An empirical evaluation of the time-series relation between price and accounting based value in imperfect markets

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I examine the extent to which accounting information is reflected in market prices at different points in time. The efficient market hypothesis implies that price always reflects (value-relevant) accounting information, based on the assumptions of rational investors and costless arbitrage. I examine the time-series relation between price and value in two studies which are motivated by potential shortcomings of these assumptions. First, there is significant debate regarding the rationality of equity investors during the late 1990s. I therefore contrast the historical time-series relation between price and value with that of the 1990s, and show that the historical tendency of price to converge towards value breaks down during this period. Second, I examine the impact of the lack of close substitutes – an arbitrage cost – on the time-series relation between price and value. I find some evidence of a positive association between this arbitrage cost and both the level and the duration of the disparity between price and value. My results provide empirical support for the hypothesis that price requires time to reflect (accounting) information and has implications for research that assumes that prices are measured without error.

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An Empirical Evaluation of the Time-Series Relation Between Price and Accounting based Value in Imperfect Markets

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2007

A Thesis Submitted As Partial Completion Of The Award Of Doctor Of Philosophy In Accounting At The University Of New South Wales, Sydney Australia.

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Abstract

I examine the extent to which accounting information is reflected in market prices at different points in time. The efficient market hypothesis implies that price always reflects (value-relevant) accounting information, based on the assumptions of rational investors and costless arbitrage. I examine the time-series relation between price and value in two studies which are motivated by potential shortcomings of these assumptions. First, there is significant debate regarding the rationality of equity investors during the late 1990s. I therefore contrast the historical time-series relation between price and value with that of the 1990s, and show that the historical tendency of price to converge towards value breaks down during this period. Second, I examine the impact of the lack of close substitutes – an arbitrage cost – on the time-series relation between price and value. I find some evidence of a positive association between this arbitrage cost and both the level and the duration of the disparity between price and value. My results provide empirical support for the hypothesis that price requires time to reflect (accounting) information and has implications for research that assumes that prices are measured without error.

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Chapter 4 contains some material that was included in my paper presented under the title "Costly arbitrage and the convergence of price to accounting fundamentals." I would like to acknowledge the helpful comments received from participants at the Stern School at New York University, and from my new colleagues at the University of Utah, especially Christine Botosan, Rachel Hayes, Matt Magilke and Marlene Plumlee.

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AN EMPIRICAL EVALUATION OF THE TIME-SERIES RELATION BETWEEN PRICE AND ACCOUNTING BASED VALUE IN IMPERFECT MARKETS

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1. Introduction

I examine the extent to which accounting information is reflected in market prices at different points in time, using accounting based value as a summary measure of value-relevant accounting information. The Efficient Market Hypothesis predicts that price reflects all available value-relevant accounting information, based on arbitrageurs being unimpeded.[1] The correcting actions of arbitrageurs therefore imply price always reflects value-relevant accounting information, making the task of fundamental valuation superfluous. When there are limits to arbitrage activity,[2] however, price is a noisy measure of intrinsic value as mispricing will take time to be corrected. Where this is the case, accounting based measures of value potentially become informative about mispricing (e.g., Lee, Myers and Swaminathan, 1999).

In a perfectly efficient market, sufficiently many arbitrageurs ensure that price reflects value. At the other end of the continuum, in a perfectly inefficient market, the lack of arbitrage activity implies price is arbitrary and does not reflect value. The equity market is neither perfectly efficient nor perfectly inefficient. In both scenarios, however, arbitrage activity is crucial to the extent to which value-relevant accounting information (hereafter, value) is expected to be reflected in price. The conditions that make the market more or less efficient are

[1] Markets are described as informationally efficient if the market incorporates all available information (see e.g., Fama 1970, 1976), and as such the term arbitrageur refers generally to sophisticated investors that are considered capable of profiting from the correction of informational inefficiencies (e.g., Lee, 2001). This definition of equity market arbitrageurs is different to the definition of pure arbitrage which is has zero risk and zero cost with a certain gain.

[2] See Shleifer and Vishny (1997) and Chapter 2 for a discussion of the limits to arbitrage.

paramount to understanding the functioning of the equity market, and therefore the extent to which accounting information is expected to be reflected in price.

Extending this logic, I investigate two research questions. First, I investigate whether the run-up in prices observed in the late 1990s is consistent with a speculative bubble, or was accompanied by an increase in value. This question is important, as the rationality of investors during this period is debatable and there is evidence that arbitrageurs were either not active or not trading against mispricing during the late 1990s market.[3] Second, I investigate whether the lack of close substitutes, which can be considered as a cost to arbitrage, helps explain variation in the run-up in prices relative to their values during this period. This question is important as arbitrageurs are expected to provide the mechanism that makes a relatively inefficient market a relatively efficient market; potential impediments to their activities are therefore of interest. These questions are addressed in Chapter 3 and Chapter 4, respectively.

Prior research, which I discuss in Chapter 2, suggests that price and value should be cointegrated. In Chapter 3, I empirically examine the time-series relation between price and value over the period 1979–2004. I contrast the historical relation, defined as the first half of the sample period (1979–1991), with the more recent period which potentially contained a speculative bubble (1992–2004). I find evidence of cointegration for the historical period, but not for the more recent period. I show that the breakdown in cointegration is due to a change in the tendency of price to 'anchor' to value.

Prior research also suggests that attempts to arbitrage equity mispricing are not without cost and risk. One arbitrage cost highlighted in prior research is the lack of close substitutes. In Chapter 4, I contrast the time-series relation between price and value for high and low arbitrage

[3] Taulli (2004) discusses anecdotal evidence that fundamental analysis based short-selling specialist funds exited the market in the late 1990s because they viewed the market as too bullish, and Brunnermeier and Nagal (2004) present evidence consistent with hedge funds tilting their portfolios towards overpriced speculative stocks in the 1990s.

cost firms. High arbitrage cost firms tend to have run-ups of price relative to value, both in the historical period (1979–1991) and during the speculative period (1992–2004). Consistent with higher arbitrage costs delaying the actions of arbitrageurs, deviations of price from value take a longer time to correct when the firm has a high level of arbitrage cost relative to a low level of arbitrage cost.

In both of these studies I acknowledge that measurement errors in the estimation of value also contribute to the disparity between price and value and potentially to their dynamic relation. First, using accounting information to measure intrinsic value requires an appropriate discount rate and appropriate assumptions for growth over a continuing horizon (Penman, 1998; Lundholm and O'Keefe, 2001). Second, accounting rules relating to the measurement tend to be conservative (e.g., Basu, 1997; Beaver and Ryan, 2005). It is likely that any limited arbitrage based explanation will be, at least in part, correlated with a measurement error explanation. The firms which are the hardest to value will also likely be the hardest to arbitrage. I therefore investigate measurement error based explanations, but find only limited support for these alternative explanations.

Taken together the results challenge the risk based explanations for two recent themes in the literature. First, in an informationally efficient market, the existence of a financial market bubble is ruled out through the rationality of investors. In contrast, I find that the cointegration between price and value breaks down in the late 1990s. I do not find compelling evidence to support risk based explanations for the dramatic run-up in prices during this period; instead, I find support for limited arbitrage based explanations. Second, in an informationally efficient market, the disparity between price and value should solely be a function of measurement errors.

In contrast, I find that the level of the disparity between price and value increases with a measure of the difficulty to arbitrage that firm.

The remainder of this thesis is structured as follows. In Chapter 2, I review the related literature. In Chapter 3, I contrast the time-series relation between price and value in the late 1990s with the historical relation. In Chapter 4, I investigate whether costs to arbitrage lead to larger and more prolonged deviations of price from value. Chapters 3 and 4 make up the majority of the thesis and can be read independently without loss of continuity. In Chapter 5, I present additional sensitivity analyses. The conclusion is contained in Chapter 6.

2. Literature review

In this chapter, I provide a review of related literature. I draw on important results in accounting, finance and econometrics research, with a focus on two themes in the literature: (i) the relation between price and accounting based value, and (ii) the limits to arbitrage.

2.1. *Fundamental valuation*

2.1.1. *Notation*

Throughout the remainder of this chapter, the following notational conventions will be used:

V_t^* = the firm-level (unobservable) intrinsic value of equity at time t,

P_t = the firm-level market value of equity at time t,

D_t = the firm-level dividends for the period t–1 to t,

X_t = the firm-level earnings for the period t–1 to t,

B_t = the firm-level book-value of equity at time t,

Θ_t = the information set available to all market participants at time t,

$\rho = (1 + re)$ = the discount rate, with re = the cost of equity capital.

In addition, the natural log of these variables (except for the discount rate) are expressed in the lower case, e.g., p_t = log(P_t), tildes (e.g., $\tilde{x}_{t+1}$) denote stochastic variables which are unobservable at time t but become observed at time t+n, asterisk denotes a variable that is never observed, e.g., v_{it}^*.

2.1.2. *Intrinsic value*

Most financial economists agree that the intrinsic value of the firm is equal to the present value of all (appropriately discounted) future dividends. The intrinsic value of a company can, be expressed in the form of a 'dividend discount' or 'present value of expected dividends' (PVED) model, which is generally written as:

$$V_t^* = \sum_{\tau=1}^{\infty} \rho^{-\tau} E_t\left(\tilde{D}_{t+\tau} \mid \Theta_t\right), \tag{2.1}$$

where, E_t refers to the expectations operator at time t and is conditional on the information set, Θ_t, at time t. To simplify the notation, this conditional expectations operator (i.e., $E_t(\bullet \mid \Theta_t)$) will be denoted simply as $E(\bullet)$.

Accounting information can be combined to form alternative accounting based valuation models using a method recently proposed by Ohlson (2005) and Ohlson and Juettner-Nauroth (2005).[4] Specifically, consider the sequence of numbers that satisfies the following condition:

$$0 = y_0 + \rho^{-1}(y_1 - \rho y_0) + \rho^{-2}(y_2 - \rho y_1) + \ldots + \rho^{-T}(y_T - \rho y_{T-1}), \tag{2.2}$$

where the sequence of y (the sequence of redundant terms) can be any sequence of variables that grows over time at a rate less than ρ such that $\rho^{-T} y_T \rightarrow 0$ as $T \rightarrow \infty$ (i.e., the transversality condition). Substituting Equation (2.2) into the discounted dividend model of Equation (2.1) yields:

$$V_t^* = y_0 + \sum_{t=1}^{\infty} R^{-t} z_t, \tag{2.3}$$

where, $z_t \equiv y_t + D_t - R y_{t-1}$. Equation (2.3) allows for alternative expressions for the present value of future expected dividends, which can be implemented using other accounting variables.

[4] Accounting based valuation formulae based on earnings and book-values have existed for a considerable time (Ohlson, 1990).

One of the most popular alternative accounting based valuation models is the residual income model. While popularised by Ohlson (1995), the residual income model has a long history,[5] and is based upon the reasoning that, in order to create wealth, a firm must earn more than its cost of capital. The residual income model uses book-value as the sequence of redundant terms and eliminates dividends from the present value equation through the clean-surplus relation: that all changes in book-value (i.e., net assets) are reflected in earnings and dividends. The clean-surplus relation is written as:

$$B_{t-1} = B_t + D_t - X_t \,. \tag{2.4}$$

Using these conditions, the residual income model is formed as the sum of book-value and discounted expected future abnormal returns. Specifically:

$$V_t^* = B_t + \sum_{\tau=1}^{\infty} \rho^{-\tau} E\left(X_{t+\tau}^a\right), \tag{2.5}$$

where B refers to the book-value of equity and X^a refers to abnormal earnings, which is defined as earnings less the required rate of return on book-value of equity:

$$X_t^a \equiv X_t - (\rho - 1) B_{t-1} \,. \tag{2.6}$$

While the residual income model and the dividend discount model are analytically equivalent, and with ideal implementation they should yield the same result, the residual income model has been found to be more highly correlated with both contemporaneous prices and subsequent market returns than alternative valuation models.[6]

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[5] Early research into residual income was scattered over many journals and books for around a century (see Peasnell, 1982). Relatively recent examples include Edwards and Bell (1961), who discussed residual income as excess realizable profit. Prior studies have also recommended residual income for internal measurement of business-segment performance (Solomons, 1965). Peasnell (1982) presents a reconciliation of the early literature's concept of residual income with accounting profits using only the assumption that the clean-surplus relation holds.

[6] A number of studies examine the association between market prices and alternative fundamental valuation models (Penman and Sougiannis, 1998; Francis et al., 2001). Summarising this literature, the residual income model performs at least as well as alternative measures of intrinsic value in measuring contemporaneous market prices. Whilst there is significant debate regarding the usefulness of these studies (Lundholm and O'Keefe, 2001a; 2001b;

Recently, Ohlson and Juettner-Nauroth (2005) developed a model based on forecast earnings and growth in those earnings.[7] The model is termed the abnormal earnings growth based model, and is based on the substitution of capitalized forward earnings into the present value equation in (2.1). Specifically:

$$V_t^* = \frac{X_{t+1}}{re} + \sum_{t=1}^{\infty} R^{-t} z_t \,, \tag{2.7}$$

where $z_t = \frac{1}{re}\left[X_{t+1} + reD_t - RX_t\right]$. It is clear that using this technology can result in numerous alternative inputs for accounting based valuation. In his discussion, however, of Ohlson (2005) and Ohlson and Juettner-Nauroth (2005), Penman (2005, p372) argues that one of the strong features of the residual income approach is the link to accounting theory through the clean-surplus requirement.

2.1.3. Linking intrinsic value to price

In this section, I draw on the literature that reconciles intrinsic value with price. In the prior section, accounting based measures of intrinsic value are formed as the present value using an appropriate discount rate. This discount rate can be expressed in terms of the expected return for holding the security:

$$\rho^{-\tau} = \prod_{n=1}^{\tau} \left[1 + E\left(\tilde{R}_{t+n}\right)\right]^{-1} , \tag{2.8}$$

where R_{t+1} refers to the security return for the period t to t+1, and is equivalent to the net simple return to holding a stock (including the payment of dividends) over the single period t to t+1. Specifically:

Penman 2001), the literature establishes an empirical basis for the residual income model as a measure of intrinsic value.

[7] For a detailed discussion of this model see Ohlson and Gao (2006).

$$E\left(\tilde{R}_{t+1}\right)=\frac{E\left(\tilde{P}_{t+1}+\tilde{D}_{t+1}\right)}{P_t}-1, \tag{2.9}$$

where the prices at t and t+1 are the ex-dividend prices at times t and t+1, respectively, which means that the return is the cum-dividend return from t to t+1. The linear present value relation can be derived under the assumption that the expected stock return is constant over time, or that $E_t(R_{t+n}) = R$, $\forall n$. It is then possible to express the dividend discount model using the following constant expected return dividend discount model (see, Campbell, Lo and MacKinlay, 1997, p255):

$$V_t^* = E\left[\sum_{i=1}^{K}\left(\frac{1}{1+R}\right)^i \tilde{D}_{t+i}\right]+E\left[\left(\frac{1}{1+R}\right)^K \tilde{P}_{t+K}\right]. \tag{2.10}$$

Equation (2.10) equates value to the present value of the sum of K future dividends plus a terminal value. If the forecast horizon K is sufficiently large, then the present value of the terminal price $\tilde{P}_{t+k}$ will tend towards zero (the transversality condition). Specifically:

$$\lim_{K\to\infty} E\left[\left(\frac{1}{1+R}\right)^K \tilde{P}_{t+K}\right]=0. \tag{2.11}$$

This condition is sufficient to rule out the existence of asset price bubbles. In particular it allows for the constant expected returns model to be expressed as a cointegration equation.[8] Specifically, the difference between price and the capitalised current (i.e., time t to t+1) dividend can be expressed as:

$$P_t-\frac{D_t}{R}=\left(\frac{1}{R}\right)E_t\left[\sum_{i=1}^{K}\left(\frac{1}{1+R}\right)^i \Delta\tilde{D}_{t+1+i}\right], \tag{2.12}$$

[8] Cointegration refers to the statistical property that the linear combination of two non-stationary variables is stationary. That is, two stochastic time-series that are non-stationary can deviate from their long-run relation for a short period of time, but the time-series will have a tendency to correct for the deviation.

where, in this model, the difference between the stock price and the capitalised current dividend equals the capitalised discounted value of future dividend changes. This model is important because while price and dividends can both be nonstationary individually, a linear combination of observable price at time t and a perpetuity of the observable current dividend for time t to $t+1$, discounted at the rate R should be stationary. This implies that price and dividends should be cointegrated with the coefficient $1/R$ as long as the expected rate of return is constant and the transversality condition in Equation (2.11) holds (i.e., dividends grow at a rate less than R). Following these assumptions, the model can also be written as: $P_t = D_t / (R - G)$ for any constant rate of growth in dividends, G (Gordon, 1962). A similar result can be shown for price and earnings when dividends are expressed in terms of the product of the payout ratio and earnings (Ohlson 1983).

The key result from this early literature, for my purposes, is that the relation between observed ex-dividend price at time t and dividends paid an instant before time t, should be one of cointegration. In addition, recent work in accounting based valuation has shown that dividends can be substituted by other accounting information, such as earnings and book-values. In the following section, I outline an important extension to this literature due to Campbell and Shiller (1988 a, b).

2.1.4. The log dividend-to-price ratio

To generalise from the situation where expected rates of return are constant, Campbell and Shiller (1988 a, b) suggest a log-linear relation between price, dividends and expected returns. This generalisation allows for expected returns to be time-varying, allowing for the more general situation where high prices must eventually be followed by either: (i) high dividends, (ii) low returns, or (iii) a combination of the two (Campbell, 1991; Campbell et al., 1997). Specifically, let log returns equal:

$$r_{t+1} = \log(P_{t+1} + D_{t+1}) - \log(P_t), \tag{2.13}$$

Then it follows that:

$$r_{t+1} = \log(P_{t+1}) - \log(P_t) + \log(1 + \exp(\log(D_{t+1}) - \log(P_{t+1}))), \tag{2.14}$$

where the expression in Equation (2.14) includes the non-linear function of the dividend-to-price ratio, $\log(1 + \exp(\log(D_{t+1}) - \log(P_{t+1})))$, which Campbell and Shiller (1988a, b) approximate using a first order Taylor expansion. Specifically:

$$r_t \approx k + \lambda \log(P_{t+1}) + (1-\lambda)\log(D_{t+1}) - \log(P_t), \tag{2.15}$$

where $k \equiv -\log(\lambda) - (1-\lambda)\log(1/(\lambda - 1))$, $\lambda \equiv 1/\left(1 + \exp\left(\overline{d-p}\right)\right)$ and $\left(\overline{d-p}\right)$ is the average log dividend-to-price ratio.

The approximation for log returns in Equation (2.15) can be used to derive a relation between the log price and log dividend. This allows for a more general version of Equation (2.12) which required the assumption of constant expected returns. Imposing the condition that:

$$\lim_{j \to \infty} \lambda^j p_{t+j} = 0, \tag{2.16}$$

where $p_t = \log(P_t)$ then:

$$p_t = \frac{k}{1-\lambda} + E_t\left[\sum_{j=0}^{\infty} \lambda^j \left[(1-\lambda)d_{t+1+j} - r_{t+1+j}\right]\right], \tag{2.17}$$

where $d_t = \log(D_t)$. Equation (2.17) can again be expressed as cointegrating relation between price and dividends, in this case, through log dividend-to-price ratio:

$$d_t - p_t = \frac{k}{1-\lambda} + E_t\left[\sum_{j=0}^{\infty} \lambda^j \left[-\Delta d_{t+1+j} + r_{t+1+j}\right]\right]. \tag{2.18}$$

In Equation (2.18) the log of the dividend-to-price ratio is expected to be stationary, which means that the log-linear cointegrating parameter is equal to one (recall that the linear model was cointegrated with the parameter $1/R$ where R is unknown ex-ante). Equation (2.18) highlights a logistic cointegrating relation between price and accounting fundamentals. This point has important empirical implications. Specifically, the relation between price and value (of cointegration), assuming that the Campbell and Shiller (1988) approximation is valid, should be tested in the log-form. The usefulness of cointegration for the relation between testing price and value is discussed in the following sections.

2.1.5. *Cointegration*

Cointegration is a statistical term that describes the long-run equilibrium tendency of variables to be tied together, or comove. More technically, Hamilton (1994, p571) defines cointegration as:

> "An (n x 1) vector time-series $\mathbf{y}_t$ is said to be cointegrated if each of the series taken individually is $I(1)$, that is, nonstationary with a unit root, while some linear combination of the series $\mathbf{a}'\mathbf{y}_t$ is stationary, or $I(0)$, for some non-zero (n x 1) vector $\mathbf{a}$."

Hamilton's definition includes terminology for the order of integration of a process. Specifically, $I(d)$ refers to a process which needs to be differenced d times before it is stationary. For example, a variable that is nonstationary in levels, but is stationary in changes is denoted as an $I(1)$ process. For cointegration to exist between two such variables, both variables, y_1 and y_2 must be nonstationary in levels, but stationary in changes, (i.e., each time-series is individually $I(1)$). The linear combination of y_1 and y_2, however, should be stationary (i.e., the time-series of the linear combination should be $I(0)$).

Consider the following simple example for two nonstationary variables x and y:

$$y_t - \beta x_t = u_t. \tag{2.19}$$

Then y and x are considered cointegrated if there exists some parameter β that makes u stationary. If both x and y are $I(1)$ and the linear combination is stationary, then x and y can be denoted as cointegrated as $CI(1,1)$. For the special cases where the variables x and y are cointegrated with the $\beta = 1$, this suggests that the variables never drift too far from being equal over time, and the log of the ratio x/y will be stationary.

2.1.6. Cointegration, the transversality condition, and mispricing

In an efficient market, prices should comove with their fundamental values (e.g., Campbell and Shiller, 1987; Lee et al., 1999). A sufficient condition for the rejection of the existence of an asset price bubble, therefore, is the cointegration of price with fundamental value. This result is due to the transversality condition:

$$\lim_{K\to\infty} E_t\left[\left(\frac{1}{1+R}\right)^K \tilde{P}_{t+K}\right] = 0. \tag{2.20}$$

Equation (2.20), the transversality condition, states that the present value of the terminal price tends toward zero as $K \to \infty$. This condition relies upon the expected stock price growing at a rate less than R. Where this is the case, there is a unique solution to the present value equation, and price equals the present value of dividends, or $P_t = V_t^*$. Specifically:

$$P_t = V_t^* \equiv \sum_{\tau=1}^{\infty} \rho^{-\tau} E_t\left[\tilde{D}_{t+\tau}\right], \tag{2.21}$$

where Equation (2.21) has a unique real solution for ρ such that price equals value. Note that while this equation is for the dividend model, the same result must hold for all equivalent accounting based valuation models. In this model the relation between price and value is one of static equality, and price is perfectly informationally efficient.

An alternative to the perfectly informationally efficient model above is to allow price to contain an error, or speculative component, ε_t^p. Specifically:

$$P_t = V_t^* + \varepsilon_t^p \tag{2.22}$$

Models of bubbles suggest that for a mispricing error to be 'rational' it must be expected to grow at a rate of R or less (e.g. Blanchard and Watson, 1982; Froot and Obstfeld, 1991). Specifically:

$$E_t\left[\varepsilon_t^p \mid \Theta_t\right] = (1+R)\varepsilon_t^p . \tag{2.23}$$

Blanchard and Watson (1982) then propose that a 'rational bubble' is the result of a 'self-fulfilling' inclusion of the speculative component into prices. The authors suggest that this rational bubble exists and pays investors an extraordinary return as compensation for the eventual capital loss if the stock is held until the bursting of the bubble. Specifically:

$$\varepsilon_t^p = \begin{cases} \left(\frac{1+R}{\pi}\right)\varepsilon_t^p + \xi_{t+1}, & pr(\pi) \\ \xi_{t+1}, & pr(1-\pi) \end{cases} \tag{2.24}$$

This specification implies that the bubble component grows at a rate greater than the expected rate of return, with an exogenous probability of collapse, π. While the bubble is growing, investors are compensated at a rate $\left(\frac{1+R}{\pi}\right)-1$ greater than R. The bubble of Blanchard and Watson (1982) also has the explosive expectation that:

$$E\left(\varepsilon_{t+i}^p\right) = (1+R)^i \varepsilon_t^p \tag{2.25}$$

This explosive condition implies that if the bubble lasts for sufficiently many periods, it will become nonstationary. This has important implications for the testing of cointegration between price and value.

If a bubble is present in the price of the stock, then the linear association between price and fundamentals will include a nonstationary component unless the bubble is included into the (researcher's) measure of fundamentals. Specifically, for the $P_t = V_t^* + \varepsilon_t^p$ case, if $V_t^* \sim I(d)$ and $\varepsilon_t^p \sim I(b)$ then $P_t \sim I(d+b)$ and for any b>0 the linear combination of price and value will not be stationary. The failure to find cointegration, however, does not in itself imply that the price contains a bubble. Instead, as Hamilton and Whiteman (1985) suggest, the finding of differing orders of integration for price and fundamentals could be due to rational agents conditioning on fundamentals not observed by the researcher.[9]

The bubble of Blanchard and Watson (1982) is based on rational expectations, implying that knowledge is common amongst all traders. This form of an asset price bubble in equilibrium has been ruled out analytically by Diba and Grossman (1982, 1987, 1988), and Tirole (1982, 1985). The reason put forward by Diba and Grossman (1982, 1987, 1988) is that for a bubble to exist at any point in time, it must have had a non-zero bubble in the price at all points in time since the asset's inception. Tirole (1982, 1985) rules out asset price bubbles in an equilibrium with common knowledge, due to the existence of rational profit-driven arbitrageurs. In summary, a bubble of this form is ruled out for any security with a finite horizon by backwards induction, and only under special conditions can an asset price bubble exist with symmetric information (Santos and Woodford, 1997).

However, subsequent work by Abreu and Brunnermeier (2002, 2003) suggests that rational arbitrageurs will delay corrective trades to remove mispricing in markets where arbitrage is limited, even when the mispricing is common knowledge. In such a market, the lack of timely arbitrage activity leads to the oscillation of price around its fundamental value and allows for the

[9] For example, if investors capitalize the value from R&D expenditure.

possibility of extended oscillations, or market bubbles, to exist in equilibrium. In such a setting, cointegration between price and value would describe the 'normal' market with a breakdown in cointegration being consistent with a market bubble.

In the following section, I briefly review the econometric literature to discuss the pertinent features of cointegration testing relevant to this thesis.

2.1.7. *The strengths and weaknesses of cointegration analysis*

The above discussion suggests that cointegration analysis is an appropriate technique for examining the time-series relation between market price and accounting based value. As with all statistical analysis, estimation of a cointegrating relation using a small dataset (in this case over a short time horizon) leads to a reduction in the power of the test. Hence, it could be claimed that in order to find a breakdown in the cointegrating relation, one can simply search for a short enough period of time and the test will be true by construction. This is a difficult objection to overcome when testing for an asset price bubble, as the nature of a bubble is to eventually correct at some point in time in the future.

An interesting econometric scenario arises where the span of a time-series can be disaggregated and measured at a higher frequency. In this section, I review prior econometric findings that discuss the effect of increasing the number of observations while holding the span of the data constant. I will focus mainly on the arguments relating to whether the use of time disaggregated data poses an econometric concern, rather than the technical aspects of the tests, which are covered in Maddala and Kim (2002).

The early work in this area is provided by Shiller and Perron (1985), Perron (1989, 1991) and Diebold and Rudebusch (1991). These authors studied whether the use of quarterly data would provide a more powerful test than the corresponding annual data and suggested that the power of the test depends more on the span of the data than the number of observations. Subsequent studies found that for a given span, the frequency of observations increases the power of the test, but at a diminishing rate (Ng, 1995), and while there can be an increase in the power of some tests, there is not a reduction in power in the tests when data is sampled more frequently (Choi and Chung, 1995). The opposite, however, is not true. Aggregating sub-interval data (such as monthly observations to annual observations) decreases the power of the test (Choi, 1992).

The above studies relate to stationarity tests; additional studies in this area have examined the power of the tests of cointegration from the vector-error correction based procedure; the Johansen trace-test and the maximum eigenvalue test (e.g., Johansen and Juselius, 1990). Again, the studies provide mixed evidence. Hooker (1993) first argued that temporal dissaggregation increased the power of the test significantly; however, Lahiri and Mamingi (1995) found no improvement in the power of the test. Hu (1996) investigated both the power of the Johansen trace test and maximum eigenvalue test for the VECM model.[10] She finds that the power of the test increases moderately with an increase in the frequency of observations, given a fixed time span. When a larger time span is available, however, the increase in the power of the test is greater. The small sample bias of under-rejecting the null of no cointegration, appears to apply only to a sample size of less than 100 observations (Podivinsky, 1998). The conclusion drawn from this literature is that for a given time-span, it is preferable to use higher frequency data when available (Maddala and Kim, 2002, p229).

[10] These models use multivariate time-series approaches and have not been implemented in the accounting literature at this point. The benefit of this method is that it allows for the modeling of processes for each variable under investigation (see Hamilton, 1994).

2.2. *Limits to arbitrage*

2.2.1. *The efficient market hypothesis*

In this section, I briefly examine the literature on the debate of market efficiency, as it is pertinent to the development of the limits to arbitrage literature which seeks to explain potential deviations from market efficiency. There is considerable evidence that accounting information is associated with future returns.[11] Collectively these studies suggest that part of the information that is made publicly available through accounting disclosures is not reflected in contemporaneous market prices. There is a lack of consensus, however, as to whether or not the market is efficient with respect to accounting information.

Market efficiency puts forth the hypothesis (hereafter, the efficient market hypothesis) that efficient market prices 'fully reflect' all available information (e.g., Fama, 1998).[12] Consider the simple example where the market price at time t, P_t is a function of the rational intrinsic value of the security at time t, V_t^*, given the information available at time t, Θ_t:

$$P_t = E[V_t^* \mid \Theta_t] = E_t V_t^* . \tag{2.26}$$

One period ahead, investors hold additional information, Θ_{t+1}. If price is an accurate reflection of the information set at time t, as $\Theta_t \subset \Theta_{t+1}$, by the law of iterated expectations, the expected (abnormal) return over the period t to t+1, is equal to zero:

$$E[P_{t+1} - P_t] = E_t[E_{t+1}[V_{t+1}^*] - E_t[V_t^*]] = 0 . \tag{2.27}$$

Fama (1970, 1991, 1998), among others, suggests that the market can be considered efficient when it is always true that market prices fully reflect available information (see also,

[11] I review this literature in the following section.

[12] While the origins of the efficient market hypothesis can be traced back to the early works of Bachelier (1900) and Cowles (1933), modern interpretation of the efficient market hypothesis stems largely from the work of Samuelson (1965, 1973) who shows that in an informationally efficient market, abnormal price changes will not be forecastable if the price incorporates the expectations of all market participants.

Campbell, Lo and MacKinlay, 1997; Lo and MacKinlay, 1999). In an efficient market, publicly available accounting information should be fully reflected in the market price, and when new information is released to the market, the price should reflect this new information in a timely and unbiased manner.

2.2.2. *Accounting information and the efficient market hypothesis*

A large body of accounting literature examines the relation between accounting information and market returns. In this section, I review some of the many studies that have investigated whether or not accounting information is associated with future returns.[13] Finding evidence against the null, many of these studies have claimed that the association between accounting information and stock returns provides evidence of market inefficiency (e.g. Campbell and Shiller, 1988 a,b; Bernard and Thomas, 1989, 1990; Hodrick, 1992; Shiller, 1984; Lakonishok, Shleifer and Vishny, 1994; Sloan, 1996; Frankel and Lee, 1998; Lee, Myers and Swaminathan, 1999). These studies generally show that publicly available accounting information can be used to predict future stock returns over long horizons.[14]

Since Ball and Brown (1968), studies have shown that while the market incorporates earnings information into market prices, some of the information in earnings appears to be incorporated into market prices with a delay (see also, Lev and Ohlson, 1982). This evidence of post-earnings-announcement drift has been interpreted as evidence against market efficiency by

[13] Kothari (2001) and Lee (2001) provide a much more detailed discussion of this voluminous literature.
[14] There are also numerous studies which examine the long-horizon predictability of stock returns using other announcements and events. The first of these studies by Fama, Fisher, Jensen and Roll (1969) found that market prices adjusted to stock-splits in a timely manner. More recent examples include: the predictable poor performance following a firm's initial public offer (e.g. Ritter, 1991) and seasoned equity offers (e.g. Loughran and Ritter, 1995). As the focus of this review, however, is on the relation between price and accounting based measures of value, I focus primarily on studies which use accounting information directly in their analysis.

authors such as Lev and Ohlson (1982) and Bernard and Thomas (1989, 1990).[15] Collectively, these studies fall into the 'mispricing based explanations' literature. Other authors such as Ball, Kothari and Watts (1988) and Foster, Olsen and Shevlin (1984) have suggested that the observation of abnormal returns around the announcement of earnings is due to misspecification of the equilibrium model of 'normal returns,' and Ball, Kothari and Watts (1993) argue that the drift is due to nonstationarity in systematic risk, which is not captured in the design of these studies. These studies fall into the 'risk based explanations' literature. Recent studies have investigated market frictions such as arbitrage cost (Mendenhall, 2004) and liquidity risk (e.g., Chordia, Goyal, Sadka, Sadka and Shivakumar, 2006) as potential reasons why post-earnings-announcement drift persists nearly 30 years after its discovery.

Evidence of an association between ratios of accounting variables to price and future returns is another area where prior literature has debated whether or not the market misprices accounting information (see, Campbell and Shiller, 1988 a, b; Fama and French, 1988, 1992, 1995; Hodrick, 1992; Shiller, 1992; Frankel and Lee, 1998). One stream of this literature illustrates that the variation in stock prices cannot be explained by the variation in fundamentals over time, suggesting that some component of stock prices is due to mispricing (LeRoy and Porter, 1981; Shiller, 1989; Lee, Myers and Swaminathan, 1999). The risk based explanation for the predictability stems mainly from the collective work of Fama and French (1992, 1995, 1996, 1998) who suggest that the association observed between market prices and the ratio of book-value-to-price is due to the book-to-market ratio being a proxy for risk.[16] Lakonishok, Shleifer

[15] Bernard and Thomas (1989, 1990) suggest an alternative hypothesis that unsophisticated investors systematically underreact to earnings information. Since then, Bartov, Radhakrishnan, and Krinsky, (2000) found that the evidence associated with post-earnings announcement drift is stronger for firms with lower institutional holdings, which the authors claim supports the unsophisticated investors hypothesis.

[16] Additional studies that investigate mispricing and risk based explanations for the association between returns and the book-to-market ratio (often termed the value-glamour literature) include Ball (1978), Banz (1981), Kothari,

and Vishny (1994) propose an alternative explanation: that ratios of earnings-to-price, cash-flow-to-price and book-to-market are proxies for overreaction to past performance. Poor performance is represented in low ratios, and strong performance is represented in high ratios. As investors overreact to past earnings growth, poor long-horizon returns are consistent with investors being surprised by the mean-reversion in earnings growth.

An important addition to the debate on market efficiency with respect to accounting information is found in Sloan (1996). His finding of a negative association between accruals and subsequent stock returns, now labeled the accrual anomaly, implies that investors overreact to accrual information, and are systematically surprised by the lack of persistence in the accrual component of earnings. Risk based explanations for the accrual anomaly, such as those presented in Khan (2006), consider low-power tests (see, Hirshliefer, Hou and Teoh, 2006). Recent explanations include the accrual anomaly is due to the limits to arbitrage (Mushruwala, Rajgopal and Shevlin, 2006) and the stock-selection preferences of institutional investors (Lev and Nissim, 2006).

2.2.3. Debate on the efficient market hypothesis

Overall, these studies suggest that either the market systematically underreacts or overreacts to publicly available accounting information. In a recent review of the literature, Fama (1998, p284) defends the efficient market hypothesis by claiming:

> A problem in developing an overall perspective on long-term return studies is that they rarely test a specific alternative to market efficiency. Instead, the alternative hypothesis is vague, market inefficiency. This is unacceptable. Like all models, market efficiency (the hypothesis that prices fully reflect available information) is a faulty description of price formation. Following the standard scientific rule, however, market efficiency can only be replaced by a better specific model of price formation, itself potentially rejectable by empirical tests.

Shanken and Sloan (1995), Kothari and Shanken (1997), Daniel and Titman (1997) and Conrad, Cooper and Kaul (2003).

> Any alternative model has a daunting task. It must specify biases in information processing that cause the same investors to under-react to some types of events and over-react to others. The alternative must also explain the range of observed results better than the simple market efficiency story; that is, the expected value of abnormal returns is zero, but chance generates deviations from zero (anomalies) in both directions.

Fama (1998) considers the market efficiency hypothesis to be robust to deviations from perfect efficiency due to 'chance' even when these 'chance' findings are robust to multiple studies and empirical settings. In contrast Hirshleifer (2001, p1534) summarises the disagreement in the literature as:

> Despite many empirical studies, scholarly viewpoints on the rationality of asset pricing have not converged. This is probably a result of strong prior beliefs on both sides. On one side, strong priors are reflected in the methodological claim that we should adhere to rational explanations unless the evidence compels rejection, and in the use of the term "risk premium" interchangeably with "mean return in excess of the risk-free rate." For those on the opposite side, risk often comes quite late in the list of possible explanations for return predictability.

Hirshleifer (2001, p1534) continues by suggesting that mispricing and risk are most likely intertwined:

> There are several potential noisy proxies for the degree of underpricing, such as price containing variables (e.g., book/market, market value, earnings/price), measures of public mood (e.g., the weather), or actions possibly taken to exploit mispricing (e.g., recent occurrence of a stock repurchase or insider purchases). Risk and mispricing effects do not necessarily take such a simple linearly separable forms ... but it is still useful to keep the two notions conceptually distinct.

These quotes from Fama (1998) and Hirshleifer (2001) highlight the impasse between risk based and mispricing based explanations for the long horizon predictability of accounting information for future stock returns. The behaviour of the market observed during the late 1990s has led to further debate. Market efficiency strictly rules out the existence of market bubbles. However, the rapid run-up and subsequent decline of the market during the late 1990s was considered by many observers to be consistent with a market bubble. In the following section, I outline some of the pertinent features of this debate.

2.2.4. *Debate on the existence of a bubble in the late 1990s*

Although the large gains and subsequent losses observed in the 1990s market have been well documented, there is a surprising lack of consensus on whether this period contained what the financial press and other observers labeled a 'financial market bubble.'[17] The term financial market bubble (or asset price bubble), has two common definitions, one using returns as the primitive and one using price as the primitive. For returns, a bubble can be defined as an ex-post description of market behaviour which displays an extended rise in price followed by a sharp decline in the price of the same asset (e.g., Kindleberger, 1978). For prices, a bubble is the ex-ante divergence of price from intrinsic value (e.g., Brunnermeier, 2001). The latter definition is of interest here.

One reason for the extended debate over the existence of a financial market bubble is due to the difficulty in measuring intrinsic value. First, intrinsic value is unobservable and is generally not considered exogenous, but rather determined endogenously in equilibrium (e.g., Brunnermeier, 2001, p47). Second, it is heavily debated as to whether or not market behaviour consistent with asset price bubbles can exist in equilibrium solutions that include rational agents (Brunnermeier, 2001, p48–55).

Abreu and Brunnermeier (2003) present a model in which a financial market bubble can exist despite actions of active arbitrageurs. They argue that the arbitrageur will be subject to the funds-flow to performance problem. The funds-flow to performance problem is discussed by Shleifer and Vishny (1997) as a limit to arbitrage, especially when the arbitrage position would require a long holding period. Building on this logic, Abreu and Brunnermeier (2003) show that

[17] Shiller (2000) and Stiglitz (2002) provide book-length coverage of many aspects of the 1990s market. An excellent discussion of the use (and misuse) of accounting information in 'bubble' markets can be found in Penman (2003).

rational arbitrageurs may delay taking their corrective actions in the short-term, in favour of 'riding out a wave' of sentiment. The authors conclude by stating:

> This paper argues that bubbles can persist even though all rational arbitrageurs know that the price is too high and they jointly have the ability to correct mispricing. Though the bubble will ultimately burst, in the intermediate term, there can be a large and long-lasting departure from fundamental values.

In a similar vein, Penman (2003, p77) argues that: "Speculative beliefs feed rising stock prices that beget even higher prices, spurred on by further speculation" and one of the roles of accounting is to "challenge speculative beliefs, and so anchor investors on fundamentals." The common theme in support of a market bubble during the late 1990s is that fundamental values did not justify the high prices observed during this period.

Much of the recent debate regarding the existence of a financial market bubble centres around the technology and internet stocks that were listed on the NASDAQ during the late 1990s. For example, Ofek and Richardson (2002, 2003) and Demers and Lev (2001) highlight the 'implausibly high' growth rates needed to justify the prices of internet stocks during the late 1990s, Bartov and Mohanram and Seethamraju (2002) show an association between internet stock prices with non-traditional (and temporary) measures of potential value and Cooper, Dimitrov and Rau (2001) show that firms who changed their name to include '.com' accrued abnormal market returns during the late 1990s. On the other side of the debate, Pastor and Veronesi (2006) present a model based on rational expectations. Their valuation model justifies the high prices on the NASDAQ during the late 1990s, based on the assumption that investors demand a growth premium due to the risk associated with future profitability, and that variance in future profitability was abnormally high for NASDAQ listed firms.[18]

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[18] Instead of this risk being embedded in prices in the traditional manner, which would mean that market participants demand a higher expected rate of return, the authors suggest that this risk is embedded in prices as higher growth

2.2.5. *Accounting information and investor rationality*

In this section, I discuss the role of arbitrageurs in linking price to fundamental value. One critique of the traditional approach to asset-pricing under the efficient market hypothesis is that it is static in nature. Specifically, if an asset is priced efficiently, then a particular piece of information (such as accounting information about value) has already been incorporated into the price, and is superfluous. A dynamic approach to this problem admits that if asset prices are being revised to reflect new information, then achieving market efficiency is a process. If efficiency is a process, then how asset prices become efficient (i.e., how they adjust to new information) cannot be separated from asset returns at any point in time. If markets are imperfect in the way they process information, then the mechanism by which the market becomes more or less efficient is crucial in understanding asset-pricing. Prior literature suggests that a crucial mechanism in the market is the existence of active arbitrageurs.

The term arbitrageur refers to sophisticated investors who are considered capable of profiting from the correction of informational inefficiencies (e.g., Lee, 2001). The traditional model which supports the equality of price and intrinsic value is centred on rational and competitive market participants.[19] Even if not all market participants are perfectly rational, the traditional argument is that sufficiently many arbitrageurs offset the positions of imperfectly rational traders. Hirshleifer (2001, p1536) argues that the case for rational prices based on arbitrage is deficient:

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rates (see also, Pastor and Veronesi, 2003). Considering how the value estimate obtained from a constant growth model increases as the difference between the discount rate and the growth rate decreases, the authors' conclusion, that their calibrated model can justify high prices, appears somewhat circular in nature.

[19] DeBondt (2005, p207) suggests that "Under complete rationality an individual is able to view the full consequences of each possible action at once and choose objectively what is best for him or her. In other words, the theory does not treat cognition as a scarce resource."

> To think about whether mispricing is viable, consider the traditional argument for rational price setting. In this account, smart traders spot dollar bills lying on the ground and grab them, which does away with mispricing. Setting aside the dynamics of wealth momentarily, the arbitrage story is incomplete in two ways. First, equilibrium prices reflect a weighted average of the beliefs of the rational and irrational traders. So long as each group has significant risk-bearing capacity, both influence prices significantly. Arbitrage is a double-edged blade: Just as rational investors arbitrage away inefficient pricing, foolish traders arbitrage away efficient pricing. Second, in some respects, *all* investors may be imperfectly rational. Even in the Olympics no one runs at the speed of light; some cognitive tasks are just too hard for any of us.

Following this view of Hirshleifer (2001), perfectly efficient markets are not feasible unless all investors are equally (and perfectly) rational. The equilibrium price, in his view, will therefore oscillate between being relatively more and relatively less efficient over the course of time. If we accept the possibility that price and value can diverge, then understanding processes and mechanisms through which prices become more or less efficient is important (e.g., Lee, 2001).

2.2.6. Limits to arbitrage

Early analytical work into the problem of how information is reflected in market prices suggests that the arbitrageur must be able to trade with stealth. Specifically, the arbitrageur cannot profit if their information is reflected the instant their order is submitted to the market (Grossman and Stiglitz, 1980; Diamond and Verrechia, 1981; Kyle, 1985). DeLong, Shleifer, Summers and Waldmann (1990a) suggest that an arbitrageur can take positions in securities following good news, on the expectation their positions induce future demand from the delayed actions of feedback traders. Following the price increase, rational arbitrageurs are able to unwind their positions at a profit as the feedback traders continue to drive prices up in the short-term.

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. This leads to risk-aversion (i.e., arbitrageurs cannot be considered risk-neutral) and capital constraints. Even when there are no trading frictions, arbitrageurs who are risk-averse and face capital constraints, arbitrageurs will require a higher return (i.e., mispricing) to hold 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, Shleifer, Summers and Waldmann, 1990b, Dow and Gorton, 1994; Shleifer and Vishny, 1997). 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.

Recently, Abreu and Brunnermeier (2002, 2003) propose that arbitrageurs are also subject to synchronization risk. Specifically, the collective actions of arbitrageurs are required to force mispricing to correct endogenously (i.e., through their actions). Without synchronization, arbitrageurs cannot successfully trade against mispricing. Due to the exit of capital from investment funds that perform poorly in the short-term regardless of their long-term potential profitability, arbitrageurs also behave myopically. They may prefer to profit from a divergence from intrinsic value rather than trade against mispricing when there is uncertainty as to when sufficiently many arbitrageurs will trade to correct the mispricing.

Arbitrageurs also face holding costs (e.g., Dow and Gorton, 1994; Pontiff, 2006), which are higher for short-sales as the capital from short-positions is not completely available to the short-seller. As arbitrageurs are assumed to be under-diversified (e.g., Shleifer and Vishny, 1997) they will seek to hedge their positions with a close substitute. This is presumably the case not just for the arbitrage positions that require a direct short position (e.g., Dechow, Hutton, Meulbroek and Sloan, 2001), but also potentially as a hedge against taking a long position. While prior literature has shown that in the cross-section most stocks are easily and cheaply shorted (e.g. Geczy, Musto and Reed, 2002; D'Avolio, 2002; Ofek, Richardson and Whitelaw, 2004; Asquith, Pathak and Ritter, 2005) margin calls will still mean that more of the arbitrageur's capital becomes restricted when taking a more volatile short-position.

2.2.7. Empirical studies on the impact of arbitrage costs

Pontiff (1996) found that the discount on closed-end funds is higher for firms with higher idiosyncratic risk, which is a proxy for arbitrage cost as it provides a measure of the difficulty to find a substitute for the mispriced stock. This measure of arbitrage costs has also been related to anomalous returns to firms that are included in the S&P 500 index (Wurgler and Zhuravskaya, 2002), the book-to-market anomaly (Ali, Hwang and Trombley, 2003), the post-earnings-announcement drift anomaly (Mendenhall, 2004), SEO underperformance (Pontiff and Schill, 2004) and the accruals anomaly (Mushruwala, Rajgopal and Shevlin, 2006). Together, these studies show that the magnitude of the anomalous return is greatest for stocks with higher levels of idiosyncratic risk, suggesting that these anomalies may persist due to the inability of arbitrageurs to remove mispricing in a timely manner.

2.3. Conclusions

Reviewing the literature, I highlight that a necessary (but not sufficient) condition for market efficiency is that price and value should be cointegrated. This implies that finding cointegration between price and value provides a sufficient condition for the rejection of an asset pricing bubble. The converse, lack of cointegration, is a necessary but not sufficient condition for the existence of a price bubble. This is because measurement error in accounting information means that identifying a price bubble using value is a joint test of the bubble and measurement error. In Chapter 3, I address the open question as to whether or not the price and value remained cointegrated during the late 1990s, which potentially included a financial market bubble.

I also highlight that a crucial mechanism, which maintains market efficiency, is the timely corrective actions of arbitrageurs whenever new information arrives to the market. Given the growing literature on the limits to arbitrage, it appears more plausible to consider the market as a process of moving towards efficiently reflecting information into prices. In such a setting, prices cannot be considered as simply equating to fundamental value by assumption. In Chapter 4, I examine whether arbitrage costs influence the relation between market price and value.

3. Can market price diverge from fundamentals for an extended period? Tests of risk and mispricing explanations[20]

3.1. Synopsis

In this chapter, I contrast the historical time-series relation between market price and accounting based value with the relation during the late 1990s, which potentially included a financial market bubble. Historically, price and value tend to converge, with convergence documented by cointegration. In contrast, during the speculative period, there is no reliable evidence that price and value are cointegrated. The breakdown in cointegration is consistent with the breakdown in the tendency of price being 'anchored' to value, which is found historically. I then contrast risk (i.e., measurement error) and mispricing based explanations for this lack of cointegration. The breakdown does not appear to be due to measurement error in either the equity premium or expected growth. The results are instead consistent with the speculative market of the late 1990s disregarding fundamental information and replacing information trading with momentum trading, displaying 'bubble-like' behaviour.

3.2. Introduction and motivation

Traditional asset-pricing theory rules out the existence of market bubbles. Instead, competition among rational investors is expected to result in bid and ask prices being at their equilibrium levels, where price reflects the present value of expected future dividends. Even if some traders are irrational, the existence of rational arbitrageurs guarantees the timely removal of

[20] This chapter is a revised (and adapted) version of a paper with the title "Can market price diverge from fundamentals for an extended period? Evidence from the late 1990s," that has previously been presented at the *Review of Financial Studies* – Indiana University Conference on the "Causes and Consequences of Recent Financial Market Bubbles," the American Accounting Association's Annual Meeting, the University of Queensland, the University of Technology Sydney, Macquarie University, Rice University, and the Melbourne University Capital Markets Symposium.

mispricing such that price is 'anchored' to intrinsic value. Lee, Myers and Swaminathan (1999) show that, at least until mid 1996, price and accounting based measures of fundamental value (hereafter, value) tend to converge over time, that is, they are cointegrated.[21] Their result provides evidence that the market is reasonably efficient, because it implies price deviates from intrinsic value only temporarily.[22] The late 1990s provides an interesting setting to re-examine the relation between price and accounting based measures of fundamental value as there is ongoing debate about the rationality of investors during that period.

Accounting provides quantifiable information about firm-value; however, during the late 1990s market participants criticized financial reporting as out of date, backward looking and without value relevance in the 'new economy.' With hindsight this conclusion appears hastily accepted and generally incorrect (see, Penman, 2003). Despite the common consensus that simple valuation metrics such as price-to-earnings ratios and market-to-book ratios were unusually high during the late 1990s, debate remains strong about whether this was due to regime shifts in expected growth and rates of return or due to market imperfections.[23] It remains an open question whether price continued to be anchored to measures of value during this more recent 'speculative' period.

I address this question by examining the time-series relation between price and measures of value which incorporate expected growth and rates of return over the period 1979–2004. Specifically, I investigate whether the cointegrating relation expected between price and value, using the estimates of value from the residual income model, breaks down in the late 1990s. The

[21] Cointegration is a statistical term that describes the long-run equilibrium tendency of two variables to be tied together. See Chapter 2.
[22] Where the term 'anchored' is used to describe the mean-reversion of the deviations back towards their equilibrium position. Penman (2003) describes the anchoring role of fundamental valuation in capital markets.
[23] For example Shiller (2000) refers to the late 1990s as a period of 'irrational exuberance' following the quip by Alan Greenspan (as quoted in Miller 2002). Alternatively, Pastor and Veronesi (2006) suggest that it was "not obvious" that there was a market bubble during this period, providing a non-traditional model which provides plausible calibrated values.

potential 'bubble period' is often considered to be the time from the Netscape IPO in August 1995 until the NASDAQ 'market correction' in March 2000. I therefore contrast the later period (1992–2004), which contains the potential bubble period, with an earlier period (1979–1991), which has been shown to be cointegrated (Lee et al., 1999).[24]

When examining the first half of the time-series, I find that during the early period of the study (1979–1991) price and value are cointegrated, consistent with the findings in Lee et al. (1999). In marked contrast, however, when examining the second half of the time-series, which includes the potential 'bubble period' (1992–2004), there is no reliable evidence of cointegration between price and value. This result, which is robust to alternative statistical methods and alternative measures of value, provides evidence that price diverged from fundamentals for an extended period in the late 1990s.

My results are conservative in nature. I bias against the finding of a breakdown in cointegration by testing over an extended time-series around the potential 'bubble period' rather than just the shorter period in question.[25] I bias against finding a bubble (significant overvaluation) in price relative to value by using analyst forecasts of growth to estimate value, where these forecasts are known to be optimistic (e.g., Sharpe, 2002). Finally, I use value-weighted tests to bias against a small firm effect, allowing the results to be dominated by large firms with a large analyst following. As such, my results suggest market-wide overvaluation during this period – not just overvaluation of small, recently-listed technology firms.

[24] I choose to contrast two equal periods of time as the power of tests for cointegration decrease as the span of the data decreases (Maddala and Kim, 2002). As I choose equal periods, ceteris paribus, the power of the test is equal between the two periods.

[25] Any potential 'bubble' must be temporary in nature. However, testing for cointegration using only a short time-span can be argued as a statement on the power of the test. Podivinsky (1998) suggests that the critical values for cointegration tests may be inappropriate for sample sizes of less than 100 observations. I split the sample, from 1979–1991 and from 1992–2004, to yield 156 observations in each sample period, so that power is less of a concern. In addition, I provide simulation results for the power of the cointegrating test, in Chapter 5.

If the results for the potential 'bubble period' are due to a change in the behaviour of market participants, then the process whereby price was revised over time to reflect new information about intrinsic value must have changed. To investigate this possibility, I decompose the relation between price and value using a vector error correction model (VECM). I find evidence consistent with a change in the behaviour of the price time-series, but no reliable evidence of a change in the behaviour of the value time-series. Specifically, I show that historically price corrected for the level of divergence from value, which is evidence of an 'anchoring' effect. This anchoring effect was not reliably evident in the speculative period. Instead, price held an increased association with lagged price, consistent with a market-wide shift that replaced information trading with momentum trading. While this implies lagged price was a better indication of intrinsic value during this period, the results are also consistent with the findings of Brunnermeier and Nagal (2004), who show hedge funds tilted their portfolios towards speculative technology stocks, rather than 'attacking' their speculative prices.

Finally, I provide further tests to add to the debate on the role of expected rates of return and expected growth rates during this period. One potential explanation for the high prices during the potential 'bubble period' is that the equity premium, and hence the discount rate, was low. I investigate this possibility by first estimating value with equity premia that range from zero to six percent. I find no compelling evidence in support of the low equity premium rate proposition. Second, I sort firms into portfolios based on their prior period market beta. I find some evidence that firms with relatively low betas were weakly cointegrated during this period, but no consistent evidence of cointegration for medium and relatively higher beta firms. These results are generally consistent with the contention that the mispricing was market-wide, and not due to the equity premium. To investigate the possible growth explanation, I sort firms into growth portfolios based on analyst forecasts of expected long-term growth. I find no compelling evidence in support of the mis-measurement of growth in the residual income model. Instead, I report evidence consistent with historical cointegration between price and value for all portfolios except the lowest growth portfolio, and no reliable evidence for any portfolio during the potential bubble period.

The consequences of speculative markets are evident at both the macroeconomic and microeconomic level. At the macroeconomic level, mispricing in speculative markets will take an extended period to correct. This has the possibility of influencing the actions of arbitrageurs, for example, in the decision to short-sell (Lamont and Stein, 2004; Abreu and Brunnermeier, 2003). At the microeconomic level, in speculative markets accounting information arguably becomes less relevant to corporate 'value.' Instead of a focus on the measurement of earnings with historical asset values, firms highlight subjective pro-forma earnings and untested measures of 'value' (Penman, 2003).

The remainder of this chapter is organized as follows. In Section 3.3, I discuss the expected time-series relation between intrinsic value and market price. In Section 3.4, I outline the measurement of variables used in the study. In Section 3.5, I present empirical tests. In Section 3.6, I present further analysis of expected growth and required rates of return explanations. In Section 3.7 I provide a brief conclusion to the chapter, with a discussion of the implications of this research for valuation and market efficiency.[26]

3.3. *Model and hypothesis*

3.3.1. *Definitions of price and value*

[26] Further sensitivity analysis is presented in Chapter 5.

In this section, I outline a model of the expected relation between price and value, for which the ratio based model of Lee et al. (1999) is a particular case.[27] This model leads to a testable hypothesis based on the cointegration of price and value. Under 'no-arbitrage' assumptions, price and intrinsic value are equal at all points in time, $P_t = V_t \ \forall t$.[28] Intrinsic value is defined as the present value of future cash-flows based on all available information. Where this is the case, price is the best reflection of intrinsic value and fundamental valuation (which uses accounting information) is superfluous. Intrinsic value, however, is not perfectly observed. Assuming that intrinsic value is unobservable (denoted V^*), a general representation of price is that it provides an unbiased estimate of intrinsic value. In log form, the expected relation between log price (p_t) and (unobservable) log intrinsic value (v_t^*) is:

$$\begin{aligned} p_t &= \gamma_1 v_t^* + u_{1t}^* \\ v_t^* &= v_{t-1}^* + \eta_t^* \end{aligned} \tag{3.1}$$

where $p_t = \log(P_t)$ and $v_t^* = \log(V_t^*)$, and both u_{1t}^* and η_t^* are assumed to be unrelated white noise processes. Unobservable log intrinsic value and log price both follow random walks, and price is a 'good' estimate of value when the parameter γ_1 is known and the variance of u_{1t}^* is small. Where the market is efficient, the expected value of γ_1 is one and price is an unbiased estimate of intrinsic value. This model implies all information about intrinsic value can be instantaneously, accurately and continuously incorporated into price (Campbell, Lo and MacKinlay, 1997). Characterizing market dynamics in this fashion, fundamental analysis is superfluous.

[27] Lee et al. (1999) impose restrictions on the relation between price and value such that the value-to-price ratio is appropriate for testing for cointegration. The generalisation of their model allows for inferences to be drawn regarding the extent to which price reverts towards value and value reverts towards price through time.

[28] 'No arbitrage' assumptions imply there are powerful economic agents who instantly spot and correct mispricing such that no arbitrage opportunities exist and price always reflects intrinsic value.

In a 'relatively' efficient market, price is bid and offered to the point that it reflects intrinsic value 'rapidly.' In practice, error in measuring prices, arising from transaction costs and trading frictions, prevents market efficiency from holding exactly at every possible date (e.g., Garman and Ohlson, 1981). Trading strategies using fundamental analysis, however, in an attempt to capture mispricing (i.e., u_{1t}^*), would not be profitable after accounting for transaction costs. This model also implies that the expected value of γ_1 is one and price is an unbiased estimate of intrinsic value. Again, characterizing market dynamics in this fashion, fundamental analysis is superfluous.

Alternatively, with impediments to arbitrage, price may temporarily diverge from intrinsic value and fundamental analysis may provide a profitable tool for identifying mispriced securities (Shleifer and Vishny, 1997; Lee et al., 1999; Lee, 2001). As arbitrage impediments are overcome, price should revert towards value. If arbitrageurs are impeded from taking actions immediately with respect to over- and under-priced securities, mispricing, u_{1t}^*, will be mean-reverting. Where this is the case, price and intrinsic value are expected to be cointegrated, suggesting that price and value will not diverge 'too far' from one another. In the following section, I discuss the Lee et al. (1999) model of cointegration between price and accounting measures of value.

3.3.2. Relating market price with value using the ratio specification

Lee et al. (1999) consider price and accounting measures of intrinsic value (both observable) as unbiased functions of unobservable intrinsic value plus a random error:

$$\begin{aligned} p_t &= v_t^* + u_{1t}^* \\ v_t &= v_t^* + u_{2t}^* \end{aligned} \tag{3.2}$$

where v_t^* represents the log of unobservable intrinsic value, such that at any point in time both the log of price, p_t, and the log of the value measure, v_t, approximate intrinsic value with error.[29] Price includes a mispricing error, u_{1t}, and value estimates include a measurement error, u_{2t}. In logs, dividing value by price yields:

$$v_t/p_t = v_t - p_t = u_{2t}^* - u_{1t}^*, \tag{3.3}$$

where v_t/p_t is value-to-price ratio which is the 'spread' between the log of price and the log of value which includes a mispricing error and a value measurement error.

The time-series of the log value-to-price ratio, $\{v_t/p_t\}_{t=1}^T$, is potentially informative about time-variation in mispricing as long as u_{2t}^* and u_{1t}^* are not perfectly correlated and price is not always equal to intrinsic value. In addition, if the measurement error, u_{2t}^*, is not entirely due to time-varying expected returns, then v_t/p_t will have some predictive ability over future market movements because impediments to arbitrage create a mispricing term, u_{1t}^*, that is expected to be removed over time.

Lee et al. (1999) find that the time-series of v_t/p_t is stationary, implying that, for the long run cointegrating relation between price and value, the cointegrating vector **a** is [1,1] (Hamilton 1994, pp582–596). Specifically:

$$\mathbf{a}[v_t - p_t]' = v_t - p_t = v_t/p_t, \tag{3.4}$$

Equation (3.4) is written as a cointegrated function between observable price and value. In this model the cointegrating vector is restricted to one, which presupposes that the long-run relation between price and intrinsic value is equality, and measurement errors are mean zero. The

[29] Alternative specifications of v_t include discounted dividend, forward EPS and EPS growth, and residual income models. This study focuses on the residual income model.

hypothesis tested under this specification can be considered as a model of 'long-run market efficiency.' Specifically, this specification provides a practical test of the hypothesis that price equals value in the long-run.[30] This specification is a particular case of the more general model that I outline in the following section.

3.3.3. *Decomposing the ratio specification*

The ratio specification imposes a static relation between the price and value variables, and restricts the expected value of the long-run relation to one. If price and value are revised over time to reflect new information about intrinsic value, then the process whereby each of these measures incorporates information should be separated. This can be done by decomposing the ratio measure into separate time-series. Separating the time-series is particularly important in this setting as a breakdown in the relation between price and value due to a 'bubble' implies that prices did not incorporate information about intrinsic value.

The first assumption I remove from the ratio based model is that the cointegrating parameter is equal to one. Consider the following specifications of price and value, respectively: $p_t = \gamma_1 v_t^* + u_{1t}^*$, and $v_t = \gamma_2 v_t^* + u_{2t}^*$. Combining p_t and v_t: $v_t = \gamma_1^{-1}\gamma_2(p_t - u_{1t}^*) + u_{2t}^*$, which can be rearranged as:

$$v_t - \beta p_t = u_{2t}^* - \beta u_{1t}^* \tag{3.5}$$

where $\beta = \gamma_1^{-1}\gamma_2$ and is not restricted to one.[31] Equation (3.5) is a generalisation of Equation (3.4) with the cointegrating vector **a** in this case equal to $[1,-\beta]$.

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[30] See for example Hamilton (1994, p572–3).

[31] Equation (3.5) is identical to the Lee et al. (1999) model in equation (3.4) with the restriction that $\gamma_1 = \gamma_2$.

For price and value to be cointegrated, either price or value (or both variables) must correct for the disparity between price and value over time. This suggests that either price or value must be related to lagged value or lagged price respectively. I consider the following model, which includes a single lag of price and a single lag of value:

$$\begin{aligned} v_t &= a_{11}v_{t-1} + a_{12}p_{t-1} + e_{vt} \\ p_t &= a_{21}v_{t-1} + a_{22}p_{t-1} + e_{pt} \end{aligned} \tag{3.6}$$

Equation (3.6) is a vector-autoregression (VAR) representation of the relation between price and value in their logs. The Granger-representation theorem states that for any set of nonstationary (in levels) variables, vector error correction models (VECM) and cointegration are equivalent representations (e.g., Enders 2004, p333). The VECM representation is found by differencing the VAR representation of the relation between price and value. Accordingly, Equation (3.6) is written in VECM form as:

$$\begin{aligned} \Delta v_t &= \alpha_v(v_{t-1} - \beta p_{t-1}) + \phi_{21}\Delta p_{t-1} + \phi_{22}\Delta v_{t-1} + e_{vt} \\ \Delta p_t &= \alpha_p(v_{t-1} - \beta p_{t-1}) + \phi_{11}\Delta p_{t-1} + \phi_{12}\Delta v_{t-1} + e_{pt} \end{aligned} \tag{3.7}$$

In this model, the cointegrating parameter in Equation (3.5) is estimated as $[1,-\beta]$ by normalizing for value (i.e., the coefficient on value is one). The parameters α_p and α_v are the error correction terms for price and value, respectively. Error correction terms display the tendency of price and value to respond to the disparity from the long-run relation between price and value. The parameters of the VECM model relate to the VAR model in Equation (3.6) as follows: $\beta = (1-a_{22})/a_{21}$, $\alpha_v = -a_{12}a_{21}/(1-a_{22})$ and $\alpha_p = a_{21}$. The autoregressive parameters, ϕ, control for autocorrelation in the changes of price and value.

It is clear from the above reconciliation that the cointegrating parameter is the ratio of the tendency of price to be related to prior prices and to prior values. Reconciling with the model of Lee et al. (1999), the cointegrating vector will be equal to one when the sum of the parameters in the price equation (a_{21} + a_{22}) equals one. The error correction parameter, α_p, also warrants further attention. In this form, it provides a measure of the tendency of price to 'anchor' to value, and is given as the association of price with lagged value (conditional on the information in lagged price). The additional parameters relative to the Lee et al. (1999) ratio specification allow for a test of whether or not price corrects for the disparity between price and value (as suggested by the expected anchoring role of fundamentals).

3.3.4. Stability of the relation between price and value

Lee et al. (1999) found that value and price were cointegrated historically. One aim of this thesis is to compare this historical period with a later period. The model presented in Equation (3.5) can be written as split into two time periods, the first T and the last m observations:

$$v_t = \begin{cases} \beta_1 p_t - \beta_1 u_{1t}^* - u_{2t}^*, & t = 1,\ldots,T \\ \beta_t p_t - \beta_t u_{1t}^* - u_{2t}^*, & t = T+1,\ldots,T+m \end{cases} \tag{3.8}$$

where β_1 can be considered the stable or equilibrium relation between price and value, with the stationary combination of $(\beta_1 u_{1t}^* - u_{2t}^*) \sim I(0)$. For the Lee et al. (1999) model, the additional restriction is that $\beta_1 = 1$. For the general case, the null and alternative hypotheses are:

$$H_0 : \begin{cases} \beta_1 = \beta_t, \forall t = T+1,\ldots,T+m \text{ and} \\ \{\beta_1 u_{1t}^* - u_{2t}^* : t = 1,\ldots,T+m\} \text{are stationary and ergodic} \end{cases}$$

$$H_1 : \begin{cases} \beta_1 \neq \beta_t, \text{for some } t = T+1,\ldots,T+m \text{ and} \\ \text{the distribution of } \{\beta_t u_{1t}^* - u_{2t}^* : t = T+1,\ldots,T+m\} \text{differs from} \\ \text{the distribution of } \{\beta_1 u_{1t}^* - u_{2t}^* : t = 1,\ldots,T\} \end{cases}$$

Under the null hypothesis (H_0), price and value are cointegrated with the relation being stable through time. Under the alternative (H_1), the cointegrating relation between price and value breaks down, as both the stable relation changes and the distribution of the errors in the second period are different when controlling for the potential regime shift (i.e., given that $\beta_0 \neq \beta_t$, for some $t = T+1,...,T+m$). The null and alternative hypotheses are a modification of the Andrews and Kim (2006) stability tests that the authors refer to as cointegration breakdown tests.

3.4. Measurement of variables

3.4.1. Accounting based valuation models

I consider seven accounting based valuation models to provide estimates of value. I consider the traditional value information provided by dividends (*D*) and book-values (*B*), and complement this with the richer valuations yielded by the residual income model and earnings based models. The residual income model has been implemented using two main approaches, the forecast approach and the linear information dynamic approach. The forecast approach uses the structure of the residual income model to incorporate explicit estimates of future dividends (e.g. Frankel and Lee, 1998; Lee, Myers and Swaminathan, 1999). Using the clean surplus relation, the forecast-based residual income model (V_f) can be used to explicitly forecast future dividends by using the following structural form with *T*–period ahead observations of forecast earnings:

$$Vf(T)_t = b_t + \frac{f(1)_t - r.b_t}{(1+r)} + \frac{f(2)_t - r.b(1)_t}{(1+r)^2} + ... + \frac{f(T)_t - r.b(T-1)_t}{(1+r)^2 r} \tag{3.9}$$

The model can then be collapsed to provide an estimate of the Gordon growth model by using a single forecast of earnings and assuming a perpetual growth rate *g*. Specifically:

$$Vf(1)_t = b_t + \frac{f(1)_t - r.b_t}{r-g}. \tag{3.10}$$

This very simple structure can be used to calibrate multiple measures of value, based on changing r and g. While less sensitive to unusual earnings, the model is similar to an earnings based approach (and equals the earnings capitalization model when g = 0). In implanting this model, I assume r equals the one-year constant yield to maturity treasury bond rate plus an equity premium of 6% and g is set equal to 3%.[32] Book-value, b_t, is the end of year book-value from the most recent fiscal year-end, $f(1)$ is the forecast of earnings and is taken from the I/B/E/S consensus forecast of one-year ahead earnings at the end of each month.

The second approach is to consider the linear information dynamics suggested in Ohlson (1995) to combine the information in earnings and book-value (e.g., Dechow et al., 1999).

$$VI(*)_t = b_t + \alpha_1 x_t^a + \alpha_2 \upsilon_t, \tag{3.11}$$

where $\alpha_1 = \omega/(1+r-\omega)$, $\alpha_2 = (1+r)/[(1+r-\omega)(1+r-\gamma)]$ and υ_t is 'other information.' Equation (3.11) is based on the following assumed linear information dynamics:

$$\begin{aligned} x_{t+1}^a &= \omega x_t^a + v_t + e_{1,t+1} \\ v_{t+1} &= \gamma v_t + e_{2,t+1} \end{aligned} \tag{3.12}$$

When other information (υ_t) is ignored, the model comprises a convex combination of book-value and abnormal earnings. If the persistence of abnormal earnings is zero, then residual income equals book-value. If the persistence of abnormal income is expected to exceed zero but be less than one, then the residual income model, again ignoring other information, can be written as:

$$VI(x)_t = b_t + \frac{\omega_t x_t^a}{(1+r_t-\omega_t)}. \tag{3.13}$$

[32] The impact of changes to assumptions regarding r and g is the focus of the further analysis presented later in this chapter. Changing the forecast horizon, by adding additional forecasts to the model, as in Lee et al. (1999) and Frankel and Lee (1998), requires assumptions regarding payout policy. Similar to these papers, I find only marginal benefit from implementing models with longer horizons.

Assuming 'other information' is zero, however, implies that the conditional expectation of abnormal earnings for t+1 based on all earnings information is equal to the conditional expectation of abnormal earnings based on prior period earnings. An alternative is to consider the following as a measure of other information based on analyst forecasts (Dechow et al., 1999 p7):

$$v_t = f(1)_t - rb_t - \omega_t x_t^a . \tag{3.14}$$

Then the residual income model can be estimated as:

$$Vl(f)_t = b_t + \frac{\omega_t x_t^a}{(1 + r_t - \omega_t)} + \frac{(1 + r_t)v_t}{(1 + r_t - \omega_t)(1 + r_t - \gamma_t)} . \tag{3.15}$$

In all of the models, the ω and γ terms are considered to be constants. Following Dechow et al. (1999) I estimate these terms using a one-period lagged AR(1) model, using the prior year's data for each industry.

The capitalization of forward earnings is also related to the linear information dynamics of Ohlson (1995) and is the result of assuming that abnormal earnings has zero persistence and other information is a random walk. The resulting earnings capitalization model (*Ve*) is:

$$Ve(1)_t = \frac{f(1)_t}{r_t} . \tag{3.16}$$

To the extent that the consensus I/B/E/S forecasts of earnings do not rely on conservative treatments of accounting, then it possible that an EPS based model, such as that proposed by Ohlson and Juettner-Nauroth (2005), may provide a better measure of value. I therefore consider the following implementation, which is similar to Gode and Mohanram (2003):

$$Ve(f)_t = \frac{f(1)_t}{r_t} + \frac{stg_t - r_t(f(1)_t - d(1)_t)}{(r_t - g_t)r_t} , \tag{3.17}$$

where the model expresses value as a function of forecast earnings per share, *feps*, the required rate of return, *r*, short-term growth in earnings, *stg*, and long-term growth, *g*. I model the short-term growth rate, *stg*, using the average of $(feps_2 - feps_1)/feps_1$ and the 'long-term' growth estimate reported by I/B/E/S. I use the same specification of the required rate of return, *r*, and long-term growth, *g*, as for the residual income model.

3.5. *Empirical analysis*

3.5.1. *Data and sample selection*

I measure fundamental value monthly, using the most recent analyst forecasts at the end of the month as expectations of future earnings. I also use a time-varying required rate of return that is based on the addition of time-varying interest rates and a market risk premium as per Lee et al. (1999). I use a sample of publicly traded firms over the period 1979–2004; I do not require that firms have available data for the whole period.[33] Instead, I include firms in the sample that are covered by *CRSP* and *Compustat*, with financial data required to estimate the residual income model and market value (using share price and shares outstanding on *CRSP*).[34] I collect financial data required for the residual income model from the *Compustat* database (annual book-values, earnings and dividends), and analyst forecast data using the mean forecast of earnings per share from the I/B/E/S database. Where I/B/E/S does not cover the firm, historical earnings replace the analyst forecasts. The final sample includes both firms with analyst

[33] Limiting the sample to firms that have available data for the whole period does not qualitatively change the results (see Chapter 5). Smaller firms have more pronounced inefficiencies and growth measurement issues, these firms have a lower weighting in value-weighted tests. Limiting the sample to firms that were covered on CRSP in 1929, or only in the top 1000 or top 50 by market capitalisation, yield qualitatively similar results.

[34] I exclude firms with a negative book-value and negative earnings. Firms with earnings less than the required rate of return are retained, but I set their residual income value equal to book-value. Changing this assumption does not change the tenor of the results. I use I/B/E/S information where available. When it is not available I replace the forecast of earnings with earnings from the most recent fiscal year-end. Limiting the sample to include only firms with analyst coverage produces qualitatively similar results.

coverage and some with no analyst coverage. I aggregate these firms monthly to construct the time-series of price and the time-series of residual income value.

3.5.2. Descriptive statistics

In Table 3.1, I present summary statistics for the sample period. In Panel A, I present descriptive statistics for the market portfolio price and fundamentals. The log of price is on average 15.03 with a standard deviation of 0.85, the fundamental value based on one-period ahead perpetuity of residual income, the log of $Vf(1)$, has a mean of 14.89 and has the closest standard deviation to price at 0.81. I also provide stationarity tests for all of the variables in Panel A, and show that all variables are nonstationary in levels but strongly stationary in first differences, suggesting that they are all $I(1)$ processes.

I present descriptive statistics of the fundamental-to-price ratios in Panel B. The residual income value-to-price ($Vf(1)P$) is defined as the aggregation of month-end residual income value divided by aggregated month-end price. The mean of the value-to-price ratio is 0.890 suggesting aggregate value is on average 89% of aggregate price. I also provide descriptive information for alternative measures of the fundamental to price ratio. The dividend-to-price ratio (DP) is on average 3%, the book-to-price ratio (BP) is on average 62.7%, the two linear-information dynamics based models have an average level of 65.1% ($Vl(x)P$) and 66.3% ($Vl(f)P$), respectively. Finally the two earnings based approaches; the earnings capitalization ($Ve(x)P$) is on average 65.3%, and the earnings growth model is on average 116.5%. Other than the earnings growth model, as all of the ratios are below one, this suggests that the value models underestimate price.

Table 3.1 Summary statistics

Panel A: Descriptive statistics for the aggregate market portfolio ($T = 312$)

					Stationarity tests of logged variables			
					In Levels		In Changes	
Variable	Mean	Std. Dev.	Min	Max	Z–rho (Z_ρ)	Z–tau (Z_τ)	Z–rho (Z_ρ)	Z–tau (Z_τ)
p	15.03	0.85	13.53	16.42	–0.7358	–0.78	–297.2*	–17.10*
e	12.41	0.52	11.61	13.52	–0.8268	–1.06	–283.7*	–21.32*
b	14.53	0.60	13.52	15.66	–0.9371	–0.48	–292.3*	–18.71*
d	11.47	0.53	10.47	12.30	0.0753	0.10	–282.6*	–18.84*
vf(1)	14.89	0.81	13.59	16.52	0.1947	0.28	–251.7*	–14.48*
vl(*x*)	14.57	0.62	13.53	15.75	0.0941	0.12	–284.9*	–18.73*
vl(*f*)	14.59	0.62	13.54	15.76	0.1037	0.14	–275.9*	–18.08*
ve(1)	14.58	0.75	13.33	16.06	0.0184	0.02	–250.4*	–15.06*
ve(*f*)	15.12	1.00	13.40	17.15	–0.5139	–0.43	–267.6*	–13.69*

Panel B: Descriptive statistics of fundamental-to-price ratios ($T = 312$)

					Autocorrelation at lag					
Variable	Mean	Std. Dev.	Min	Max	1	12	24	36	48	60
DP	0.030	0.010	0.012	0.050	0.979	0.816	0.748	0.648	0.506	0.384
BP	0.627	0.174	0.298	1.056	0.974	0.778	0.714	0.610	0.458	0.352
Vf(1)*P*	0.890	0.190	0.488	1.450	0.966	0.590	0.217	–0.110	–0.300	–0.323
Vl(*x*)*P*	0.651	0.170	0.317	1.076	0.972	0.764	0.698	0.592	0.436	0.329
Vl(*f*)*P*	0.663	0.169	0.328	1.080	0.972	0.757	0.691	0.583	0.427	0.320
Ve(1)*P*	0.653	0.135	0.324	1.056	0.951	0.554	0.285	0.092	–0.097	–0.119
Ve(*f*)*P*	1.165	0.488	0.555	2.835	0.981	0.664	0.168	–0.164	–0.228	–0.180

Panel C: Correlation among fundamental-to-price ratios ($T = 312$)

	DP	*BM*	*Vf*(*1*)*P*	*Vl*(*x*)*p*	*Vl*(*f*)*p*	*Ve*(*1*)*p*	*Ve*(*f*)*p*
DP	1						
BP	0.8795*	1					
Vf(*1*)*P*	0.3333*	0.2945*	1				
Vl(*x*)*P*	0.9441*	0.923*	0.5323*	1			
Vl(*f*)*P*	0.9461*	0.9184*	0.5396*	0.9996*	1		
Ve(*1*)*P*	0.6012*	0.5289*	0.9204*	0.7388*	0.7476*	1	
Ve(*f*)*P*	–0.245*	–0.297*	0.7969*	–0.054	–0.048	0.5627*	1

(Table 3.1 continued on the following page…)

Panel D: Descriptive statistics for fundamental-to-price ratios by period

(i) Historical period (January 1979 – December 1991, T = 156)

					Autocorrelation at lag					
Variable	Mean	Std. Dev.	Min	Max	1	12	24	36	48	60
DP	0.038	0.005	0.027	0.050	0.888	0.173	0.251	0.153	−0.032	−0.072
BP	0.767	0.116	0.533	1.056	0.918	0.312	0.373	0.266	0.058	−0.080
Vf(1)*P*	0.909	0.102	0.683	1.236	0.864	−0.099	0.136	0.180	0.019	−0.069
Vl(*x*)*P*	0.786	0.114	0.550	1.076	0.915	0.281	0.361	0.266	0.057	−0.082
Vl(*f*)*P*	0.795	0.113	0.560	1.080	0.914	0.260	0.349	0.256	0.061	−0.074
Ve(1)*P*	0.701	0.095	0.535	1.056	0.852	−0.045	0.018	0.155	−0.034	−0.001
Ve(*f*)*P*	0.976	0.156	0.610	1.453	0.826	0.008	−0.221	−0.076	0.050	0.045

(ii) Potential bubble period (January 1992 – December 2004, T = 156)

					Autocorrelation at lag					
Variable	Mean	Std. Dev.	Min	Max	1	12	24	36	48	60
DP	0.022	0.006	0.012	0.033	0.974	0.680	0.481	0.209	−0.020	−0.269
BP	0.487	0.090	0.298	0.638	0.960	0.636	0.314	−0.072	−0.379	−0.500
Vf(1)*P*	0.872	0.247	0.488	1.450	0.983	0.699	0.206	−0.188	−0.384	−0.396
Vl(*x*)*P*	0.518	0.095	0.317	0.683	0.960	0.634	0.287	−0.103	−0.405	−0.506
Vl(*f*)*P*	0.532	0.097	0.328	0.696	0.963	0.648	0.298	−0.101	−0.405	−0.512
Ve(1)*P*	0.604	0.151	0.324	0.928	0.980	0.681	0.189	−0.206	−0.407	−0.425
Ve(*f*)*P*	1.354	0.617	0.555	2.835	0.988	0.693	0.178	−0.157	−0.255	−0.264

Notes: In panel A, I report the log of the aggregate price (p), dividends (d), earnings (e), book-values (b), and accounting based value estimates where; $Vf(1)$ is the one-period forecast-based residual income model, $Vl(x)$ is the residual income model based on the linear information dynamics of Ohlson (1995), $Vl(f)$ is the linear information dynamics model with analyst forecasts as 'other information' from Dechow et al. (1999), $Vl(e)$ is the earnings capitalization, and $Ve(f)$ is the Ohlson-Juettner model of earnings growth. The stationarity estimates are performed using the Phillips-Perron estimates, variables that are nonstationary in levels and stationary in changes are denoted $I(1)$. In Panels B–D, the variables are not logged, i.e., they are raw ratios of the fundamental-to-price. All variables are measured monthly. T is the sample size in monthly observations. $^*p<0.01$, $^{\delta}p<0.1$

Panel B also shows that all of the first-order autocorrelations for the fundamental ratios are high, suggesting they are either nonstationary or slowly mean-reverting. The first-order autocorrelation coefficient reported in Table 3.1 for the value-to-price ratio of 0.966 is marginally higher than that reported in Lee et al. (1999) at 0.93. With the first-order autocorrelation close to one, formal statistical tests are required to determine if the value-to-price ratio is nonstationary.

In Panel C, I present correlations between the fundamental-to-price ratios. As is expected most of the ratios are significantly and positively associated. The one exception is the earnings growth model (*Ve(f)P*) which has negative correlations with the dividend-to-price ratio and the book-to-market ratio. This is possibly due to the construction of the model, which requires positive expected growth.

It is likely that in the recent speculative period prices diverged more from accounting based valuation. In Panel D, I split the sample into two equal periods, 1979–1991 and 1992–2004. Excluding the earnings growth model,[35] the means of the fundamental-to-price ratios are all lower in the 1992–2004 period than in the 1979–1991 period. The mean of the value-to-price ratio (*Vf*(1)*P*) decreases from 90.9% to 87.2%. In addition, the value-to-price ratio was less stable in the more recent period, with a large increase in the standard deviation and both a lower minimum and a higher maximum.

3.5.3. Graphical display of the time-series variation in price and fundamentals

In Figure 3.1, I present the time-series variation in the relation between value and price. Graphed in its inverted form, (i.e., price-to-value) the dramatic increase in price relative to value in the late 1990s is evident. In addition, prior to 1996, the ratio regularly fluctuated around its

[35] The earnings growth model is significantly higher than prices following the crash.

long-run mean, which is represented as a horizontal dotted line. At the end of the Lee et al. (1999) time-period (June 1996), the value-to-price ratio (coincidentally) is roughly equal to the long run mean. The extent to which the ratio increases from June 1996 until March 2000 is unprecedented in this time-series, as is the steepness of the decline in the value-to-price ratio starting at April 2001. The magnitude and duration of the deviation of the value-to-price ratio from its long-run mean is consistent with the belief that, during this period, prices moved away from fundamental values.

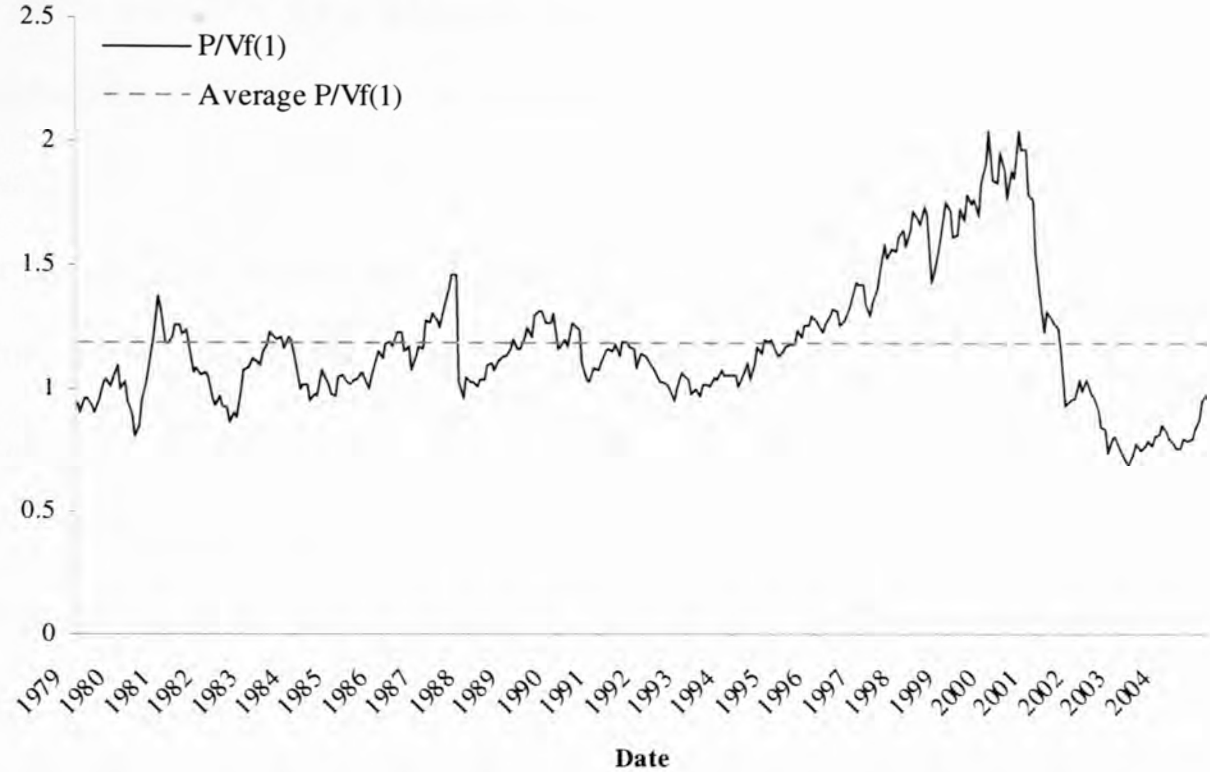

Figure 3.1 The time-series behaviour of the ratio of price-to-value

This graph displays the ratio of price (P) to value (Vf(1)), where value is measured using the sum of book-value and a perpetuity of one-period ahead residual income value growing at 3%. The ratio is measured at the value-weighted portfolio level and includes all firms with sufficient data during the period 1979 to 2004. The horizontal dashed line displays the sample mean of the ratio which is equal to 1.178.

In Figure 3.2, I provide further evidence of a movement of price away from fundamentals during the late 1990s. By graphing the time-series of price and value separately, I provide insight regarding the properties of the series underlying the ratio. Prior to 1995, both time-series drift upwards at approximately the same rate. After 1995, the rate of growth in price increases sharply, with price diverging substantially from value. The deviation between price and value continues for approximately five years before the trend in price reverses and rapidly returns toward value. The growth in value appears to approximate a continuation of its historical growth. In the early 2000s, the effect of low interest rates on the value of residual income is clearly visible on the level of value, which exceeds price (a 6% equity premium, however, may understate the equity premium following the prior turbulence during the late 1990s).

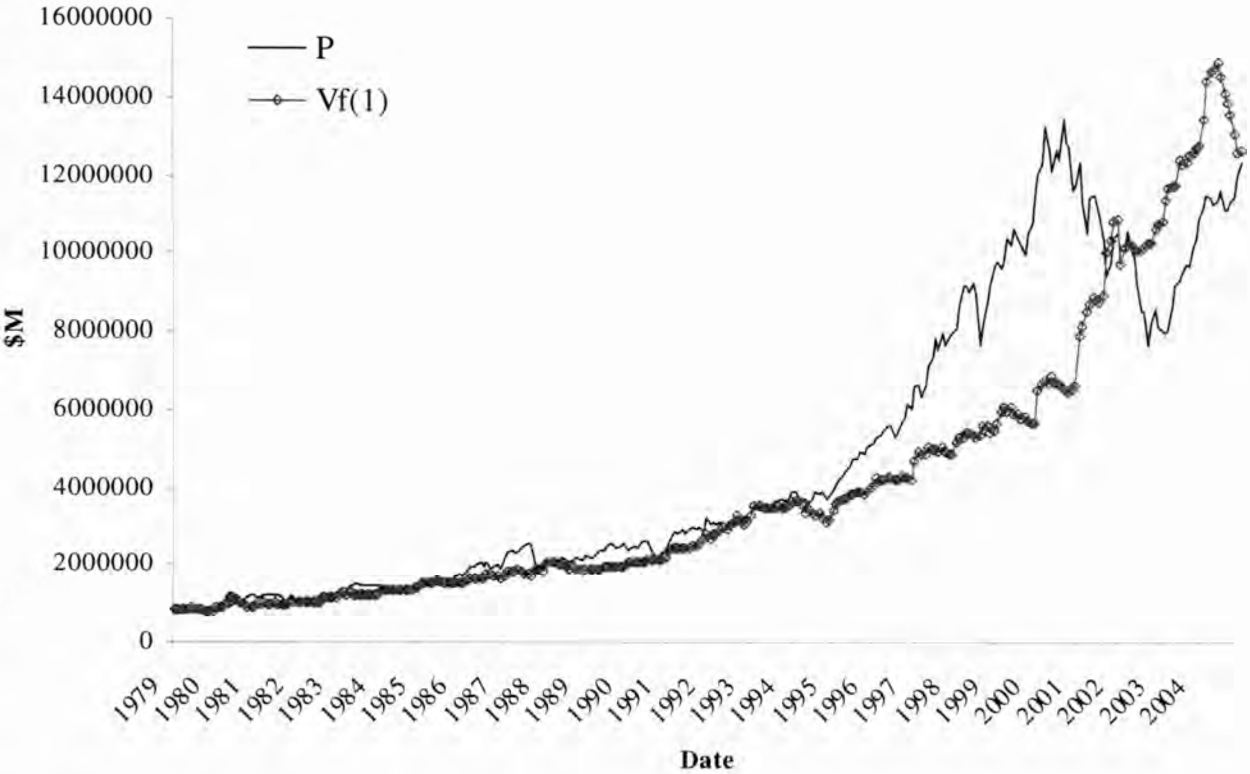

Figure 3.2 The time-series behaviour of price and value

This graph displays the time-series of aggregate market value (P) and the aggregate accounting based measure of value (Vf(1)), where value is measured using the sum of book-value and a perpetuity of one-period ahead residual income value growing at 3%. The time-series are measured at the portfolio level in millions of US Dollars and includes all firms with sufficient data during the period 1979 to 2004.

3.5.4. Testing for stationarity in the log value-to-price ratio

Testing for stationarity in the log ratio of fundamental value-to-price ratio (vp_t) provides a statistical test for cointegration, under the restriction that the cointegrating parameter is one. I test whether the time-series of the vp_t is stationary using the Phillips-Perron unit-root tests (Hamilton, 1994). The null hypothesis for these tests is that the series contains a unit root. Formally, I test for both a unit root with a stochastic drift and the alternative of stationarity around a time-trend.[36] I use the Phillips-Perron test instead of the augmented Dickey-Fuller test because the former test corrects for heteroscedastic and serially correlated residuals. The tests are conducted using up to 12 lags,[37] and are specified as:

$$\Delta vp_t = a + (b-1)vp_{t-1} + \sum_{i=1}^{12} c_i \Delta vp_{t-i} + u_t, \tag{3.18}$$

$$\Delta vp_t = a + (b-1)vp_{t-1} + \delta t + \sum_{i=1}^{12} c_i \Delta vp_{t-i} + u_t, \tag{3.19}$$

where Equation (3.18) includes a non-zero mean, a, and excludes a time-trend and Equation (3.19) includes both a non-zero mean, a, and a time-trend, δt. These equations test for cointegration between price and value, under the restriction that the cointegrating vector is $\mathbf{a}=[1,-1]'$ (from Equation 3.4). The test is against the null of a unit root in the time-series (b=1).

I report the results of the Phillips-Perron tests of stationarity for $vf(1)p$ in Table 3.2, along with tests of the alternative fundamental-to-price ratios. In Panel A, I present the tests with an intercept but without a time-trend; in Panel B the tests include a time-trend. For the monthly time-series between 1979 and 2004, the null hypothesis of no cointegration cannot be rejected for $vf(1)p$. This result is consistent with either a breakdown in cointegration subsequent to the

[36] I do not consider Phillips-Perron tests for a unit root without an intercept as the means of the time-series considered are non-zero.

[37] Tests on smaller lags give similar results.

Lee et al. (1999) sample period, or the inclusion of additional firms biasing the test towards the null of no cointegration. I address the concern of including additional firms by presenting the tests for two equal sub-periods, 1979–1991 and 1992–2004. The first period overlaps with that of Lee et al. (1999) and the subsequent period includes the potential bubble period. I confirm the existence of cointegration between price and value during the earlier time-period (1979–1991). Examining the effect of adding the late 1990s, using the period 1992–2004, I find that there is no reliable evidence of cointegration.

In Panel B, I report stationarity tests that include a time-trend term. It is not surprising, given the prominent upward trend displayed graphically in Figure 3.1, that the test statistics are generally higher for these tests. Interpretation of these tests, however, requires caution, as the inclusion of a time-trend allows for the possibility that price and value are diverging from each other over the sample period but can still be statistically cointegrated. The results are generally consistent with the results in Panel A, although for some of the fundamental-to-price ratios there is some weak evidence of cointegration over the full sample period. Again, there is no reliable evidence of cointegration between price and value for the period 1992–2004, which includes the potential bubble period.

To summarize the results outlined in this section, the finding of cointegration between price and value is robust to the inclusion of additional firms outside the Dow 30, during the time-period 1979–1992, consistent with Lee et al. (1999). After 1992, and in particular from 1995 to at least 2000, the value-to-price ratio is nonstationary, and I therefore cannot reject the null of no cointegration using the ratio based approach.

Table 3.2 Tests for cointegration for fundamental-to-price ratios

Panel A: Phillips-Perron tests without time-trend

	(i) 1979 – 2004 T=312		(ii) 1979 – 1991 T=156		(iii) 1992 – 2004 T=156	
	Z-rho (Z_ρ)	Z-tau (Z_τ)	Z-rho (Z_ρ)	Z-tau (Z_τ)	Z-rho (Z_ρ)	Z-tau (Z_τ)
dp	–4.17	–1.53	–15.47*	–2.84*	–3.24	–1.61
bm	–6.71	–2.11	–10.08	–2.27	–5.32	–1.81
vf(1)p	–10.93	–2.37	–22.91*	–3.55*	–2.69	–1.12
vl(x)p	–7.40	–2.20	–10.92	–2.37	–5.42	–1.79
vl(f)p	–7.47	–2.20	–11.13$^{\delta}$	–2.39	–4.96	–1.71
ve(1)p	–15.01*	–2.87$^{\delta}$	–22.86*	–3.57*	–3.25	–1.25
ve(f)p	–6.21	–1.72	–25.48*	–3.53*	–2.62	–1.13

Panel B: Phillips-Perron tests including a time-trend

	(i) 1979 – 2004 T=312		(ii) 1979 – 1991 T=156		(iii) 1992 – 2004 T=156	
	Z-rho (Z_ρ)	Z-tau (Z_τ)	Z-rho (Z_ρ)	Z-tau (Z_τ)	Z-rho (Z_ρ)	Z-tau (Z_τ)
dp	–24.04*	–3.50*	–23.12*	–3.48*	–4.79	–1.54
bm	–20.39$^{\delta}$	–3.18$^{\delta}$	–23.07*	–3.46*	–4.87	–1.61
vf(1)p	–10.71	–2.32	–28.68*	–3.90*	–3.26	–1.30
vl(x)p	–19.91$^{\delta}$	–3.15$^{\delta}$	–23.76*	–3.52*	–5.00	–1.64
vl(f)p	–19.97$^{\delta}$	–3.15$^{\delta}$	–23.68*	–3.51*	–4.54	–1.54
ve(1)p	–16.46	–2.89	–29.73*	–4.01*	–3.32	–1.29
ve(f)p	–8.63	–3.45	–25.92*	–3.59*	–2.10	–1.35

Notes: This table summarises the results of the Phillips-Perron unit root tests on variables used in this study. The unit root tests are performed without a time-trend in Panel A and with a time-trend in Panel B: $\Delta vp_t = a + (b-1)vp_{t-1} + \sum_{i=1}^{12} c_i \Delta vp_{t-i} + u_t$, $\Delta vp_t = a + (b-1)vp_{t-1} + \delta t + \sum_{i=1}^{12} c_i \Delta vp_{t-i} + u_t$,

where, vp_t refers to the log value-to-price ratio. The test is against the null of a unit root in the time-series (b=1). The two test statistics from the Phillips-Perron test are an adjusted regression coefficient Z-rho (Z_ρ) and an adjusted t-statistic Z-tau (Z_τ). Additional tests are provided for the following variables; bp_t refers to the book-to-price ratio, ep_t refers to the earnings-to-price ratio, dp_t refers to the dividend-to-price ratio, Δv_t refers to the monthly change in value, Δp_t refers to the monthly change in price.

**p<0.001, *p<0.05, $^{\delta}p$<0.1

3.5.5. *Cointegration tests between log intrinsic value and log price time-series*

In the prior section, I found evidence that supports the proposition that price and value diverged for an extended period. In this section, I decompose the value-to-price ratio to facilitate further understanding of why the time-series relation in the late 1990s was distinct from the earlier period. Equation (3.6) can be written compactly in matrix form (see, Greene, 2003, pp586–588) as:

$$\Theta y_t = \alpha + \Psi y_{t-1} + \omega_t , \tag{3.20}$$

where $y_t = (v_t, p_t)'$, Θ is a matrix of contemporaneous parameters with diagonal elements equal to one and off-diagonals being the inverse slope coefficients that relate to the elements of y_t, and Ψ is a matrix of autoregressive coefficients. Expressing Equation (3.20) as $y_t = \Gamma y_{t-1} + \omega_t$ (where $\Gamma = \Theta^{-1}\Psi$) and taking first differences produces a vector error correction model (see, Greene, 2003). The general VECM model considered in this case is written as:

$$\Delta y_t = \Pi y_{t-1} + \sum_{i=1}^{n-1} \Phi_i \Delta y_{t-i} + \varepsilon_t , \tag{3.21}$$

where $y_t = (v_t, p_t)'$ and $\Pi = \alpha(\beta_1, \beta_0)$.[38] Including k variables in the matrix y, the rank of the matrix Π provides the following information: (1) if Π has full rank ($r = k$), all components of y_t are $I(0)$ stationary, or (2) if the rank of the matrix is $r<k$ then there are r stationary cointegrating relations. Where value and price are cointegrated there will be a single linear combination such that the reduced rank of the matrix equals one. Following the Johansen procedure, the null hypothesis of $r = 0$ would be rejected and the null of $r \leq 1$ would not be rejected.

[38] The model is normalised for v_t and the cointegrating vector is represented as a = [-1,-β_0 -β_1].

In Table 3.3, I present both the estimation of the VAR[39] in levels model of Equation (3.20) and the related tests for cointegration and the cointegrating parameters of the VECM model of Equation (3.21) for the *vf*(1) model of value. In Panel A, I show evidence of a relation between price and lagged value during the 1979–1991 period, which is not reliably observed in the 1992–2004. Instead, during the 1992–2004 period the relation between price and lagged price increases from 0.852 to 0.991. From Panel B: (1) I find evidence to reject the null of no cointegration for the 1979–1991 period (r=1), and (2) I do not find evidence to reject the null of no cointegration for the 1992–2004 period (r=0).

Within a set of cointegrated variables at least one will respond to the disequilibrium in the system. In Panel B, I report the estimation of the VECM model; the additional parameters of this model provide insights into why the cointegration appears to break down. The evidence in Panel B of Table 3.3 during the 1979–1991 period suggests price responds to the disparity from the long-run association between price and value. The response parameter for price, α_p, is 0.170 and is statistically significant at the 5% level. The response parameter for value α_v, which is −0.014, is not statistically significant at conventional levels. The statistically significant coefficient on price is evidence of an 'anchoring' of price to fundamental value. The evidence for the 1992–2004 suggests that the same anchoring mechanism was not present during the speculative period. One interpretation is that the behaviour of price changed causing a prolonged disparity between price and value. The estimated long-run relation between price and value is 0.87 with a 95% confidence interval 0.985 to 0.761, which rejects the ratio-based cointegrating vector of **a**=[1,–1]' at the 5% level. The constant term is small and negative (–1.67), in contrast

[39] The system is estimated using a single lag of price and value. Statistical tests for the selection of the lag length can be found following the minimum Schwarz Bayesian Information Criterion (BIC). The selection of a single lag was determined using the minimum BIC following a simulation including up to 12 lags.

with its value in the 1992–2004 (10.80), consistent with a considerable divergence of price and value during the potential bubble period.

This section presents a complementary test of cointegration between price and value. While the possibility that measurement error in value drives the divergence between price and value during the period 1992–2004 cannot be ruled out, I find evidence consistent with a significant price correction, consistent with fundamental value historically providing an anchor to price. During the late 1990s, however, there is no evidence of anchoring. Rather there is evidence of increased price momentum, one interpretation of which is that there was additional speculation in price.

Table 3.3 Tests of the time-series behaviour of value and price

Panel A: Vector–autoregressive model of price and value

	(i) 1979 – 2004 $T = 312$		(ii) 1979 – 1991 $T = 156$		(iii) 1992 – 2004 $T = 156$	
	p_t	v_t	p_t	v_t	p_t	v_t
(a_{10}, a_{20})	0.019	0.001	–0.274	0.055	0.125	–0.004
$(a_{11}, a_{21})\ p_{t-1}$	0.981**	0.010	0.852**	0.012	0.991**	0.010
$(a_{12}, a_{22})\ v_{t-1}$	0.019	0.990**	0.169**	0.985**	0.001	0.990**

Panel B: Vector-correction model and cointegration tests between price and value

	(i) 1979 – 2004 $T = 312$		(ii) 1979 – 1991 $T = 156$		(iii) 1992 – 2004 $T = 156$	
Long–run parameters	β_0	β_1	β_0	β_1	β_0	β_1
	1.08	1.05	–1.67	0.87	10.80	1.67
Error–correction parameters	α_p	α_v	α_p	α_v	α_p	α_v
	0.019^{δ}	–0.010	0.170*	–0.014	0.007	–0.005
Cointegration test statistics	λ_{eigval}	λ_{trace}	λ_{eigval}	λ_{trace}	λ_{eigval}	λ_{trace}
$\hat{\lambda}$ (r>0)	4.88	4.88	16.16*	16.20*	1.88	2.46
$\hat{\lambda}$ (r>1)	0.001	0.001	0.044	0.044	0.58	0.58

Notes: This table summarises the results of tests for cointegration between price and value based on the bilinear autoregressive model presented as Equation (3.20) and Equation (3.21). In Panel A, the parameter estimates of the following VAR model are presented: $y_t = a + \Psi y_{t-1} + u_t$, where $y_t = (v_t, p_t)'$. Following the Johansen method, this model is differenced, allowing for the identification of the cointegrating parameters and the error correction terms. In Panel B, the following model is estimated: $\Delta y_t = \alpha(\beta_1', \beta_0)(y_{t-1}', 1) + \Phi_1 \Delta y_{t-1} + \varepsilon_t$, where, the vector α contains the adjustment parameters and the β parameters are the cointegrating parameters (β_0 is the intercept). If the matrix $\alpha(\beta', \beta_0)$ has full rank ($r = k$), all k components of y are I(0) stationary. Alternatively, if the rank of the matrix is $r<k$ then there are r stationary cointegrating relations. In this case, $k = 2$, so price and value are cointegrated where the null hypothesis of r=0 is rejected. In all cases the null hypothesis of $r \leq 1$ is not rejected. In Panel B, the maximum eigenvalue test and the trace test are reported, along with their 95% critical values.

** p<0.001, * p<0.05, $^{\delta}$ p<0.1

3.6. Further analysis

3.6.1. Measurement of expected rates of return

The results in the prior section are consistent with a breakdown in the late 1990s of the cointegrating relation between value and price that had been found previously. In this section, I investigate the possibility that a large decrease in the price of risk contributed to the divergence between value and price. To estimate required rates of return in the prior tests of cointegration between price and value, I used the 3-year constant yield to maturity T-Bill plus a constant equity risk premium of 6%. If the average risk premium during the potential bubble period was much lower than 6%, then the measure of value would be understated during the 1992–2004 period.

To address this potential measurement error, I present stationarity tests of the value-to-price ratio when estimating value using different expected rates of return. First, I report estimates based on constant expected rates of return equal to 6% and 12%. Second, I report estimates based on the 3-year constant yield to maturity T-Bill plus a constant equity risk premium that ranges from 0% to 5%.

The estimates are reported in Panel A of Table 3.4. The constant expected returns models, using 6% and 12%, are not cointegrated in any of the periods, suggesting a constant discount rate may not be appropriate. Furthermore, the results reported earlier using time-varying interest rates plus a risk-premium of 6% are robust to the assumption of lower risk premium. Specifically, there is strong support for cointegration between price and value during the period 1979–1992 for all levels of the risk premium and there is no reliable evidence that price and value are cointegrated during the period 1992–2004 for any of them.

Table 3.4 Further analysis of expected rates of return

Panel A: Phillips-Perron tests with alternative values for the equity risk premium

	(i) 1979 – 2004 T=312		(ii) 1979 – 1991 T=156		(iii) 1992 – 2004 T=156	
	Z-rho (Z_ρ)	Z-tau (Z_τ)	Z-rho (Z_ρ)	Z-tau (Z_τ)	Z-rho (Z_ρ)	Z-tau (Z_τ)
vf(1)p Constant R						
6%	–3.72	–1.65	–2.85	–0.96	–3.99	–1.40
12%	–3.72	–1.65	–15.74	–2.57	–2.85	–0.96
vf(1)p Varying R						
Tbill + 0%	–6.67	–1.71	–24.61*	–3.61*	–2.34	–1.01
Tbill + 1%	–9.05	–2.05	–25.77*	–3.76*	–2.33	–0.99
Tbill + 2%	–10.87	–2.31	–25.36*	–3.77*	–2.45	–1.01
Tbill + 3%	-11.6^{δ}	–2.45	–23.99*	–3.67*	–2.60	–1.04
Tbill + 4%	-11.5^{δ}	–2.48	–22.16*	–3.52*	–2.75	–1.08
Tbill + 5%	–10.97	–2.47	–20.22*	–3.35*	–2.88	–1.11

Panel B: Phillips-Perron tests for firms sorted into portfolios by their market beta

	(i) 1979 – 2004 T=312		(ii) 1979 – 1991 T=156		(iii) 1992 – 2004 T=156	
	Z-rho (Z_ρ)	Z-tau (Z_τ)	Z-rho (Z_ρ)	Z-tau (Z_τ)	Z-rho (Z_ρ)	Z-tau (Z_τ)
vf(1)p						
Q1 – Low Beta	–18.21*	–3.05*	–9.15	–2.12	-12.65^{δ}	–2.57
Q2	–32.98*	–4.12*	–18.65*	–3.10*	-14.89^{δ}	-2.78^{δ}
Q3	–20.63*	–3.30*	–34.21*	–4.49*	–6.83	–1.84
Q4	-12.97^{δ}	–2.56	–25.86*	–3.68*	–3.44	–1.31
Q5 – High Beta	–10.64	–2.30	–19.16*	–3.11*	–2.72	–1.07

Notes: This table summarises the results of the Phillips-Perron unit root tests on variables used in this study. The unit root tests are performed without a time-trend: $\Delta vp_t = a + (b-1)vp_{t-1} + \sum_{i=1}^{12} c_i \Delta vp_{t-i} + u_t$, where, $vf(1)p$ refers to the log value-to-price ratio and is calculated as $b_t + (f(1) - r_i b_t)/(r_i + g)$, where r_i is changed for each of the models in Panel A, and equals the treasury bill rate plus 6% in Panel B. Beta is estimated using the association between the firm's return and the value-weighted index on CRSP, both less the risk-free (T-bill) rate and estimated over the prior 12 months using monthly data. Firms are assigned to beta quintile portfolios annually in January. In both panels, the test is of the null of a unit root in the time-series (b=1). The two test statistics from the Phillips-Perron test are an adjusted regression coefficient Z-rho (Z_ρ) and an adjusted t-statistic Z-tau (Z_τ).

$^{**}p<0.001$, $^{*}p<0.05$, $^{\delta}p<0.1$

In Panel B, I report a complementary test for potential measurement error in the expected rate of return by forming quintile portfolios of value and price based on the level of the firm's beta. I estimate beta over the prior 12 months, and assign firms annually to their quintile portfolios in January of each year. If the lack of cointegration during the potential bubble period is due to the misspecification of firm-specific rates of return, then by aggregating firms into portfolios of similar risk, only the extreme portfolios (i.e. high and low beta firms) will not reject the null of no cointegration. Estimating these portfolios separately, therefore, will allow the middle portfolios (which have firms with market betas closer to one) to reject the null of no cointegration. When examining the potential bubble period, I do not find this expected pattern generally, with only some very weak evidence that low beta firms are cointegrated during the 1992–2004 period. Thus risk differences do not appear to have any role.

3.6.2. Measurement of expected growth

In this section, I present tests to examine whether the divergence between price and value is due solely to the potential measurement error in expected growth. In both the residual income model and the earnings capitalisation model, I assume that growth rates for all firms converge to 3% in the long-run. In these models, the information in analysts' forecasts of long-term growth in earnings is ignored. If firm-specific differences in the relation between price and value are due to a failure to account for growth, then sorting firms into portfolios based on analysts' forecasts of expected growth will provide insights into the effect of growth on the disparity between price and value.

In Table 3.5, I report tests for potential measurement error in expected growth by forming quintile portfolios of value and price based on the level of the firm's consensus analyst forecasts of long-term growth in earnings per share. The first portfolio includes firms with no analyst forecasts of growth. In the other portfolios, I assign firms annually to their quintile portfolios in January of each year. If the lack of cointegration during the potential bubble period is due to the misspecification of firm-specific growth, then by aggregating firms into growth based portfolios, it is likely that only low growth firms will show evidence of cointegration during this period.

In Panel A I provide estimates of the residual income model and in Panel B I present estimates of the earnings capitalisation model. The results in both panels are remarkably consistent. Specifically, there is evidence of cointegration for all portfolios during the period 1979–1991 (except for the firms with no analyst estimates of growth) but no reliable evidence of cointegration between price and value for any of the portfolios during the period 1992–2004.

While it is possible that price diverged from fundamentals due to the combination of low discount rates and high growth expectations, the results reported in this section do not support either of these explanations.

Table 3.5 Further analysis of growth expectations

Panel A: Phillips-Perron tests for firms sorted into portfolios by their long-term growth

	(i) 1979 – 2004 T=312		(ii) 1979 – 1991 T=156		(iii) 1992 – 2004 T=156	
	Z-rho (Z_ρ)	Z-tau (Z_τ)	Z-rho (Z_ρ)	Z-tau (Z_τ)	Z-rho (Z_ρ)	Z-tau (Z_τ)
vf(1)p						
Q1 – No Growth	−9.23	−2.31	−7.38	−2.11	−3.34	−1.25
Q2 – Low Growth	−6.83	−1.88	−12.14$^\delta$	−2.68$^\delta$	−2.13	−0.95
Q3	−6.50	−1.74	−14.42*	−2.92*	−1.84	−0.85
Q4	−11.57$^\delta$	−2.42	−19.18*	−3.25*	−4.72	−1.52
Q5 – High Growth	−11.03	−2.31	−17.73*	−3.05*	−4.70	−1.46

Panel B: Phillips-Perron tests for firms sorted into portfolios by their prior year change in earnings

	(i) 1979 – 2004 T=312		(ii) 1979 – 1991 T=156		(iii) 1992 – 2004 T=156	
	Z-rho (Z_ρ)	Z-tau (Z_τ)	Z-rho (Z_ρ)	Z-tau (Z_τ)	Z-rho (Z_ρ)	Z-tau (Z_τ)
ve(x)p						
Q1 – No Growth	−12.05	−2.55	−14.58*	−2.80$^\delta$	−3.85	−1.35
Q2 – Low Growth	−7.84	−2.01	−21.25*	−3.48*	−2.57	−1.10
Q3	−6.71	−1.77	−22.17*	−3.47*	−1.88	−0.86
Q4	−11.26	−2.40	−30.42*	−4.13*	−4.48	−1.47
Q5 – High Growth	−12.40$^\delta$	−2.54	−21.00*	−3.33*	−6.47	−1.78

Notes: This table summarises the results of the Phillips-Perron unit root tests on variables used in this study. The unit root tests are performed without a time-trend in Panel A and with a time-trend in Panel B: $\Delta vp_t = a + (b-1)vp_{t-1} + \sum_{i=1}^{12} c_i \Delta vp_{t-i} + u_t$, where, vp_t refers to the log value-to-price ratio. The test is of the null of a unit root in the time-series (b=1). The two test statistics from the Phillips-Perron test are an adjusted regression coefficient Z-rho (Z_ρ) and an adjusted t-statistic Z-tau (Z_τ).
$^{**}p<0.001$, $^{*}p<0.05$, $^{\delta}p<0.1$

3.7. Conclusions

Until the mid 1990s, the relation between price and value had been one of cointegration, implying that price is anchored to the fundamentals of the firm and mispricing is short-lived (Lee et al., 1999). I investigate whether or not this cointegrating relation was still present in the late 1990s, as there is ongoing debate regarding the existence of an asset price bubble during this period. Using value-weighted tests, I find that cointegration between price and value was not reliably observed for the period 1979–2004 when including the late 1990s. The result is due to the breakdown of the historical tendency of price to correct for any deviation of price from value, which implies that the market sustained a significant deviation of price from value for an extended period. The results are robust to alternative measures of value, and are not explained by varying the equity risk premium or forecast growth rates.

The results are consistent with price historically being 'anchored' to value, with accounting information providing a challenge to speculative beliefs (e.g., Penman, 2003). In the 1990s market, however, my results suggest that speculation caused price to wander arbitrarily high (e.g., Abreu and Brunnermeier, 2003). These results challenge the traditional risk based explanation for the substantial run-up in asset prices in the 1990s; in particular, my results suggest that asset prices can sustain substantial deviations from value for an extended period. In the following chapter, I investigate whether limits to arbitrage help explain this phenomenon.

4. The duration of the disparity between price and value: Tests of the arbitrage cost explanation[40]

4.1. Synopsis

In this chapter, I compare the time-series relation between price and accounting based value (hereafter, value) for portfolios of firms based on the level of arbitrage cost, where high arbitrage cost firms are those with a lack of close substitutes. I find some evidence consistent with the disparity between value and price being associated with arbitrage costs. First, for relatively overpriced firms, the level of overpricing is higher for firms that lack close substitutes; there is no significant difference for firms that are relatively underpriced. Second, the duration of the disparity between value and price is on average longer for high arbitrage cost firms. These results are consistent for the historical and the potential bubble period, though the size of the disparity is much greater for high arbitrage cost firms during the bubble period. The results suggest that arbitrage costs influence the extent to which accounting information is reflected in price.

4.2. Introduction and motivation

Arbitrage plays a crucial role in the analysis of the relation between price and value relevant accounting information – its effect is to align price with intrinsic value. Traditionally, arbitrage activity has been assumed to be without cost or risk. Transaction costs aside, an arbitrage trade will only be without risk when a perfect substitute for the mispriced asset exists.[41]

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[40] This chapter is a revised (and adapted) version of a paper which has previously been presented under the title "Costly arbitrage and the convergence of price to accounting fundamentals" at the University of Utah and New York University.

[41] See Shleifer and Vishny (1997) and Abreu and Brunnermeier (2002) for a discussion of limits to arbitrage in the equity markets. Papers that have examined the lack of close substitutes, which I will discuss in the following section,

As markets are incomplete, would-be arbitrageurs will rarely be able to fully hedge their arbitrage positions, and will require a premium for holding unhedged risk. The more difficult it is for the arbitrageur to find a close substitute, the higher the premium required to induce arbitrage. The largest disparities between price and value, therefore, should be found in those firms which are the most difficult to arbitrage.

Recently, Abreu and Brunnermeier (2002) proposed that when arbitrageurs are capital constrained, the actions of a single arbitrageur will not always be sufficient to correct mispricing. With uncertainty as to when sufficiently many arbitrageurs will exploit a common arbitrage opportunity, and thereby correct mispricing, arbitrage activity by individual arbitrageurs can be delayed. In addition, the authors suggest that as arbitrageurs are risk-averse they will not only attempt to minimize the period that they hold a risky arbitrage position, but their delay in taking positions is increasing in the risk associated with the arbitrage position.

The late 1990s provides an interesting setting to empirically evaluate this model, as many traders believed that the prices of firms during this period were over-valued, but did not trade accordingly (Ofek and Richardson, 2002; Brunnermeier and Nagal, 2004). To investigate the role of limits to arbitrage, I use accounting based value as a summary measure of (value-relevant) accounting information. I use idiosyncratic risk from the market model to measure arbitrage cost, as it proxies for the relative difficulty to find a close substitute for the stock (Pontiff 1996; 2006). The higher the level of idiosyncratic variation of the mispriced asset, the more difficult it is to hedge this risk with another asset, making the risk premium needed to induce arbitrage higher.

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include Pontiff (1996), Wurgler and Zhuraskaya (2001), Ali, Hwang and Trombley (2003), Mendenhall (2004), Mushruwala, Rajgopal and Shevlin (2006) and Pontiff and Schill (2004).

The main results can be summarised as follows. First, I show that the relative overpricing (i.e., a low value-to-price ratio) is greater for high arbitrage cost firms relative to low arbitrage cost firms. No significant differences are found for relatively underpriced firms. Second, I find that the persistence of the value-to-price ratio is greater for higher arbitrage cost firms. The results are consistent with limits to arbitrage affecting the timeliness of price reflecting accounting information.

The remainder of this chapter is set out as follows. In Section 4.3 I discuss the implications of the Abreu and Brunnermeier (2002) model and form testable hypotheses. In Section 4.4 I discuss the measurement of the variables used in the study. In Section 4.5 I present the empirical analysis, and in Section 4.6 I conclude the study.

4.3. Model and hypotheses

4.3.1. Model set-up

In this section, I outline the key result in Abreu and Brunnermeier (2002). They assume that: (i) behavioural traders exist in the market with their initial absorption capacity (i.e. excess demand) denoted as κ_0, (ii) arbitrageurs become sequentially aware of mispricing, and (iii) to correct mispricing requires the collective actions of a critical mass of informed arbitrageurs. The authors show that in equilibrium the expected time arbitrageurs delay their correcting trades (denoted τ^*) can be written as follows:

$$\tau^* = \bar{\tau} - \frac{(\bar{\tau} + \eta\kappa_0)}{\lambda\eta\kappa_0}\ln\left(\frac{c}{c - \lambda\beta}\right), \quad (4.1)$$

where, $\bar{\tau}$ is the maximum duration of mispricing, η is the length of the "awareness window" (which measures how quickly the actions of arbitrageurs are synchronized), κ_0 is the initial absorption capacity of behavioural traders, λ is the arrival rate of arbitrageurs, β is the level of mispricing, and c is the level of holding cost. The comparative statics of Equation (4.1) suggest that the expected duration of mispricing is increasing in the length of the awareness window (η), initial absorption capacity (κ_0) and holding cost (c). Conversely, the expected duration of mispricing is decreasing in the arrival rate of arbitrageurs (λ) and the level of expected mispricing (β).

4.3.2. *Hypotheses*

In this section, I combine the duration function from Abreu and Brunnermeier (2002) with a value-to-price relation similar to that in Chapter 3 to form testable hypotheses. In the Abreu and Brunnermeier (2002) model, arbitrageurs are expected to trade immediately when they become aware of the mispricing, if the cost to do so is below

$$c > \left(\lambda / \left(1 - e^{-\lambda \eta \kappa_0}\right)\right)\beta, \tag{4.2}$$

which implies that $\beta > c\left(\lambda / \left(1 - e^{-\lambda \eta \kappa_0}\right)\right)^{-1}$, or that mispricing is bounded by an increasing function of arbitrage costs. In this format, however, there is no role for accounting information about firm value. Instead, mispricing is assumed to be observed directly by the would-be arbitrageur.

In Lee, Myers and Swaminathan (1999) and Chapter 3, price and accounting based measures of value are observable but noisy functions of unobservable intrinsic value (v_{it}^*). For each firm, i, the log of price and of measured value (using accounting information) at time t can be written as: $p_{it} = v_{it}^* + u_{1it}^*$ and $v_{it} = v_{it}^* + u_{2it}^*$, where, u_{1it}^*, is the 'mispricing' error and u_{2it}^* the value measurement error.

Expressing the relation between log price and log value as a ratio yields: $v_{it}/p_{it} = v_{it} - p_{it} = u^*_{2it} - u^*_{1it}$, where v_{it}/p_{it} is the log value-to-price ratio, or 'spread' between the log of price and the log of value that includes both the mispricing error and the value measurement error for firm i at time t. Note that in this model, the components of mispricing and measurement error cannot be identified directly.

While measurement error and mispricing are theoretically distinct, in practice increased noise in price and increased noise in measurement of value are likely to lead to the same outcome. Specifically, stocks with the highest sensitivity to speculation will have the most subjective valuations and will be the most risky to arbitrage. Clearly, the disparity between price and value (both under- and over-valuation) will be increasing with the difficulty to take an arbitrage position in the firm. Stated formally:

H2: The level of the disparity between price and accounting based value increases with the difficulty to arbitrage.

All else equal, from the Abreu and Brunnermeier (2002) model and the discussion above, arbitrage is likely to take longer when the firm is more difficult to arbitrage. In other words, the persistence of the disparity between price and value will increase with the difficulty to take an arbitrage position in the firm. Stated formally:

H3: The duration of the disparity between price and value increases with the difficulty to arbitrage.

4.4. Measurement of variables

4.4.1. Measuring arbitrage costs

To test the hypotheses outlined above, I require a measure of the difficulty to take an arbitrage position that is available for a broad cross-section of firms. Pontiff (2006) argues that the greatest cost to the arbitrageur is their exposure to idiosyncratic risk. Idiosyncratic risk is a measure of how difficult it is to find a close substitute for the potentially mispriced firm. I follow the model used by Wurgler and Zhuravskaya (2002) to measure this arbitrage cost. Specifically, I use the standard deviation of prior idiosyncratic errors from the market model. As the majority of the analysis in this thesis is value-weighted, I use the value-weighted *CRSP* index in the market model. I estimate the idiosyncratic variance using daily returns data over the prior 12 month period, stating in January and ending in December of the prior calendar year.

Ideally, a measure of expected difficulty to arbitrage would be used to measure expected arbitrage cost. To the extent would-be arbitrageurs consider prior idiosyncratic risk to be a good indicator of future exposure to idiosyncratic risk, this measure suffices. Other studies that use this variable to measure arbitrage costs include Ali, Hwang and Trombley (2003), Mendenhall (2004) and Mushruwala, Rajgopal and Shevlin (2006). These studies examine the book-to-market effect, post-earnings announcement drift and the accruals strategy, respectively. These studies all show that this proxy is not significantly affected by changing the index used in the market model.

4.4.2. Measuring persistence

I measure persistence using the autocorrelation of the log value-to-price ratio (*vp*). In a dynamic setting, the disparity between price and value can be corrected for by either changes in measurement error or by changes in mispricing. A shock to the time-series of *vp* must be due to a differential change over time in either the measurement error or the mispricing error. This implies that the time-series of *vp* is related to the past differences in the measurement errors and mispricing errors. Specifically, *vp* can be represented as:

$$v_{it}/p_{it} = \sum_{j=0}^{\infty} a_j \Delta\left(u^*_{2it-j} - u^*_{1it-j}\right). \tag{4.3}$$

In Equation (4.3), if the differences between the errors are removed over time, then the value of a_j approaches zero as j increases. At the other extreme, if the differences in the errors have a permanent effect on the level of the value-to-price ratio, then *vp* follows a random walk. Rewriting Equation (4.3) as a parsimonious, single-lag autoregressive process (AR(1)) model yields:

$$v_{it}/p_{it} = \phi\left(v_{it-1}/p_{it-1}\right) + e_{it}, \tag{4.4}$$

where ϕ refers to the AR(1) parameter and can be considered as a measure of the persistence in the value-to-price ratio and e_{it} is an error term. The AR(1) parameter can then be used in the following measures of persistence.

First, the time-domain measures of persistence include the cumulative response function (CIR), which can be estimated as $1/(1-\phi)$. The CIR measures the cumulative effect of a shock to *vp* on the level of *vp*. A low CIR means that the system is not highly persistent as shocks to the disparity between price and value are removed relatively quickly (see, for example, the discussion in Andrews and Kim, 1994). I also report the half-life of the shock to the series, calculated as $\ln(0.5)/\ln(\phi)$. The benefit of the half-life is that it is expressed in months, and as such, the difference between the half-lives of two series gives an indication of the economic significance of the difference in timeliness of the two series. Both of these measures relate directly to the autocorrelation coefficient.

I also present a measure of persistence based on the frequency domain, from spectral analysis. Phillips (1991) suggests that the frequency at spectrum zero (F(0)), which is estimated as $\sigma_e^2/(1-\phi)^2$, provides a measure of persistence (where σ_e^2 is the sample standard deviation of the error term in Equation (4.4)). Specifically, F(0) measures the low-frequency autocovariance of the time-series. One benefit of using F(0) to compare the time-series is that, for a given model, the persistence estimate increases with the variance of the deviations from the model structure. In comparing the time-series, for a more noisy *vp*, a lower F(0) means that while there are larger and more volatile shocks to the time-series of *vp*, they are less persistent than the shocks to a less noisy *vp* process.[42]

4.5. *Empirical analysis*

4.5.1. *Descriptive statistics and correlations*

I use the same base sample as in Chapter 3, except I exclude firms that do not have available market prices over the prior 12 months. This requirement is due to the measurement of idiosyncratic risk over the prior 12 months, used to proxy for the lack of close substitutes. In Panel A of Table 4.1, I present the average of the yearly estimates of the mean, standard deviation, 25th and 75th percentiles of the firm-specific fundamental-to-price ratios (i.e., the number of observations for the summary statistics is $T = 26$). I use the December fundamental-to-price ratios in each year, where each ratio is estimated using the month-end estimate of value

[42] The OLS estimates for ρ and σ_e^2 will not yield consistent unbiased estimates (potential ways of overcoming this problem have been suggested by e.g., Andrews and Kim 1994; Perron and Ng, 1998). Hamilton (1994, p165) notes that the maximum likelihood estimates of ρ and σ_e^2 are sufficient for the estimation of the population spectrum. This is true even if the model (AR(1) in this case) is incorrectly specified as long as the autocovariances of the true process are reasonably close to those estimated by the model.

divided by the closing market value (i.e., price times the number of shares outstanding on *CRSP*). The means of the fundamental-to-price ratios are all below one, except for the *Ve(f)P* model. In addition, the bulk of the distribution of the ratios lies below one, with only the *Vf*(1)*P* and *Ve(f)P* ratios having their average 75th percentile above one.[43]

[43] In Chapter 3, I calculated these statistics at the aggregate portfolio level. Comparing the firm-specific statistics to the aggregate level statistics reported in Chapter 3 shows that the means are similar (except for the dividend-to-price ratio which is lower due to the equal weighting on zero dividend-to-price ratios) and the standard deviations are higher for all of the ratios. The standard deviation is expected to be higher, as the standard deviation reported in this chapter is based on equal-weighting (versus value-weighting in Chapter 3), and the deviations of smaller firms are likely to be greater.

Table 4.1 Summary statistics

Panel A: Descriptive statistics for annual firm-level fundamental-to-price ratios

Variable	Mean	Std. Dev.	Q25	Q75
DP	0.016	0.027	0.000	0.023
EP	0.064	0.078	0.012	0.096
BM	0.638	0.648	0.175	0.865
Vf(1)P	0.876	0.785	0.369	1.096
Vl(x)P	0.662	0.659	0.197	0.887
Vl(f)P	0.657	0.561	0.226	0.891
Ve(1)P	0.574	0.419	0.190	0.789
Ve(f)P	1.168	1.341	0.337	1.413

Panel B: Descriptive statistics for proxies for difficulty to arbitrage and difficulty to value

Variable	Mean	Std. Dev.	Q25	Q75
ArbCost	0.028	0.015	0.016	0.036
MVE (Mil)	1,532	9,446	32	611
TA	2555	19123	40	744
FROE	0.124	0.097	0.039	0.175
Disp	0.063	0.087	0.013	0.067
Loss	0.231	0.422	0.000	0.000
SpecItem	0.305	0.460	0.000	1.000
DivPay	0.459	0.498	0.000	1.000
Beta	0.695	0.989	0.286	1.024
Price	22.660	620.479	3.474	25.172
Volume	37,327	227,152	1,433	20,093

(Table 4.1 continued on the following page...)

Panel C: Spearman and Pearson correlations between select variables

	Vf(1)P	*ArbCost*	*MVE*	*Sales*	*FROE*	*Disp*	*Loss*	*SpecItem*	*DivPay*	*Beta*	*Price*	*Volume*
Vf(1)P	1	–0.071	–0.261	0.103	0.298	–0.282	–0.249	–0.203	–0.303	0.021	–0.010	0.158
ArbCost	–0.050	1	–0.447	–0.457	–0.290	–0.078	0.124	–0.338	0.246	0.014	0.082	–0.590
MVE	–0.063	–0.089	1	0.793	0.021	0.768	0.353	0.748	0.095	0.175	0.038	0.316
Sales	0.180	–0.135	0.518	1	0.231	0.569	0.264	0.623	–0.094	0.202	0.048	0.392
FROE	0.022	–0.017	0.051	0.037	1	–0.126	0.001	0.157	–0.053	0.028	0.017	0.309
Disp	–0.070	–0.008	0.593	0.262	–0.004	1	0.482	0.628	0.209	0.197	0.083	0.010
Loss	–0.106	0.048	0.042	0.034	–0.007	0.132	1	0.326	0.306	0.062	0.049	–0.064
SpecItem	–0.154	–0.307	0.351	0.304	0.001	0.302	0.252	1	0.083	0.125	–0.008	0.206
DivPay	–0.087	0.109	0.017	–0.018	–0.078	0.049	0.057	0.033	1	0.083	0.086	–0.242
Beta	0.036	0.014	0.090	0.107	0.001	0.102	0.029	0.132	0.043	1	0.167	–0.012
Price	–0.011	0.060	0.002	0.010	0.000	0.049	0.027	–0.010	0.071	0.167	1	–0.065
Volume	0.173	–0.484	0.091	0.158	0.005	–0.005	–0.041	0.220	–0.126	–0.012	–0.065	1

Notes: Panels A and B report averages of the yearly mean, standard deviation, 25th and 75th percentiles (T=26) for the yearly cross-section of firms. In Panel C, Spearman correlations are reported on the upper triangular half and Pearson correlations are reported on the lower triangular half. The correlations are pooled across the sample period of 1979–2004 for all firms with available variables (N=73,144). *DP* is the dividend-to-price ratio, *EP* is the earnings-to-price ratio, *BM* is the is the book-to-market ratio, *Vf*(1)*P* is the one-period forecast-based residual income model to price ratio, *Vl*(*x*)*P* is the residual income model based on the linear information dynamics of Ohlson (1995) to price ratio, *Vl*(*f*)*P* is the linear information dynamics model with analyst forecasts as 'other information' from Dechow et al. (1999) to price ratio, *Vl*(*e*)*P* is the earnings capitalization to price ratio, and *Ve*(*f*)*P* is the Ohlson-Juettner model of earnings growth to price ratio, *ArbCost* is the standard deviation of residuals from the firm-specific market model regression, *MVE* is the market value of equity using December closing prices and shares outstanding from the prior year, *TA* is total assets, *FROE* is the analyst forecast of earnings divided by book-value, *Disp* is the average of the standard deviation of analyst forecasts for all months in the prior year, *Loss* is an indicator for firms incurring a loss in the prior financial year (using *Compustat* data 18), *SpecItem* is an indicator variable for non-zero special items (using *Compustat* data 17), *DivPay* is an indicator variable for the payment of common share dividends (using *Compustat* data 21), *Beta* is the slope coefficient from the firm-specific market model, *Price* is the CRSP closing price per share for the year, *Volume* is the average monthly volume measured over the prior calendar year.

In Panel B, I report the mean of pooled annual firm characteristics. The variable of interest, *ArbCost*, is the standard deviation of idiosyncratic risk. The mean of *ArbCost*, is 0.028 with a standard deviation of 0.015. Consistent with Campbell, Lettau, Malkiel and Xu (2001), the mean of *ArbCost* increases over the sample period (not tabulated). The summary statistics for the other firm characteristics display patterns as expected. Not surprisingly, there is a large amount of skewness in the size of the firms included in the sample, using either market value or total assets as a measure of size, and also in the volume traded. In the pooled sample, 30.5% of firm-years report a special item, 23.2% report a loss and 45.9% pay a dividend.

In Panel C, I report the correlations between selected variables. I report the Spearman rank-correlations in the upper triangular half of the panel and the Pearson correlations in the lower. *ArbCost* is negatively correlated with market value (*MVE*), sales (*Sales*), trading volume (*Vol*) and firms reporting special items (*SpecItem*). Firms reporting losses (*loss*) and paying dividends (*DivPay*) have a positive correlation with *ArbCost*. The counter-intuitive correlation between *ArbCost* and firms reporting special items appears to be due to the positive correlation between firm size and firms reporting special items (0.768, $p<0.01$). Excluding the positive correlation with dividend paying firms, the results are generally consistent with firms that are the most difficult to arbitrage also being the most difficult to value.

4.5.2. A graphical display of high and low arbitrage cost portfolios

In Figure 4.1, I present the time-series variation in the price-to-value ratio (using *Vf*(1) to measure value) for the high and low arbitrage cost quintile portfolios. In this form, the dramatic increase in price relative to value in the late 1990s for the high arbitrage cost group is evident. A smaller increase in the low arbitrage cost group is also evident. In addition, while there are periods where the high and low cost portfolios are roughly equal, in general the high arbitrage cost portfolio tends to have both larger positive deviations (or bubbles) and larger negative deviations (or crashes). In general, the three quintiles not graphed in Figure 4.1 display similar trends, and in terms of magnitude, tend to fall in between the extreme quintile portfolios.

While the time-trends for other data are not displayed, some of the time-series trends in firm characteristics are worthy of discussion. First, there is an upward trend in the proportion of loss firms in the sample, consistent with Collins, Maydew and Weiss (1997). A similar trend is seen in the proportion of firms reporting special items over time. Consistent with prior research, the proportion of firms paying dividends appears to be 'disappearing,' with the proportion of the firms in the sample paying a common share dividend decreasing over the sample period (e.g., Fama and French, 2001; DeAngelo, DeAngelo and Skinner, 2004). Unlike the price-to-value ratio, however, the trends in these series do not appear to spike as significantly around the potential bubble period.

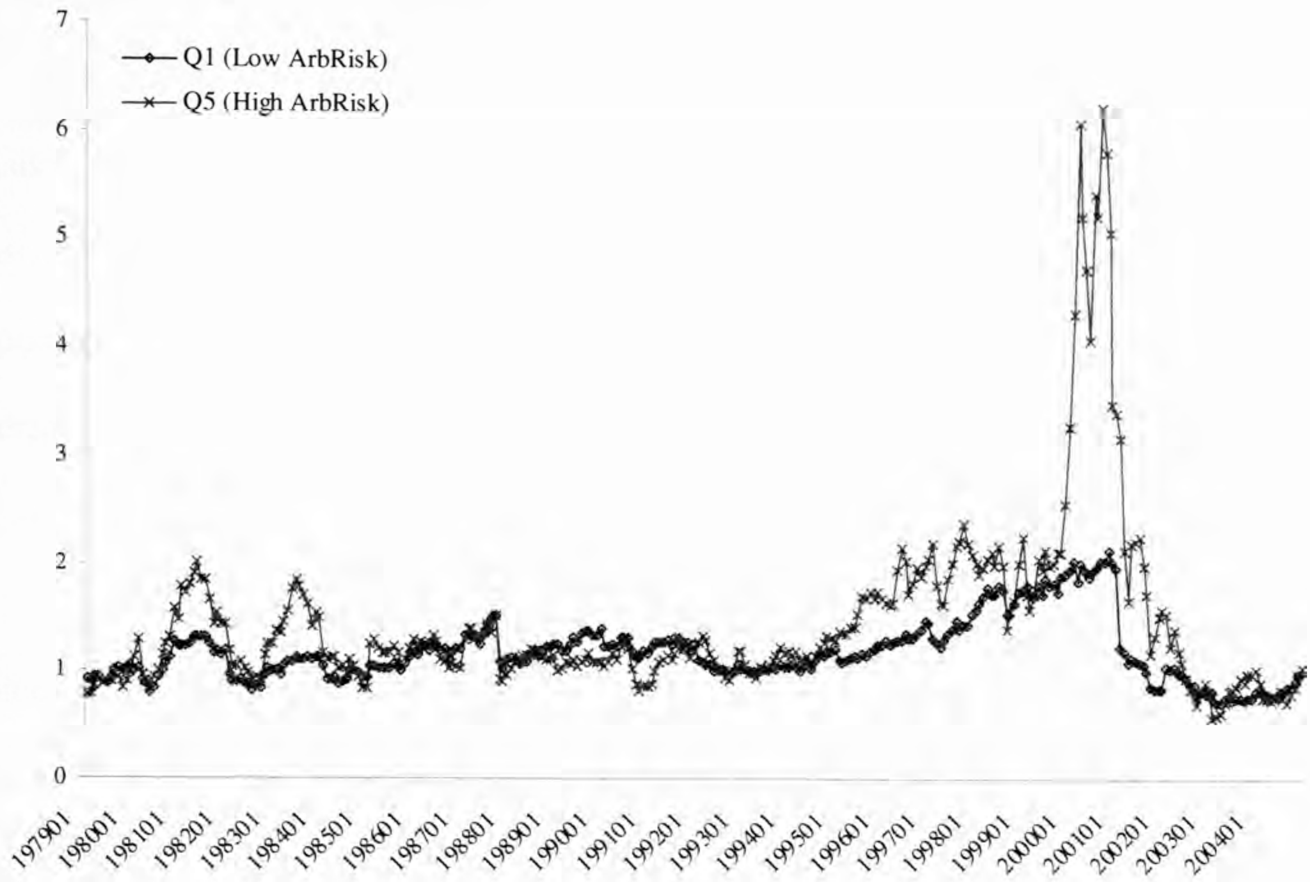

Figure 4.1 A comparison of high and low arbitrage cost price-to-value ratios

This graph displays the time-series of aggregate market value (P) and the aggregate accounting based measure of value (Vf(1)), where value is measured using the sum of book-value and a perpetuity of one-period ahead residual income value growing at 3%. The time-series are measured at the portfolio level in millions of US dollars and includes all firms with sufficient data during the period 1979–2004.

4.5.3. Characteristics of quintile portfolios

In Table 4.2, I present further descriptive statistics for quintile portfolios formed on the basis of market value of equity, the value-to-price ratio (using *Vf*(1)*P*) and the level of arbitrage costs. Firms are assigned to quintiles annually using the calendar year-end values. In Panel A, I report the results for size (using market value of equity, *MVE*) and, as expected, I find that beta increases with the market value of the firm, and that the market-adjusted returns over 12, 24 and 36 months for small firms are bigger than those for large firms.

In Panel B, I repeat the analysis for portfolios of firms sorting by the level of *vp*, the value-to-price ratio, (measured as *Vf*(1)*P*). I find that the average beta for the portfolio of high *vp* firms is significantly lower than the average beta for low *vp* firms. Consistent with Frankel and Lee (1998), I find that the portfolio of high *vp* firms outperforms the portfolio of low *vp* firms, based on 12, 24 and 36 month market-adjusted returns. This result implies that the theoretical hedge of taking a long position in under-valued firms and a short position in over-valued firms nets positive abnormal returns. This is important because if subsequent returns are an ex-post indicator of the removal of mispricing, the *vp* ratio may be a useful measure of ex-ante mispricing.

In Panel C, I present quintile portfolios based on the level of arbitrage cost (*ArbCost*). High arbitrage firms are generally smaller firms with lower fundamental-to-price ratios, which provides preliminary support for Hypothesis 2. I do not find, however, that high arbitrage cost firms have significantly different abnormal returns to low arbitrage cost firms. This is interesting, as it implies that while arbitrage costs are associated with the level of *vp*, unlike *vp*, *ArbCost* does not appear to provide an ex-ante measure of mispricing.

Table 4.2 Characteristics of quintile portfolios formed by size, value-to-price and arbitrage cost

Panel A: Market value of equity (MVE) portfolios (within sample)

	Q1 (low MVE)	Q2	Q3	Q4	Q5 (high MVE)	Q5–Q1 Diff.
MVE	18.6	78.7	240.6	736.0	6501.6	–
BP	0.875	0.796	0.689	0.593	0.518	–0.357**
Vf(1)P	1.148	1.050	0.926	0.821	0.754	–0.394**
Vl(x)P	0.906	0.823	0.712	0.615	0.539	–0.367**
Vl(f)P	0.872	0.806	0.707	0.619	0.549	–0.323**
Ve(1)P	0.655	0.659	0.624	0.585	0.566	–0.089*
Ve(f)P	1.180	1.362	1.348	1.249	1.106	–0.074
Beta	0.357	0.555	0.698	0.845	1.015	0.658**
Ret12	0.079	0.036	0.011	0.004	0.009	–0.07**
Ret24	0.154	0.083	0.023	0.007	0.017	–0.137**
Ret36	0.215	0.142	0.054	0.020	0.019	–0.196**

Panel B: Value-to-price (VP) portfolios (within sample)

	Q1 (low VP)	Q2	Q3	Q4	Q5 (high VP)	Q5–Q1 Diff.
MVE	1466.5	2442.4	1824.8	1204.2	532.9	–933.6*
BP	0.198	0.427	0.461	0.825	1.441	1.243**
Vf(1)P	0.334	0.568	0.644	1.090	1.939	–
Vl(x)P	0.213	0.441	0.479	0.849	1.494	1.281**
Vl(f)P	0.230	0.449	0.492	0.858	1.408	1.178**
Ve(1)P	0.214	0.458	0.476	0.780	1.039	0.825**
Ve(f)P	0.524	1.071	0.954	1.414	2.048	1.524**
Beta	0.711	0.789	0.739	0.658	0.584	–0.127*
Ret12	0.007	0.011	0.006	0.032	0.070	0.063**
Ret24	0.002	0.004	0.023	0.081	0.157	0.155**
Ret36	0.001	0.006	0.032	0.128	0.253	0.252**

(Table 4.2 continued on the following page...)

Panel C: Arbitrage cost (*ArbCost*) portfolios (within sample)

	Q1 (low *ArbRsk*)	Q2	Q3	Q4	Q5 (high *ArbRsk*)	Q5–Q1 Diff.
MVE	4061.5	1944.7	938.2	475.1	157.2	–3904**
BP	0.747	0.715	0.542	0.675	0.630	–0.117*
Vf(1)P	0.989	0.947	0.765	0.926	0.891	–0.098
Vl(x)P	0.771	0.739	0.566	0.700	0.657	–0.114**
Vl(f)P	0.751	0.726	0.565	0.695	0.660	–0.091*
Ve(1)P	0.656	0.636	0.493	0.607	0.560	–0.096**
Ve(f)P	1.245	1.288	1.029	1.291	1.099	–0.146
Beta	0.566	0.700	0.756	0.777	0.670	0.104*
*Ret*12	0.012	0.015	0.009	0.038	0.043	0.031
*Ret*24	0.030	0.037	0.031	0.078	0.070	0.040
*Ret*36	0.046	0.061	0.060	0.117	0.097	0.051

Notes: *MVE* is the market value of equity using December closing prices from the prior year, *BP* is the book-to-market ratio, *Vf*(1)*P* is the one-period forecast-based residual income model to price ratio, *Vl*(*x*)*P* is the residual income model based on the linear information dynamics of Ohlson (1995) to price ratio, *Vl(f)P* is the linear information dynamics model with analyst forecasts as 'other information' from Dechow et al. (1999) to price ratio, *Vl(e)P* is the earnings capitalization to price ratio, and *Ve(f)P* is the Ohlson-Juettner model of earnings growth to price ratio, *ArbCost* is the standard deviation of residuals from the firm-specific market model regression, *Beta* is the slope coefficient from the firm-specific market model, *Ret*12, *Ret*24, *Ret*36 are market-adjusted returns over 12, 24 and 36 months, respectively, with a correction for delisting using the method of Shumway (1997).

4.5.4. *Independent sorts of arbitrage cost and value-to-price*

In the analysis above, I compared fundamental-to-price ratios for portfolios of high and low arbitrage cost portfolios and found evidence that some of the fundamental-to-price ratios are smaller for the high *ArbCost* portfolio relative to the low *ArbCost* portfolio. This provides some support for Hypothesis 2, but it implies that the disparity between value and price is affected asymmetrically by arbitrage cost. I investigate this possibility further in this section. I do this by independently assigning firms to both a value-to-price ranked quintile portfolio (using *Vf*(1)*P*) and an *ArbCost* ranked quintile portfolio, on a yearly basis within an industry. This procedure yields 25 (5 by 5) portfolios, which potentially have different cell sizes. The advantage of sorting the firms independently (rather than conditionally) is that it allows for the identification of clustering of firms in certain categories of *vp* and *ArbCost*.

In Table 4.3, I report the average value-to-price ratio for the 25 independently sorted *vp* and *ArbCost* portfolios. There are a total of 73,144 observations assigned to the categories. The equal assignment of firms to each of the 25 portfolios would be 4%, which is approximately 2,926 observations. Consistent with the asymmetry observed in Table 4.2, in Column 1 of Table 4.3 there is evidence of a clustering of firms in the high value-to-price (i.e., most overpriced) and high arbitrage cost category ($N = 5{,}096$, approx. 7%). On the other hand, there are relatively few observations that are both high value-to-price and low arbitrage cost ($N = 1{,}713$, approx. 2.3%). The difference between these portfolios is –0.061, which is significant using a conventional *t*-test. In Column 2, the level of clustering is not as evident, but the difference between the high cost and low cost portfolios (–0.54) is still significant.

Table 4.3 Average *vf*(1)*p* ratios for bi-dimensional portfolios

	Value-to-price quintiles					
	Q1 (Low *Vf(1)P*) Mean (*n*)	Q2 Mean (*n*)	Q3 Mean (*n*)	Q4 Mean (*n*)	Q5 (high *Vf(1)P*) Mean (*n*)	All firms Mean (*n*)
Arbitrage cost quintiles						
Q1 (low arbitrage cost)	0.372	0.592	0.781	1.049	1.964	0.989
	(1713)	(2823)	(3444)	(3186)	(2688)	(13854)
Q2	0.356	0.577	0.777	1.068	1.963	0.947
	(2079)	(3369)	(3606)	(3372)	(2684)	(15110)
Q3	0.345	0.566	0.475	1.095	1.925	0.765
	(2556)	(2981)	(3378)	(3223)	(2816)	(14954)
Q4	0.327	0.555	0.792	1.125	1.920	0.926
	(3502)	(2804)	(2749)	(2973)	(2933)	(14961)
Q5 (high arbitrage cost)	0.311	0.539	0.789	1.134	1.939	0.891
	(5096)	(1972)	(1878)	(2157)	(3162)	(14265)
All firms	0.334	0.568	0.644	1.090	1.939	
	(14946)	(13949)	(15055)	(14911)	(14283)	
Q5 – Q1 Diff.	–0.061**	–0.054*	0.008	0.085	–0.025	
	[–3.19]	[–2.06]	[0.20]	[1.31]	[–0.12]	

Notes: The 25 portfolios reported are formed independently based on value-to-price and arbitrage cost. I use the December year-end level of idiosyncratic risk (*ArbCost*) over the prior 12 months and December year-end level of the value-to-price ratio. The value-to-price ratio is measured as *Vf*(1)*P*, which is the one-period forecast-based residual income model to price ratio. The cell means reported are the averages of the 26 yearly means of the value-to-price ratio, the cell sizes are the sum of the number of firms assigned to the 26 yearly portfolios. The *t*-tests are based on the 26 yearly means for each of the five *vp* portfolios accommodating different variances. **p<0.01, *p<0.05, $^{\delta}$p<0.1.

In contrast, in the remaining three columns I do not find any difference between high and low arbitrage cost portfolios. Interestingly, excluding the over-valued (*vp* Q5) portfolio, the difference between high and low arbitrage cost portfolios within each *vp* quintile is monotonically increasing across the *vp* quintiles Q1 to Q4, consistent with Hypothesis 2. The most under-valued firms in *vp* Q5, have a negative difference, but this difference is clearly not significant. These results suggest that the arbitrage cost increases the disparity in the *vp* ratio when the firm is relatively over-valued, and are consistent with the lack of close substitutes being a greater friction when arbitraging the mispriced firm would require a short position.

4.5.5. Multivariate analysis

In this section, I test for an association between *vp* and the level of arbitrage cost (Hypothesis 2) controlling for other potentially correlated variables. I measure the disparity between price and value as the absolute value of *vf*(1)*p* (i.e., the absolute value of the log of the *Vf*(1)*P* ratio). Taking the log removes the bias in the weighting of the regression parameters model towards overvaluation (i.e., a firm can be over-valued by greater than 100%) and I take the absolute value of this measure to accommodate the expected U-shape distribution caused by over- and under- valuation. I estimate the following multivariate regression:

$$\begin{aligned}|vp_i| = a + b_1 ArbCost_i + b_2 \log(Vol_i) + b_3 \log(Sales_i) + b_4 Beta_i + b_5 DivPay_i + b_6 Loss_i \\ + b_7 SpecItem_i + e_i\end{aligned} \tag{4.5}$$

To support Hypothesis 2, the coefficient b_1 is expected to be positive and significant. I include volume (*Vol*) as a proxy for liquidity (e.g., Chordia, Roll and Subrahmanyam, 2000) as low liquidity is potentially a trading cost. I include sales as a proxy for size and beta as a measure of the variation in the expected rate of return (which is excluded from the measure *vf*(1)*p*). I also include indicator variables for dividend paying firms (*DivPay*), which is expected to be negatively related to |*vp*|, and for firms reporting losses and special items, which are expected to be positively related to |*vp*| as they proxy for the difficulty to value the firm (e.g., Pastor and Veronesi, 2003).

I report the results in Table 4.4. In Panel A, I report results for the full sample (1979–2004) and the two equal subperiods 1979–1991 and 1992–2004. In Panel B, I report the full sample results by the value-to-price portfolios. In Column (1) of Panel A, I show that *ArbCost* is positively and significantly associated with the disparity between price and value. As expected, when including other correlated variables the significance level of the association declines. In Column (2), I include volume, size, and beta, and in Column (3) I report the full model. In both cases the association between *ArbCost* and |*vp*| is positive and significant. In both models I find a positive and significant association of |*vp*| with beta. However, there is no significant association between |*vp*| and volume and I find only weak and inconsistent associations with the indicator variables, *DivPay*, *Loss* and *SpecItem*.

In Columns (4) and (5), I present the parameter estimates for the sub-period 1979–1991. The results for arbitrage cost are on the whole similar to those for the full period. The coefficient on *SpecItem* is negative and significant, implying a lower level of disparity between price and value for special item reporting firms. I present the results for the second subperiod (1992–2004) in Columns (6) and (7). In the full model (Column 7), I find that during this period, *ArbCost* is not significantly associated with the disparity between price and value. I find instead a strong positive association with *Beta*, *SpecItem* and *DivPay*. These results are partially counter-intuitive and do not support Hypothesis 2.

I investigate why this may be the case in Panel B, where I split the firms into quintile portfolios based on the value-to-price ratio (not the absolute value). The results in Panel B show that the model has a substantially better fit, in terms of Adjusted R^2, for firms in quintiles one (i.e., the most over-valued) and five (i.e., the most under-valued). Also the association of the absolute level of disparity with *ArbCost* is counter to the predicted direction for the most under-valued firms. As all of the firms are pooled in the prior analysis, this explains why the results are weaker. For the lowest quintile portfolio of value-to-price, the disparity between price and value is positively and significantly associated with *ArbCost* and *Beta* as predicted, and negatively and significantly associated with *Vol* and *DivPay* as expected. Both *Sales* and *SpecItem* are significant in the opposite direction to that predicted. The positive association with *Sales* may be due to a 'glamour story,' where investors over-extrapolate prior success (Lakonishok, Shleifer and Vishny, 1994).

For the highest quintile of value-to-price (i.e., the most under-valued firms), contrary to my prediction the disparity between price and value is negatively and significantly associated with *ArbCost*. As the result is the opposite of the result for low *vp* firms, it is clear that pooling these firms resulted in a lower power test. Both *Loss* and *DivPay* are also significantly associated in the opposite direction.

Overall, the multivariate analysis provides qualified support for Hypothesis 2. Specifically, the lack of close substitutes is associated with an increase in the disparity between price and value for relatively over-valued firms. This is consistent with the relative importance of requiring a close substitute to hedge a short position in comparison to a long position.

Table 4.4 Cross-sectional determinants of the level of disparity between price and value

Panel A: All firms

			1979 – 2004			1979 – 1991	1992 – 2004	
Model	Predicted sign	Coefficient (*t*-statistic) (1)	Coefficient (*t*-statistic) (2)	Coefficient (*t*-statistic) (3)	Coefficient (*t*-statistic) (4)	Coefficient (*t*-statistic) (5)	Coefficient (*t*-statistic) (6)	Coefficient (*t*-statistic) (7)
Intercept		0.453**	0.596**	0.576**	0.371**	0.487**	0.535**	0.665**
		(20.97)	(10.04)	(15.10)	(15.30)	(9.73)	(34.52)	(14.06)
ArbCost	Positive	5.160**	1.484*	1.498*	6.814**	2.686**	3.506**	0.311
		(7.24)	(1.73)	(2.40)	(6.13)	(2.92)	(5.39)	(0.42)
Log(*Volume*)	Negative		0.005	0.005		0.013 $^{\delta}$		–0.002
			(1.00)	(1.06)		(1.63)		(–0.36)
Log(*Sales*)	Negative		–0.033**	–0.036**		–0.021*		–0.051**
			(–4.13)	(–3.65)		(–1.90)		(–3.26)
Beta	Positive		0.114**	0.119**		0.029 $^{\delta}$		0.210**
			(4.41)	(4.34)		(1.61)		(5.48)
DivPay	Negative			0.030		–0.056		0.116**
				(0.92)		(–1.26)		(3.17)
Loss	Positive			–0.024		–0.051 $^{\delta}$		0.002
				(–1.23)		(–1.63)		(0.10)
SpecItem	Positive			0.027*		–0.035**		0.088**
				(1.79)		(–3.55)		(6.22)
Adjusted R^2		0.019	0.038	0.050	0.031	0.046	0.008	0.054

(Table 4.4 Continued on the following page…)

Panel B: Full period estimates by value-to-price ranked portfolios

Model	Predicted sign	Coefficient (*t*-statistic) (VPQ1)	Coefficient (*t*-statistic) (Q2)	Coefficient (*t*-statistic) (Q3)	Coefficient (*t*-statistic) (Q4)	Coefficient (*t*-statistic) (VPQ5)
Intercept		0.887**	0.434**	0.189**	0.205**	1.070**
		(16.59)	(15.24)	(10.67)	(6.85)	(7.80)
ArbCost	Positive	5.668**	0.217*	0.090	–0.180 $^{\delta}$	–3.003**
		(6.03)	(1.86)	(1.23)	(–1.36)	(–2.10)
Log(*Volume*)	Negative	–0.014*	0.001	0.001	0.001*	–0.001
		(–2.32)	(0.08)	(–0.16)	(1.95)	(–0.06)
Log(*Sales*)	Negative	0.025**	0.004**	0.002**	–0.003**	–0.126**
		(3.23)	(4.13)	(2.52)	(–3.20)	(–6.99)
Beta	Positive	0.023 $^{\delta}$	0.011**	0.001	–0.001	0.207**
		(1.62)	(3.11)	(0.76)	(–0.79)	(3.60)
DivPay	Negative	–0.142**	–0.014**	–0.004*	–0.003	0.618**
		(–11.5)	(–5.17)	(–2.47)	(–1.18)	(8.65)
Loss	Positive	0.013	0.012	–0.005	0.008*	–0.129**
		(0.54)	(0.95)	(–0.94)	(1.97)	(–3.68)
SpecItem	Positive	–0.015*	–0.004	0.001	0.004**	0.143**
		(–1.72)	(–1.6)	(0.65)	(2.56)	(3.69)
Adjusted R^2		0.115	0.014	0.006	0.010	0.130

Notes: Coefficients are the average of the 26 coefficicents from yearly cross-sectional regressions. *Vf*(1)*P* is the one-period forecast-based residual income model to price ratio, *ArbCost* is the standard deviation of residuals from the firm-specific market model regression, *Volume* is the average monthly volume measured over the prior calendar year, *Sales* is the prior fiscal year sales (*Compustat* data 12), *Beta* is the slope coefficient from the firm-specific market model, *Loss* is an indicator variable for firms incurring a loss in the prior financial year (using *Compustat* data 18), *SpecItem* is an indicator variable for non-zero special items (using *Compustat* data 17), *DivPay* is an indicator variable for the payment of common share dividends (using *Compustat* data 21).
** p<0.01, * p<0.05, $^{\delta}p$<0.01.

4.5.6. *Tests of hypothesis 3*

Hypothesis 3 predicts that firms that are more difficult to arbitrage will have a more persistent disparity between price and value, as the actions of arbitrageurs are expected to be delayed by arbitrage costs. I use the *vf*(1)*p* model of value-to-price ratio and estimate the parameters of an AR(1) model for each of the arbitrage cost quintile portfolios. The model is estimated using maximum likelihood and the time-series are pre-whitened to remove the means of the times-series. For each of the five portfolios, firms are assigned annually in January based on arbitrage cost over the prior 12 months (ending in December of the prior calendar year).

In Table 4.5 I report the results. The first five columns report the estimates of the model for each of the arbitrage cost portfolios, starting with the low arbitrage cost portfolio and ending with the high arbitrage cost portfolio. The final column reports the difference between the high arbitrage cost parameter and the low arbitrage cost parameter. As the series are pre-whitened, the intercepts are all mechanically zero and are not included in the model. I report the first-order autocorrelation coefficient, ϕ, and the standard deviation of the error term, σ_e^2, from the AR(1) maximum likelihood estimate. From this output, I calculate the cumulative impulse response function, the half-life and the frequency at spectrum zero.

Table 4.5 Estimates of the persistence of the value-to-price ratio, by arbitrage cost portfolio

Panel A: Full sample (1979–2004) estimates of persistence by arbitrage cost quintile ($T = 312$)

	Q1	Q2	Q3	Q4	Q5	Q5–Q1
ϕ	0.920	0.937	0.932	0.953	0.956	0.036
σ_e^2	0.061	0.088	0.129	0.158	0.226	0.166
CIR	12.54	15.85	14.63	21.06	22.81	10.268
Half-life	8.34	10.64	9.79	14.25	15.46	7.119
F(0)	0.58	1.95	3.57	11.03	26.70	26.119

Panel B: Early period (1979–1991) estimates of persistence by arbitrage cost quintile ($T = 156$)

	Q1	Q2	Q3	Q4	Q5	Q5–Q1
ϕ	0.834	0.830	0.845	0.866	0.903	0.069
σ_e^2	0.067	0.088	0.137	0.104	0.112	0.045
CIR	6.04	5.89	6.46	7.46	10.33	4.289
Half-life	3.83	3.72	4.12	4.81	6.81	2.977
F(0)	0.16	0.27	0.78	0.60	1.33	1.166

Panel C: Late period (1992–2004) estimates of persistence by arbitrage cost quintile ($T = 156$)

	Q1	Q2	Q3	Q4	Q5	Q5–Q1
ϕ	0.945	0.960	0.953	0.949	0.952	0.007
σ_e^2	0.051	0.085	0.116	0.194	0.300	0.249
CIR	18.25	25.23	21.31	19.70	20.78	2.523
Half-life	12.30	17.14	14.42	13.30	14.05	1.749
F(0)	0.85	4.58	6.11	14.58	38.75	37.900

Panel D: Differences in the persistence estimates between the early and late periods

	Q1	Q2	Q3	Q4	Q5	Q5–Q1
ϕ	0.111	0.130	0.108	0.083	0.049	–0.062
σ_e^2	–0.016	–0.003	–0.021	0.09	0.188	0.204
CIR	12.21	19.34	14.85	12.24	10.45	–1.766
Half-life	8.47	13.42	10.30	8.49	7.24	–1.228
F(0)	0.69	4.31	5.33	13.98	37.42	36.734

Notes: This table reports the estimates of persistence of the log value-to-price ratio, *Vf*(1)*P* (i.e., the one-period forecast-based residual income model to price ratio), for each arbitrage cost quintile portfolio. Each portfolio is made up of all firms which meet the available data requirements, with firms re-assigned on an annual basis. In Panel A, the model is estimated over the entire period 1979–2004, in Panel B for the period 1979–1991 and in Panel C, the potential 'bubble period' of 1992–2004. The AR(1) model is estimated using maximum likelihood with a pre-whitened time-series, (i.e., ϕ is estimated centred using the sample mean). F(0) is the frequency at spectrum zero and is estimated as $\sigma_\varepsilon^2/(1-\phi)^2$, where σ_ε^2 is the sample standard deviation of the model error term, CIR is the cumulative response function, estimated as $1/(1-\phi)$ and Half-life is calculated as $\ln(0.5)/\ln(\phi)$ and is measured in months.

Panel A reports the estimates of persistence using the full time-series (1979–2004). Consistent with Hypothesis 3, the difference in the persistence estimates between the high and low arbitrage cost portfolios is positive, implying that following a shock the disparity between price and value takes longer to return to its mean for high arbitrage cost firms. The difference in the half-life estimate for the high and low arbitrage cost portfolios is approximately seven months. Similar results are shown for the spectrum at frequency zero, where the estimate considers the volatility of the errors in the model.

In Panels B and C of Table 4.5, I split the sample into two equal time periods. In Panel B, I examine the persistence of *vf*(1)*p* for the period 1979–1991 and in Panel C, I investigate the period 1992–2004. Comparing the portfolios within periods yields further interesting results. First, in the earlier period (1979–1991), excluding the low arbitrage cost firms, the half-life of the value-to-price ratio is monotonically increasing for the remaining four arbitrage cost portfolios. The difference in the half-life of the high and low arbitrage cost portfolios is only 2.977 months. Adjusting the persistence measures by the volatility of the errors, i.e., using the spectrum at frequency zero, the difference is only 1.166. These results support Hypothesis 3, but suggest that when price and value are cointegrated, having a lack of close substitutes has only a minor impact on the timeliness of the removal of mispricing.

Surprisingly, in the 1992–2004 period the difference between the half-life of the high and low arbitrage cost portfolios declines to 1.749 months. This is due to the relatively small increase in the half-life of the high arbitrage cost portfolio. When comparing the portfolios using the spectrum at frequency zero, however, the difference in the high and low arbitrage cost portfolios is much larger in the later than in the earlier period. Taken together, these results provide some support for Hypothesis 3.

Consistent with a speculative bubble in the later period, the measures of persistence are all substantially higher in the later period (1992–2004) relative to the earlier period (1979–1991). This is due to the substantial shift upwards in the autocorrelation coefficient ϕ, towards one. This shift upwards in the persistence of value-to-price was not limited to the high-arbitrage cost portfolio. Instead, the results are consistent with the market-wide breakdown in cointegration reported in Chapter 3. There is an asymmetric shift in the volatility of the errors, σ_ε^2, with the lowest three quintiles of arbitrage cost having a lower level of σ_ε^2 in the 1992–2004 period than in the 1979–1991 period. The top two quintiles have a substantial increase in the level of σ_ε^2 in 1992–2004 relative to the earlier period. Recall that σ_ε^2 measures the lack of mean-reversion conditional on the parameter of the model. As such, these results imply that the run-ups in prices and idiosyncratic volatility in the late 1990s do not appear to be due to fluctuations in the value of the firm, which would be captured by a mean-reversion model.

4.6. Conclusions

The arbitrage mechanism conditions the extent to which accounting information is reflected in prices. If arbitrageurs are able to counter speculation then price deviations from the fundamentals of the firm will be short-lived. I investigate whether the lack of close substitutes (arbitrage cost) is associated with the level and duration of the disparity between price and value. I find that the disparity between price and value is associated with arbitrage cost only when the firm is relatively over-valued. This implies that arbitrage costs appear to matter most for firms which would require the arbitrageur to adopt a short position. I then show that the duration of the disparity between price and value is highest for the high arbitrage cost firms. The results are consistent with limits to arbitrage helping explain why markets can sustain a substantial run-up in price that is unrelated to the stock's fundamental value.

The results in this chapter are based on a single measure of arbitrage costs. The conclusions drawn from these tests are limited by the possibility that idiosyncratic variation does not provides a measure of arbitrage costs. In addition, to the extent that other market frictions, such as trading costs, price impact of a trade, the costs to short-selling, and liquidity costs, are not incorporated into this measure, the results may understate the importance of arbitrage costs. A potentially important avenue for future research is to examine alternative measures of arbitrage costs.

5. Robustness analysis

5.1. Synopsis

I present robustness analyses in two sections to address two potentially significant design differences between my thesis and the work of Lee et al. (1999). First, I compared shorter time-series for which the power of the Johansen test to detect cointegration may be biased downwards. I address this potential concern using a Monte Carlo simulation. Second, I used a different sampling procedure to that of Lee et al. (1999), as I included all firms with available data. I address the potential concern that the changing composition of firms biases the Johansen test from detecting cointegration by examining the sensitivity of the results across subsamples.

The remainder of this chapter is set out as follows. In Section 5.2 I examine the power of the Johansen tests for cointegration using a Monte Carlo simulation technique. In Section 5.3 I examine the sensitivity of the results across subsamples. I conclude in Section 5.4.

5.2. Analysis of the power of the cointegration test

5.2.1. Background

In this section, I examine the power of the test for cointegration, as the conclusion drawn in Chapter 3 (that there was a breakdown of the value-price relation in the late 1990s) could potentially be due to a low power test. Cointegration is often considered to be a phenomenon of long time-series. Any results that relate to a shorter period of time are a joint description of the power of the test as well as market behaviour. Hence, it could be claimed that in order to find a breakdown in the cointegrating relation, one can simply search for a short enough period of time and a breakdown will be found. This is a difficult objection to overcome, as the nature of a bubble is to eventually correct, or to crash, at some point in time in the future. In financial economics, a 'long time' might be if available information is not embedded within a day or earlier of information arriving (e.g., Chordia, Roll and Subrahmanyam, 2005). In the macro-economics literature the time-span for a cointegration model can cover multiple decades. The issue is that the power of the test to detect cointegration declines as the period examined is shortened.

The period examined in this study is from 1979 until 2004, a period of 26 years in total. This time span, when sampled at a monthly frequency, comprises 312 observations, or when split in half, two 13 year periods of 156 observations each. The main finding, presented in Chapter 3, is that for these equal periods, the earlier 'historical' period held the tendency to converge, while the later 'potential bubble' period did not. In this section, I examine the power of the cointegrating test by varying the number of observations (i.e., the span of the time period) included in the test.

5.2.2. Simulation set-up

The simulation set-up is as follows. I simulate two nonstationary time-series, v and p:

$$\begin{aligned} v_{mt} &= a_{10} + a_{11} p_{mt-1} + a_{12} v_{mt-1} + \psi_{1mt} , \\ p_{mt} &= a_{20} + a_{21} p_{mt-1} + a_{22} v_{mt-1} + \psi_{2mt} , \end{aligned} \tag{5.1}$$

where m indexes the simulation number and t is time. The set of parameters for the simulated process $\{a_{10}, a_{20}, a_{11}, a_{12}, a_{21}, a_{22}\}$ are the estimated values from Equation (3.20), for each of the three time periods or 'data-generating processes' (i.e., 1979–2004, 1979–1991 and 1992–2004). The innovations are simulated from the multivariate normal with a variance-covariance (Σ) equal to the estimate from the data-generating process (i.e., $\Psi = [\psi_{1t}\ \psi_{2t}]'$ and $\Psi \sim N(0, \Sigma)$). Each simulated time-series, m, is started with p_{m0} and v_{m0} both set equal to zero, and is simulated for a time-series length of t=1 to t = {76, 156, 312, 624} with an additional 20 observations discarded from the beginning of each simulated process.[44] For each of the four time-series lengths, the process is simulated 5000 times (i.e., m = 1 ... 5000), with independent draws from the multivariate normal distribution.

The simulated time-series are then used to simulate the distributions of the test-statistics and parameters of the following VECM:

$$\Delta y_{mt} = \Pi_m \, y_{mt-1} + \sum_{i=1}^{n-1} \Phi_i \Delta y_{mt-i} + \varepsilon_t \,, \tag{5.2}$$

where $y_{mt} = (v_{mt}, p_{mt})'$ and $\Pi_m = \alpha_m (\beta_{m1}, \beta_{m0})$ for the mth iteration of the bivariate time-series represented in Equation (5.1), where the VECM is estimated with a single lag of Δy_{mt}.

5.2.3. Small-sample estimates of the test of no cointegration

In Chapter 3, I presented evidence that the null of no cointegration was rejected for the historical period (1979–1991) but not rejected for the 'potential bubble' period (1992–2004). This breakdown in the cointegrating relation between value and price was presented as evidence that during the speculative period of the 1990s, the general level of market mispricing was not removed in a timely manner. The 13-year period for which the breakdown in the cointegrating relation is found could be considered by some readers as 'short.' In this section, I present simulation results that investigate the power of the test for the length of time required to find cointegration.

I use the Monte Carlo set-up described in the prior section. First, I consider the 'true' model to be the historical model, and I iterate the bi-variate time-series based on the parameters

[44] i.e., t_0 is preceded by 20 observations $\{t_{-20}, t_{-19}, \ldots , t_{-1}\}$ and the retained part of the process is for periods of 6, 13, 26 and 52 years, respectively, with 12 monthly observations per year.

of the historical (1979–1991) VAR model and covariance of the innovations from that model. The aim of the analysis is to estimate the statistical properties of a 'short' time-series, given that the 'true' model is cointegrated. I run the model for time-series lengths (spans) of T equal to (i) T=72 (6 years) and (ii) T=156 (13 years). I then repeat the process with the potential bubble period (1992–2004) and the full period (1979–2004). The aim of these tests is to determine the frequency of failing to reject the null of no cointegration. When the model is expected to be cointegrated, the simulations should reject the 95% critical values 95% of the time. A low power test would reject the null less often than 95% and an upwardly biased test would reject the null more often. For the two models that are not expected to be cointegrated, I provide two tests. First, the frequency of rejecting the critical values (for which they are expected to reject about 5% of the time) and second, the frequency that the test statistic is higher than the empirical test statistics reported in Chapter 3.

Table 5.1 Monte Carlo evaluation of the power of the cointegration test

		Simulation period		
		(i) 1979–1991	(ii) 1992–2004	(iii) 1979–2004
$T = 72$	P[λ(*trace*)>λ^c]	0.548	0.079	0.078
	P[λ(*max e.v.*) >λ^c]	0.549	0.078	0.077
	P[λ((*T–np*)/*T max e.v.*) >λ^c]	0.465	0.059	0.056
$T = 156$	P[λ(*trace*) >λ^c]	0.993	0.158	0.095
	P[λ(*max e.v.*) >λ^c]	0.995	0.180	0.112
	P[λ((*T–np*)/*T max e.v.*) >λ^c]	0.994	0.167	0.094
$T = 312$	P[λ(*trace*) >λ^c]		0.477	0.210
	P[λ(*max e.v.*) >λ^c]		0.549	0.230
	P[λ((*T–np*)/*T max e.v.*) >λ^c]		0.532	0.219
$T = 156$	P[λ(*trace*) > 1.88]		0.959	
	P[λ(*max e.v.*) > 2.46]		0.999	
$T = 312$	P[λ(*trace*) > 4.88]			0.994
	P[λ(*max e.v.*) > 4.88]			0.983

Notes: The results in this table are based on simulated, not actual, data. I generated 5,000 replications of the time-series of value and price and estimated Equation (5.2) for each of the simulated time-series. T is time-series length of the simulated data (i.e., the number of simulated observations). P[] is the percentage of simulated test statistics that are greater than the asymptotic critical value, *trace* is the Johansen trace-test statistic for which the asymptotic critical value $\lambda^c = 18.69$, *max e.v.* is the maximum eigenvalue test statistic for which the asymptotic critical value $\lambda^c = 11.59$. For $T = 156$ and $T = 312$, I report the percentage of simulated test statistics that are greater than the reported test statistics for the potential bubble period (1992–2004) and full period (1979–2004), respectively. I generate data based on the parameters and variance-covariance matrix of the innovations estimated using the actual data for three distinct time periods, (i) 1979–1992, (ii) 1992–2004, and (iii) 1979–2004. The estimates of the VAR model based on actual data are reported in Table 3.3, Panel A.

I report the results of the simulations in Table 5.1. In each of the three columns I report the tendency of the model to reject the null of no cointegration at the 95% level of significance. The 'no bubble period' model reported in Column (1), is expected to be cointegrated, as the underlying model reported in Chapter 3 was cointegrated. The simulated results should therefore reject the null approximately 95% of the time. In the Columns (2) and (3) I report the simulations for the 'potential bubble period' and the 'full period' respectively. Neither model is found to be cointegrated in Chapter 3. The simulated data should, therefore, reject the null approximately 5% of the time. If the tests are of low power, then they should all under-reject the null.

The first set of simulations reported are for the time-series length of T = 72. Not surprisingly, in Column (1), for the no bubble period, there is a tendency to under-reject the null with only 54.9% of the series rejecting the null of no cointegration using the maximum eigenvalue statistic. This result is consistent with the small size distortions reported in Podivinsky (1998) for samples of less than 100. The bubble and full periods, however, appear to overreject the null based on the assumption that the model is not cointegrated. Using the maximum eigenvalue statistic, I find that 7.8% and 7.7% of the time-series reject the null in the bubble and full periods, respectively. Similar results are found for the simulated trace statistic. I also report an adjusted maximum eigenvalue statistic which applies a correction for the degrees of freedom, $(T-np)/T$, where T is the sample size, n is the number of parameters (i.e., 6) and p is the number of lags included in the model (i.e., 1). Not surprisingly, the proportion of simulations that reject the null declines relative to the maximum eigenvalue statistic. Specifically, I find that 5.9% and 5.6% of the time-series reject the null in the bubble and full periods, respectively.

I then report the simulated results for the T = 156 and T = 312 sample sizes. Importantly, in Column (1), for the no bubble period, there is a no tendency to under-reject the null with

99.5% of the series rejecting the null of no cointegration using the maximum eigenvalue statistic. The potential bubble and full periods, again overreject the null based on the assumption that the model is not cointegrated, with 18% and 11.2% of the simulated time-series rejecting the null of no cointegration according to the maximum eigenvalue statistic, for the potential bubble and full periods respectively. For T = 312, I only report the results for the full period. I again find an overrejection of the null, with 23% of the simulated time-series rejecting the null according to the maximum eigenvalue statistic.

In addition to the statistics reported above, I also report the Monte Carlo rank p-value for the trace and maximum eigenvalue statistics for the results in Chapter 3. If the tests are of low power, then the asymptotic critical value is potentially inappropriate, being too high a hurdle to infer cointegration between price and value. I therefore report the ranking of the actual trace-test and maximum eigenvalue statistics relative to the simulated distribution. The rank measure can be interpreted in the same way as a normal p-value (i.e., p<0.05 shows that the statistic is in the far right tail of the distribution, and implies 'significance'). Comparing the maximum eigenvalue statistic estimated from the potential bubble period to the distribution of simulated maximum eigenvalue statistics for T = 156, the Monte Carlo p-value is 0.999, which is clearly not significant, and instead lies in the far left tail of the distribution. A similar result is found by comparing the full-period to the distribution of T = 312, where the Monte Carlo p-value is equal to 0.983, again implying that it is not significant.

Taken together, these results imply that the power of the test does not drive the results reported in Chapter 3. Specifically, the Monte Carlo based results reported here imply that the test should have sufficient power to detect cointegration when it is present.

5.2.4. Monte Carlo based tests for the difference in parameters

In the prior section, I did not find that the results are driven by low power tests. This implies that the differences between the periods are due to the data generating processes that underlie the different time-periods. In this section, I display histograms of the long-run association between price and value and the error correction parameter for price. I graph the distributions of these parameters from the extended time-period simulations of $T = 624$. Similar results are found for smaller simulated time-series lengths, but the distributions are more dispersed.

In Figure 5.1, I display the long-run relation between price and value, β_1, for the three periods. I also indicate the estimates of β_1 from Chapter 3 (no bubble = 0.87, bubble = 1.67 and full = 1.05 and the implied β_1 of one from the ratio), as vertical dotted lines. The shapes of the distributions for the three time periods are strikingly different. The estimates of β_1 for the no bubble case are tightly clustered around the mean, with the bulk of the distribution being greater than one, implying that the estimated beta was biased downwards for the no-bubble period. The estimates simulated from the full period are flatter, but the mean of the distribution is closer to the estimated mean. Again the bulk of the distribution is above one, but there is a substantial tail below one. The estimates from the bubble period are very spread out, with a considerable mass in the tails of the distribution which stretches from below one to above three. These graphical results suggest significant differences between the bubble and no bubble periods.

In Figure 5.2, I display the error correction parameter for price, α_{mp}. There is a stark difference between the cointegrated no bubble period and the bubble and full periods. When price is correcting for the disparity between value and price, the bulk of the distribution lies above zero. Clearly this is only true for the simulated time-series based on the no bubble period.

As the bulk of the distribution lies below zero for the bubble period, this further supports the contention that the bubble was due to an increase in momentum trading at the expense of information trading.

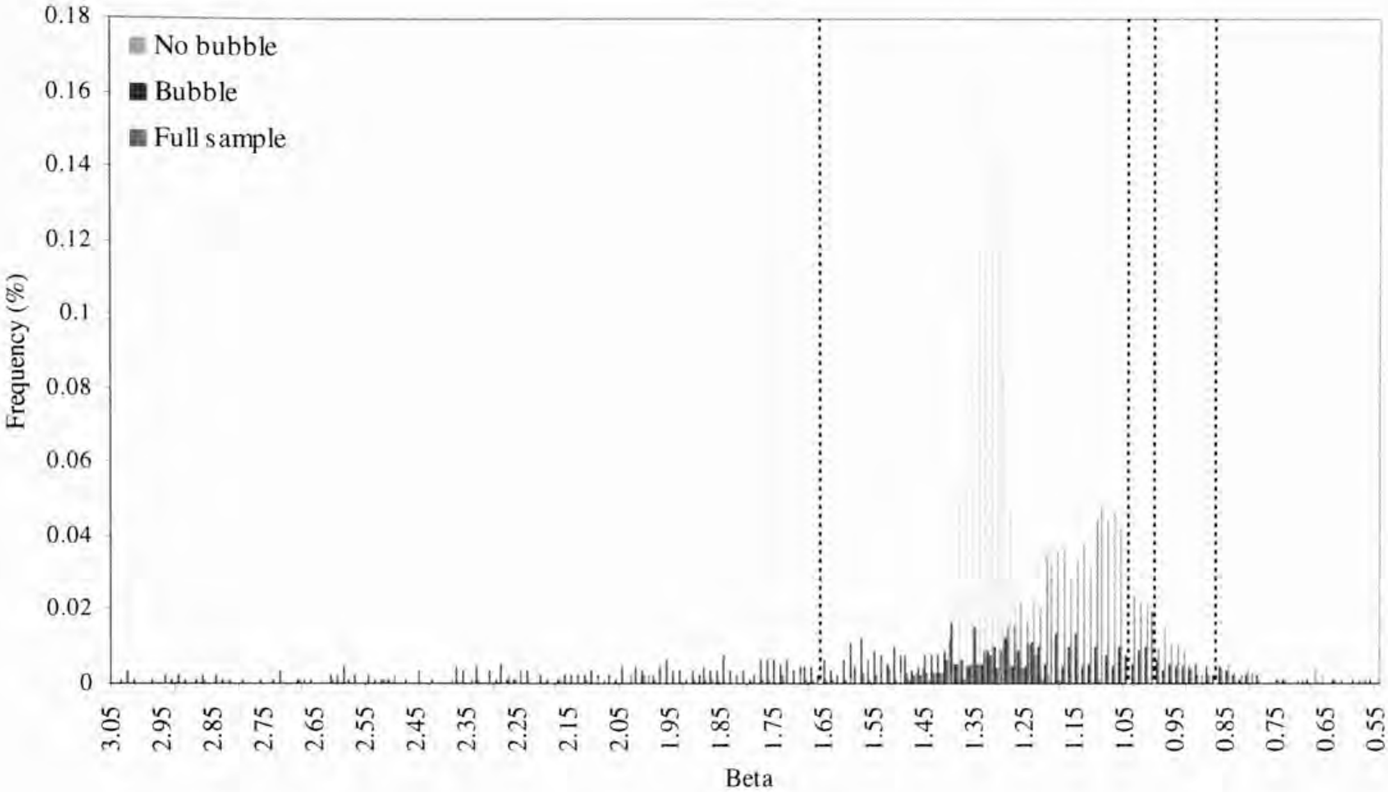

Figure 5.1 Histogram of simulated long-run association between value and price

This histogram graphs simulated, not actual, data. I generated three samples of value and price time-series (see Equation (5.1)) with 5,000 replications for each sample for a time-series length of T = 624. For each time-series I estimated Equation (5.2), and report the distribution of the estimates of β_{m1} above. The three samples are based on the parameters of the VAR model estimated with actual data for the three time periods 1979–1991, 1992–2004 and 1979–2004 (see Table 3.3, Panel A). I label the periods as (i) the no bubble period (1979–1991), (ii) the potential bubble period (1992–2004), and (iii) the full-sample period (1979–2004). The dotted vertical lines show the estimates of β_1 reported in Table 3.3 and one, which is the restricted value of β_1 under the ratio case.

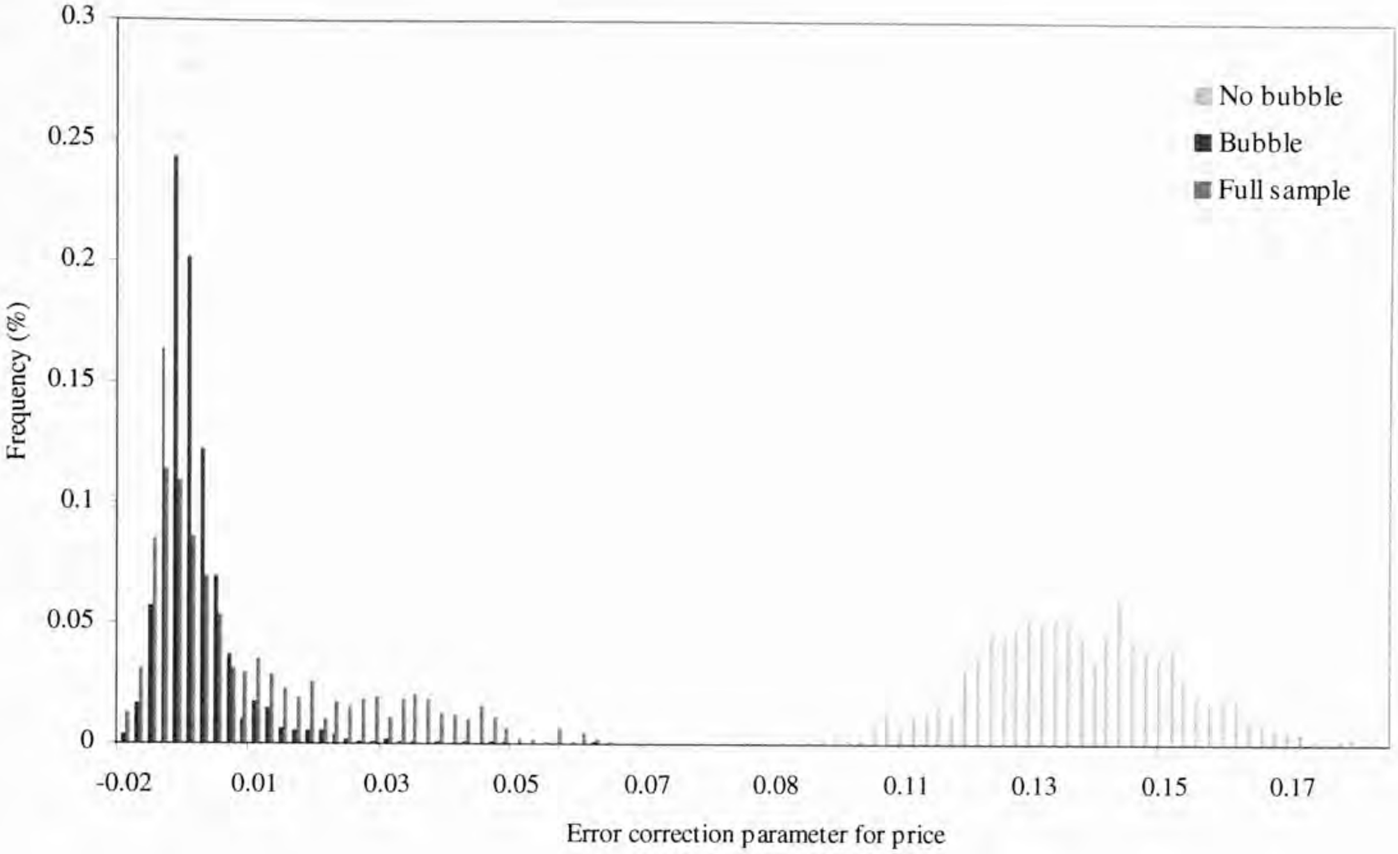

Figure 5.2 Histogram of simulated error-correction parameter for price

This histogram graphs simulated, not actual, data. I generated three samples of value and price time-series (see Equation (5.1)) with 5,000 replications for each sample for a time-series length of T = 624. For each time-series I estimated Equation (5.2), and report the distribution of the estimates of α_{mp} above. The three samples are based on the parameters of the VAR model estimated with actual data for the three time periods 1979–1991, 1992–2004 and 1979–2004 (see Table 3.3, Panel A). I label the periods as (i) the no bubble period (1979–1991), (ii) the potential bubble period (1992–2004), and (iii) the full-sample period (1979–2004). For the model to be cointegrated by price anchoring to value, the distribution of the error-correction parameters should lie above one.

5.3. *Robustness of the results to alternative sampling procedures*

In the analysis presented in Chapter 3, I used the most general sample available. This could give rise to the concern that the results are due to the changing composition of the sample. In this section, I report results for alternative sampling procedures. First, I display the robustness of the aggregation procedure by reporting the results for a constant sample of firms. Second, I examine the generality of the results to the S&P 500 and compare the firms listed on the NASDAQ with those listed on the NYSE/AMEX.

5.3.1. Constant sample results

In this section, I examine whether the lack of cointegration between price and value is due to the changing composition of the sample, by estimating the cointegration tests using a constant sample. The use of a constant sample has the following features: (i) there is no variation in the price-value relation due to the entry and exit of firms over the testing period, however, (ii) this method imposes a survivorship bias.[45]

I find that during this period there are 778 firms with available data for the entire 1979–2004 period (in Chapter 3, the average is over 3,000 firms). In Figure 5.3, I graph the price-to-value and market-to-book ratios. The price-to-value ratio appears almost identical to that displayed in the graph presented in Chapter 3.

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[45] While I hold the sample of firms constant, it also true that firms evolve over time so that some change over 26 years is inescapable.

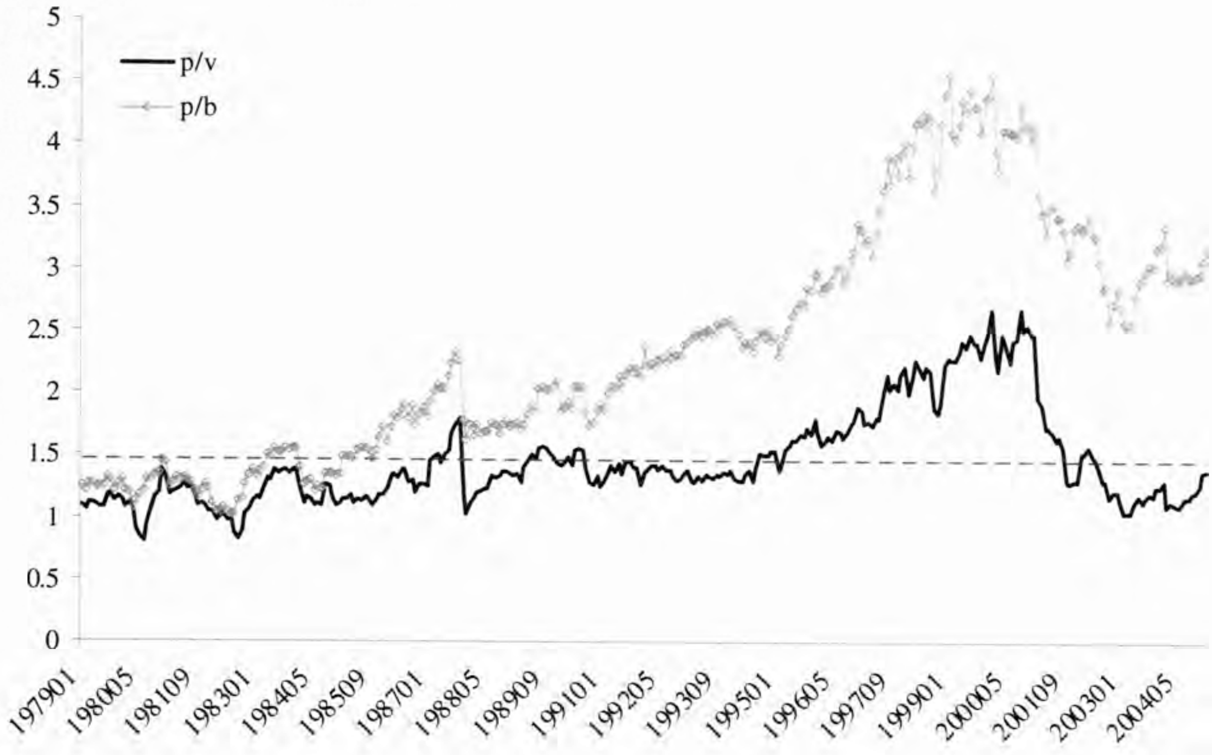

Figure 5.3 The ratio of market to residual income value and the ratio of market to book-value for a constant sample of firms.

This graph shows the market to residual income value and the market to book ratio for all firms with available data that also survived the entire period from January 1979 to December 2004. The monthly estimates of residual income value and market value are aggregated to form a portfolio measure, the ratio is based on the aggregate market value divided by the aggregate accounting-value using month end data. The horizontal line shows the sample mean of the market to residual income value ratio of 1.46.

In Table 5.2, I present similar results for the constant sample to those reported in Table 3.2 of Chapter 3. The results show the same marked difference in the time-series of the price and value relation. Specifically, there is strong evidence of cointegration in the historical period (1979–1991) and no reliable evidence of cointegration for either (i) 1992–2004 or (ii) 1979–2004.

Table 5.2 Tests for cointegration using a constant sample of firms

Panel A: Phillips-Perron tests without time-trend

	(i) Jan 1979 – Dec 1991 T=156		(ii) Jan 1992 – Dec 2004 T=156		(iii) Jul 1979 – Dec 2004 T=312	
	Z-rho (Z_ρ)	Z-tau (Z_τ)	Z-rho (Z_ρ)	Z-tau (Z_τ)	Z-rho (Z_ρ)	Z-tau (Z_τ)
vp_t	−16.79*	−2.96*	−3.74	−1.36	−9.78	−2.28
(p-value)	(0.02)	(0.04)	(0.56)	(0.60)	(0.14)	(0.18)
bp_t	−2.87	−0.98	−3.90	−1.73	−2.33	−1.25
(p-value)	(0.67)	(0.76)	(0.22)	(0.32)	(0.74)	(0.66)

Panel B: Phillips-Perron tests including a time-trend

	(i) Jan 1979 – Dec 1991 T=156		(ii) Jan 1992 – Dec 2004 T=156		(iii) Jul 1979 – Dec 2004 T=312	
	Z-rho (Z_ρ)	Z-tau (Z_τ)	Z-rho (Z_ρ)	Z-tau (Z_τ)	Z-rho (Z_ρ)	Z-tau (Z_τ)
vp_t	−31.44**	−4.08*	−3.78	−1.39	−13.01	−2.49
(p-value)	(p<0.01)	(0.01)	(0.90)	(0.86)	(0.26)	(0.33)
bp_t	−21.07*	−3.31	−3.31	−1.30	−10.81	−2.20
(p-value)	(0.05)	(0.07)	(0.92)	(0.88)	(0.38)	(0.49)

Notes: This table reports the results of the Phillips-Perron unit root tests on the value-to-price ratio using a constant sample of firms. The unit root tests are performed without a time-trend in Panel A and with a time-trend in Panel B:

$\Delta vp_t = a + (b-1)vp_{t-1} + \sum_{i=1}^{12} c_i \Delta vp_{t-i} + u_t$, $\Delta vp_t = a + (b-1)vp_{t-1} + \delta t + \sum_{i=1}^{12} c_i \Delta vp_{t-i} + u_t$, where, vp_t refers to the log value-to-price ratio. The test is of the null of a unit root in the time-series (b=1). The two test statistics from the Phillips-Perron test are an adjusted regression coefficient Z-rho (Z_ρ) and an adjusted t-statistic Z-tau (Z_τ). ** p<0.01, * p<0.05, $^\delta p$<0.1

5.3.2. A comparison of NYSE and NASDAQ listed firms on the S&P 500

In this section, I examine whether the breakdown in cointegration is due solely to the inclusion of NASDAQ listed firms in the value-weighted index. These tests are important, as a commonly held belief is that only those firms listed on the NASDAQ contained bubbles (see, for example, Pastor and Veronesi, 2006). I show that while the disparity between price and value was highest for NASDAQ listed firms, the breakdown is more general in nature.

In Figure 5.4, I display the dramatic level of disparity for the NASDAQ listed firms relative to the NYSE listed firms. As a guide, the average price-to-value ratio is 4.62 for NASDAQ listed firms in the S&P 500 and 2.60 for the NYSE listed firms. This difference is significant at the $p<0.001$ level using either a standard t-test for the means with unequal variance or a non-parametric test based on the medians. The maximum peak for the time-series of price-to-value ratio for the NASDAQ portfolio is 21.42, which was reached in December 1999. The NYSE portfolio also reached its maximum peak in December 1999, which was 5.31. For the period 1979–1991, the price-to-value ratio had a mean (standard deviation) of 2.46 (0.86) and 1.73 (0.39) for the NASDAQ and NYSE portfolios respectively. Ex-ante, using the historical period, this implies that the peaks of the NASDAQ and NYSE diverged 21.94 and 9.19 standard deviations from their historical means, respectively.

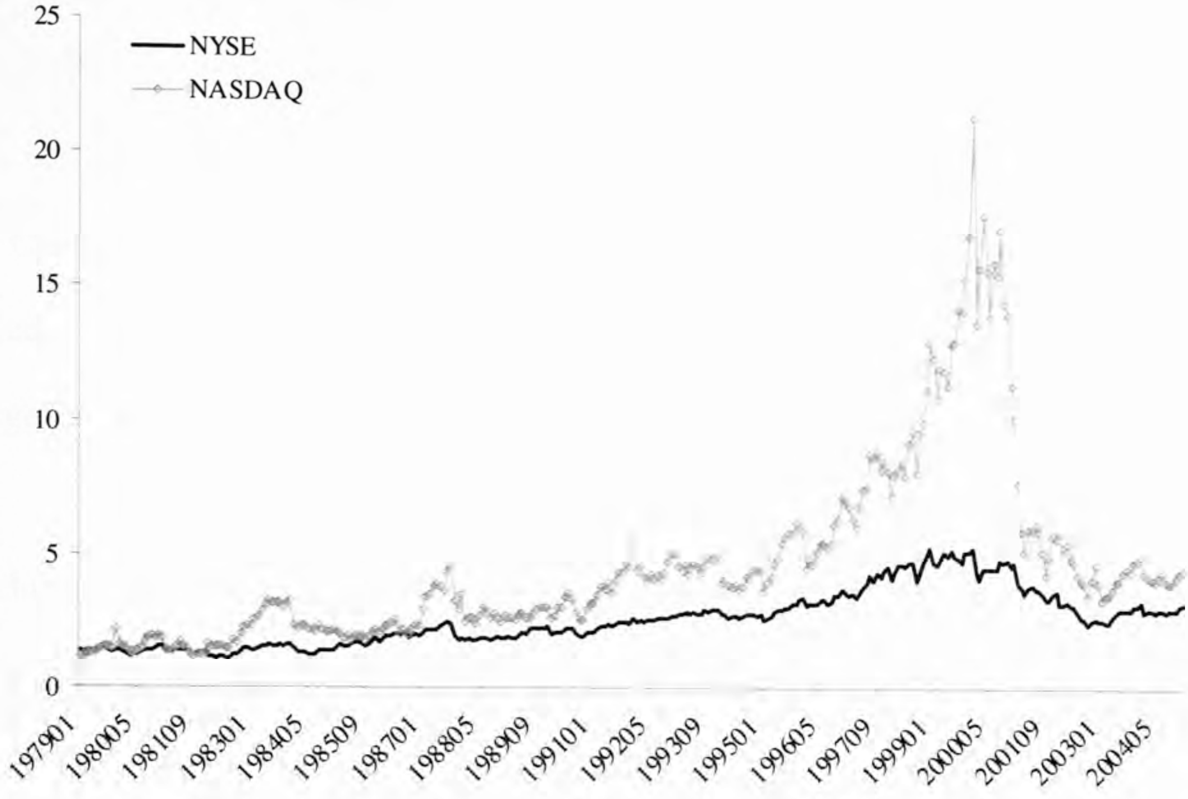

Figure 5.4 The price-to-value ratio for the S&P 500 comparing NYSE and NASDAQ firms

This graph displays the market to residual income value and the market to book ratio for all firms with available data in the S&P 500 from January 1979 to December 2004. The firms are assigned to portfolios based on where their shares are listed (NYSE or NASDAQ). The monthly estimates of residual income value and market value are aggregated to form a portfolio measure. The portfolio's ratio is the aggregate market value divided by the aggregate accounting-value using month end data.

In Table 5.3, I report the cointegration test for the S&P 500 firms and a comparison of NASDAQ and NYSE listed S&P 500 firms. I find that the NYSE listed portfolio does not reject the null of no cointegration in the 1990s, suggesting that the bubble-like behaviour documented in Chapter 3 is not solely due to firms listed on the NASDAQ having significantly inflated prices. Instead these results point to the bubble prices being generally consistent across the largest companies listed on either market.

During the period 1979–1991, both the NYSE listed and the NASDAQ listed firms included in the S&P 500 index reject the null of no cointegration between price and value, when including a time-trend (Panel B). The results are weaker when a time-trend is not included (Panel A), with the NYSE not displaying reliable evidence of cointegration for any of the periods, and the NASDAQ displaying some weak evidence of cointegration over the full period. In the potential bubble period (1992–2004), both the NYSE and the NASDAQ listed firms consistently do not reject the null of no cointegration between price and value regardless of the specification of the time-trend.

Table 5.3 Tests for cointegration for S&P 500 firms, S&P 500 listed on the NYSE, and S&P 500 firms listed on the NASDAQ

Panel A: Phillips-Perron tests without time-trend

	(i) Jan 1979 – Dec 1991 T=156		(ii) Jan 1992 – Dec 2004 T=156		(iii) Jul 1979–Dec 2004 T=312	
	Z-rho (Z_ρ)	Z-tau (Z_τ)	Z-rho (Z_ρ)	Z-tau (Z_τ)	Z-rho (Z_ρ)	Z-tau (Z_τ)
S&P 500						
vp_t	−11.38$^\delta$	−2.38	−3.42	−1.32	−6.87	−1.94
(p-value)	(0.09)	(0.15)	(0.60)	(0.62)	(0.28)	(0.31)
NYSE						
vp_t	−11.56$^\delta$	−2.41	−3.22	−1.26	−7.45	−2.01
(p-value)	(0.09)	(0.14)	(0.63)	(0.65)	(0.25)	(0.28)
NASDAQ						
vp_t	−12.37$^\delta$	−2.68$^\delta$	−6.24	−1.78	−11.94$^\delta$	−2.74$^\delta$
(p-value)	(0.07)	(0.08)	(0.32)	(0.39)	(0.08)	(0.07)

Panel B: Phillips-Perron tests including a time-trend

	(i) Jan 1979 – Dec 1991 T=156		(ii) Jan 1992 – Dec 2004 T=156		(iii) Jul 1979–Dec 2004 T=312	
	Z-rho (Z_ρ)	Z-tau (Z_τ)	Z-rho (Z_ρ)	Z-tau (Z_τ)	Z-rho (Z_ρ)	Z-tau (Z_τ)
S&P 500						
vp_t	−30.72*	−4.06*	−3.35	−1.31	−10.06	−2.17
(p-value)	(0.01)	(0.01)	(0.92)	(0.88)	(0.43)	(0.51)
NYSE						
vp_t	−30.30*	−4.03*	−3.38	−1.33	−10.03	−2.17
(p-value)	(0.01)	(0.01)	(0.92)	(0.88)	(0.43)	(0.51)
NASDAQ						
vp_t	−20.57*	−3.37$^\delta$	−6.16	−1.71	−19.08$^\delta$	−3.13
(p-value)	(0.05)	(0.06)	(0.73)	(0.74)	(0.08)	(0.10)

Notes: This table reports the results of the Phillips-Perron unit root tests on the value-to-price ratio using samples of firms categorised by market of listing being either the (i) NYSE or (ii) the NASDAQ. The unit root tests are performed without a time-trend in Panel A and with a time-trend in Panel B: $\Delta vp_t = a + (b-1)vp_{t-1} + \sum_{i=1}^{12} c_i \Delta vp_{t-i} + u_t$, $\Delta vp_t = a + (b-1)vp_{t-1} + \delta t + \sum_{i=1}^{12} c_i \Delta vp_{t-i} + u_t$, where, vp_t refers to the log value-to-price ratio. The test is of the null of a unit root in the time-series (b=1). The two test statistics from the Phillips-Perron test are an adjusted regression coefficient Z-rho (Z_ρ) and an adjusted t-statistic Z-tau (Z_τ). $^{**}p$<0.01, $^{*}p$<0.05, $^\delta p$<0.1

5.4. Conclusions

Robustness analyses were performed to examine the two significant design differences between my thesis and the work of Lee et al. (1999). First, Monte Carlo results suggest that the Johansen tests have sufficient power to detect cointegration for the time-series lengths considered in Chapter 3. Second, by examining the sensitivity of the results across subsamples, the results are robust to variation in the sampling technique used.

6. Conclusion

I examine the extent to which accounting information is reflected in market prices at different points in time, using accounting based value as a summary measure of value-relevant accounting information. The Efficient Market Hypothesis predicts that price reflects all available value-relevant accounting information, based on rational arbitrageurs being unimpeded. I present two complementary empirical studies which are motivated by potential shortcomings in this view.

The traditional asset pricing literature establishes that if investors rationally price a firm based on it's intrinsic value, then the expected time-series relation between price and value is one of cointegration. Even if some investors are not rational, competition amongst arbitrageurs is expected to align price with intrinsic value and maintain market efficiency. Prior literature has heavily debated whether or not the market is efficient, with the debate intensifying over the potential existence of an asset price bubble in the late 1990s. Recent research has shown that arbitrage is not without risk in the equity market (e.g., Shleifer and Vishny, 1997) and LeRoy and Porter (1981), Shiller (1989) and Lee, Myers and Swaminathan (1999) illustrate that the variation in stock prices cannot be completely explained by the variation in fundamentals over time, suggesting that some component of stock prices is potentially due to mispricing.

A review of prior literature highlights two hitherto unexplored questions which I address in my thesis. First, "Did price and accounting based value continue to be cointegrated during the late 1990s, where there was a potential asset price bubble?" Second, "Do costs to arbitrage affect the time-series relation between price and value?"

I present the main empirical analyses in Chapters 3 and 4. In Chapter 3, I contrast the time-series relation of the late 1990s market with the historical relation. The late 1990s is an interesting period to investigate as the rationality of investors during this period has been questioned. I show that there was indeed a breakdown in cointegration in the late 1990s. I then present an alternative model of cointegration which allows for the identification of the 'equilibrium correction parameters' to identify the driver of cointegration. Historically, price tended to correct for the disparity from fundamentals but during the late 1990s this correction was not reliably observed. I then test for potential measurement error in the expected rate of return and the growth rate in fundamentals.

I analyse the role of costs to arbitrage in Chapter 4. Specifically, I compare the time-series relation between price and value for portfolios of firms ranked on the relative availability of close substitutes, necessary for hedging the arbitrage trade. Based on the model of Abreu and Brunnermeier (2002), I hypothesize that arbitrage costs increase the level of mispricing and the persistence of mispricing. As the firms which are the most difficult to arbitrage will also likely be the most difficult to value, I compare the level and persistence of arbitrage cost portfolios of value-to-price. I find results consistent with the Abreu and Brunnermeier (2002) model, where the persistence is greater for high arbitrage cost firms. I find asymmetric results for association between the difficulty to arbitrage and the level of value-to-price, as only over-valued firms appear to be affected by arbitrage costs. These results are consistent with limits to arbitrage influencing the extent to which accounting information is reflected in price, especially when a short position is required.

In Chapter 5 I provide robustness tests. Simulating the value-to-price relation suggests that finding a cointegration breakdown is not due to the power of the test. I then show that the results in Chapter 3, which include all firms with available data, are not due to the sample selection method, either. Both a constant sample and the S&P 500 index display similar breakdowns in cointegration during the late 1990s. I also show that contrary to popular opinion, the breakdown in cointegration is not limited to NASDAQ firms, but was also evident for firms listed on the NYSE.

The evidence in this thesis suggests that the speculative bubble was market-wide, not simply limited to small firms with a small analyst following, or just to the technology sector. Evidence of a market-wide bubble has significant implications for the asset management industry as well as academics. Asset allocation is premised upon the assumption that markets are efficient, hence, investing in the index is considered 'safe' as equities are less risky in the long-run. Future research could further examine this important asset allocation problem when the market index can be mispriced.

The results suggest several avenues for future research. In asset pricing, a descriptively accurate model will require regimes when returns are not an accurate reflection of the compensation for risk, but instead contain fluctuations in the level of mispricing. In the value-relevance literature, a more descriptively complete model will require a role for time-varying levels of mispricing. Finally, as arbitrageurs are expected to align price with fundamental value, further study of both the accounting and market features that aid or hinder the activities of arbitrage potentially are important areas for future research.

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