}{2} \left( \hat{\sigma}_i^2 - \hat{s}_i^2 \right)$, $\hat{s}_i^2 = \frac{1}{T} \sum_{t=1}^{T} \mu_{it}^2$, and $\hat{\sigma}_i$ is the standard error of $\mu_{it}$. Pedroni (1995) derives asymptotic distributions for each of these test statistics for both homogeneous and heterogeneous panels, together with Monte Carlo-based finite sample test statistics. He finds that the finite sample distributions quickly approach asymptotic distributions, even in relatively small panels. As in the single series case, large positive values of the panel variance ratio test statistic indicate rejections of the null hypothesis, while for the panel autocorrelation and $t$-tests large negative values indicate rejections of the null of no cointegration.

We first test for cointegration in the heterogeneous model where both the cointegrating vector (slope) and intercept term can differ between individual panel members. The results of applying the Pedroni tests to this panel are given in the first two columns of figures in Table 2. The first column reports test statistics when the model allows for the possibility of time specific

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aggregate effects ($\gamma_t \neq 0$ in equation (8)). There is clear evidence that stock prices and dividends cointegrate in this framework since three of the four test statistics are significant at standard confidence levels. This conclusion is unchanged when time effects are excluded from the model (column 2), indicating that common cross-sectional effects may not be important.[13]

*Table 2*
*Panel Cointegration Tests*

| | Heterogeneous Slopes | | Homogeneous Slope | |
|---|---|---|---|---|
| | $\gamma_t \neq 0$ | $\gamma_t = 0$ | $\gamma_t \neq 0$ | $\gamma_t = 0$ |
| Variance Ratio | 39.63 | 39.96 | 21.68 | 38.05 |
| Autocorrelation Test | -193.22 | -190.04 | -196.47 | -179.22 |
| Phillips-Perron *t*-test | -47.84 | -55.56 | -65.18 | -52.64 |
| Dickey-Fuller *t*-test | -24.91 | -24.43 | -20.79 | -20.98 |
| OLS Point Estimate | | | 16.55 | 17.40 |
| FMOLS Point Estimate | | | 17.04 | 17.23 |

*Notes*: The first four rows give the test statistics of the null of no cointegration between real stock prices and real dividends detailed in the text. For the variance ratio, autocorrelation and *t*-tests respectively, the one (five) and [ten] percent critical values for a T=25, N=50 panel are 89.74 (82.38) [78.64], -53.58 (-49.75) [-47.82], and -15.74 (-14.99) [-14.62].

We have strong evidence that cointegration exists between stock prices and dividends if we allow the cointegrating vector (i.e. the slope coefficient or, equivalently, the inverse of the discount rate) to differ across the panel members. It is plausible for expected dividend streams to be discounted at different rates if companies have different risk characteristics. Estimates of these slopes range from 0.9 to 76 with a mean value of 20.36 (an implied discount rate of

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[13] This strong finding of cointegration suggests that the dividend series are indeed non stationary. Since the left hand side variable, the stock price, is unambiguously *I*(1), this must be combined with another integrated series for stationary residuals to occur.

4.91%).[14]

If risk is ignored, then for the present value model estimated for stocks traded within one financial centre we might expect the discount rate to be equal for all companies.  Thus we estimate the following equation

$$y_{it} = \alpha_i + \gamma_t + x_{it}\beta + e_{it} \tag{10}$$

Pedroni (1995) shows that for a cointegrating relationship such as this with homogeneous slopes, the asymptotic critical values for the tests described in equation (9) will be identical to those of the raw panel unit root tests under the additional assumption that the regressors are strictly exogenous.  This is almost certainly not true in our case and so we proceed with caution.  When testing the residuals from (10) we shall also consider the higher critical values obtained for the heterogeneous model.  These results are reported in the final two column of Table 2.  Again, and irrespective of whether $\gamma_t$ is restricted to zero, even using the more conservative critical values we have strong evidence of cointegration.  Furthermore, the OLS point estimate of the cointegrating vector implies a very plausible real discount rate of 5.75% (6.04% with time dummies).  Pedroni (1996) shows how to obtain unbiased, fully-modified OLS estimates of the common slope coefficient in a panel cointegrating regression.  These are reported in the final cells of Table 2 and imply an equally realistic discount rate of 5.81%, or 5.87% when time effects are incorporated.

While the average result is fairly realistic, the overall conclusion from the analysis is disappointing.  A large proportion of the implied discount rates are either very low or unusually high.  This finding may suggest that other variables are missing from the analysis of the

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[14]  The authors attempted to relate these discount rates to (i) estimated beta values and (ii) gearing levels but no significant relationships were found.

relationship between dividends and stock prices; further work using other important variables may yield more realistic discount rates for a larger proportion of the sample.


## 5    Conclusions

One of the most important puzzles within the finance literature is the relationship between share prices and share dividends.  Behavioural research based on interviews with finance directors suggests that a great deal of care and attention is paid when determining the dividend level which is paid to shareholders each year.  In addition, the results of market-based event studies indicate that news of a dividend change appears to cause share prices to alter although the direction of the share price change remains a matter of contention among different empirical studies.  However, attempts to investigate the relationship between share prices and share dividends using time series analysis have generally been unsuccessful.  Studies document either no significant relationship between the two variables (Campbell and Shiller, 1987; Phillips and Ouliaris, 1986) or only detect evidence of a significant link between the two measures once additional variables such as the payout ratio are included in the analysis (MacDonald and Power, 1995).

One possible explanation for these disappointing time series results is that the power of previous tests is relatively low because of the long data set needed to undertake such an analysis and because the underlying character of most companies varies over such a long time span because of mergers, takeovers and other forms of corporate restructuring.  Another reason for the relatively unimpressive findings is that most of the analysis to date has concentrated on one US data set - the S&P composite index.  Both of these limitations hinder researchers' ability to determine whether their results have uncovered important evidence about the validity of the present value model (and the efficiency of the stock market).

This present paper overcomes some of these limitations by employing the newly

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developed panel cointegration technique which provides a much stronger test using a smaller time series of data. Specifically, the paper investigates the relationship between share prices and share dividends for a panel of 56 UK firms over the recent time period from January 1968 to December 1996. The results indicate that there is a significant cointegrating relationship between share prices and share dividends for our sample of firms. More importantly, the cointegrating vector implies a mean real discount rate of 5.8% which is similar to the average real rate of interest over the time period studied; previous US investigations have obtained estimates of the discount rate which have been much lower in value. Of course, implied discount rates for individual securities varied widely around this mean level, and further work is needed on why this variation occurs.

Overall, the findings of this paper support the present value model which links dividend payments with share prices, although the actual nature of the relationship needs to be further examined. The work of the present paper also highlights a number of avenues for further research in the future. For example, a larger sample of companies would facilitate an industry comparison of the discount rates which appear to be applied to corporate dividends when arriving at share prices. In addition, the role of growth prospects in affecting the long-run discount rate linking dividends and pries might yield valuable insights into the operation of the equity markets. Nevertheless, the empirical evidence in this paper accords with basic finance theory and suggests that these two variables are related in a systematic fashion.

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