

Dynamics of the Relation between Prices and Fundamentals[†]

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Comments Welcome.

ABSTRACT: Motivated by the debate on whether prices reflect fundamentals, we investigate the tendency of aggregate prices to be cointegrated with measures of fundamental value, where fundamental value is measured using accounting and analyst inputs (hereafter, fundamentals). We document significant time-series variation in the tendency of prices and fundamentals to be cointegrated. We next examine whether the tendency of prices and fundamentals to be cointegrated is associated with macro-economic indicators, arbitrage costs and accounting quality. We identify that evidence of cointegration is lower during periods of high consumer confidence. We find mixed results for other explanations based on macro-economic news (such as GDP growth and unemployment), costly arbitrage, and accounting quality.

Keywords: Fundamental valuation; market price dynamics; limits to arbitrage; macro-finance.

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1. Introduction

The observation of extreme price movements in recent “bubbles,” “crashes,” and “financial crises” amplify the debate about whether prices reflect the present value of future cash-flows to shareholders (e.g., Shiller, 2000; Stiglitz, 2003; Hunter et al., 2005). Central to this debate is whether significant price movements appear to be accompanied by significant movements in the underlying fundamental value of the stock (Campbell and Shiller, 1987; Brunnermeier 2001, 47; Lee et al., 1999). Results informing the debate are important as evidence of a disconnect between price movements and their underlying fundamentals have social welfare implications for tactical asset allocation and policy implications for the regulation of financial markets. The results of this study contribute to the debate by documenting evidence of disconnects between price movements and their underlying fundamentals using cointegration analysis and by providing evidence of factors that are associated with a reduction in the tendency of prices and fundamentals to be cointegrated.¹

Prior theoretical research describes situations where price movements do not reflect changes in fundamentals (Shiller 1984; Summers 1986; Black 1986). In recent models, prices diverge from their underlying fundamentals when arbitrage is costly (Shleifer and Vishny 1997) and this divergence is expected to last longer in periods of high expected growth, because arbitrage actions are expected to be delayed (Abreu and Brunnermeier 2001; 2002). A central theme in these models is that a fundamental value can be established, and then a level of speculation, or mispricing, can be inferred from the difference between price and fundamental value. Part of the difficulty with assessing whether equity prices reflect fundamental values is

¹ Cointegration is a statistical term that describes the long-run equilibrium tendency of two variables to be tied together (Hamilton 1994). When examining nonstationary (i.e., not mean-reverting) variables, cointegration is an appropriate method; variance ratios, which have been used in prior research, are not appropriate as variance is undefined for nonstationary variables (Kleidon 1986).

that there has been limited success in documenting accounting variables that are cointegrated with equity prices.²

All valuation methods based on the present value of future cash-flows should yield the same valuation regardless of the inputs (e.g. Lundholm and O’Keefe 2001). Prior empirical literature, however, suggests that given the inputs available to researchers, the residual income model provides a superior valuation model relative to other discounted cash-flow models in terms of 1) their association with current price levels (e.g., Penman and Sougiannis, 1998; Dechow et al., 1999; Penman, 2001), 2) their association with future returns (Frankel and Lee, 1998; Dechow et al., 1999), and 3) their tendency to be cointegrated with prices (Lee et al. 1999).

The residual income model is calculated by combining the information in book values and expected earnings, and can incorporate time-varying discount rates and analysts’ forecasts of future earnings (Lee et al., 1999). Lee et al. (1999) document that estimates of a residual income model that incorporate these features is cointegrated with prices for the Dow 30; evidence that differences between prices and fundamentals are temporary. Similar results for the Dow 30 and other market indices are documented in Ritter and Warr (2002), Bakshi and Chen (2005) and Brown and Cliff (2004; 2005).³ Taken together, these studies suggest that there is time-series variation in the strength of cointegration between aggregate prices and fundamentals, a result we explicitly test in our analysis. Directly extending this literature, our study is the first to investigate whether macro-economic factors are associated with the tendency of prices and fundamentals to be cointegrated.

² For example, ratios of dividends-, bookvalues- and earnings-to-price are typically nonstationary, suggesting that these fundamentals do not appear to be cointegrated with stock prices (LeRoy and Porter, 1981; Shiller, 1989; Hodrick, 1992; Lee et al., 1999; Goyal and Welch 2003; Lettau and Ludvigson 2005).

³ Using more recent data, Curtis (2012) documents a lack of cointegration between prices and fundamentals.

We explicitly measure the time-series variation in the strength of cointegration between aggregate prices and fundamentals, using test statistics from rolling window cointegration tests.⁴ Similar to Curtis (2012) we document that statistically significant evidence of cointegration between prices and fundamentals is limited to periods prior to the late 1990s, at least according to traditional levels of significance. We extend prior research by examining the test statistics from a rolling window allowing us to identify the changes in the strength of cointegration on a monthly basis. This innovation in the measurement of the strength of cointegration allows us to examine factors that are associated with periods where aggregate prices reflect more or less of the information in fundamentals.

In a cointegration framework, the tendency of cointegration between prices and fundamentals is a function of how well each of these variables tracks intrinsic value (e.g., Lee et al., 1999). The tendency, or strength, of cointegration is lower when either measurement errors in fundamentals are more persistent, or the level of mispricing is more persistent. For example, if a change in accounting lead to a more persistent under-recognition of assets relative to the prior accounting treatment then the measurement error in accounting would be more persistent. Similarly, if mispricing is removed less quickly by arbitrageurs in a recent period relative to the prior period, then the persistence of the mispricing error is more persistent in the recent period. Measurement error and mispricing are not mutually exclusive, instead they are likely correlated. For example, accounting does not capture expected growth, and expected growth is the most likely part of value to be subject to speculation (e.g., Penman, 2003; 2011). In this case, when firms in the Dow 30 experience a period where there is higher expected growth, investors could over-extrapolate current growth into the future (increasing mispricing) and accounting will not

⁴ Specifically, we suggest that the rolling window estimates of the Phillips-Perron test statistics (Z-Rho and Z-Tau) for cointegration and their associated *p*-values can serve as both continuous and time-varying measures of the strength of the relation between prices and fundamentals.

capture the implications of current growth for future growth (increasing measurement error). In this case, the errors in prices and fundamentals will both increase the disparity between prices and fundamentals.

Part of the reason that evidence of cointegration between prices and accounting-based measures of fundamental value is weak could be due to accounting inputs being measured with an emphasis on verifiability, limiting the ability of these inputs to incorporate changes in market conditions and expected growth. While the residual income model is able to incorporate some of the changes in market conditions and expected growth through the use of time-varying discount rates and analysts' forecasts, these features are not sufficient to produce valuation estimates that are cointegrated with prices in the post 2000 period (Curtis, 2012). We investigate macro-economic news as a source of information on changes in market conditions and expected growth that are expected to affect the tendency of prices and fundamentals to be cointegrated. Consistent with our expectations, consumer confidence, a proxy for macro-economic news on future expected consumption and investor sentiment (e.g., Lemmon and Portniaguina 2006)⁵, is associated with a lower tendency of cointegration.

Another reason that the tendency of prices and fundamental values to be cointegrated is reduced is that limits to arbitrage prevent prices from immediately reflecting intrinsic value. Shleifer and Vishny (1997) argue that the separation of the knowledge required to implement arbitrage positions and the resources required for their implementation leads to an imperfect market for arbitrage resources, and horizon risk. Horizon risk is the risk that mispricing does not revert before the arbitrageur needs to report their performance to their principal (e.g., DeLong et al., 1990a, 1990b; Dow and Gorton, 1994). Significant horizon risk would lead to more

⁵ Lemmon and Portniaguina (2006) suggest that consumer confidence proxies for both expectations of future expenditure and consumer sentiment. They next provide evidence that consumer sentiment affects stock prices, consistent with consumer sentiment and investor sentiment being correlated.

permanent mispricing, for example, Abreu and Brunnermeier (2002, 2003) show that in equilibrium, arbitrageurs wish to trade with (rather than against) mispricing due to the horizon risk constraint along with the assumption that arbitrageurs require the coordinated actions of other arbitrageurs to correct mispricing. These studies suggest that arbitrageurs will not undertake profitable long-term arbitrage opportunities when there is a risk that price will diverge further from fundamental value in the short- to intermediate-term. We find mixed evidence consistent for a limited arbitrage based explanation as variation in the tendency of prices and fundamentals to be cointegrated is lower when idiosyncratic risk,⁶ a measure of the difficulty to arbitrage, is higher, but only when the model does not include consumer confidence.

Accounting quality – how well accounting reflects true economic position and performance of the firm will also affect the tendency of prices and fundamentals to be cointegrated. We find evidence that the degree of matching, measured as the correlation between revenues and expenses (following Dichev and Tang 2008), is positively associated with the tendency of cointegration. Other measures, such as the volatility of revenues, expenses and net income provide mixed evidence.

Our paper adds to the relatively recent literature that focuses on identifying and explaining deviations of prices from fundamentals. Prior work includes that of Mitchell et al. (2007) who document that convertible bond arbitrage prices deviate from fundamentals and Mitchell and Pulvino (2011) who provide similar results for the over-the-counter derivatives market. Similarly, Fleckenstein et al. (2010), show that the price of Treasury bonds can differ substantially from the price of “synthetic” Treasury bonds created via inflation-swapped “TIPS.” Finally, Pasquariello (2011) creates a measure for “financial market dislocations” that is derived

⁶ This measure is based on the lack of close substitutes to proxy for the amount of unhedged risk that an undiversified arbitrageur would be exposed to (Pontiff, 1996; 2006).

from violations of three textbook arbitrage opportunities. Curtis (2011) provides evidence of a link between price deviations from fundamentals and the speed to which accounting fundamentals are reflected in stock prices for a cross-section of firms.

Our study makes the following contributions to the literature. First, we measure the strength of cointegration, or the extent to which equity prices reflect fundamentals, using a time-varying test. Second, we present evidence that there are significant changes over time in the strength of cointegration. Third, we identify macro-economic factors that are associated with the strength of cointegration between aggregate prices and fundamentals.

The remainder of the study is as follows. Section 2 presents our model and hypotheses. Section 3 discusses our sample and variable measurement. Section 4 presents our empirical analysis. Section 5 presents further analysis. Section 6 concludes.

2. Model and hypotheses

2.1. The cointegration framework

The appropriate framework for time-series analysis of prices and fundamentals is the cointegration based framework discussed in Lee et al. (1999) among others. Consistent with prior literature, we make the assumption that intrinsic value is unobservable and equal to the present value of all future dividends:

$$V_t^* = \sum_{\tau=1}^{\infty} E_t(R_t^{-\tau} D_{t+\tau}), \quad (1)$$

where V_t^* is the unobservable intrinsic value of the stock assumed to be a random walk (Samuelson 1965), and $R_t^{-\tau}$ is one plus the discount rate, which is time-varying and assumed to be greater than zero (Campbell and Shiller 1988). $D_{t+\tau}$ equals the real dividend paid to the owner

of the stock between $t-1$ and t , and E_t denotes expectations conditional on information available at time t .

The next assumption required in the cointegration based framework is that both prices, P_t , and measures of fundamentals, F_t , are noisy estimates of intrinsic value. When written in log form, the expected relation between log price (p_t) and (unobservable) log intrinsic value (v_t^*) and between log fundamentals (f_t) and (unobservable) log intrinsic value (v_t^*) are:

$$p_t = v_t^* + u_{1t}^*, \quad (2)$$

$$f_t = v_t^* + u_{2t}^*, \quad (3)$$

where $p_t = \log(P_t)$, $v_t^* = \log(V_t^*)$, u_{1t}^* is unobservable mispricing error, $f_t = \log(F_t)$, and u_{2t}^* is unobservable measurement error.

In this framework, price and accounting fundamentals are measures of intrinsic value with error, and they share the unobservable intrinsic value as a common trend. Combining the equations the difference between fundamentals and price can be considered as a function of the difference between the two unobservable errors u_{1t}^* and u_{2t}^* :

$$p_t - f_t = u_{1t}^* - u_{2t}^*, \quad (4)$$

As Equation (4) is expressed in logs, the difference between price and fundamentals can also be expressed as the log ratio of price and fundamentals. In a cointegrated framework, the ratio of price-to-fundamentals (p_t/f_t) is stationary, which provides statistical evidence of cointegration between prices and fundamentals.

To test for stationarity we estimate standard tests for stationarity using the Phillips and Perron (1988) regression models, which include a constant and time-trend:

$$\Delta(p_t/f_t) = a_0 + (\rho - 1) * (p_{t-1}/f_{t-1}) + e_t, \text{ and} \quad (5)$$

$$\Delta(p_t/f_t) = a_0 + \delta t + (\rho - 1) * (p_{t-1}/f_{t-1}) + e_t, \quad (6)$$

where Equation (5) includes a constant and Equation (6) includes both a constant and a time-trend. The null in both regressions is that the price-to-fundamental ratio has a unit root (i.e., it is nonstationary) when $\rho = 1$. Estimated values of $\rho > 1$ imply explosive behavior in the ratio of price to fundamentals. Alternatively, $\rho < 1$, would indicate evidence of cointegration between accounting fundamentals and price, as the changes in errors, on average, are reducing the prior level of the disparity (as $(\rho - 1) < 0$) in this case). There are two formal test statistics from the regression models: 1) an adjusted regression coefficient labeled Z-rho, which is calculated by adjusting the estimate of $T * (\rho - 1)$, where T is number of time-series observations used in the estimation, and 2) Z-tau, which is an adjusted t -statistic related to the $(\rho - 1)$ coefficient estimate.⁷ If $\rho < 1$, both Z-rho and Z-tau are negative, and are more negative for lower estimates of ρ where lower values of ρ imply less persistent, or faster decaying, errors.

We measure the strength of cointegration between price and fundamentals using the p -values associated with the Z-Rho statistic which are based on the distribution provided in Phillips and Perron (1988).⁸ Specifically, we estimate Equations (5) and (6) using rolling windows of 60 observations, such that our measure at the end of each month provides a test of cointegration based on the prior five years of movements in price and fundamentals. As the null in Equations (5) and (6) is that the ratio is nonstationary, higher p -values for an estimate of Z-Rho suggest that the strength of cointegration between price and fundamentals for the five year period is lower. Specifically, the p -value measures the degree of confidence that can be placed on the hypothesis that price and fundamentals are not cointegrated. Lower p -values therefore suggest that during the past five year period, the strength of cointegration between prices and fundamentals is stronger.

⁷ The adjustments to these statistics are functions of T , the determinants of the cross-product matrix and the difference in the squared sum less the squared errors of the model (Maddala and Kim, 2002 80-81).

⁸ Similar results are found when we use either the estimates of Z-Rho or the p -values associated with Z-Tau statistic.

As we are interested in the strength of cointegration, the ratio of prices-to-fundamentals is not a sufficient statistic. Specifically, if prices diverge more from intrinsic value, then u_{1t}^* will increase, however, if that increase in mispricing is perfectly transitory then there are no expected effects on the strength of cointegration. The same logic follows for increases in u_{2t}^* , if shocks to growth expectations are perfectly transitory then there are no expected effects on the strength of cointegration. Alternatively, if that increase in mispricing or measurement error in fundamentals is more persistent, then the autocorrelation coefficient ρ will approach 1, which implies that the coefficient $(\rho - 1)$ is closer to 0. If the changes in the price-to-fundamental ratio have no association with the prior level of the price-to-fundamental ratio, the implication is that there is no tendency of either prices or fundamentals to correct for the prior differences in the levels of price and fundamentals. In this extreme case, there is no evidence of cointegration between prices and fundamentals. By examining the test statistics from estimating Equations (5) and (6) using rolling windows we are able to examine the level of persistence of mispricings and measurement errors for different periods of time.

Equation (4) leads to two distinct, but unobservable channels through which the strength of cointegration will decline. Specifically, if there is an increase in the persistence of either mispricing or measurement error in a period of time, then the strength of cointegration is lower. As both of these channels are unobservable, and not mutually exclusive, we investigate factors that are likely to be more associated with one channel than the other. Our first hypothesis relates to macro-economic news. Macro-economic news relates to the growth of the economy, rather than directly to the growth of the firms which are part of that economy (e.g., GDP growth, Consumer Confidence etc). Markets incorporate information in macro-economic news on the days of their scheduled releases, while the effect on future cash-flows recorded by accounting

will not incorporate these changes until the resulting earnings are verifiable⁹. This leads to a lag in the incorporation of current macro-economic news in accounting-based measures of fundamentals. We assume that current macro-economic news is incorporated into prices with respect to the expected implications for future cash-flows.¹⁰ We predict that proxies for macro-economic news will lower the cointegration between prices and fundamentals. We refer to this as the “macro-economic news” hypothesis, or H1:

H1: Macro-economic news variables are significantly associated with a lower strength of cointegration.

In addition to macro-economic news variables influencing the tendency of prices and values to be cointegrated, accounting quality is also expected to affect the strength of cointegration. Higher quality accounting is likely to have lower errors in the measurement of assets, liabilities and income, which implies that the errors in the valuation model will not be as persistent. Accounting quality is distinct from macro-economic news, as measurement errors associated with accounting quality will influence the inputs into our valuation models (whereas macroeconomic news is aimed at identifying factors excluded from our models). This leads to a related hypothesis to H1 – a set of hypotheses aimed at identifying measurement errors in accounting – we call this hypothesis the accounting quality hypothesis, which we label *HIA*:

HIA: Accounting quality variables are significantly associated with a higher strength of cointegration.

⁹ This limitation of accounting measurement extends across numerous other sources such as changes in the expensing of research and development costs (or other conservative accounting treatments). We focus on macro-economic news as we are interested in the aggregate pricing of accounting fundamentals. Future research could investigate whether specific accounting measurement limitations, when aggregated, affect the pricing of aggregate fundamentals.

¹⁰ To the extent that current macro-economic news is not viewed by market participants as having implications for future cash-flows, our tests are biased towards finding no result. Similarly, if the macro-economic news is incorporated into analysts' forecasts in the same way it is incorporated into prices, then we will find no result.

Our second hypothesis relates to the permanence of mispricings. Theoretical models based on limited arbitrage suggest that the persistence of mispricing is a function of the difficulty of arbitrage (Abreu and Brunnermeier, 2001; 2002). Following the prior empirical literature (e.g., Pontiff, 1995; 2006; Wurgler and Zhuravskaya, 2002) we predict that the level of idiosyncratic risk in the firms that comprise our index are associated with a lower strength of cointegration. We refer to this as the “limited arbitrage” hypothesis, or H2:

H2: The level of idiosyncratic risk in the individual stocks comprising the index is associated with a lower strength of cointegration.

While our hypotheses are motivated from the identification of factors that are likely to be more associated with one channel than the other, these channels are not mutually exclusive. For example, prior research has attributed consumer confidence to both mispricing and measurement error explanations (Brown and Cliff 2004; 2005; Lemmon and Portniaguina 2006). In our setting, H1 is not unambiguously due to the lack of accounting to incorporate growth (as reflected in consumer confidence) as the valuation models we use include analysts’ forecasts. It is also possible that investors speculate that good current economic conditions will lead to greater future cash-flows to individual firms than would be expected due to competitive pressures. In such a case, both the mispricing error (more positive) and the measurement error (more negative) in fundamentals lead to an increase in the divergence of prices and fundamentals.

3. Sample and Variable Measurement

3.1. Sample and Measurement of the Dow 30 Market Value

The Dow Jones Industrial Index, or Dow 30, is one of the most widely followed equity indices in the world. We construct the Dow 30 index using the Dow Jones Industrial Average Historical Components brochure. Similar to Lee et al. (1999), not all firms are available in the earlier section of our time period. In each month, the minimum number of firms included in the aggregation is 23, most months include all 30 firms included in the Dow 30 index at the end of the month. Following Lee et al. (1999), we aggregate available firm-data at the monthly frequency to construct time-series of the Dow 30 aggregate market value. We aggregate all available firms listed in the Dow 30 at the end of each month. Specifically, the aggregate market value of the Dow 30 is the sum of the market value of Dow 30 stocks, where market value is the product of the end of month CRSP price and CRSP shares outstanding.

3.2. Measurement of fundamental value

In our construction of fundamental values, we use annual book-values of equity and earnings from Compustat. We also require forecasts of earnings which we source from I/B/E/S. To calculate time-varying rates of return we use the Fama and French monthly risk-free interest rate available from WRDS. Our equity risk premia, which we discuss in further detail below, are generally measured using data from either CRSP or from the Fama and French database.

We use the residual income model to derive an estimate of fundamental value (e.g., Frankel and Lee, 1998; Lee et al., 1999). Using the clean-surplus relation, the residual income model can be used to rewrite the dividend discount model as a function of end-of-year book-values (B_t) and a discounted expectation of the amount of future earnings (X_{t+i}) that exceed the required rate of return (r_t):

$$F_t = B_t + \sum_{i=1}^{\infty} \frac{E_t[X_{t+i} - r_t B_{t+i-1}]}{(1+r_t)^i}. \quad (7)$$

Equation (7) can also be written in terms of return on equity (ROE), where ROE is defined as earnings for period t divided by opening book-value ($t-1$):

$$F_t = B_t + \sum_{i=1}^{\infty} \frac{E_t[(ROE_{t+i}-r_t)B_{t+i-1}]}{(1+r_t)^i}. \quad (8)$$

Book values are constructed using the most recent annual book-value of common equity (Compustat variable CEQ). We use analysts' forecasts of annual earnings from the Thomson Reuters I/B/E/S unadjusted summary database. We retain annual book-values for 12 months to maintain consistency with the use of discounting of future annual net income amounts.¹¹ Equations (7) and (8) are written as a discounted sum of an infinite stream of future residual income, in order to practically implement this model we require an explicit forecast horizon. We follow Lee et al. (1999) in estimating the model using a three-stage forecast: (1) the first two years are forecast explicitly using analysts' forecasts, (2) we linearly fade expected earnings to the industry ROE over a medium horizon, ending in period 3, and (3) value beyond period 3 is estimated as a perpetuity of residual income at time 3. Specifically, we compute fundamental values for each firm based on the following finite horizon model:

$$F_t = B_t + \frac{[(ROE_{t+1}-r_t)B_t]}{(1+r_t)} + \frac{[(ROE_{t+2}-r_t)B_{t+1}]}{(1+r_t)^2} + TV, \quad (9)$$

Where TV is the terminal value, which is estimated as the sum of the linear fade to the industry ROE to period $t+K-1$ and a perpetuity to capture the present value of all cash-flows post period K :¹²

$$TV = \sum_{i=3}^K \frac{[(ROE_{t+i}-r_t)B_{t+i-1}]}{(1+r_t)^i} + \frac{[(ROE_{t+1}-r_t)B_{t+i-1}]}{(1+r_t)^i r_t}. \quad (10)$$

¹¹ While book values change on a quarterly basis due to the incorporation of earnings in the quarter, analysts' forecasts also incorporate this information. Using quarterly book values with analysts' forecasts would result in an inconsistency in the calculation of future residual income as the horizon of book values are increased while expected earnings are held constant.

¹² Similar to Lee et al. (1999) our results are robust to the implementation of the model with shorter forecast horizons over 3 and 12 months.

where, $K = 3, 12,$ and 18 . The target industry ROE is the prior median of all firms in the same industry based on the 48 industry classifications in Fama and French (1997). We use 10 years of prior Compustat data to calculate the target industry ROE median. Similar to Lee et al. (1999) this procedure attempts to capture the expected mean reversion in ROE and assumes that there is no expected growth in residual income after period K .¹³

The residual income model requires a discount rate based on the riskiness of future cash-flows to shareholders. Lee et al. (1999) find that the inclusion of a time-varying risk-free rate is important when calculating the time-series of residual income. The discount rate can be further broken down into a time-varying risk-free rate and time-varying equity premium. Lee et al. (1999) provide evidence that short-term rates, such as the T-Bill rate, provide the best estimates of the time-varying risk-free rate. Lee et al. (1999) also note that the models are not sensitive to the choice of a constant equity premium (of 4 to 7%) versus the use of a time-varying equity premium based on prior industry returns to Fama-French (1997) industry portfolios, or prior returns to the NYSE/AMEX value-weighted portfolio. In our primary analysis, we calculate the discount rate as the sum of the monthly annualized one-month T-Bill rate plus a constant equity premium of 5%. We consider models which include time-varying measures of the equity premium in our robustness analysis.

Our estimates of fundamentals are calculated at the per-share value on a monthly basis. We aggregate fundamentals by summing the 30 firms' per share fundamental values to be consistent with our aggregated price for the Dow 30 which is a share price weighted index. We limit the valuation models to only include information available at the end of each month. We label the aggregate fundamental models using the following conventions. We use the term "V" to

¹³ As our sample begins in 1979 the majority of the firms in the Dow 30 have analyst coverage, in the few rare cases where the analysts forecasts are unavailable we follow the procedure in Lee et al. (1999) to estimate a forecast of future ROE using the parameters of a pooled autoregression of ROE on lagged ROE for the Dow 30 firms.

refer to the fundamental value model, followed by the horizon K , we then add an acronym for the method of the discount rate that we apply to the model. For example, we label the 12-year horizon model with the time-varying T-Bill rate and an equity premium of 5% as V12(TBe), where “TB” refers to the Treasury Bill rate and “e” refers to the constant equity premium of 5%. In our main analysis we report results for the V12(TBe) model, we also report some descriptive results for alternative horizon models.¹⁴

3.3. Measurement of macro-economic news

To test hypothesis 1, we investigate macroeconomic variables that have been used in prior literature which are generally considered as indicators of economic growth. Specifically, we investigate the index of consumer sentiment, (ICS), (Brown and Cliff, 2004; 2005; Lemmon and Portniaguina 2006), the GDP growth rate, and the consumption-wealth variable, (CAY), (Lettau and Ludvigson, 2005). We also examine the default spread, Treasury Bill yields, the dividend yield of the broad market index, labor income, unemployment, and inflation (Lemmon and Portniaguina 2006).

Lemmon and Portniaguina (2006) provide evidence that the consumer confidence index is highly correlated with macroeconomic growth and is a reasonably good predictor of business cycles. The ICS covers a large number of US households and includes questions regarding both consumers’ expectations of future economic performance and their future expectations of personal consumption for major household items. To the extent that these expectations of future consumption are priced by the market, we expect the ICS serve as a measure of omitted expected

¹⁴ In future versions we will report more of the robustness results using models that differ in terms of the calculation of the discount rate.

growth in the expected earnings of Dow 30 firms. In our analysis we log the ICS, and refer to this variable as *ics*.

We measure the default spread, *def*, as the difference between the yield to maturity of Moody's Baa-rated and Aaa-rated bonds. We use the yield from three-month Treasury Constant Yield to Maturity T-Bill to measure the treasury yield, *yld*. The dividend yield, *dp*, is the total ordinary cash dividend from the CRSP value-weighted index over the last twelve months divided by the current month's index value default spread. The change in GDP, *gdp*, is the quarterly change in the natural log of GDP (in 1996 dollars), multiplied by 100. Labor income, *labor*, is the quarterly change in the natural logarithm of labor income (per capita total personal income minus dividend income, deflated by the personal consumption expenditures deflator), both multiplied by 100. The unemployment rate, *urate*, is the seasonally-adjusted three-month average unemployment rate. The consumer price index, *cpiq*, is the consumer price index, and *cay* is the consumption-to-wealth ratio from Lettau and Ludvigson (2001). We collect all variables except *dp* and *cay* from the Federal Reserve Economic Data, *dp* is collected using the CRSP-Compustat merged file and *cay* is collected from Professor Lettau's website.

3.4. Measurement of accounting quality variables

To test hypothesis 1A, we measure accounting quality factors expected to be associated with the strength of the cointegration between aggregate prices and fundmanetals. Following Dichev and Tang (2008), we examine several accounting quality measures. *Match* measures the quality of the contemporaneous matching between a firm's revenues and expenses. For each firm, we measure *Match* as the Pearson correlation coefficient between revenues and expenses measured over the firm's prior five yearly accounting observations. From Compustat, we

measure revenues as net revenues, earnings as earnings before extraordinary items, and expenses as the difference between earnings and revenues. We deflate earnings, revenues, and expenses by the firm's average total assets over the prior five yearly observations. Because the immediate expensing of R&D investments might also result in poor matching, we also examine R&D directly. We measure R&D investments as R&D expenditures scaled by net sales. Like Dichev and Tang, we also examine broader measures of accounting volatility, *vol*, based on the standard deviation of reported earnings, revenues, and expenses measured over the prior five yearly observations.

3.5. Measurement of arbitrage costs

To test hypothesis 2, we measure arbitrage costs at the firm-level similar to the method described in Pontiff (1996) and aggregate these estimates. Specifically, we estimate daily market model regressions using the prior month of data for each of the individual stocks in the Dow 30. We then save the residuals and calculate the standard deviation of these residuals for each firm, to calculate idiosyncratic risk. We next value-weight the average arbitrage cost based on the share price at the end of the month to be consistent with our weighting procedure for the price and fundamental indices.

4. Empirical Analysis

4.1. Descriptive statistics

We provide graphical evidence of the time-series variation in the ratio of aggregate price to aggregate fundamentals using the V12(TBe) model in Figure 1. The ratio exhibits much higher volatility in the second half of the sample, including more extreme highs and lows. The

fluctuations in the ratio correspond roughly with the large increases in price in the late 1990s generally labeled the “internet bubble” and the subsequent “internet crash” or “market correction” in 2001, and the large price declines beginning in 2007 largely considered to be due to the “financial crisis.” The price-to-fundamental ratio does appear to drop during recessions, as identified by the NBER. The drop during the 2007-2008 recession, however, is unprecedented in the time-series, during this “financial crisis” period the Dow 30 appears to be significantly undervalued.

[Insert Figure 1 about here]

We present summary statistics in Panel A of Table 1. Specifically, we report the levels of the aggregate price index of the Dow 30 and the aggregate fundamentals models measured over three, 12 and 18 month horizons. The descriptive statistics reported in Panel A suggest that there is very little difference from increasing the forecast horizon from 12 to 18 months. We also report autocorrelations, and note that all of the first-order autocorrelations are high, implying that these variables are nonstationary in levels.

4.2. Static cointegration analysis

We first reconcile our results to prior research by estimating static cointegration tests of the periods 1979-2008, 1979-1993, and 1994-2008 following Curtis (2012) in Panel B of Table 1. Consistent with Curtis (2012), we find that evidence of cointegration between prices and fundamentals is not supported by the data for the full sample period 1979-2008 and the latter half of the time-series 1994-2008. In Panel B, however, we find evidence of cointegration in the earlier period, 1979–1993, which overlaps with Lee et al. (1999). These results suggest that there is significant variation in the strength of cointegration between prices and fundamental values

estimated using the residual income model. The lack of cointegration implies that the model of fundamental value is misspecified in recent periods or that in recent periods prices did not reflect fundamental values or both. This implies that investors conditioned their trades on information other than estimates of fundamental values, such as macro-economic news or possibly these price movements reflect speculation not removed by arbitrageurs.

[Insert Table 1 about here]

4.3. Descriptive statistics for dynamic cointegration tests

In Figure 2, we graph the strength of cointegration using the Phillips-Perron Z-Rho statistic based on monthly rolling window estimates of Equation (5) using 60 prior monthly observations, which tests for stationarity with a constant over each prior five-year period. We use data from 1979 which means our first estimate with 60 data points is December of 1983. Early in the sample, the Z-Rho statistic is negative, which suggests evidence of cointegration between price and fundamentals. In the lead up to the October 1987 crash the strength of cointegration weakens substantially until October 1987. The strength of cointegration then slowly increases reaching a statistically significant level around 1995, where it remains “strong” until mid way through 1996, at which point it weakens again. After this period, the strength of cointegration measure remains weak, and the estimates in the early 2000s, following the crash, exhibit “explosive behavior.” In this case, the Z-Rho coefficient is positive, suggesting that the changes in the price-to-fundamentals ratios moved further away from the level of the prior price-to-fundamental ratio, widening the gap between prices and fundamentals during the prior five-years.

[Insert Figure 2 about here]

In Figure 3, we graph the time-series of the one minus the p -value associated with Z-Rho in Figure 2. One minus the p -value of Z-Rho can be interpreted as the level of confidence that can be placed on the hypothesis that aggregate prices and fundamentals are cointegrated. For example, in the beginning of the time-series, the figure documents over 90% confidence that prices and fundamentals are comoving over the past five years. This level of confidence then declines to around 20% in the lead up to the October 1987 crash. Consistent with the findings in Curtis (2012), in Figure 3 we do not find any evidence of cointegration between price and fundamentals at conventional levels ($p < 0.10$) after 1996.

[Insert Figure 3 about here]

We report descriptive statistics for the data used in Figures 2 and 3 in Panel A of Table 2. From the 301 overlapping rolling window estimates the average Z-Rho is -6.149 with an average p -value of 0.443. These results suggest that the average strength of cointegration is low for the sample. The standard deviations, however, are high, for example, the standard deviation for the p -value is 0.337. The high standard deviations suggest that there is a significant amount of time-series variation in the strength of cointegration. The minimum and maximum p -values also suggest that there are large differences in the time-series, with strongly cointegrated periods (minimum p -value = 0.011) and with almost zero evidence of cointegration (maximum p -value = 0.999). For the time-series, 11.96% of the overlapping monthly windows have p -values less than 5%, with 19.93% (or 7.97% more) of the overlapping monthly observations having p -values less than 10%. In addition, 8.97% of the overlapping monthly observations have Z-Rho estimates greater than 0, implying explosive behavior during these periods. These estimates suggest that evidence of strong cointegration is rarer than would be expected from the results of static cointegration tests.

In Panel B, we report the descriptive statistics for the macro-economic factors we use in the study. All of the variables display fairly large ranges relative to their standard deviations. For example, the minimum Treasury yield is 2.08% and the maximum is 10.98% consistent with the significant changes in fiscal policy observed in the past three decades. Similarly the unemployment rate ranges from a low of 4.47% to a high of 8.38%. Given the large ranges for these variables observed in the time-series and the large ranges observed for the strength of cointegration, it is possible that some of these variables help explain the strength of cointegration.

[Insert Table 2 about here]

In Table 3, we report the correlations between the strength of cointegration variables (Z-Rho, ρ , and the p -value associated with Z-Rho, $p(\rho)$) and macro-economic variables. Both H1 and H2 predict a positive correlation between macro-economic factors and either ρ or $p(\rho)$ which is consistent with the factor being associated with a reduction of the strength of cointegration. We find that consumer confidence, ics (0.65), arbitrage costs, arb (0.45), inflation, cpi (0.74), and GDP growth, gdp (0.17), have significant positive associations with $p(\rho)$ and ρ . Alternatively, the factors that are negatively associated with ρ or $p(\rho)$ (i.e., factors that are associated with an increase in the strength of cointegration), include the default spread, def (-0.47), Treasury yield, yld (-0.65), dividend yield, dp (-0.77), consumption, $cons$ (-0.49), labor income, $labor$ (-0.50), the unemployment rate, $urate$ (-0.68), and the consumption-to-wealth ratio, cay (-0.47).

Consistent with prior literature, consumer confidence is highly correlated with the other macro-economic variables, including gdp (0.58). In addition, we also find a positive correlation between consumer sentiment and arbitrage costs of 0.65. Many of the macro-economic news

variables are highly correlated, with the absolute values of the correlations being above 0.80 in many cases. Consumer confidence is also highly correlated at 0.65, with arbitrage costs.

These results provide some preliminary support for our hypotheses. Specifically, in H1 we predict that variables that are associated with future consumption are expected to decrease the strength of cointegration, some of the variables above display this expectation including consumer confidence, GDP growth, unemployment rates, the consumption-to-wealth ratio, and inflation. The evidence is mixed, however, as labor income and current consumption are negatively associated with the strength of cointegration. In preliminary support of H2, arbitrage costs are negatively correlated with the strength of cointegration.

[Insert Table 3 about here]

4.4. Factors associated with the strength of cointegration

In Table 4, we examine the strength of cointegration with individual macro-economic and accounting quality variables using regression analysis. We estimate the regressions using the Newey-West correction for standard errors as our dependent variables overlap for 59 of the 60 monthly observations. The regressions are estimated as:

$$StrCI_t = a_0 + b_1(Factor_t) + e_t, \quad (11)$$

where, $StrCI_t$ is the strength of cointegration measures (ρ or $p(\rho)$), and $Factor_t$ refers to one of 11 macro-economic news variables or 5 earnings quality variables.

In Panel A, we report these results using Z-Rho as the dependent variable, and in Panel B, we report these results using the p-value associated with Z-Rho as the dependent variable. Consumer confidence is associated with a lower strength of cointegration. For example, in Panel B, we report an association of 2.695 with an associated t -statistic of 5.75 ($p < 0.01$), supporting

H1. Consistent with H2, we also find evidence that arbitrage costs are associated with a lower strength of cointegration. In Panel B, arbitrage costs have an association of 55.751 with an associated t -statistic of 2.16 ($p < 0.05$).

When controlling for a constant, there is no statistical evidence of an association between GDP growth and either Z-Rho or the p -value of Z-Rho, which questions whether the associations we find relating to consumer confidence are related to current levels of growth. Similarly, we find a negative association for current consumption and labor income, inconsistent with H1. Overall the results in Table 4 provide clear evidence in support of H2, but the results are mixed for H1.

In Table 4, we also examine preliminary evidence of the association between factors for accounting quality and the strength of cointegration. Prior research suggests that conceptual differences between fundamental valuation models are not as important as the ability to project a firm's future fundamentals. Dichev and Tang (2008) examine several accounting quality measures that they suggest are associated with the quality of projecting future fundamentals, including measures of the matching between revenues and expenses and the volatility of earnings components.

In Table 4, only our measures of for revenue-expense matching are strongly associated with the strength of cointegration. *match*, which measures the correlation between a firms' net revenues and expenses over the past five years, is positively associated with Z-Rho or the p -value of Z-Rho suggesting that better matching reduces the cointegration between market values and fundamentals. Similarly, R&D scaled by sales is also a proxy for a weak match between economic revenues and expenses and accounting revenues and expenses. The association

between R&D and Z-Rho or the p -value of Z-Rho is negative, also suggesting that better matching increases the strength of cointegration.

[Insert Table 4 about here]

In Table 5, we report multivariate analysis of the factors associated with the strength of cointegration. As many of the macro-economic variables are highly correlated, we consider a subset of the variables. Our regression model is a multivariate version of Equation (11), and again we estimate this correcting for the 59 overlapping observations in the dependent variables using the Newey-West method:

$$StrCI_t = a_0 + \sum_{k=1}^K b_k(Factor_{k,t}) + e_t, \quad (12)$$

In Panel A, we report these results using Z-Rho as the dependent variable, and in Panel B, we report these results using the p -value associated with Z-Rho as the dependent variable. In these models, we include arbitrage costs and macro-economic news variables. We find that arbitrage costs no longer continue to be significantly positively associated with the strength of cointegration variables when including consumer confidence (columns 1 and 2). In this model, consumer confidence subsumes the explanatory power of arbitrage costs, providing support for H1 but not for H2. Inconsistent with this result, we find that arbitrage costs remain a significant and positive explanatory variable when including other macro-economic news variables (columns 3 and 4). In these models cay and the $ylid$ continue to be significantly negatively associated with the strength of cointegration variables while current levels of consumption, $cons$, and $labor$ are not significant in these models. These results are consistent with H2 but not with H1.

[Insert Table 5 about here]

4.5. Accounting quality factors associated with the strength of cointegration

In Table 6, we report very preliminary multivariate analysis of the accounting quality factors associated with the strength of cointegration. We focus on aggregate-level measures of accounting quality following Dichev and Tang (2008). Our regression model is set up in the same way as Equation (12) and again we estimate this correcting for the 59 overlapping observations in the dependent variables using the Newey-West method. As in Table 5, in Panel A, we report these results using Z-Rho as the dependent variable, and in Panel B, we report these results using the p-value associated with Z-Rho as the dependent variable. In these models, we include arbitrage costs and consumer confidence variables.

We find that matching is significantly and positively associated with Z-Rho and the p-value of Z-Rho across all specifications. We find inconsistent results for the volatility of expenses, revenues and net income when we include or exclude arbitrage costs and the consumer sentiment index. Similarly, RD to Sales is not significant when we include arbitrage costs and the consumer confidence index. Including accounting quality measures in the tests reduces the explanatory power of the consumer confidence, suggesting that consumer confidence is highly correlated with accounting quality.

[Insert Table 6 about here]

4.6. Analysis of robustness to alternative measurement of fundamentals

In untabulated results we examine additional measures of fundamental value. We find that limiting the models to only include historical earnings information, i.e., excluding analysts' forecasts from the model, does not qualitatively change our results. We also investigate the possibility that alternative assumptions regarding the equity premium (which we held at a constant rate of 5%) impact our results. Again we find no qualitative changes in our results for

the following specifications: including a stock-specific beta times 5% premium, including an equity premium based on 12 to 60 month lagged market returns, including an equity premium based on forward looking (actual) market returns over the following 12 to 24 months, and similar industry-based prior returns as the discount rate.

4.7. Limitations and caveats

Our results must be interpreted with the important caveat that our main results are limited by our choices of how to measure firm-fundamentals. While the residual income model has strong theoretical and empirical support in prior literature (Ohlson, 1995; Frankel and Lee, 1998; Lee et al., 1999) the model is always implemented with error. As such, it is possible that our implementation of the residual income model excludes an important component of intrinsic value and our results should be interpreted with this caveat in mind. We also recognize that consumer confidence is a variable that could proxy for either sentiment or expected growth in consumption. In future drafts we aim to examine the association between consumer confidence and future consumption in order to assess whether the dominant reason for consumer confidence to be associated with a lower strength of cointegration is due to sentiment based explanations or due to the “rational” prediction of future consumption.

5. Conclusion

We investigate factors associated with the strength of cointegration between price and fundamentals. Our study is motivated by the importance of understanding the reasons why prices diverge from fundamentals. We propose a measure of the strength of cointegration between prices and fundamentals based on rolling window stationarity tests. We document significant

time-series variation in the strength of cointegration between prices and fundamentals, suggesting that prices reflect varying amounts of the information in accounting fundamentals at different points in time.

We then examine factors that we expect to be associated with the strength of cointegration between prices and fundamentals based on the mispricing and the measurement errors in fundamentals. We identify growth and arbitrage costs as reasons why prices should diverge from accounting-based measures of fundamentals and arbitrage costs as a reason that prices diverge from intrinsic values. Our main proxy for growth, consumer confidence, is reliably associated with a lower strength of cointegration. Our proxy for arbitrage costs is not reliably associated with the strength of comovement, being subsumed by consumer confidence in a multivariate setting. Our tests provide mixed evidence on other macro-economic news variables that are expected to provide information on growth.

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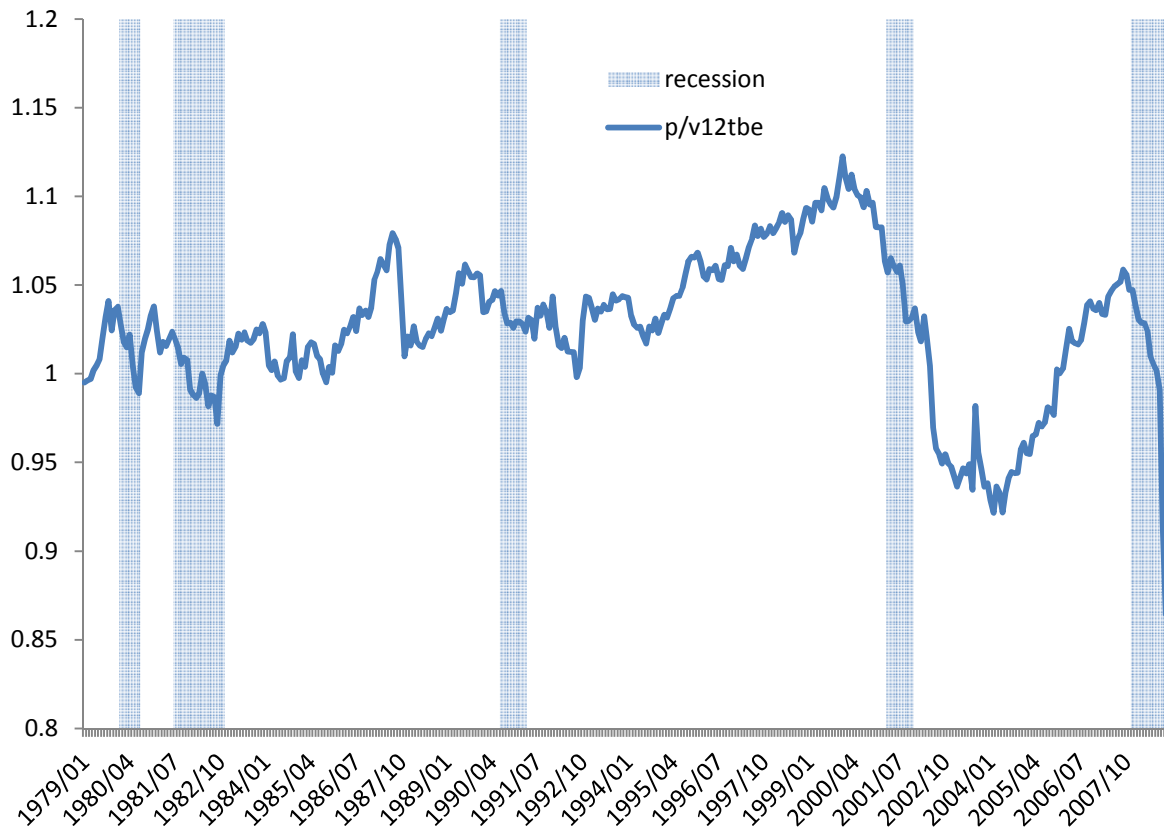


Figure 1. Price-to-fundamental ratio

This graph displays aggregate price and fundamentals where aggregate fundamentals are measured as the V12(TBe) model. The model includes 12 forecasts of earnings based on analysts' forecasts for the first two years and a fade rate to the industry median ROE for the remainder, plus a terminal value based on the industry median ROE. The discount rate, r , is the three-year constant yield to maturity T-bill rate plus an equity premium of 6%. Recession is from NBER.

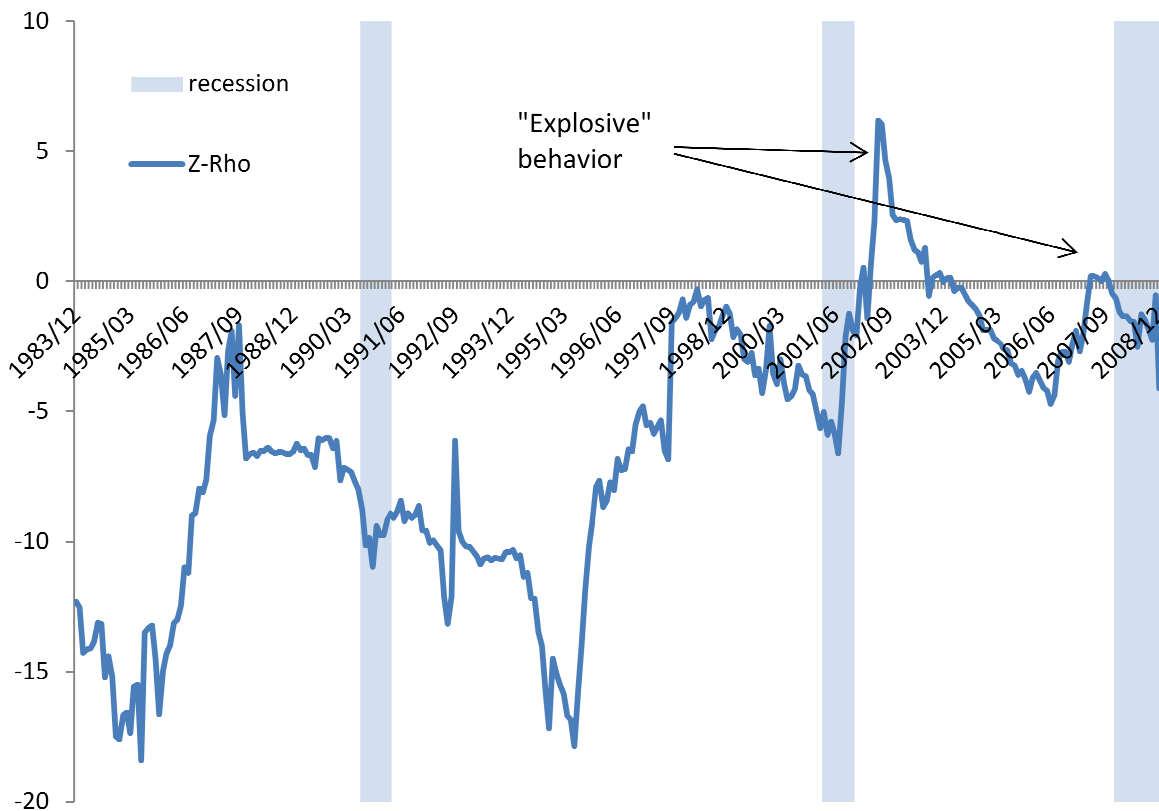


Figure 2. Strength of cointegration between prices and fundamentals using Z-Rho

This graph displays evidence of explosive behavior between aggregate price and fundamentals where aggregate fundamentals are measured as the V12(TBe) model. Explosive behavior is found when the estimate of Z-Rho is positive as it implies the opposite of mean-reversion. Z-Rho is the test statistic on the adjusted regression coefficient from the Phillips-Perron regressions which test for stationarity. The V12(TBe) model includes 12 forecasts of earnings based on analysts' forecasts for the first two years and a fade rate to the industry median ROE for the remainder, plus a terminal value based on the industry median ROE. The discount rate, r , is the three-year constant yield to maturity T-bill rate plus an equity premium of 6%. Recession is from NBER.



Figure 3. Strength of cointegration between prices and fundamentals using $p(Z-Rho)$

This graph displays the strength of cointegration between aggregate price and fundamentals where aggregate fundamentals are measured as the V12(TBe) model. The strength of cointegration is measured as 1 less the p-value associated with Z-Rho, where Z-Rho is the test statistic on the adjusted regression coefficient from the Phillips-Perron regressions which test for stationarity. The V12(TBe) model includes 12 forecasts of earnings based on analysts' forecasts for the first two years and a fade rate to the industry median ROE for the remainder, plus a terminal value based on the industry median ROE. The discount rate, r , is the three-year constant yield to maturity T-bill rate plus an equity premium of 6%. Recession is from NBER.

Table 1

Descriptive statistics and static cointegration tests

Panel A summarizes the descriptive statistics for the log of aggregate price and the log of aggregate fundamentals measured over three-month 12-month and 18-month horizons. Panel B summarizes the results of the Phillips-Perron unit root tests on the value-to-price ratios. The stationarity regression statistics presented in the columns with a constant is based on:

$$\Delta\epsilon_t = a_0 + (\rho - 1)\epsilon_{t-1} + e_t$$

The stationarity regression statistics presented in the columns with a constant and a trend is based on:

$$\Delta\epsilon_t = a_0 + \delta t + (\rho - 1)\epsilon_{t-1} + e_t$$

The test is against the null of a unit root in the time series ($\rho = 1$). The test statistics from the Phillips-Perron regressions are an adjusted regression coefficient Z-rho (Z_ρ) and an adjusted t -statistic Z-tau (Z_τ). ** $p < 0.001$, * $p < 0.05$, and $\delta p < 0.1$. T is the number of time-series observations.

<i>Panel A: Descriptive statistics for price and fundamentals (T=360)</i>									
Variable	Mean	Std. Dev	Autocorrelation at lag						
			1	12	24	36	48	60	
P	20.70	1.06	0.993	0.913	0.820	0.726	0.626	0.536	
V3(TBe)	21.28	1.35	0.981	0.839	0.732	0.639	0.570	0.473	
V12(TBe)	21.02	1.23	0.979	0.829	0.717	0.620	0.545	0.441	
V18(TBe)	21.06	1.23	0.979	0.832	0.724	0.629	0.553	0.449	
<i>Panel B: Static cointegration tests</i>									
Variable	Constant		Constant and trend		T				
	Z-rho (Z_ρ)	Z-tau (Z_τ)	Z-rho (Z_ρ)	Z-tau (Z_τ)					
Time period = 1979–2008.									
VP3(TB)	-10.1165	-1.89	-15.0155	-2.52	360				
VP12(TB)	-11.8774 ^δ	-2.11	-13.3426	-2.33	360				
VP18(TB)	-12.3298 ^δ	-2.15	-14.1950	-2.41	360				
Time period = 1979–1993.									
VP3(TB)	-27.5677**	-3.78**	-37.4750**	-4.54**	180				
VP12(TB)	-36.3634**	-4.48**	-36.5208**	-4.48**	180				
VP18(TB)	-38.7360**	-4.63**	-39.4492**	-4.68**	180				
Time period = 1994–2008.									
VP3(TB)	-3.0018	-0.91	-6.4710	-1.74	180				
VP12(TB)	-3.0716	-0.93	-7.1729	-1.88	180				
VP18(TB)	-3.1457	-0.94	-7.1674	-1.87	180				

Table 2**Descriptive statistics for the strength of cointegration and macro-economic factors**

This table summarizes the results of the rolling window and expanding window estimations of the Phillips-Perron unit root tests on the value-to-price ratios. In Panel A, each estimate includes 60 time-series observations. The stationarity regression statistics presented in the columns with a constant is based on:

$$\Delta \epsilon_t = a_0 + (\rho - 1)\epsilon_{t-1} + e_t$$

The test is against the null of a unit root in the time series ($\rho = 1$). The test statistics from the Phillips-Perron regressions are an adjusted regression coefficient Z-rho (Z_ρ) and an adjusted t -statistic Z-tau (Z_τ). % explosive is the proportion of estimates where the Z-rho is greater than 0, % $p < 0.05$ is the proportion of Z-rho with a p-value less than 0.05 and % $p < 0.1$ is the proportion of Z-rho with a p-value less than 0.10. In Panel B, the descriptive statistics for the rolling window averages are presented, each variable represents the average over the prior five-years. *ics* is the level of the Michigan Consumer Confidence index, *arb* is the value-weighted average arbitrage cost for the stocks in the Dow 30 measured over the prior month, *def* is the difference between the yield to maturity of Moody's Baa-rated and Aaa-rated bonds, *yld* is the yield from three-month Treasury Bill, *dp* is the total ordinary cash dividend from the CRSP value-weighted index over the last twelve months divided by the current month's index value default spread, *gdp* is the quarterly change in the natural log of GDP (in 1996 dollars), multiplied by 100, *labor* is the quarterly change in the natural logarithm of labor income (per capita total personal income minus dividend income, deflated by the personal consumption expenditures deflator), both multiplied by 100, *urate* is the seasonally adjusted three-month average unemployment rate, *cpiq* is the consumer price index, and *cay* is the consumption-to-wealth ratio from Lettau and Ludvigson (2001).

Panel A: Dynamic cointegration tests using rolling window estimates (T=301)

Variable	Z-rho (Z_ρ)	p-value
Average	-6.149	0.443
Standard deviation	5.157	0.337
Minimum	-18.4142	0.011
Maximum	6.1812	0.999
% explosive	8.97%	
% $p < 0.05$	11.96%	
% $p < 0.1$	19.93%	

Panel B: Rolling window independent variables (T=301)

	Mean	Std. Dev.	Min	Max
<i>ics</i>	90.34	6.98	71.30	103.09
<i>ln ics</i>	4.50	0.08	4.26	4.63
<i>arb</i>	0.02	0.00	0.01	0.02
<i>def</i>	1.04	0.34	0.64	1.81
<i>yld</i>	5.54	2.36	2.08	10.98
<i>dp</i>	2.94	1.17	1.37	5.23
<i>gdp</i>	0.77	0.17	0.40	1.20
<i>cons</i>	1.58	0.29	1.11	2.34
<i>labor</i>	1.46	0.32	0.95	2.37
<i>urate</i>	6.07	1.15	4.47	8.38
<i>cpi</i>	4.33	0.95	2.64	6.04
<i>cay</i>	0.01	0.01	-0.02	0.03

Table 3
Correlations between the strength of cointegration and macro-economic factors

The Pearson correlations for the rolling window averages are presented, each variable represents the average over the prior five-years. *ICS* is the level of the Michigan Consumer Confidence index, *arb* is the value-weighted average arbitrage cost for the stocks in the Dow 30 measured over the prior month, *def* is the difference between the yield to maturity of Moody's Baa-rated and Aaa-rated bonds, *yld* is the yield from three-month Treasury Bill, *dp* is the total ordinary cash dividend from the CRSP value-weighted index over the last twelve months divided by the current month's index value default spread, *gdp* is the quarterly change in the natural log of GDP (in 1996 dollars), multiplied by 100, *labor* is the quarterly change in the natural logarithm of labor income (per capita total personal income minus dividend income, deflated by the personal consumption expenditures deflator), both multiplied by 100, *urate* is the seasonally adjusted three-month average unemployment rate, *cpi* is the consumer price index, and *cay* is the consumption-to-wealth ratio from Lettau and Ludvigson (2001).

	<i>p(rho)</i>	<i>rho</i>	<i>ics</i>	<i>ln ics</i>	<i>arb</i>	<i>ln arb</i>	<i>def</i>	<i>yld</i>	<i>dp</i>	<i>gdp</i>	<i>cons</i>	<i>labor</i>	<i>urate</i>	<i>cpi</i>	<i>cay</i>
<i>p(rho)</i>	1.00														
<i>rho</i>	0.95	1.00													
<i>ics</i>	0.65	0.72	1.00												
<i>ln ics</i>	0.65	0.73	1.00	1.00											
<i>arb</i>	0.46	0.45	0.65	0.61	1.00										
<i>ln arb</i>	0.40	0.39	0.62	0.57	0.99	1.00									
<i>def</i>	-0.47	-0.46	-0.53	-0.55	-0.31	-0.29	1.00								
<i>yld</i>	-0.65	-0.62	-0.45	-0.48	-0.19	-0.12	0.81	1.00							
<i>dp</i>	-0.77	-0.74	-0.65	-0.65	-0.51	-0.44	0.79	0.92	1.00						
<i>gdp</i>	0.17	0.28	0.58	0.58	0.21	0.26	-0.02	0.19	0.09	1.00					
<i>cons</i>	-0.49	-0.45	-0.35	-0.38	-0.17	-0.12	0.88	0.94	0.85	0.30	1.00				
<i>labor</i>	-0.50	-0.47	-0.32	-0.36	-0.15	-0.09	0.74	0.94	0.81	0.32	0.94	1.00			
<i>urate</i>	-0.68	-0.68	-0.71	-0.72	-0.50	-0.42	0.82	0.81	0.92	0.07	0.79	0.70	1.00		
<i>cpi</i>	0.74	0.69	0.46	0.48	0.29	0.21	-0.72	-0.94	-0.96	-0.21	-0.85	-0.81	-0.85	1.00	
<i>cay</i>	-0.47	-0.45	-0.15	-0.14	-0.13	-0.04	-0.31	0.17	0.28	0.11	-0.10	0.01	0.18	-0.39	1.00

Table 4

Macro-economic factors associated with the strength of cointegration

This table summarizes the results of the regression estimates of:

$$StrCI_t = a_0 + b_1(Factor_t) + e_t$$

Standard errors are estimated using the Newey-West correction for the standard errors as our dependent variables overlap for 59 observations. $StrCI_t$ is the strength of cointegration measures (ρ or $p(\rho)$), and $Factor_t$ refers to the accounting quality variables and macro-economic news variables. ics is the level of the Michigan Consumer Confidence index. arb is the value-weighted average arbitrage cost for the stocks in the Dow 30 measured over the prior month. def is the difference between the yield to maturity of Moody's Baa-rated and Aaa-rated bonds. yld is the yield from three-month Treasury Bill. dp is the total ordinary cash dividend from the CRSP value-weighted index over the last twelve months divided by the current month's index value default spread. gdp is the quarterly change in the natural log of GDP (in 1996 dollars) multiplied by 100. $labor$ is the quarterly change in the natural logarithm of labor income (per capita total personal income minus dividend income, deflated by the personal consumption expenditures deflator), both multiplied by 100. $urate$ is the seasonally adjusted three-month average unemployment rate. $cpiq$ is the consumer price index. cay is the consumption-to-wealth ratio from Lettau and Ludvigson (2001). $Match$ measures the quality of the contemporaneous matching between a firm's revenues and expenses over the firm's prior five yearly accounting observations. The volatility of net revenues $vol(revenue)$, earnings before extraordinary items $vol(earnings)$, and the difference between earnings and revenues $vol(expenses)$, is the standard deviation of the prior five yearly observations. $R\&D/Sales$ is aggregate R&D expenditures scaled by net sales.

<i>Panel A: Dependent variable = Z-Rho (ρ)</i>				
<i>Independent Variable</i>	a	t-statistic	b	t-statistic
<i>ln ics</i>	-208.09	-6.27	44.94	6.13
<i>arb</i>	-19.67	-2.90	813.76	2.34
<i>def</i>	1.14	0.34	-6.90	-3.48
<i>yld</i>	1.27	0.70	-1.32	-7.53
<i>dp</i>	3.28	3.05	-3.17	-6.33
<i>gdp</i>	-12.35	-1.96	8.16	1.24
<i>cons</i>	6.17	1.03	-7.74	-2.57
<i>labor</i>	4.54	1.01	-7.23	-3.18
<i>urate</i>	11.86	3.10	-2.95	-4.49
<i>cpi</i>	-21.76	-6.60	3.63	5.43
<i>cay</i>	-5.10	-3.58	-166.10	-2.11
<i>match</i>	-58.60	-2.64	57.75	2.42
<i>vol(revenue)</i>	-9.87	-1.33	27.35	0.66
<i>vol(expenses)</i>	-9.37	-1.31	25.68	0.61
<i>vol(earnings)</i>	-5.96	-1.72	11.15	0.14
<i>R&D/sales</i>	-27.97	-2.08	1,139.05	1.78
<i>Panel B: Dependent variable = p-value of Z-Rho ($p(\rho)$)</i>				
<i>Independent Variable</i>	a	t-statistic	b	t-statistic
<i>ln ics</i>	-11.67	-5.54	2.70	5.75
<i>arb</i>	-0.49	-0.98	55.75	2.16
<i>def</i>	0.94	4.44	-0.47	-4.07
<i>yld</i>	0.96	7.60	-0.09	-6.22
<i>dp</i>	1.10	14.24	-0.22	-6.60
<i>gdp</i>	0.19	0.45	0.33	0.75
<i>cons</i>	1.35	3.57	-0.57	-3.11
<i>labor</i>	1.21	4.13	-0.52	-3.61
<i>urate</i>	1.65	5.93	-0.20	-4.40
<i>cpi</i>	-0.69	-3.59	0.26	6.66
<i>cay</i>	0.52	5.25	-11.72	-2.21
<i>match</i>	-2.20	-1.65	2.91	2.03
<i>vol(revenue)</i>	0.40	0.86	0.55	0.83
<i>vol(expenses)</i>	0.44	0.98	0.35	0.13
<i>vol(earnings)</i>	0.47	1.97	0.34	0.06
<i>R&D/sales</i>	0.515	5.25	-11.716	-2.21

Table 5
Multivariate analysis of macro-economic factors associated with the strength of cointegration

This table summarizes the results of the regression estimates of:

$$StrCI_t = a_0 + \sum_{k=1}^K b_k(Factor_{k,t}) + e_t$$

Standard errors are estimated using the Newey-West correction for the standard errors as our dependent variables overlap for 59 observations. $StrCI_t$ is the strength of cointegration measures (ρ or $p(\rho)$), and $Factor_t$ refers to the 11 macro-economic news variables as follows: $ln\ ics$ is the natural log of the level of the Michigan Consumer Confidence index, $ln\ arb$ is the natural log of the value-weighted average arbitrage cost for the stocks in the Dow 30 measured over the prior month, where arbitrage cost is the standard deviation of the residuals from the daily market model over the prior month, def is the difference between the yield to maturity of Moody's Baa-rated and Aaa-rated bonds, yld is the yield from three-month Treasury Bill, dp is the total ordinary cash dividend from the CRSP value-weighted index over the last twelve months divided by the current month's index value default spread, gdp is the quarterly change in the natural log of GDP (in 1996 dollars), multiplied by 100, $labor$ is the quarterly change in the natural logarithm of labor income (per capita total personal income minus dividend income, deflated by the personal consumption expenditures deflator), both multiplied by 100, $urate$ is the seasonally adjusted three-month average unemployment rate, $cpiq$ is the consumer price index, and cay is the consumption-to-wealth ratio from Lettau and Ludvigson (2001).

<i>Panel A: Dependent variable = Z-Rho (ρ)</i>				
	coefficient	t-statistic	coefficient	t-statistic
<i>Intercept</i>	-221.149	-3.17	28.860	2.41
<i>ln ics</i>	46.483	4.69		
<i>ln arb</i>	-1.488	-0.19	10.798	4.47
<i>cay</i>			-37.682	-0.76
<i>yld</i>			-3.424	-4.17
<i>cons</i>			12.580	1.44
<i>labor</i>			6.099	0.97
<i>Panel B: Dependent variable = p-value of Z-Rho ($p(\rho)$)</i>				
	b	t-statistic	b	t-statistic
<i>Intercept</i>	-10.951	-1.98	3.326	4.24
<i>ln ics</i>	2.610	3.63		
<i>ln arb</i>	0.0816	0.13	0.721	3.99
<i>cay</i>			-6.169	-2.26
<i>yld</i>			-0.167	-3.31
<i>cons</i>			0.263	0.48
<i>labor</i>			0.436	0.93

Table 6
Multivariate analysis of macro-economic factors associated with the strength of cointegration

This table summarizes the results of the regression estimates of:

$$StrCI_t = a_0 + \sum_{k=1}^K b_k(Factor_{k,t}) + e_t$$

Standard errors are estimated using the Newey-West correction for the standard errors as our dependent variables overlap for 59 observations. $StrCI_t$ is the strength of cointegration measures (ρ or $p(\rho)$), and $Factor_t$ refers to the accounting quality variables and macro-economic news variables as follows: $\ln ics$ is the natural log of the level of the Michigan Consumer Confidence index, $\ln arb$ is the natural log of the value-weighted average arbitrage cost for the stocks in the Dow 30 measured over the prior month, where arbitrage cost is the standard deviation of the residuals from the daily market model over the prior month. $Match$ measures the quality of the contemporaneous matching between a firm's revenues and expenses, over the firm's prior five yearly accounting observations. We measure the volatility of net revenues $vol(revenue)$, earnings before extraordinary items $vol(earnings)$, and the difference between earnings and revenues $vol(expenses)$, using the prior five yearly observations. $R\&D/Sales$ is aggregate R&D expenditures scaled by net sales.

<i>Panel A: Dependent variable = Z-Rho (ρ)</i>				
	coefficient	t-statistic	coefficient	t-statistic
<i>Intercept</i>	266.16	7.06	-83.30	-1.04
<i>match</i>	385.26	5.52	235.49	6.69
<i>vol(revenue)</i>	-660.04	-6.94	206.70	2.89
<i>vol(expenses)</i>	609.41	3.76	-475.31	-6.25
<i>vol(earnings)</i>	397.09	0.97	505.81	3.95
<i>R&D/Sales</i>	-242.86	-8.25	118.14	0.13
<i>ln ics</i>			-14.32	-0.86
<i>ln arb</i>			13.87	2.04
<i>Panel B: Dependent variable = p-value of Z-Rho ($p(\rho)$)</i>				
	b	t-statistic	b	t-statistic
<i>Intercept</i>	18.50	9.39	16.08	4.87
<i>match</i>	25.91	6.83	17.11	2.31
<i>vol(revenue)</i>	-46.90	-10.85	-37.78	-4.35
<i>vol(expenses)</i>	44.62	5.57	38.73	3.90
<i>vol(earnings)</i>	21.48	0.90	-10.44	-0.17
<i>R&D/Sales</i>	-15.68	-10.20	-0.31	-0.30
<i>ln ics</i>			0.74	1.20
<i>ln arb</i>			-8.17	-1.28