Dividends, Prices and the Present Value Model: Firm-Level Evidence

John Goddard
School of Business and Regional Development
University of Wales, Bangor

David G. McMillan
School of Management
University of St Andrews

John O.S. Wilson
School of Management
University of St Andrews

January 2006

Abstract

Recent stock price movements have led to a re-examination of the present value model. Typically, empirical studies have employed a long span of US stock market index data, and have attributed a failure to detect cointegration to the presence of bubbles. This study considers UK firm-level data, and implements panel unit root and cointegration tests. Recent panel tests that allow for cross-sectional dependence control for factors such as bubbles that may result in temporary deviations from the long-run price-dividend relationship. The panel test results largely support the present value model, yielding evidence of cointegration between log real prices and dividends.

Keywords: Stock Prices, Present Value Model, Firm-Level Data
JEL: C22, G12

Address for correspondence: Professor John Wilson, School of Management, University of St Andrews, The Gateway, North Haugh, St Andrews, Fife KY16 9SS, UK. Tel: +44 (0)1334 462803. Fax: +44 (0)1334 462812. e-mail: jsw7@st-and.ac.uk

1. Introduction.

The notion that current security prices depend upon the present value of discounted future dividends, where the discount rate is equivalent to the required rate of return, ranks as one of the core principles of finance theory. Recently, this relationship has been the subject of renewed attention in the empirical literature.[1] This recent interest is primarily the result of the large rise in stock prices at the end of the last century, and the subsequent fall. Currently, the general consensus view is that fundamentals were basically unchanged throughout this period: the market became significantly overvalued, and fundamentals subsequently reasserted themselves. Campbell and Shiller (2001) were among the most prominent proponents of this view during the growth phase of the late-1990s bubble. An alternate view was that the rise in stock prices reflected investors' permanently revised expectations of higher future earnings and dividends, due primarily to productivity gains originating from technological change. Furthermore, following a period of relatively low nominal interest rates and inflation during the 1990s, there may have been a tendency for investors to reduce the rates at which they discount future dividends.[2]

Previous empirical analysis of present value models, and of the long-run relationship between prices and dividends, is based predominantly upon two cointegration approaches. First, assuming a constant discount rate, real prices and real dividends should cointegrate, that is, exhibit a stable long-

1. A non-exhaustive list includes Balke and Wohar (2001, 2002), Psaradakis, Sola and Spagnola (2004), Caporale and Gil-Alana (2004), Bohl and Siklos (2004) and Nasseh and Strauss (2004).

2. The so-called 'new era' explanations are based on the view that the technological development primarily in the telecommunications and information technology sectors led to an expectation on the part of investors of permanently higher future earnings and dividends (Greenwood and Jovanovic, 1999; Browne, 1999). An alternative theory was that investors reduced the rate at which they discount future dividends, perhaps as a result of increased ease of trading (Siegel, 1999; Heaton and Lucas, 1999; Fama and French, 2002).

run relationship (Campbell and Shiller, 1987). In this case, the cointegrating parameter depends upon the discount rate. Second, if allowance is made for a time-varying discount rate, the difference between log dividends and log prices should exhibit stationarity (Campbell and Shiller, 1988a,b). However, the empirical evidence is mixed. Specifically, in Campbell and Shiller's (1987) original paper, as well as subsequent work based on similar methodology (Diba and Grossman, 1988; Brooks and Katsaris, 2003; Kapetanios et al., 2003), the results of tests for cointegration between prices and dividends are ambiguous. Similarly, support for the present value model from tests for the stationarity of the log price-dividend ratio is in short supply, with several studies reporting evidence of non-stationarity (Froot and Obstfeld, 1991; Lamont, 1998; Balke and Wohar, 2002).

While the majority of extant studies of the relationship between prices and dividends have examined the long-term relationship between a stock price index and a dividend index for the country of interest,[3] in this paper the empirical analysis is based on pooled firm-level price and dividend time-series data. To our knowledge, empirical analysis of the present value model has previously been conducted at firm level only by Nasseh and Strauss (2004), using US stock market data. The use of firm-level data enables us to observe patterns and relationships in the data that may be obscured at the aggregate stock market level by smoothing induced through aggregation. Cohen et al. (2001), Vuolteenaho (2002) and Jung and Shiller (2005) suggest that the present value model is more likely to hold at the level of the individual firm than at the aggregate (stock market index) level. Information on the cash flows and future prospects of individual firms is well understood by investors. "In contrast, there would seem not to be the same kind of clarity in the market about changes in the aggregate dividend or earnings flow for the stock market of a country. Changes in

3. The S&P 500 index has been the most extensively analysed. In addition to the sources cited above, see Lee (1995), Sung and Urrutia (1995), Timmerman (1995), Crowder and Wohar (1998).

these flows for the aggregate stock are less dramatic than for individual firms, because the aggregate averages out the individual stories of the firm and the reasons for changes in the aggregate are more subtle and harder for the investing public to understand, having to do with national economic growth, stabilizing monetary policy, and the like. If changes in aggregate dividends are harder to predict, we might expect that factors other than information about fundamentals ... would swamp out the effect of information about future dividends in determining price" (Jung and Shiller, 2005, p.221-222).

A major preoccupation of the recent econometrics literature on testing for unit roots and cointegration in dynamic panel data sets has been the development of tests that control for cross-sectional dependence, or non-zero covariance between the disturbance terms of the autoregressions that describe the time-series behaviour of the variable in question, for some or all of the panel members. In the present case, this correction for cross-sectional dependence would control for the effects of factors such as bubbles or other non-fundamental disturbances, which may result in temporary deviations from the long-run relationship between stock prices and dividends, and which may operate in a similar (but not necessarily identical) fashion on the price and dividend series of some or all of the sample member firms simultaneously. In other words, the application of panel unit root and panel cointegration tests that control for cross-sectional dependence offers a novel and powerful method for circumventing the difficulties created for tests of the present value model at the aggregate level by bubbles or other non-fundamental disturbances, that result in temporary deviations from the long-run equilibrium relationship between prices and dividends.

Specifically, this paper employs the cross-sectionally augmented panel unit root test that is developed by Pesaran (2005). This test deals with the issue of cross-sectional dependence by including the cross-sectional means of the current and lagged first-differences, and of the lagged levels, alongside the (more familiar) individual-level lagged first-difference terms, in the augmented

Dickey-Fuller autoregressions that are used to test for the stationarity of the series for each panel member. Because the cross-sectionally augmented test has not been used extensively in the previous accounting and finance literature, we also apply the more widely known first-generation panel unit root tests of Levin et al. (2002) and Im et al. (2003), which do not allow for cross-sectional dependence, to the same data, enabling us to draw some direct comparisons between the two sets of test results. For cointegration testing, we adopt a similar approach. A cointegration analogue of the Pesaran (2005) cross-sectionally augmented panel unit root test (as described above) is employed, alongside the more familiar portfolio of panel cointegration tests that were developed in a series of papers by Pedroni (1999, 2001, 2004), and which are included primarily for purposes of comparison.

The sample for the present study comprises 104 UK non-financial companies, for which complete and continuous share price and dividend series were available over a 34-year observation period from 1970 to 2003 (inclusive). As far as we are aware, this is the first firm-level investigation of the validity of the present value model based on UK data. The remainder of the paper is structured as follows. Section 2 outlines the theoretical motivations for the paper. Section 3 provides the technical details of the panel unit root and panel cointegration tests. Section 4 reports and interprets the results of these tests. Finally, Section 5 summarises and concludes.

2. The present value model.

The rational expectations present value model that relates the real stock price to the discounted value of expected future real dividends can be written as follows:

(1) $$P_t = \sum_{i=1}^{\infty} \delta^i E_t D_{t+i} .$$

where P is the real stock price, D is the real dividend, and $\delta = (1+R)^{-1}$ is the (constant) discount rate. E_t is the expectations operator conditioned on information up to t. This formula represents the fundamental value for prices and, using the transversality condition ($\lim (k^n E_t D_{t+n}) = 0$, as $n \to \infty$), ensures a unique price. Campbell and Shiller (1987) show that if the present value model is valid, the real stock price and real dividend are cointegrated, with a cointegrating vector $(1, R^{-1})$:

$$(2) \qquad P_t - R^{-1} D_t = R^{-1} E_t [\sum_{i=0}^{\infty} (1+R)^{-i} \Delta D_{t+1+i}].$$

Therefore tests of the present value model involve tests for cointegration between real prices and real dividends. Empirical evidence for the presence of cointegration is ambiguous, however (Campbell and Shiller, 1987; Diba and Grossman, 1988; Kapetanios et al., 2003).

Relaxing the restriction of a constant discount rate to allow for time-variation in the discount rate, complicates the analysis through the introduction of non-linear terms. However, Campbell and Shiller (1988a,b) suggest a log-linear approximation that has subsequently been used widely:

$$(3) \qquad p_t = [k/(1-\rho)] + E_t [(1-\rho) \sum_{i=0}^{\infty} \rho^i d_{t+i+1} - \sum_{i=0}^{\infty} \rho^i r_{t+i+1}]$$

where lower-case symbols (p, d, r) denote the natural logarithms of prices, dividends and the discount rate, respectively. k and ρ are linearisation parameters: $k = -\ln(\rho) + (1-\rho)\ln(\rho - 1)$, and $\rho = 1/[\exp(\overline{d-p})]$. Imposing the transversality condition as before, which rules out explosive behaviour, (3) can be rewritten in terms of the log price-dividend ratio:

(4) $$p_t - d_t = -k(1-\rho)^{-1} + E_t \sum_{i=0}^{\infty} \rho^i (\Delta d_{t+i+1} + r_{t+i+1})$$

The log price-dividend ratio is stationary provided changes in dividends and the discount rate are stationary. By implication, log prices and log dividends are cointegrated, with a cointegrating vector $(1, -1)$. Therefore empirical verification of (4) requires only testing for the stationarity of the log price-dividend ratio, and does not require estimation of the (unknown) cointegrating parameter, or deflation of nominal values using any price index. Intuitively, (4) states that if dividends are expected to grow, the current price will be high and the price-dividend ratio will be high; and if the future discount rate is expected to be high, the current price will be low and the price-dividend ratio will be low.

As before, however, empirical support this model is limited, with several researchers reporting evidence of non-stationarity in respect of the log price-dividend ratio (Froot and Obstfeld, 1991; Lamont, 1998; Balke and Wohar, 2002). Crucially, both versions of the present value model depend upon the transversality condition to ensure a unique price, namely the fundamental stock price. Evidence of non-stationarity in respect of the price-dividend ratio, or failure to detect cointegration between real prices and real dividends, could be due to the presence of non-fundamental components in stock prices, which lead to violations of the transversality condition, invalidating the (linear) present value models described above.

In seeking to explain this failure, the extant literature has focused upon the presence of non-fundamental components that result in temporary deviations from the equilibrium relationship, arguing that such deviations notwithstanding, the linear present value model is valid as a long-run relationship: a stable long-run attractor point between prices and dividends exists. Typically, the non-fundamental component has been ascribed to the presence of a rational bubble, such as the speculative, periodically collapsing bubbles of Evans (1991). This approach takes the view that the price contains both the fundamental value as in (1), and a non-fundamental or bubble component:

(5) $$P_t = \sum_{i=1}^{\infty} \delta^i E_t D_{t+i} + B_t$$

B_t is the bubble component, which must satisfy two criteria: first, B_t is non-stationary, such that log prices and log dividends cannot be cointegrated $(1, -1)$; and second, B_t is non-negative. However, the empirical detection of bubbles is difficult (Hall et al., 1999). More importantly, serious theoretical objections to the notion of rational bubbles have been raised by Campbell et al. (1997, p259-260). Nevertheless, several authors have identified modelling approaches that may be consistent with such bubble dynamics, typically using non-linear specifications (van Norden, 1996; van Norden and Vigfusson, 1998; Bohl and Siklos, 2004; Psaradakis et al., 2004; Kanas, 2005). The empirical results reported in this literature, as well as a number of studies that test for linear cointegration using index data (cited in the Introduction), are generally supportive of the hypothesis of bubbles.

3. Panel tests for stationarity and cointegration.

This section describes the panel unit root tests and panel cointegration test procedures that are used in this study to test for a cointegrating relationship between the log real price and log real dividend series, and for the stationarity of the log price-dividend ratio series.

In order to investigate the possibility of a cointegrating relationship, the log real price and log real dividend series are initially tested for non-stationarity using univariate panel unit root tests. We implement the Levin et al. (2002) and Im et al. (2003) tests, subsequently referred to as the LLC and IPS tests, respectively. We also implement the cross-sectionally augmented version of the IPS test that has recently been developed by Pesaran (2005), subsequently referred to as the CIPS test. The CIPS test provides a computationally straightforward adjustment for the presence of cross-sectional dependence between the disturbance terms of the augmented Dickey Fuller (ADF) autoregressions for the individual series. In the present case, and as noted in the Introduction, this adjustment controls for the effects of bubbles or other non-fundamental disturbances that result in temporary deviations from the long-run equilibrium price-dividend relationship, and which may operate in a similar (but not necessarily identical) manner on the log real price and/or log real dividend series of some or all of the sample firms simultaneously.

After failing to reject the null hypothesis of non-stationarity in the majority of these tests on the log real price and log real dividend series, we test for a cointegrating relationship between these two series, using panel analogues of the Engle and Granger (1987) two-step method. Evidence of a cointegrating relationship between the log real price and log real dividend series is obtained from Pedroni's (1999, 2001, 2004) portfolio of panel cointegration tests. We also report a CIPS-type panel cointegration test, which incorporates an adjustment for cross-sectional dependence, and in which the test statistic is the cross-sectional mean of the individual t-statistics from cross-sectionally augmented ADF-type autoregressions involving the individual cointegrating 'levels' regressions. This test provides further support for the existence of a cointegrating relationship between prices and dividends.

Finally, the stationarity of the log price-dividend ratio series is investigated using the same portfolio of panel unit root tests that is applied to the log real price and log real dividend series individually. The details of these tests are described in subsection 3.1, and the details of the panel cointegration tests are described in subsection 3.2.

3.1. Panel unit root tests.

The LLC and IPS panel unit root tests are based on the following ADF autoregression:

$$(6) \qquad \Delta y_{i,t} = a_i + b_i t + c_i y_{i,t-1} + \sum_{j=1}^{p_i} d_{ij} \Delta y_{i,t-j} + e_{i,t}$$

The deterministic time trend may be excluded from or included in (6); if the trend is excluded, b_i=0. The homogeneous panel unit root test developed by LLC imposes the restriction c_i=c for all i. The null hypothesis H_0:c=0 is tested against the alternative H_1:c<0. The LLC test procedure can be described as follows. First, generate the residuals from regressions of $\Delta y_{i,t}$ on $\Delta y_{i,t-j}$ and the deterministic variables (in this case the fixed effects a_i and the deterministic trend) for each i; and generate the residuals from regressions of $y_{i,t-1}$ on $\Delta y_{i,t-j}$ and the deterministic variables for each i. Then normalise the two sets of residuals using the standard errors of the ADF autoregressions for each i, letting $\tilde{\varepsilon}_{i,t}$ and $\tilde{v}_{i,t}$ denote the two sets of normalised residuals. Then test δ=0 in the pooled cross section time series regression $\tilde{\varepsilon}_{i,t} = \omega \tilde{v}_{i,t-1} + \upsilon_{i,t}$, using an adjusted t-statistic on the OLS estimate of ω. The mean and standard deviation adjustments are generated by Monte Carlo simulation, and are tabulated by LLC.

The heterogeneous panel unit root test developed by IPS does not impose equality restrictions upon c_i in (6). The null hypothesis H_0:c_i=0 for all i is tested against the alternative H_1:c_i<0 for some (but not necessarily all) i. The IPS test statistic is the arithmetic mean (across i) of the N individual ADF t-statistics on c_i. The IPS test statistic follows a normal distribution. Numerical values for the mean and variance, conditional on c_i=1, are generated by Monte Carlo simulation, and are tabulated by IPS.

First-generation panel unit root tests, including LLC and IPS, are based on an assumption that the individual time series in the panel are cross-sectionally independent. Originally, cross-sectional de-meaning of the series prior to the application of these tests was suggested as a possible remedy for the problem of cross-sectional dependence between the disturbance terms of the individual series. However, it has subsequently been shown that cross-sectional de-meaning is not effective in the most general case, where the pairwise cross-sectional covariances of the disturbance terms differ between the series of the individual panel members.

Recently, a number of tests have been proposed that use orthogonalization procedures to create transformations of the original series, from which the cross-sectional dependence is asymptotically eliminated prior to the application of the standard tests to the transformed series. This literature is reviewed by Pesaran (2005), who also suggests an alternative and computationally perhaps more straightforward approach, whereby standard ADF autoregressions are augmented using the cross-sectional means of the levels and first differences of the individual series. For the first time as far as we are aware, in this paper the Pesaran cross-sectionally augmented panel unit root test procedure is applied to a panel of individual company share price and dividends series.

Pesaran's test procedure is based on the following cross-sectionally augmented ADF (denoted CADF) autoregression:

(7) $$\Delta y_{i,t} = a_i + b_i t + c_i y_{i,t-1} + \sum_{j=1}^{p_i} d_{ij} \Delta y_{i,t-j} + f_i \overline{y}_{t-1} + \sum_{j=0}^{p_i} g_i \Delta \overline{y}_{t-j} + \nu_{i,t}$$

where $\overline{y}_t = N^{-1} \sum_{i=1}^{N} y_{i,t}$ and $\Delta \overline{y}_t = \overline{y}_t - \overline{y}_{t-1}$. As before, the deterministic trend may be included in or excluded from (7). The terms in $\bar{y}_{t-1}$ and $\Delta\bar{y}_{t-j}$ control for the presence of common non-fundamental disturbances to the time-series behaviour of the series for each panel member, and the coefficients on these terms are permitted to vary between individual panel members. The CADF statistic for firm i is the t-statistic on the OLS estimate of c_i. The Pesaran cross-sectionally augmented IPS test statistic (denoted by CIPS) is the arithmetic mean (across i) of the N individual CADF t-statistics on the estimated c_i in (7). Critical values for the CIPS statistic for the values of N and T that are specific to this study are generated by the authors by Monte Carlo simulation.[4]

3.2. Panel cointegration tests.

The panel cointegration tests that are implemented below are all based on extensions of the Engle and Granger (1987) two-step approach to testing for a cointegrating relationship between two or more time series variables. In general, in the two-variable case, the first step involves estimation of the cointegrating regression:

$$(8) \qquad y_{i,t} = \hat{\alpha}_i + \hat{\beta}_i t + \hat{\delta}_i x_{i,t} + \hat{e}_{i,t}$$

The deterministic time trend may be excluded from or included in the cointegrating regression; if the trend is excluded, $\hat{\beta}_i$ =0. The second step involves examination of the residuals $\hat{e}_{i,t}$ for stationarity, using some variant of the following autoregression:

$$(9) \qquad \hat{e}_{i,t} = \gamma_i \hat{e}_{i,t-1} + \sum_{j=1}^{p} \theta_{ij} \Delta\hat{e}_{i,t-j} + \nu_{i,t}$$

4. These critical values have been checked for accuracy and consistency with the critical values for elective values of N and T that are tabulated by Pesaran (2005).

Pedroni (1999, 2001, 2004) distinguishes between the two classes of 'within dimension' and 'between dimension' panel cointegration tests. Within dimension tests are constructed by summing the numerator and denominator components of the cointegration test statistic over the N (cross-section) dimension separately, and then computing the ratio of these summations. Effectively, within dimension tests pool the autoregressive coefficient across different firms, and are equivalent to homogeneous panel cointegration tests. The null hypothesis $H_0:\gamma_i=1$ for all i is tested against an alternative hypothesis of $H_1:\gamma_i=\gamma<1$ for all i. Between dimension tests are constructed by dividing the numerator component by the denominator component for each firm individually, prior to summing these ratios over the N-dimension. Between dimension tests are based on averages of the individually estimated autoregressive coefficients for different firms, and are equivalent to heterogeneous panel cointegration tests. The null hypothesis $H_0:\gamma_i=1$ for all i is tested against an alternative hypothesis of $H_1:\gamma_i<1$ for some i.

Pedroni develops four alternative within dimension tests (T1-T4), and three between dimension tests (T5-T7). Among the within dimension tests, T1 is a non-parametric variance ratio statistic. T2 and T3 are analogous to the non-parametric Phillips and Perron (1988) rho-statistic and t-statistic, respectively. T4 is analogous to the parametric ADF t-statistic. T4 is also closely analogous to the LLC panel unit root test, applied to the estimated residuals from the cointegrating regression. Among the between dimension tests, T5 and T6 are analogous to the Phillips and Perron rho- and t-statistics, respectively. T7 is analogous to the ADF t-statistic. T7 is also closely analogous to the IPS panel unit root test, applied to the estimated residuals from the cointegrating regression. The asymptotic distribution for each of these test statistics is expressed in the form

$v^{-1/2}(\tau_{N,T} - mN^{1/2}) \sim N(0,1)$, where $\tau_{N,T}$ is the panel test statistic; and m and v are mean and variance adjustments generated by Monte Carlo simulation and tabulated by Pedroni (1999). A large positive value of T1 is required in order to reject the null of no cointegration, and large negative values of T2 to T7 are required in order to reject the null of no cointegration.

The panel cointegration analogue of the Pesaran (2005) cross-sectionally augmented panel unit root test is based on the following CADF-type autoregression for $\hat{e}_{i,t}$:

(10) $$\Delta\hat{e}_{i,t} = \rho_i \hat{e}_{i,t-1} + \sum_{j=1}^{p} \theta_{ij} \Delta\hat{e}_{i,t-j} + \psi_i \bar{e}_{t-1} + \sum_{j=0}^{p} \vartheta_{ij} \Delta\bar{e}_{t-j} + \nu_{i,t}$$

where $\bar{e}_t = N^{-1} \sum_{i=1}^{N} \hat{e}_{i,t}$, $\Delta\bar{e}_t = \bar{e}_t - \bar{e}_{t-1}$, and $\hat{e}_{i,t}$ are obtained from (8) with the time trend either excluded or included. The cointegration analogue of the CADF statistic for firm i is the t-statistic on the estimated ρ_i in (10) and the cointegration analogue of the CIPS statistic is the arithmetic mean (across i) of the N individual CADF-type t-statistics on ρ_i. Critical values for this panel cointegration statistic for the values of N and T that are specific to this study are generated by the authors using Monte Carlo simulation.

4. Data and empirical results.

Annual time series data on price and dividends were obtained for 104 UK non-financial companies for which complete and continuous share price and dividend series were available over a 34-year observation period from 1970 to 2003 (inclusive). All data were obtained from the *DataStream* non-financials list.

Table 1 provides a summary of the trends in the sample data at an aggregated level. For the construction of Table 1, each firm's share price series was normalized using factors that produce an identical initial value (1970=100) for every firm. Each firm's dividend series was normalized using the same factors, so as to preserve the correct price-dividend ratio in the normalized series. Table 1 reports the annual cross-sectional means (across the 104 sample firms) of the normalized share price and dividend series; the means of same two series in real terms (converted using a price inflation index); and finally, the price-dividend ratio. The growth in the average share price series exceeds the growth in market indexes such as the FTSE-100 over the same period.[5] This presumably reflects the atypical nature of a sample comprising firms that survived over a period of more than 30 years, whose average performance exceeded that of the market index by some distance.

Tables 2 and 3 report the panel unit root test results for the log real price and log real dividend series, respectively. In each of these tables, Section 1 reports the results of the LLC test based on two specifications: first, with intercept only; and second, with intercept and deterministic trend. In both cases, test results are reported for a range of lag augmentations, from P=0 to P=3 (inclusive). Section 2 of Tables 2 and 3 report the numbers of firms for which the null hypothesis of non-stationarity is rejected in the individual ADF autoregressions, at significance levels of 1%, 5% and 10%. The individual ADF autoregressions are based on the same two specifications (intercept only, and intercept and trend) and the same range of lag augmentations (P=0 ... 3). For each firm, a preferred lag augmentation, denoted P^*, is selected using the Akaike Information Criterion (AIC). The numbers of firms for which the null is rejected in the set of ADF autoregressions based on each firm's selected value of P^* are also reported. Section 2 also

5. Between December 1970 and December 2003, the value of the FTSE-100 index increased from 326.9 to 4476.9.

reports the IPS t-bar$_{N,T}$ statistics based on each set of ADF autoregressions. Section 3 of Tables 2 and 3 reports the results for the CADF and CIPS tests, displayed in the same format as the ADF and IPS tests in Section 2. Finally, in order to check for the significance of the cross-sectional augmentation in the auxiliary regressions used for the CADF tests, Section 3 of Tables 2 and 3 reports the numbers of firms (out of 104) for which the AIC from the auxiliary regression for the CADF test is smaller than the AIC from the corresponding auxiliary regression (with the same lag augmentation) for the ADF test.

For the log real price series reported in Table 2, in the model with intercept only the LLC test fails to reject the null hypothesis of non-stationarity at the 5% significance level for P=1,2,3; but the same test does reject this null hypothesis for P=0. In the model with intercept and trend, the null hypothesis is rejected for all values of P. The individual ADF tests produce a similar pattern of results: in the large majority of cases these tests fail to reject the null when the trend is excluded; but there are significantly more rejections when the trend is included. Accordingly, the IPS tests fail to reject the null in the former case, but invariably reject the null (at the 1% level) in the latter case.

Overall, there is some ambiguity in the results of the LLC and ADF/IPS tests on the log real price series: the diagnosis of stationarity or non-stationarity depends upon whether or not the time trend is included. However, this inconsistency is much reduced by the CADF/CIPS tests, which produce very similar results regardless of the exclusion or inclusion of the trend. Effectively, the augmentation to allow for cross-sectional dependence eliminates the effects of any trend component that is common across firms, thereby reducing the sensitivity of the test results to the exclusion or inclusion of the deterministic trend. According to the AIC, the cross-sectional augmentation contributes significantly to the explanatory power of the auxiliary regressions for virtually all of the 104 sample firms.

For the log real price series, the individual CADF tests produce numbers of rejections of the null of non-stationarity that are only slightly higher than would be expected (due to Type I error) if the null were true in all cases. Similarly, the CIPS test results are borderline, with the test significant at the 5% level (but not at the 1% level) for the lag augmentation P=1, and at the 10% level (but not at the 5% level) for several other lag augmentations. However, the CIPS test based on the set of preferred lag augmentations P^* is significant at the 1% level.

For the log real dividend series reported in Table 3, there is similar ambiguity in the LLC and ADF/IPS test results, where the diagnosis of non-stationarity or stationarity is again sensitive to the exclusion or inclusion of the deterministic trend. For the model with intercept only, the panel tests fail to reject the null hypothesis of non-stationarity for all firms; but for the model with intercept and trend, the panel tests reject the same null hypothesis. According to the AIC, the cross-sectional augmentation contributes significantly to the explanatory power of the auxiliary regressions for a large majority of the 104 firms, but in this case there is a (non-negligible) minority of firms for which the cross-sectional augmentation is insignificant. As before, the sensitivity of the unit root test results to the exclusion or inclusion of the trend is smaller with the CADF/CIPS tests than it is with the ADF/IPS tests. The CIPS tests predominantly fail to reject the null hypothesis of non-stationarity for all firms. The sole exception is the CIPS test for the model with trend, based on the set of preferred lag augmentations P^*, which is significant at the 1% level. Overall, however, support for the null hypothesis of non-stationarity appears to be stronger for the log real dividend series than it is in the case of the log real price series.

Given that the panel unit root tests yield some evidence of trend-stationarity, at least for some sample firms, in Table 4 we proceed to the tests for cointegration between the log real dividend and log real price series with this caveat in mind. For cointegration testing, Nasseh and Strauss (2004) discuss the important issue of the choice of normalization: in other words, the choice between the log real price series and the log real dividend series as the regressand (dependent variable) in (8), the cointegrating regression. As shown by Ng and Perron (1997), the choice of normalization is important when one of the two I(1) variables contains a large temporary or stationary component. Nasseh and Strauss find that the innovations in their dividend series for a majority of their sample firms contain a significant moving average component, while the moving average component in the innovations in their price series is insignificant for virtually all firms. This finding is consistent with standard conditions for market efficiency, which require that price adjustments are unpredictable, and only unanticipated dividends at time t should cause price adjustments at time t. If dividend innovations contain a large temporary (or moving average) component, estimation of (8) with $y_{i,t} = \ln(p_{i,t})$ and $x_{i,t} = \ln(d_{i,t})$ will result in an underestimation of δ_i, and a bias towards non-rejection of the null hypothesis of no cointegration. Accordingly, $y_{i,t} = \ln(d_{i,t})$ and $x_{i,t} = \ln(p_{i,t})$ is the preferred normalization in (8).

Section 1 of Table 4 reports the results of applying the portfolio of seven Pedroni cointegration tests to the residuals from the cointegrating levels regression for all 104 sample firms, for the cases where the cointegrating regression excludes and includes a deterministic time trend. The null of no cointegration is rejected when the T1 test statistic is large and positive, and when the T2-T7 test statistics are large and negative. In six of the seven cases, and regardless of the exclusion or inclusion of the deterministic trend, the Pedroni tests suggest the presence of a cointegrating relationship between the log real price and log real dividend series. T1 is the only test in the portfolio that fails to reject the null hypothesis of no cointegration.

Sections 2 and 3 of Table 4 report the results of CADF/CIPS-type cointegration tests for the stationarity of the residuals of the cointegrating regression. In Section 2, these tests are based on the data for all 104 sample firms. In Section 3, the same tests are repeated, based on the data only for those sample firms for which the CADF tests using the preferred lag augmentation (reported in Tables 2 and 3) failed to reject the null hypothesis of non-stationarity for both the log real price and the log real dividend series. Accordingly, in Section 3, we exclude those firms for which the CADF tests suggest that either or both of the series are stationary. This results in the exclusion of 26 firms for the specification with no trend (29 firms for the specification with trend). The cointegration tests are based on the remaining 78 firms (75 firms), for which there is evidence that both of the series are non-stationary.[6]

In Section 2 of Table 4, in the set of CADF-type tests based on the preferred lag augmentation, the null hypothesis of no cointegration is rejected at the 5% level for 50 firms (48.1%) when the trend is excluded from the cointegrating regression, and for 55 firms (52.9%) when the trend is included. These proportions are sufficiently high for the CIPS-type tests to reject the null hypothesis of no cointegration for every firm at the 1% level. Trimming the sample as described above turns out to make little difference to this pattern of results. In Section 3, in the CADF-type tests based on the preferred lag augmentation, the null is rejected at the 5% level for 41 firms (52.6%) when the trend is excluded, and for 40 firms (53.3%) when the trend is included. Again, the CIPS-type tests reject the null at the 1% level.[7]

---

6. Cointegration tests based on the residuals of levels regressions involving series that are themselves stationary are likely to produce a misleading diagnosis supporting cointegration.

7. As a check on the sensitivity of the cointegration test results to the normalization (with log dividend as the regressand and log price as the regressor in the cointegrating regression), the tests reported in Table 4 were repeated with equation (8) estimated using log price as regressand and log dividend as regressor. In this case, the CIPS tests consistently rejected the null hypothesis of no cointegration at the 1% level, but the absolute values of the test statistics were smaller than their counterparts reported in Table 4. The numbers of firms for which the ADF/CADF test statistics were significant were smaller than the numbers reported in Table 4. This pattern is consistent with the conclusions and recommendations of Nasseh and Strauss (2004) concerning the implications for the

In view of the relatively low power of the individual CADF-type cointegration tests in a sample with a relatively short time-series dimension, it is likely that the proportion of firms for which the log price and log dividend series are actually cointegrated is higher than the proportions of rejections (around 50%) based on these tests as reported in Table 4. However, while the decisive rejection of the null in the CIPS-type tests implies that the hypothesis of no cointegration for all firms can be confidently rejected, this finding can not be construed as evidence that a cointegrating relationship exists for every firm. The tests employed in the present study do not allow us to investigate this latter (and much stronger) hypothesis. It seems unlikely that such a hypothesis would be supported: visual inspection of the data suggests for a number of firms the price-dividend relationship behaves idiosyncratically in some manner that would categorically rule out a diagnosis of cointegration. Nevertheless, we interpret the results reported in Table 4 as supportive of cointegration for a large proportion (at least 50%, and perhaps more) of the sample firms.

Further evidence pointing in a similar direction is obtained from Table 5, which reports the results of the application of the panel unit root test procedures directly to the log price-dividend ratio series. In Section 1, the LLC test rejects the null hypothesis of non-stationarity for all lag augmentations, regardless of the exclusion or inclusion of the deterministic trend. In Section 2, the null hypothesis of non-stationarity is rejected by the individual ADF tests in relatively large numbers. In the ADF tests based on the preferred lag augmentation for each firm, for example, the null is rejected at the 5% level in 61 cases (58.7%) for the model without the trend, and 68 cases (65.4%) for the model with the trend. Accordingly, all of the IPS panel unit root tests are unanimous in rejecting the null at the 1% level. In Section 3, according to the AIC the cross-sectional augmentation contributes significantly to the explanatory power of the auxiliary regressions for virtually all of the

---

cointegration tests of the choice of normalization.

sample firms. The null hypothesis of non-stationarity is rejected by the individual CADF tests in smaller, but still non-negligible numbers. The corresponding numbers of rejections are 31 (29.8%) for the model without the trend, and 24 (23.1%) for the model with the trend. As before, the CIPS tests are able to reject the null in the majority of cases, including both of the tests based on the preferred lag augmentation.

Our comments regarding the interpretation of these results are similar to those above. The panel unit root results constitute evidence in support of the present value model, in the form that does not rely on the fixed discount rate assumption, for at least 20% (and in view of the low power of the individual CADF tests, perhaps considerably more) of the sample firms.

5. Summary and conclusion.

A long-standing principle of finance theory, known as the present value model, is that current security prices depend upon the present value of discounted future dividends, where the discount rate is equivalent to the required rate of return. Recently, this relationship has been subject to renewed empirical scrutiny, partly as a result of the large rise in stock prices at the end of the last century, and the subsequent precipitous fall. The consensus view of this period is that the market became overvalued for a time, but fundamentals subsequently reasserted themselves: the stock market boom of the late-1990s is now widely regarded as a form of bubble phenomenon.

The majority of extant empirical studies of the relationship between stock prices and dividends use US stock market data, and focus upon the time-series properties of aggregate price and dividend indexes. In this paper, we analyse the stock price and dividend data of a sample of 104 UK non-financial companies that reported data continuously throughout the period 1970 to 2003. The paper exploits recent developments in dynamic panel econometrics in order to focus on firm-level data rather than aggregate stock market data. We argue that the use of firm-level data permits the observation of patterns that may otherwise be obscured through averaging in the aggregation process. Moreover, investors are primarily concerned with information on the cash flows and future prospects of individual firms, and not of aggregate stock market indexes. To our knowledge, there has only been one previous firm-level empirical analysis of the present value model, which was based on US data. Therefore this study provides the first such analysis for the UK.

Alongside the well-known first-generation LLC and IPS panel unit root tests, we employ a test recently developed by Pesaran (2005), which allows for cross-sectional dependence between the disturbance terms of the ADF autoregressions on which the earlier tests were based. In the context of the present value model, the cross-sectional augmentation allows for factors such as bubbles which may result in temporary deviations from the long-run relationship between stock prices and dividends, and which tend to operate in a similar (but not necessarily identical) fashion on the data of some or all of the sample firms simultaneously. For cointegration testing, this paper employs the portfolio of panel cointegration tests developed by Pedroni (1999, 2001, 2004), alongside a cointegration analogue of the Pesaran (2005) cross-sectionally augmented panel unit root test. The latter also includes an adjustment for cross-sectional dependence.

The panel unit root and panel cointegration tests are ambivalent concerning the stationarity or non-stationarity properties of the firm-level log real price and log real dividend series. In both cases, there is evidence of non-stationarity for the majority of firms, but in neither case is an alternative hypothesis of trend-stationarity ruled out for every firm in the sample. The evidence favouring non-stationarity (rather than trend-stationarity) is generally slightly stronger for the dividend series than for the price series. However, panel cointegration tests, either based upon all sample firms, or restricted to those firms for which diagnoses of non-stationarity are obtained in respect of both the price and dividend series individually, provide consistent evidence (across a variety of tests and model specifications) supporting a cointegrating relationship between price and dividend for a large proportion of the sample firms, in accordance with the present value model. Similarly, panel unit root tests applied directly to the log price-dividend ratio are consistent in rejecting a null hypothesis of non-stationarity in respect of all firms: a finding which provides support for the formulation the present value model that allows for a time-varying discount rate. Overall our empirical results, based on firm-level data, appear to be somewhat more supportive of the present value model than those of several previous studies that were based on aggregated stock price and dividend index data.

Table 1 Sample mean share price, dividend, real price, real dividend and price-dividend ratio, 1970-2003

Year	Mean price	Mean dividend	Mean real price	Mean real dividend	Mean price-dividend ratio
1970	100.0	3.4	100.0	3.4	37.1
1971	146.4	3.8	133.0	3.4	46.1
1972	200.4	4.7	172.1	4.0	47.2
1973	179.8	4.9	141.0	3.8	40.0
1974	72.9	5.1	48.8	3.4	15.5
1975	132.4	5.9	70.3	3.1	26.1
1976	108.4	6.7	51.0	3.2	17.7
1977	239.1	7.7	95.6	3.1	34.8
1978	278.6	9.1	103.4	3.4	46.8
1979	321.2	11.9	103.0	3.8	27.4
1980	408.7	14.1	112.1	3.9	27.1
1981	451.8	15.5	111.8	3.8	25.8
1982	661.4	17.1	150.5	3.9	32.7
1983	765.7	19.4	167.3	4.2	33.8
1984	897.2	22.8	187.6	4.8	33.9
1985	1299.0	26.9	254.1	5.3	40.5
1986	1653.3	32.2	315.9	6.1	42.0
1987	1928.9	39.6	353.0	7.3	41.9
1988	2044.5	49.4	357.0	8.6	37.9
1989	1990.6	62.0	321.1	10.0	31.3
1990	1961.7	69.2	288.3	10.2	26.4
1991	2574.9	72.1	358.6	10.0	32.9
1992	2605.6	75.9	349.7	10.2	31.9
1993	3282.9	79.5	434.7	10.5	47.0
1994	3212.5	86.6	415.6	11.2	44.0
1995	3698.6	96.0	462.2	12.0	37.9
1996	4399.2	104.9	537.8	12.8	38.9
1997	4641.9	114.2	549.1	13.5	37.3
1998	4566.6	123.3	522.0	14.1	32.5
1999	4972.2	138.2	561.1	15.6	31.4
2000	5114.8	149.5	558.9	16.3	30.8
2001	4885.0	161.1	525.2	17.3	27.4
2002	4493.5	158.9	476.0	16.8	26.4
2003	5126.2	170.1	526.8	17.5	29.5

Table 2 Panel unit root test results: real log price series

Dependent variable: $y_{i,t} = \ln(p_{i,t}/rpi_t)$

	Model with intercept only					**Model with intercept and trend**				
	P=0	P=1	P=2	P=3	P=P*	P=0	P=1	P=2	P=3	P=P*
1. LLC test										
LLC t-stat.	-2.10**	-1.40	-0.34	-0.04	n/a	-6.05***	-6.38***	-11.9***	-6.64***	n/a
2. ADF and IPS tests										
Number of firms with ADF p-value as follows:										
$p \le 0.01$	5	3	3	3	4	14	14	22	13	23
$0.01 < p \le 0.05$	2	4	7	1	8	24	15	19	16	33
$0.05 < p \le 0.1$	2	2	6	5	5	14	12	13	12	18
$p > 0.1$	95	95	88	95	87	52	63	50	63	30
IPS t-bar$_{N,T}$	-1.64	-1.51	-1.66	-1.41	-1.68	-3.22***	-3.13***	-3.39***	-3.07***	-3.58***
3. CADF and CIPS tests										
Number of firms with CADF p-value as follows:										
$p \le 0.01$	0	3	1	2	4	3	2	1	2	4
$0.01 < p \le 0.05$	6	2	5	6	10	7	9	9	8	13
$0.05 < p \le 0.1$	8	6	11	7	11	5	8	6	6	14
$p > 0.1$	90	93	87	89	79	89	85	88	88	73
CIPS t-bar$_{N,T}$	-2.09**	-2.14**	-2.05*	-2.03*	-2.32***	-2.56*	-2.65**	-2.52*	-2.49	-2.91***
Firms with $AIC_2 < AIC_1$	103	103	103	102	n/a	104	103	101	101	n/a

Notes

Critical values for the ADF test for T=30, calculated by the authors using stochastic simulation, are as follows.
Intercept only: -3.72 (1% significance level), -2.99 (5% level), -2.62 (10% level).
Intercept and trend: -4.33 (1% level), -3.33 (5% level), -2.95 (10% level).

Critical values for the IPS t-bar$_{N,T}$ statistic for N=104, T=30 are as follows.
Intercept only: -1.73 (1% level), -1.67 (5% level), -1.64 (10% level).
Intercept and trend: -2.37 (1% level), -2.31 (5% level), -2.28 (10% level).

Critical values for the CADF test for N=104, T=30 are as follows.
Intercept only: -4.11 (1% level), -3.33 (5% level), -2.95 (10% level).
Intercept and trend: -4.68 (1% level), -3.87 (5% level), -3.48 (10% level).

Critical values for the CIPS t-bar$_{N,T}$ statistic for N=104, T=30 are as follows.
Intercept only: -2.17 (1% level), -2.08 (5% level), -2.02 (10% level).
Intercept and trend: -2.66 (1% level), -2.57 (5% level), -2.51 (10% level).

*** = test statistic significant at 1% level; ** = 5% level; * = 10% level.

Firms with $AIC_2 < AIC_1$ is the number of firms (out of 104) for which the Akaike Information Criterion (AIC) for the auxiliary regression for the CADF test is smaller than the AIC for the corresponding auxiliary regression (with the same lag length) for the ADF test.

Table 3 Panel unit root test results: real log dividend series

Dependent variable: $y_{i,t} = \ln(d_{i,t}/rpi_t)$

	Model with intercept only					Model with intercept and trend				
	P=0	P=1	P=2	P=3	P=P*	P=0	P=1	P=2	P=3	P=P*
1. LLC test										
LLC t-stat.	6.09	2.70	2.78	1.81	n/a	-2.50**	-3.35***	-6.14***	-4.31***	n/a
2. ADF and IPS tests										
Number of firms with ADF p-value as follows:										
$p \leq 0.01$	2	1	0	0	2	10	6	6	6	9
$0.01 < p \leq 0.05$	2	2	2	0	1	6	9	8	10	16
$0.05 < p \leq 0.1$	3	4	1	4	4	9	9	10	5	17
$p > 0.1$	97	97	101	100	97	79	80	80	83	62
IPS t-bar$_{N,T}$	-0.77	-1.04	-1.03	-1.01	-1.09	-2.39***	-2.52***	-2.52***	-2.46***	-2.68***
3. CADF and CIPS tests										
Number of firms with CADF p-value as follows:										
$p \leq 0.01$	5	0	1	4	8	6	3	2	5	11
$0.01 < p \leq 0.05$	7	9	9	7	8	8	6	7	8	8
$0.05 < p \leq 0.1$	8	6	3	4	7	12	7	7	5	11
$p > 0.1$	86	89	90	89	81	78	88	88	86	74
CIPS t-bar$_{N,T}$	-1.84	-1.74	-1.65	-1.69	-1.96	-2.50	-2.37	-2.31	-2.36	-2.77***
Firms with $AIC_2 < AIC_1$	82	77	74	75	n/a	78	74	68	74	n/a

Note

For critical values, see notes to Table 2.

*** = test statistic significant at 1% level; ** = 5% level; * = 10% level.

Firms with $AIC_2 < AIC_1$ is the number of firms (out of 104) for which the Akaike Information Criterion (AIC) for the auxiliary regression for the CADF test is smaller than the AIC for the corresponding auxiliary regression (with the same lag length) for the ADF test.

Table 4 Cointegration test results

Dependent variable: $\hat{e}_{i,t}$ from $\ln(d_{i,t}) = \hat{\alpha}_i + \hat{\beta}_i \ln(p_{i,t}) + \hat{\delta}_i t + \hat{e}_{i,t}$ ($\hat{\delta}_i$=0 for models with intercept only; $\hat{\delta}_i \neq 0$ for models with intercept and trend).

1. Pedroni tests

	T1	**T2**	**T3**	**T4**	**T5**	**T6**	**T7**
Intercept only	0.83	-10.0^{***}	-9.07^{***}	-8.00^{***}	-8.80^{***}	-10.7^{***}	-10.2^{***}
Intercept and trend	0.64	-2.68^{***}	-5.96^{***}	-4.14^{***}	-2.61^{***}	-7.55^{***}	-6.95^{***}

2. CADF/CIPS-type cointegration tests: all sample firms

	Intercept only					**Intercept and trend**				
	P=0	P=1	P=2	P=3	$P=P^*$	P=0	P=1	P=2	P=3	$P=P^*$
Number of firms with ADF/CADF p-value as follows:										
p≤0.01	14	10	9	4	20	20	12	11	7	27
0.01<p≤0.05	31	20	13	10	30	29	23	22	13	28
0.05<p≤0.1	13	17	17	20	13	14	18	12	19	13
p>0.1	46	57	65	70	41	41	51	59	65	36
IPS/CIPS $t\text{-bar}_{N,T}$	-2.87^{***}	-2.63^{***}	-2.41^{***}	-2.28^{***}	-2.92^{***}	-3.03^{***}	-2.82^{***}	-2.61^{***}	-2.51^{***}	-3.12^{***}

3. CADF/CIPS-type cointegration tests: firms with non-stationary log real price and log real dividend series according to CADF unit root tests

	Intercept only					**Intercept and trend**				
	P=0	P=1	P=2	P=3	$P=P^*$	P=0	P=1	P=2	P=3	$P=P^*$
Number of firms with ADF/CADF p-value as follows:										
p≤0.01	13	9	8	3	19	14	7	9	4	20
0.01<p≤0.05	26	16	8	7	22	20	15	16	10	20
0.05<p≤0.1	10	12	14	16	10	12	14	8	13	11
p>0.1	29	41	48	52	27	29	39	42	48	24
IPS/CIPS $t\text{-bar}_{N,T}$	-2.99^{***}	-2.66^{***}	-2.42^{***}	-2.28^{***}	-2.97^{***}	-3.03^{***}	-2.82^{***}	-2.72^{***}	-2.53^{***}	-3.20^{***}

Notes

Critical values for the CADF-type test for N=104, T=30 are as follows.
Intercept only: -3.75 (1% significance level), -3.02 (5%), -2.67 (10%).
Intercept and trend: -3.78 (1%), -3.06 (5%), -2.71 (10%).

Critical values for the CIPS-type $t\text{-bar}_{N,T}$ statistic for N=104, T=30 are as follows:
Intercept only: -1.79 (1%), -1.72 (5%), -1.68 (10%).
Intercept and trend: -1.84 (1%), -1.77 (5%), -1.73 (10%).

Critical values for the CADF-type test for T=30 are as follows:
Intercept only (N=78): -3.73 (1% level), -3.02 (5%), -2.67 (10%)
Intercept and trend (N=75): -3.78 (1%), -3.06 (5%), -2.72 (10%).

Critical values for the CIPS-type $t\text{-bar}_{N,T}$ statistic for T=30 are as follows:
Intercept only (N=78): -1.84 (1%), -1.75 (5%), -1.71 (10%).
Intercept and trend (N=75): -1.88 (1%), -1.80 (5%), -1.75 (10%).

*** = test statistic significant at 1% level; ** = 5% level; * = 10% level.

Table 5 Panel unit root test results: log price-dividend ratio series

Dependent variable: $y_{i,t} = \ln(p_{i,t}/d_{i,t})$

	Model with intercept only					**Model with intercept and trend**				
	P=0	P=1	P=2	P=3	P=P*	P=0	P=1	P=2	P=3	P=P*
1. LLC test										
LLC t-stat.	-17.6^{***}	-12.7^{***}	-14.0^{***}	-7.53^{***}	n/a	-16.6^{***}	-11.3^{***}	-14.0^{***}	-7.6^{***}	n/a
2. ADF and IPS tests										
Number of firms with CADF p-value as follows:										
p≤0.01	25	14	17	12	32	23	7	15	12	26
0.01<p≤0.05	30	21	27	16	29	31	31	37	22	42
0.05<p≤0.1	12	15	15	17	9	16	18	12	18	10
p>0.1	37	54	45	59	34	34	48	40	52	26
IPS t-bar$_{N,T}$	-3.11^{***}	-2.69^{***}	-2.91^{***}	-2.54^{***}	-3.26^{***}	-3.49^{***}	-3.04^{***}	-3.34^{***}	-3.00^{***}	-3.71^{***}
3. CADF and CIPS tests										
Number of firms with CADF p-value as follows:										
p≤0.01	10	3	1	0	10	3	1	2	1	5
0.01<p≤0.05	14	11	7	6	21	14	9	5	5	19
0.05<p≤0.1	13	15	12	10	20	10	9	2	5	10
p>0.1	67	75	84	88	53	77	85	95	93	70
CIPS t-bar$_{N,T}$	-2.72^{***}	-2.43^{***}	-2.24^{***}	-2.10^{**}	-2.80^{***}	-2.97^{***}	-2.68^{***}	-2.45	-2.31	-3.06^{***}
Firms with $AIC_2<AIC_1$	103	103	99	99	n/a	101	103	99	98	n/a

Note

For critical values, see notes to Table 2.

*** = test statistic significant at 1% level; ** = 5% level; * = 10% level.

Firms with $AIC_2<AIC_1$ is the number of firms (out of 104) for which the Akaike Information Criterion (AIC) for the auxiliary regression for the CADF test is smaller than the AIC for the corresponding auxiliary regression (with the same lag length) for the ADF test.

References

Balke, N.S. and Wohar, M.E. (2001). 'Explaining stock price movements: Is there a case for fundamentals?' *Federal Reserve Bank of Dallas Economic and Financial Review*, Third Quarter, 22-34.

Balke, N.S. and Wohar, M.E. (2002). 'Low frequency movements in stock prices: A state-space decomposition', *Review of Economics and Statistics,* 84, 649-667.

Bohl, M.T. and Siklos, P.L. (2004). 'The present value model of US stock prices redux: A new testing strategy and some evidence', *Quarterly Review of Economics and Finance*, 44, 208-223.

Brooks, C. and Katsaris, A. (2003), 'Rational speculative bubbles: An empirical investigation of the London stock exchange', *Bulletin of Economic Research*, 55, 319-346.

Browne, L.E. (1999). 'US economic performance: Good fortune, bubble, and new era', *New England Economic Review*, May/June, 4-20.

Campbell, J.Y., Lo, A.W. and MacKinlay, C. (1997). *The Econometrics of Financial Markets*. Princeton University Press: Princeton.

Campbell, J.Y. and Shiller, R.J. (1987). 'Cointegration and tests of present value models', *Journal of Political Economy*, 95, 1062-1088.

Campbell, J.Y. and Shiller, R.J. (1988a). 'The dividend-price ratio and expectations of future dividends and discount factors', *Review of Financial Studies*, 1, 195-227.

Campbell, J.Y. and Shiller, R.J. (1988b). 'Stock prices, earnings, and expected dividends', *Journal of Finance*, 43, 661-676.

Campbell, J.Y. and Shiller, R.J. (2001). 'Valuation ratios and the long-run stock market outlook: an update', *NBER Working Paper 8221*.

Caporale, G.M. and Gil-Alana, L.A. (2004). 'Fractional cointegration and tests of present value models', *Review of Financial Economics*, 13, 245-258.

Cohen, R., Polk, C. and Vuolteenaho, T. (2001). 'The value spread', *NBER Working Paper 8242*.

Crowder, B. and Wohar, M.E. (1998). 'Stock price effects of permanent and transitory shocks', *Economic Inquiry*, 36, 540-552.

Diba, B.T. and Grossman, H.I. (1988). 'Explosive rational bubbles in stock prices?' *American Economic Review*, 78, 520-530.

Engle, R.E. and Granger, C.W.J. (1987). 'Cointegration and error-correction: representation, estimation and testing', *Econometrica*, 55, 251-276.

Evans, G.W. (1991). 'Pitfalls in testing for explosive bubbles in asset prices', *American Economic Review*, 81, 922-930.

Fama, E.F. and French K.R. (2002). 'The equity premium', *Journal of Finance*, 57, 637-659.

Froot, K. and Obstfeld, M. (1991). 'Intrinsic bubbles: The case of stock prices', *American Economic Review*, 81, 1189-1214.

Greenwood, J. and Jovanovic, B. (1999). 'The information technology revolution and the stock market', *American Economic Review*, 89, 116-122.

Hall, S.G., Psaradakis, Z., and Sola, M. (1999), 'Detecting periodically collapsing bubbles: A Markov-switching unit root test', *Journal of Applied Econometrics*, 14, 143-154.

Heaton, J. and Lucas, D. (1999). 'Stock prices and fundamentals', *NBER Macroeconomic Annual*, 213-262.

Im, K.S., Pesaran, M.H., and Shin, Y. (2003). 'Testing for unit roots in heterogeneous panels' *Journal of Econometrics*, 115, 53-74.

Jung, J. and Shiller, R.J. (2005). 'Samuelson's dictum and the stock market', *Economic Inquiry*, 43, 221-228.

Kanas, A. (2005). 'Non-linearity in the stock price-dividend relation', *Journal of International Money and Finance*, 24, 583-606.

Kapetanios, G., Shin, Y. and Snell, A. (2003), 'Testing for cointegration in nonlinear STAR error-correction models', Queen Mary, University of London, Department of Economics, Working Paper No. 497 (forthcoming in *Econometric Theory*).

Lamont, O. (1998). 'Earnings and expected returns', *Journal of Finance*, 53, 1563-1587.

Lee, B.S. (1995). 'The response of stock prices to permanent and temporary shocks to dividends', *Journal of Financial and Quantitative Analysis*, 30, 1-22.

Levin, A., Lin, C. F., and Chu, C. (2002). 'Unit root tests in panel data: asymptotic and finite-sample properties', *Journal of Econometrics*, 108, 1-24.

Nasseh, A. and Strauss, J. (2004). 'Stock prices and the dividend discount model: Did their relation breakdown in the 1990's?' *Quarterly Review of Economics and Finance*, 44, 191-207.

Ng, S. and Perron, P. (1997). 'Estimation and inference in nearly unbalanced nearly cointegrated systems', *Journal of Econometrics*, 79, 53-81.

Pedroni, P. (1999). 'Critical values for cointegration tests in heterogeneous panels with multiple regressors', *Oxford Bulletin of Economics and Statistics*, 61, 653-670.

Pedroni, P. (2001). 'Purchasing power parity tests in cointegrated panels', *Review of Economics and Statistics*, 83, 727-731.

Pedroni, P. (2004). 'Panel cointegration: asymptotic and finite sample properties of pooled time series tests with an application to the PPP hypothesis', *Econometric Theory*, 20, 597-625.

Pesaran, M. H. (2005). 'A simple panel unit root test in the presence of cross section dependence', Cambridge Working Papers in Economics 0346, revised version, University of Cambridge.

Phillips, P. and Perron, P. (1988). 'Testing for a unit root in time series regression', *Biometrika*, 75, 335-346.

Psaradakis, Z., Sola, M. and Spagnolo, F. (2004). 'On Markov error-correction models, with an application to stock prices and dividends', *Journal of Applied Econometrics*, 19, 69-88.

Siegel, J.J. (1999). 'The shrinking equity premium: Historical facts and future forecasts', *Journal of Portfolio Management*, Fall, 10-17.

Sung, H. and Urrutia, J. (1995). 'Long-term and short-term causal relations between dividends and stock prices: A test of Lintner's dividend model and the present value of stock prices', *Journal of Financial Research*, 18, 171-188.

Timmerman, A. (1995). 'Cointegration tests of present value models with a time-varying discount factor', *Journal of Applied Econometrics*, 10, 17-31.

van Norden, S (1996), 'Regime switching as a test for exchange rate bubbles', *Journal of Applied Econometrics*, 11, 219-251.

van Norden, S and Vigfusson, R (1998), 'Avoiding the pitfalls: Can regime-switching tests reliably detect bubbles?' *Studies in Nonlinear Dynamics and Econometrics*, 3, 1-22.

Vuolteenaho, T. (2002). 'What drives firm-level stock returns?' *Journal of Finance*, 62, 233-264