Excessive Stock Price Dispersion:

A Regression Test of

Cross-Sectional Volatility

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We apply a regression test of the volatility of asset prices to a cross-section data set of stock prices each year between 1932-71. We compute the ex post rational stock price for each firm, and show that the rejection of the present value model in the time series domain carries over to data on individual share prices within a particular year, even allowing for cross-sectional dependence between stocks. We further examine the relationship between mis-pricing and past dividends. Assuming that dividend yields proxy for growth expectations we find that investors are unduly optimistic about high growth stocks and too pessimistic about low growth stocks.
I Introduction

Following the seminal papers of Shiller (1981) and LeRoy and Porter (1981) there now exists a substantial body of work which has examined whether stock prices are excessively volatile, and Gilles and LeRoy (1991) in a thorough survey of this literature, concludes that there is strong statistical evidence in favour of this documented excess volatility. This work has focused exclusively on the time series behaviour of an aggregate stock price index. In this paper we examine another dimension of volatility and study the cross-section dispersion of individual company share prices. We investigate whether at a particular date, the stock prices of a large sample of firms are excessively dispersed compared with ex post rational or “fundamental” stock prices calculated from the subsequent dividend realisations paid out by these same companies.

Time series tests are inevitably contingent upon assumptions made about the time series properties of the data. Whilst appropriate tests have been developed under different plausible assumptions [Campbell and Shiller (1987), West (1988), and LeRoy and Parke (1992)], it is nevertheless advantageous to side-step these problems and this is possible in a cross-section test. The economic merit of this cross-section approach is that evidence about the dispersion of company share prices should contribute to a better understanding of the structure of the documented excess volatility. In particular it allows us to identify whether the reported excess volatility is purely a macro-phenomena, scaling all share prices by a similar amount, or whether there is also a micro-component affecting the pricing of individual shares, superimposed on the aggregate effect. The original papers identifying excess volatility of a stock market index carries no implications for the dispersion of individual stock prices.

In section II we discuss the methodology of a cross-section volatility test, which allows for the cross-sectional dependence between firms, using regression tests introduced for time series data by

\[1/\text{In fact Kleidon (1986), argues that given a non-stationary dividend series, a cross section volatility test is the only valid test criteria. Board, Bulkley and Tonks (1993) apply a cross-section variance bounds test to a sample of US firms 1926-1970 and find that before 1956 the bound is satisfied, but post 1957 is violated in every year. Of course these arguments rest on the variance of each firm’s price at any particular time as being finite in our sample. As a mere technicality this must be true given the finite history of these companies, and further an overwhelming proportion of firms in each cross-section in our sample have been in existence for a relatively short time. The large size of the cross-section then ensures that there is sufficient cross-section variation relative to time series variance to achieve reliable estimates, and justifies using the size of the cross-section as the asymptote in our statistical inference.}\]
Scott (1985), and Campbell and Shiller (1988). In fact Campbell and Shiller (1988) point out that these regression tests are equivalent to Fama and French (1988) regressions on the predictability of long run returns. Having described the data and discussed the calculation of the ex post rational price for each firm in Section III, in Section IV, we report the evidence from these cross-section regressions that in most years stock prices are more dispersed than the efficient markets paradigm would suggest. In section V we further investigate the structure and possible determinants of the excess dispersion that we have identified.\(^2\) Our results indicate that stocks with low dividend yields are overpriced and stocks with high dividend yields are underpriced. This can be understood in the context of the Gordon Growth model which says that dividend yields proxy for the conditional expectation of the future rate of growth of dividends. These results can be thought of as a cross-sectional analogue to the work of Campbell and Shiller (1988) and Fama and French (1988) who identify dividend yields as a predictor of long run returns in an aggregate time series dataset. Focusing on the mis-pricing allows us to give an interpretation to these excess returns in terms of the present value model.

Data on individual securities has been used extensively to analyse the predictability of stock returns. There is already a sizable literature on the own- and cross-autocorrelation properties of individual stocks [Lo and MacKinlay (1988, 1990), Jegadeesh (1990), Campbell, Grossman and Wang (1993)], contrarian strategies [De Bondt and Thaler (1985, 1987), Lehmann (1990)], momentum strategies [Jegadeesh and Titman (1993), Chan, Jegadeesh and Lakonishok (1996)], and value strategies [Lakonishock, Shleifer and Vishney (1993), La Porta, Lakonishock, Shleifer and Vishney (1996)] that demonstrate that expected returns on individual securities are not constant.

Our work complements this research in two respects. First, the focus of our study is mis-pricing, rather than excess returns. That is, we compare actual prices relative to fundamental prices, rather than examine whether expected returns are predictable: the existence of positive or negative excess monthly returns does not directly map into a measure of mis-pricing. To take a specific example, suppose that small firm stocks are perpetually overpriced and large firm stocks always underpriced. Time series analysis whether aggregate or on a firm by firm basis would not be able to identify any

\(^2\) For example Bulkley and Tonks (1989, 1992), Barsky and De Long (1993), Timmermann (1993, 1996), and Donaldson and Kamstra (1996) all suggest that violations of the present value model may be due to agents incorrectly estimating the dividend process.
predictability in excess returns because permanent mis-pricing of this nature does not induce time series predictability. The time series excess returns approach would therefore conclude that stock prices were not excessively volatile. On the other hand focusing on cross-sectional mispricing as we do in this paper would reveal excess dispersion. Second, we test whether this mis-pricing may be an over-reaction to information about a particular aspect of fundamental values, namely expected dividend growth.

II Regression Tests of Cross-Sectional Volatility

The standard definition of the realised one period return on share $i$ in time $t$, $r_{i,t}$, is the return accrued from purchasing the share at price $p_{i,t}$ at date $t$, selling it at date $t+1$ for $p_{i,t+1}$ and receiving a dividend of $d_{i,t}$ at the end of the holding period. We follow Campbell (1993) who solves an intertemporal representative agent optimisation problem to derive a single factor asset pricing model with the property that the one step ahead conditional expectation of asset returns is approximately constant. Using this approximation we may write expected returns as

$$E_{i}(r_{i,t}) = \frac{E_{i}[d_{i,t} + p_{i,t+1}]}{p_{i,t}} - 1 \equiv r_{i}$$

where $r_{i}$ is a risk adjusted firm specific constant discount rate. The information set at time $t$ in equation (1) includes all current dated variables except for dividends $d_{i,t}$, which are realised between time $t$ and $t+1$.

Multiplying the approximate equality in (1) by $p_{i,t}$ and solving forward gives the present value model for stock prices

$$p_{i,t} = E_{i}\left[\sum_{k=0}^{\infty} \delta_{i,k}^{t} d_{i,t+k}\right] \equiv E_{i}[p_{i,t}^{*}]$$

where $\delta_{i} = [1+r_{i}]^{-t}$ and $p_{i,t}^{*}$ is the present discounted value of realised future dividends, which is sometimes termed the ex post rational price for stock $i$.

\(^{3}\) Campbell's analysis depends on a log-linearisation of the intertemporal budget constraint assuming that the consumption-wealth ratio is approximately constant. We take the risk free rate to be constant, so that $r_{i}$ in (1) has no time subscript.
Under the present value model in equation (2) the forecast error, $p^*_{i,t} - p_{i,t}$ should be uncorrelated with any information available at date $t$ including $p_{i,t}$. This orthogonality restriction can be tested using the regression equation

$$p^*_{i,t} = \theta + \gamma p_{i,t} + v_{i,t}$$

where under the null hypothesis of the present value model, $\theta = 0$, $\gamma = 1$ and $v_{i,t}$ is the forecast error. The time series literature has taken asset $i$ to be a single stock represented by the market portfolio, and Scott (1985) tests these restrictions on a time series of the Standard and Poor's 500 market index. We propose to estimate equation (3) on a cross-section dataset of individual share prices, and this will be estimated repeatedly for 40 successive years. In the next section we discuss the computation of the set of ex post rational prices for individual shares constructed from the CRSP data tape. However there is a problem which arises in an OLS cross-section test using data on individual stock prices. It is unlikely that the prices of individual firms are independent, and hence we encounter autocorrelation in the error term of equation (3); in general we should expect $\text{Cov}(v_{i,t}, v_{j,t}) \neq 0 \forall i, j \neq j$. Of course in large samples autocorrelation is not a problem for parameter consistency unless it generates a correlation between regressors and the error term. However we will now show that in the context of shares prices in the same economy this exactly this type of correlation that is likely to occur. Our solution to this autocorrelation problem is to follow standard practice in finance and model the error covariance structure using a factor structure: we assume there are macro economic shocks which affect the return on individual shares by an amount depending on the covariance of each share's return with the factor.

Returning to the definitions of the ex post rational price and actual price, we may redefine $p^*_{i,t}$ and $p_{i,t}$ as

$$p^*_{i,t} = \delta_i [d_{i,t} + p^*_{i,t+1}]$$

and

$$p_{i,t} = \delta_i [E_i d_{i,t} + E_i (p_{i,t+1})]$$

Combining (4) and (5), the forecast error $v_{i,t}$ in any period $t$ can be written as

\footnote{Note that time series regressions of (3) also had to deal with the problem of serial correlation.}
\[ v_{it} = p_{i,t}^* - p_{i,t} = \delta_i \{ d_{i,t} - E_i d_{i,t} \} + \delta_i \{ p_{i,t+1}^* - E_i (p_{i,t+1}) \} \]

From the definition of a realised return and its expected value we may obtain an expression for unanticipated returns on an asset. Rearranging this definition to explain unanticipated dividends gives

\[ d_{i,t} - E_i d_{i,t} = p_{i,t} \{ r_{i,t} - E_i r_{i,t} \} - \{ p_{i,t+1} - E_i (p_{i,t+1}) \} \]

Substituting for unanticipated dividends from (7) into (6), and substituting recursively for \( [p_{i,t+k}^* - p_{i,t+k}] \) ultimately yields

\[ p_{i,t}^* - p_{i,t} = \sum_{k=0}^{\infty} \delta_i^{k+1} p_{i,t+k} \{ r_{i,t+k} - E_i (r_{i,t+k}) \} \]

Equation (8) is the forecast error in the present value model, which depends on a future stream of unanticipated returns. As we have already noted these unanticipated returns will in general be correlated in the cross-section, and may be modelled using a factor structure. To make this procedure explicit, for the special case of a single factor the returns generating process for share \( i \) can be written as

\[ r_{i,t} = \alpha_i + \beta_i f_i + \omega_{i,t} \]

where \( \alpha_i \) and \( \beta_i \) are the factor model parameters, \( f_i \) is the factor and \( \omega_{i,t} \) is an identically and independently distributed error term with zero mean which represents firm specific shocks. In this factor model the risk of any individual share has two distinct elements; factor risk, to which all firms are subject, but to a varying degree measured by the covariance of the share's return with the factor \( \beta_i \); and firm specific risk, \( \omega_{i,t} \) which reflects the risk associated with an individual firm's operations. Unanticipated returns in this factor model are

\[ r_{i,t} - E_i r_{i,t} = \beta_i \{ f_i - E_i f_i \} + \omega_{i,t} \]

\(^5\) To maintain consistency with Campbell's intertemporal model, which we described in section II, we adopted a single factor structure, where the factor was specified to be the rate of return on the market. In the subsequent empirical work we examined the sensitivity of our results to the adoption of a four factor model with the factors specified as the inflation rate, term premium, default premium and growth in industrial production as in Chen, Roll and Ross (1986). We found little qualitative change in the results.
Equation (10) links unanticipated returns at a single date to unanticipated factor movements and firm specific shocks at that date. Using equation (10) in (8) and taking prices over to the right hand side gives

\[ p_{i,t}^* = p_{i,t} + \sum_{k=0}^{\infty} \delta_i^{k+1} p_{i,t+k} \left( \beta_i \left( f_{1,t+k} - E_{t+k} f_{t+k} \right) \right) + \sum_{k=0}^{\infty} \delta_i^{k+1} p_{i,t+k} \omega_{i,t+k} \]

where, as was the case with dividends earlier, the information set at time \( t+k \) does not include the current value \( f_{t+k} \). Equation (11) demonstrates explicitly the exact structure of the error term in equation (3). The error is composed of two blocks of summation terms. The first block represents the future weighted sum of forecast errors associated with the factor, and the second block is the weighted sum of future firm-specific idiosyncratic errors.

We may view equation (11) as a regression equation. It can be seen that the first block of terms in the error (11) is autocorrelated in the cross-sectional dimension. More seriously a regression of \( p_u^* \) on \( p_u \) runs into the problem that \( p_u \) itself is correlated with the first block in the cross-section. In estimating equation (11) there are two alternative methods which may be adopted to overcome these problems. The first approach includes future values of \( \delta_i^{k+1} p_{i,t+k} \beta_i \) up to a truncation point, as regressors. In practice these future terms were found to be highly collinear, and therefore instead we approximated the first block of terms in equation (11) by the first term in this block \( \delta_i p_{i,t} \beta_i \).\(^6\)

We should note that the latter regressor and \( p_u^* \) are constructed using time series estimates of \( \delta_i \) and \( \beta_i \). Our data span in the time series is relatively large, and consequently these estimates will be close to their true values, regardless of the number of cross-section observations used in the regression equation (11). This is convenient because it means that the usual time series problems of generated regressors/errors in variables [Pagan (1983)] does not arise in our cross-sectional analysis. Finally we used heteroscedasticity consistent standard errors which allows inference to proceed along standard lines in equation (11) provided the cross-section is large, which it is.

\(^6\) The extra terms require instruments because they are correlated with the error term. An Arrelano-Bond (1991) dynamic IV procedure was tried but it gave estimates that were qualitatively similar to the simple approximation in the sense that the results indicated a rejection of the present value model. The standard errors in the IV case were of course much higher than in the simple case that we present.
Our second method recognises the importance of the future terms in this first block, but rather than freely estimating the regression, where we would obtain future factor shocks as estimated coefficients, we instead construct a prior estimate of this discounted sum of terms. Using data on the factor we estimate the weighted sum of the forecast errors in this first block up to a suitable truncation point, and deduct this weighted sum from the dependent variable $p^\pi_{i,t}$. This corrected variable may now be regressed on price, so that under the null of the present value model the slope coefficient will give a consistent estimate of unity and an intercept of zero.

### III Data

We constructed a set of ex post rational stock prices, as implicitly defined in equation (2) for all shares which were quoted for at least two years on the New York or American Stock Exchanges within the period 1926-1992. All share data was obtained from the Centre for Research in Security Prices (CRSP) Tape and this resulted in a maximum number of securities of 2333, though the actual number that we use in any single year will be somewhat smaller. The ex post rational stock price for each share at any date was computed as the present discounted value of all real cash payments. That is, we include not just dividends but all other payments received by shareholders due to capital exchanges, re-organisations mergers or takeovers. All cash payments are included which accrue as the result of the purchase of a single share on January 1st 1926, or the first date the share was listed, whichever is the later. All prices and cash payment data were converted to real values by the consumer price index.

The discount rate applicable to share $i$ was taken as the average real return over the lifetime of the stock which is consistent with the formulation in equation (1). Whenever there is a capital change we ensure that after the initial investment we record total cash payments on the resulting composite bundle of shares. The ex post rational price calculations require a terminal value from which to start the discounting procedure. For dividend realizations of those companies in existence after 1992 we can do no better than follow the time series literature, and use their expectations, which under the rational expectations hypothesis is the share price in 1992. For companies that leave the

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7/ For our purposes truncating the sum at $k=20$ was thought to provide a reasonable approximation.

8/ Ackert and Smith (1993) criticise the aggregate time series tests for neglecting non-dividend payments to shareholders, and showed that inclusion of such payments could rationalise actual stock price movements.
data base for any reason we assign the terminal price as the recorded payout to the equity holders at the date they exit the tape.

Since we want the main component of the ex post rational prices to be constructed from actual realizations, we restrict attention to years sufficiently far back in time that the discounted value of the terminal price is relatively small. We choose 1971 as the cut-off for our tests, when the terminal price should on average be only 20% of the value of the ex post rational prices under an 8% annual real discount rate.

**IV Results for the Regression Tests**

Initially, to relate our cross-section data to the more familiar times series dataset, we calculated the cross-section means from our dataset of the actual and ex post rational stock prices each year 1926-92. Figure 1 plots a time series of these two sets of cross-section means: each observation is the cross-sectional unweighted average value of actual company real share prices at January 1st each year, and of the company ex post rational real stock prices at the same date. It can be seen that the relative movements in the unweighted means of the actual and ex post rational prices, are broadly similar to movements in the aggregate indices obtained from Standard and Poor's data [cf figure 1 in Grossman and Shiller (1981)].

We ran a series of cross-section regressions of equation (3), using OLS with the first modification implied by equation (11), in which we added the additional explanatory variable $\delta_p, \beta_f$ to allow for the subsequent forecast error associated with the factor. In the first two columns of table 1 we report the intercept and slope coefficients of these regressions for each year 1932-71.\(^9\) These results demonstrate a striking rejection of the null hypothesis of the present value model. It can be seen that the intercept term is consistently significantly positive and the slope coefficient is significantly less than unity for 34 out of the 40 years. As noted above, extra terms in the forward sum in the error term in (11) were added and instrumented, and the regressions were repeated. The

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\(^9\) In section V of the paper, we need five years of lagged dividends, and therefore in order to be consistent throughout the paper we work with the 1932-71 dataset of 40 years. Inclusion of the earlier five years sets of regressions does not affect the thrust of the results.
slope coefficients on the current price in this case were somewhat lower so these results still indicate rejection of the null of the present value model.

To get a single test statistic of the null hypothesis that the slope coefficients in all the cross-section regressions were unity, we compute the average coefficient value over the 40 cross-sections (years) and divide it by an estimate of its standard deviation. If the coefficients were independent across time estimating the standard deviation of their average would be straightforward, but this is unlikely to be the case. However if we are able to assume that they are stationary through time and satisfy a general mixing condition [see McCabe and Tremayne (1993)], then we may use Bartlett's estimator of long run variance to get a statistic that is asymptotically standard normal [see Kwiatoski, Phillips, Schmidt and Shin (1992), p. 164]. We refer to our normalised average coefficient as the "overall t-ratio".

We compute the average intercept and slope coefficients as 0.284 and 0.540, with the overall t-ratio as 6.493 and 4.54 respectively. These summary statistics show a convincing rejection of the null over the whole sample period.

The second modification applied to implement equation (11) adjusts the dependent variable \( p_{i,t}^* \) and requires estimates of both the factor loadings and the one-step-ahead forecast errors of the factor. The former parameter values were obtained from the time series regression in equation (9) using all the monthly observations that were available for each stock. In keeping with the assumption in equation (1) that rates of return are approximately unforecastable, we estimated the series of annual one-step-ahead forecast errors as the difference between the market rate series and its average value over the years 1926-1992. As before we are using a relatively large time span to obtain estimates of \( \delta_i \), \( \beta_i \) and market forecast errors. These time-series based quantities will be close

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\[ ^{10} \text{In implementing Bartlett's estimator of long run variance we used a lag truncation parameter of four. Changing the order of lag truncation makes little difference to the test statistic or its significance.} \]

\[ ^{11} \text{In reality market rates of return may be predictable from their past [Pesaran and Timmermann (1995)], but in our sample the autocorrelation structure of the market rate is quite weak. For example the } \chi^2_2 \text{ statistic for the significance of the first two partial autocorrelations is 4.15. Nonetheless a sensitivity analysis where we allow for an AR(2) structure in the market rate did not qualitatively alter our results.} \]
to their true values when used in the cross-sectional analysis, regardless of the number of cross-sectional observations.

The results of the forty annual regressions for each year 1932-1971 are reported in table 2. Up until 1954 the intercept terms are predominantly negative although only just over half of these are significantly different from zero. After 1954 nearly all of the intercepts are significantly positive. As before, the slope coefficient which should be unity under the null hypothesis of rational expectations, is significantly less than unity on all 40 occasions. In fact the rejections of the null is so dramatic in this table, that surprisingly on occasions in the earlier years the slope coefficient is negative. We cannot offer an argument as to why this might have occurred, but note that once the null has been rejected there may be other omitted variables that can explain the forecast error, and given their omission the estimated parameters will be biased.

Computing the overall t-ratios again for the significance of the intercept from zero and the slope coefficient from unity in (3) gives mean values of -0.0285 and 0.407, and associated overall t-ratios as 8.706 and 8.089 respectively. These results confirm the general significance of cross-sectional mis-pricing during the 40 years of our sample.

We also undertook a number of additional sensitivity tests to assess the robustness of our results. First we divided the cross-sectional sample alphabetically into four equal groups and performed the cross-sectional tests on each sample. The results for each of the sub-samples were almost identical to those for the full sample. To motivate a second sensitivity test note that the construction of the ex post rational series is dependent on the discount rate applied to each firm’s dividend series. As we explained in the previous section we took the average realised real return over the companies lifetime as the appropriate firm specific discount rate. This annualised value averaged across stocks was 11.55%, and is rather high in comparison with the average real return of 8.8% on the S&P500 over the same period. The reason for the difference is probably because our sample average is unweighted, and gives undue weight to small stocks which on average have performed well over our sample period. To determine whether our results are sensitive to the value of the firm specific discount rate we re-calculated the ex post rational price series for each firm using a 10% lower discount rate for each stock, and re-estimated the cross-sectional regressions. The qualitative results were again not greatly affected by this new discount rate.
Broadly speaking, the results in tables 1 and 2 all indicate that the slope coefficient in the cross sectional regression of $p_{i,t}^*$ on $p_{i,t}$ is less than unity and for the later (earlier) years the intercept is typically positive (negative). A slope coefficient less than unity coupled with a positive intercept implies that prices are excessively dispersed, indicating a micro-component to excess volatility. Stocks with high $p_{i,t}^*$s are overpriced, and have prices higher than are warranted by the subsequent dividend realisations. Similarly a stock with a low $p_{i,t}^*$ is underpriced, and has a price lower than is in fact warranted. The negative intercepts before 1955 combined with the low value of the slope coefficient would appear to suggest that all stocks were overpriced for these years.

In these tests an interpretation of our principal result that the slope coefficient on price is less than unity is that there is an additive stock specific component in the stock price that drives a wedge between the observed price and its fundamental value. Provided that this wedge is uncorrelated across firms, then its existence will cause downward bias on the slope coefficient in the regression of $p_{i,t}^*$ on $p_{i,t}$ which is in fact what we find.

In summary what our results in this section show is that across a sample of firms the stock price is not always an unbiased predictor of subsequent dividend realisations. The time series literature obtained this result by looking at successive observations on the stock price index. We show it holds true also when the data set consists of a large number of firms in a single year.

V Predictability of Mis-pricing

We now investigate the source of the mis-pricing identified in the previous section, by relating it to key elements in the information set. We measure mis-pricing as the difference between $p_{i,t}^*$ and $p_{i,t}$. The ex post rational price is the present value of dividend realisations and therefore the difference in the ex post rational and the actual stock price relative to the actual stock price can be interpreted as a long run rate of return [Campbell and Shiller (1988)]. Under the null hypothesis of the present value model, we would expect this variable to be unforecastable using current information, but the results of section IV show that this is not so: high price stocks are overpriced and low price stocks are underpriced. Here, we investigate the hypothesis proposed by Lakonishok, Shleifer and Vishny (LSV) (1994) that stock mis-pricing is due to excessively dispersed earnings growth forecasts. LSV (1994) argued that the correlation between market-to-book and subsequent
returns could be explained as a consequence of high earnings growth expectations, resulting in high market-to-book, typically being over-optimistic. In our context this hypothesis would imply stocks with high earnings growth expectations would have market prices above their ex post rational values. We can test this hypothesis if we assume that current dividend yields are a proxy for earnings growth expectations [for example via the Gordon Growth model], and then examining whether the percentage mis-pricing \((p^*_t - p_{it})/p_{it}\) is explained by the dividend yield. From (3) and including \(d_{i,t-1}/p_{i,t}\) as a regressor we have

\[
\frac{p^*_t - p_{it}}{p_{it}} = \lambda + \pi \frac{d_{i,t-1}}{p_{i,t}} + \frac{v_{i,t}}{p_{i,t}}
\]

Under the null hypothesis of efficient markets \(\pi = \lambda = 0\), but under the LSV hypothesis \(\pi\) should be positive. It is important to note that the error term in (12) \(v_{i,t}\) is the same error as before except for the price term in the denominator, and as a result we use the same correction as for \(p^*_t\) in the second modification to equation (11).

The results of running the regression in (12) for each year between 1932-1971 are given in table 3. It can be seen that the coefficient \(\pi\) on dividends is positive in all but the first year and is significantly different from zero in 27 years. The intercept coefficient \(\lambda\) is significantly negative in most years. As a supplementary exercise, five additional lagged dividends over price terms were added to the right hand side of equation (12). The results given in table 4 appear to have no systematic sign or magnitude across years. We may examine the long run effect of dividends on mis-pricing implied by this regression, by summing the individual lagged dividend coefficients. Table 5 shows that the long run effect is significantly positive in nearly all years. In fact the long run effects are quite stable and qualitatively similar to their static counterparts (the estimates of \(\pi\)) given in table 3.

Combining the positive estimates of \(\pi\) in (12) with the finding of a negative intercept \(\lambda\) means that when dividend yields are high (low expected dividend growth) stocks are undervalued: \(p_{it}\) tends to be less than \(p^*_{i,t}\). When dividend yields are low stocks are overvalued: \(p_{it}\) tends to be greater than \(p^*_{i,t}\). Interpreting the dividend yield as a proxy for expected earnings growth then these results imply that investors appear to be unduly optimistic about high expected growth stocks.
and too pessimistic about low expected growth stocks. Put another way it appears that investors' beliefs about future company growth prospects are too widely dispersed!

Fama and French (1988) document a similar positive relationship between market dividend yields and a measure of long run excess returns, but in the time series domain. They explain this finding in terms of a time varying risk premium. It is difficult to see how this explanation of mis-pricing can apply to our cross-section results. We would expect macroeconomic phenomena such as a time varying risk premium to affect all stocks in the cross-section. However we identify both underpricing of some stocks and overpricing of others in the same time period. Therefore unlike previous work based on aggregate time series analysis, our rejection of efficient markets may not be circumscribed by falling back on an aggregate time-varying risk premium: a microeconomic theory is required to explain our results.

VI Conclusions

The result that there are movements in stock prices which are excessive relative to movements in fundamentals, has proved remarkably robust with respect to a number of tests, applied to a time series index of stock prices. In this paper we have applied a regression test of volatility to a cross-section data set, specifically testing Campbell's constant discount rate present value model. We have shown that the rejection of the present value model carries over to a data set consisting of observations on a cross-section of individual share prices within a particular year, and we have referred to this phenomena as “excess dispersion”. In nearly all of the years over the period 1932-1971 we have found that stock prices were excessively dispersed: firms with high $p^*$s have prices higher than are warranted by the subsequent dividend realisations; firms with a low $p^*$s have stock prices that are lower than are in fact warranted. This finding is consistent with the existence of a firm specific bubble, driving a wedge between the values of $p_t^*$ and $p_t$.

When we amended our regression test to allow for the cross-correlation of security prices, and for the existence of firm specific time varying discount rates, the dramatic rejection of the null hypothesis still obtained.

We went on to examine the relationship between the mis-pricing and market fundamentals which we took to be related to past dividends. Assuming that dividend yields proxy for growth expectations we found that investors are unduly optimistic about high growth stocks and too pessimistic about low expected growth stocks. Hence the mis-pricing we have identified is not just
a macroeconomic phenomena whereby all shares are either underpriced or overpriced by a similar amount, and cannot be explained away by the existence of time varying risk premia. Our results suggest that there is a microeconomic source of mis-pricing. Within the same time period those stocks with high dividend yields tend to be undervalued, and when dividend yields are low the stocks are overvalued. We would agree with Lakonishok, Shleifer and Vishney (1994) that these results can be explained by market participants having a preference for glamour or high expected growth stocks pushing up their prices, and a corresponding reluctance by investors to hold low growth securities, which depresses the prices of these stocks.
References


