On the estimation of the asymmetric timeliness of earnings: Inference and bias corrections*

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ABSTRACT

Recent studies have raised concerns about the validity of cross-group comparisons of regression coefficients capturing the asymmetric timeliness of earnings, a form of accounting conservatism. In the context of an econometric model, we show that, while regression coefficients correctly identify whether or not earnings are conservative, the inclusion of future rents in equity values causes the regression coefficients to mismeasure the degree of conservatism. This mismeasurement can bias cross-sectional comparisons of the degree of conservatism; however, a test based on a ratio of regression coefficients allows for valid cross-sectional comparisons. We empirically validate our predictions and show that the ratio measure is fairly stable over time and is positively associated other conservatism measures that capture asymmetric timeliness. We also note that the use of coefficient ratios may control for bias in other settings, as well.

I. INTRODUCTION

A large and growing body of accounting research studies the causes and consequences of the tendency of accounting earnings to recognize losses earlier than gains (Watts 2003a, 2003b). Researchers often measure this tendency, also referred to as asymmetric timeliness or conditional conservatism, using Basu's (1997) regression test of earnings on returns. A number of empirical findings depend on the validity of this measure of conservatism (hereafter 'asymmetric timeliness coefficient'). In particular, many studies compare asymmetric timeliness between time periods or between firm characteristics.[1]

Givoly, Hayn and Natarajan (2007, p. 66) provide evidence that "inferences regarding the variation in conservatism ...cannot be reliably made" using the regression coefficient as in Basu (1997). Roychowdhury and Watts (2007) maintain that book values reflect separable assets while returns incorporate information about both separable assets and rents. This induces a bias in the asymmetric timeliness coefficient and produces a negative association between the coefficient and market-to-book ratios, a measure of the cumulative effect of accounting conservatism. In this study, we demonstrate that a simple test that uses a ratio of regression coefficients provides valid cross-sectional comparisons of asymmetric timeliness.

We derive the properties of both Basu (1997) regressions and the ratio measure of asymmetric timeliness within a clearly stated econometric model.[2] The model allows for skewed distributions, which Dietrich, Muller and Riedl (2007) have demonstrated may bias asymmetric timeliness tests, and allows for earnings to contribute to the information reflected in returns. The

[1] Examples of time series comparisons include Basu (1997) and Givoly and Hayn (2000). Comparisons based on firm characteristics include cross-country comparisons (e.g. Ball, Kothari and Robin, 2000), large versus small audit firms (e.g. Basu, Hwang and Jan, 2001), and high- versus low-tech firms (e.g. Chandra, Wasley and Waymire, 2004).

[2] See the contemporaneous paper by Ball and Kothari (2007) for an alternative econometric model.

model maps the book and market values of a firm, where market values are composed of separable assets and the expected value of future rents and growth opportunities.[3] The book value incorporates the recognized value of separable assets. Within the context of our model, we show that regression tests of asymmetric timeliness correctly identify the presence of conservatism, but incorrectly measure the degree of conservatism. This results in biased cross-sectional comparisons of asymmetric timeliness using regression coefficients; however, we show that researchers can nonetheless perform valid cross-sectional comparisons using a ratio of coefficients.

Our econometric model implies that the effect of rents on returns dampens the association between earnings and returns, consistent with the intuition of Roychowdhury and Watts (2007). Thus, firms with high market-to-book ratios, which presumably have a large portion of value and returns unrelated to separable assets, have low asymmetric timeliness coefficients. In contrast, we find that such firms have higher asymmetric timeliness when we measure it with ratio statistics. We find that firms with market-to-book ratios in the top third of our sample have a higher ratio coefficient than firms in the bottom third (significant at 1%) even though they have asymmetric timeliness coefficients that are lower (significant at 1%). This suggests that the ratio is effective in controlling for the bias induced by the effect of rents on returns.

Because we make several assumptions in developing our econometric model, we derive and test several additional predictions in order to assess its empirical validity. We predict and find that interacting returns with the book-to-market ratio increases asymmetric timeliness coefficients, which is due to a predicted bias. We predict a positive bias on the ratio measure, as well, but the test of that prediction is statistically insignificant. We predict and verify that

[3] This assumption relates to Roychowdhury and Watts' (2007) development of Watts' (2003a) theory that accounting conservatism applies to the book recognition of the value of a firm's separable net assets.

scaling bias accounts for the negative asymmetric timeliness coefficient in the unscaled regressions in Dietrich, Muller and Riedl (2007). Roychowdhury and Watts (2007) state that the inclusion of goodwill violates the assumption that book values only reflect separable assets. We predict and verify that goodwill biases asymmetric timeliness coefficients upward. We also predict a similar effect on the ratio measure, but fail to find statistically significant evidence in support of that prediction.

We perform further tests to validate the ratio measure. In order to address concerns that the ratio measure may have poor econometric properties, we report confidence intervals based on both the delta method and Fieller's Theorem (Fieller, 1954), which is a more robust means to determine confidence intervals for ratios of regression coefficients, particularly in small samples. Our results suggest that, at least in large samples, the delta method confidence intervals approximate the Fieller confidence intervals very well. We also verify that the ratio measure exhibits a predicted positive association with the market-to-book ratio, which partly reflects the cumulative effects of asymmetric timeliness, and a negative association with Givoly and Hayn's (2000) cumulative discretionary accrual measure, which declines with the cumulative effects of accruals that asymmetrically anticipate bad news. In contrast, Givoly, Hayn and Natarajan (2007) find that the asymmetric timeliness coefficient has the opposite associations with these measures. Finally, we find that the ratio measure exhibits much more time series stability than the asymmetric timeliness coefficient, with a coefficient of variation of 0.124 for years after 1975 versus 0.428 for the coefficient. Thus, the ratio appears to better reflect the stability of accounting practices than the asymmetric timeliness coefficient and is well-behaved in the sense that it has an intuitive association with other measures that reflect asymmetric timeliness.

Pope and Walker (1999) derive a ratio-based test of asymmetric timeliness in a permanent income model with risk neutral investors.[4] In their context, the ratio of coefficients controls for cross-sectional differences in discount rates, which they use in a cross-country analysis.[5] Our model is distinct from theirs in that we allow for the distinction between separable assets and future returns and a more general relation between returns and earnings that does not rely on Pope and Walker's (1999) permanent income feature critiqued by Basu (1999). We also derive and test several implications of our model in order to assess its empirical validity. Furthermore, we verify the econometric properties of the ratio measure by comparing asymptotic, delta method confidence intervals to Fieller method confidence intervals, which addresses the econometric concerns raised by Basu (1999) in his discussion of Pope and Walker (1999).[6]

Roychowdhury and Watts (2007) propose extending the returns measurement horizon to, for example, three years as a means to mitigate the effect of rents on measures of asymmetric timeliness. Beatty (2007) notes that this approach reduces sample sizes due to data availability requirements and that the requirement of negative returns for a multi-year period reduces the number of bad-news observations. Moreover, Givoly, Hayn and Natarajan (2007) discuss how the aggregation of both positive and negative news events reduces the power of asymmetric timeliness tests. The use of multi-year returns exacerbates this effect. Lastly, while extending the measurement horizon may reduce bias, Roychowdhury and Watts (2007) do not discuss whether or not this approach allows for valid cross-sectional comparisons. The aggregation effect also impacts the annual regressions we use in the ratio test; however, the integral approach to interim reporting causes book values to impound information based on their impact on annual

[4] Their use of a constant discount rate implies either risk neutrality or constant consumption.
[5] Also see Francis, LaFond, Olsson and Schipper (2004) for a cross-country analysis using a ratio measure.
[6] Pope and Walker (1999) do not assess the statistical significance of their ratio measures.

earnings more so than on multi-period earnings (Rangan and Sloan, 1998). Furthermore, the test we recommend only requires the estimation of a traditional Basu (1997) regression and thus does not exacerbate the aggregation effect discussed in Givoly, Hayn and Natarajan (2007).

The paper proceeds as follows. Section II develops our econometric model and predictions. Section III describes our data. Section IV presents our primary tests and Section V discusses the empirical properties of the ratio measure. Section VI concludes.

II. ECONOMETRIC MODEL

Model Setup

In developing the econometric model, we make several assumptions, some of which provide consistency with prior literature while others facilitate the derivation of the asymptotic values of statistical tests. We first assume that the firm is priced by risk-neutral investors with a zero discount rate. This implies that the value v_t of the firm at time *t* equals the expected value of future dividends, conditional on information available at time *t* so that $v_t = \mathrm{E}_t[\sum_{j=1}^{\infty} d_{t+j}]$.

Prior literature that uses the Basu (1997) regression implicitly makes this assumption by using zero returns to partition good from bad news. If investors have a non-zero discount rate due to risk aversion or time preferences, the one-year no-news return, and therefore the demarcation between good and bad news, is the one-year risk-free rate.[7]

Given returns $r_t = (v_{t+1} + d_{t+1} - v_t) / v_t$, expected returns are identically zero and reflect

[7] See, for example, Proposition 7.7 in Christensen and Feltham (2003). Receiving no news during period *t* implies that investors' time *t* – 1 information set revealed the time *t* consumption and dividend. For example, if investors are risk-neutral with discount factor $\beta < 1$, then $v_t = \mathrm{E}_t[\sum_{j=1}^{\infty} \beta^j d_{t+j}]$ and the gross period t return is $1/\beta$ if investors receive no news during period t. Thus, returns less than $(1/\beta) - 1$ imply that investors received bad news.

only the arrival of information about the firm's cash flows. Returns thus equal unexpected returns, which are orthogonal to investors' $t-1$ information set. The time t information set may include earnings so that we can decompose returns into earnings news and non-earnings news, which is the approach implicitly taken by Dietrich, Muller and Riedl (2007) who state their analysis in terms of the price impact of earnings and non-earnings news. We take a different approach by explicitly modeling the dynamics of accounting book values in order to derive the association between returns and reported earnings, rather than the price impact of earnings.

The firm value equals the sum of the value n_t of separable net assets n_t and the value f_t of expected future abnormal profits (hereafter 'future rents'). We assume that accounting book value only attempts to reflect the value of separable assets n_t so that earnings only represent the recognition of changes in n_t (Watts, 2003a and Roychowdhury and Watts, 2007). Our objective in this study is to analyze tests of the degree to which the accounting for n_t exhibits asymmetric recognition of changes in value. Because the value of the firm $v_t = n_t + f_t$, the increase in the cum-dividend value of the firm $r_t v_{t-1} = \Delta n_t + \Delta f_t$. We denote the portion of the change in the cum-dividend value v_t attributable to revisions of beliefs about separable net assets as λ_t. The remaining portion $1-\lambda_t$ pertains to future rents.

Eventually, the firm either realizes future rents f_t and converts them to separable assets, or the rents are not realized and therefore become worthless. A negative shock $(1-\lambda_t)r_t v_{t-1}$ to expected future rents will reflect revisions of expectations for rents that are never realized. We denote the value of realized rents in period *t* by $\delta_t f_{t-1}$.[8] The market's valuation of net assets and

[8] For example, if a firm's market power at time *t* provides economic rents of $10 next year that decline by 10%

future rents therefore has the following dynamics, which satisfy $\Delta n_t + \Delta f_t = r_t v_{t-1}$:

$$n_t = n_{t-1} + \lambda_t r_t v_{t-1} + \delta_t f_{t-1} \qquad\qquad f_t = (1-\delta_t) f_{t-1} + (1-\lambda_t) r_t v_{t-1} \tag{1}$$

We model accounting book value b_t as a reflection of separable assets as determined by accounting rules. We assume that book value immediately recognizes the realization $\delta_t f_{t-1}$ of expected rents converted into separable assets during period *t*; however, book value only recognizes a fraction $\theta_t \in [0,1]$ of the time *t* shock $\lambda_t r_t v_{t-1}$ to investors' valuation of separable assets. A greater θ_t for negative shocks to separable assets relative to positive shocks is a manifestation of accounting conservatism.

Because book value fails to immediately recognize some changes in the value of separable assets, we denote the cumulative unrecognized gains and losses on separable net assets by 'goodwill' $g_t = n_t - b_t$. The accounting system eventually recognizes the gains and losses or expectations are revised, eliminating the gains and losses. These gains and losses are recognized via, for example, their impact on recognized profit or conversions of assets into cash. We represent the recognition of prior gains and losses by the parameter $\omega_t \in [0,1]$ so that the period *t* earnings include $\omega_t g_{t-1}$ from the recognition of gains and losses on separable assets.

The dynamics of the cum-dividend book value b_t and goodwill g_t are therefore:

$$b_t + d_t = b_{t-1} + \theta_t \lambda_t r_t v_{t-1} + \delta_t f_{t-1} + \omega_t g_{t-1} \tag{2}$$

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per year, the value of the future rents is $f_t = \$100 = \$10 \div 10\%$ using the formula for a growing perpetuity and our assumption of a zero discount rate. Each year $t+k$, the firm realizes $\$10 \times 0.9^{k-1}$ of the rents and expects an offsetting $\$10 \times 0.9^{k-1}$ depreciation in the value of future rents. In this example, $\delta_{t+1} f_t = \$10$ and $(1-\delta_{t+1}) f_t = \$90$. The period *t* + 1 shock to future rents is $(1-\lambda_{t+1}) r_{t+1} v_t = f_{t+1} - (1-\delta_{t+1}) f_t = \$9/0.1 - \$90 = \$0$, which is zero because no news was conveyed during period *t* + 1 and investors have a zero discount rate.

$$g_t = n_t - b_t = (1-\omega_t)g_{t-1} + (1-\theta_t)\lambda_t r_t v_{t-1} \tag{3}$$

Under the assumption of clean surplus accounting, earnings x_t therefore equal:

$$x_t = b_t + d_t - b_{t-1} = \theta_t \lambda_t r_t v_{t-1} + \delta_t f_{t-1} + \omega_t g_{t-1} \tag{4}$$

Earnings consist of the recognition of value previously impounded into the stock price (the terms involving f_{t-1} and g_{t-1}) and the recognition of value impounded into stock price in the current period (the term involving r_t). An empiricist can directly observe only earnings x_t, returns r_t and the prior value v_{t-1} so that (4) includes noise, from the empiricist's perspective. Relation (4) does not imply that returns 'cause' earnings; rather, returns are coincident with earnings.

We assume that the parameter $\lambda_t \in [0,1]$ so that the sign of the shock $\lambda_t r_t v_{t-1}$ to the value of separable assets has the same as the sign of the shock $(1-\lambda_t)r_t v_{t-1}$ to future rents.[9] We make this assumption in order to maintain consistency with prior literature that uses zero returns as the indication of bad news. While we expect that that this may be violated in some cases, we conjecture that the sign of shocks to net assets and firm value typically match. For example, we do not expect to observe decreases in the value of equipment in the same period that the market determines an increase in the expected value of the future rents made from selling the product produced by the equipment.[10]

We also assume that λ_t is independent of (θ_t, r_t) so that the firm's returns and accounting for shocks to separable assets are jointly independent of the relative value of the

[9] Under the maintained assumption that book values only attempt to measure separable assets, regression tests that follow Basu (1997) by using negative returns to proxy for bad news implicitly assume that returns are negative if and only if there is bad news about separable assets.

[10] While, for any given piece of equipment, we expect the equipment's value to move in tandem with the value of future rents, this analogy may not hold when one views the firm as a portfolio of projects. For example, if a firm has two projects and an overall positive return, the return to net assets can be negative only if the project having a negative return has a much higher impact on net assets λ_t than the project with the positive return. To the extent that firms typically hold portfolios of similar projects, we expect our assumption to hold for most firms.

firm's future rents. The independence of λ_t and θ_t is consistent with firms facing similar accounting standards regardless of the extent to which their market value reflects expected rents that are not reflected in the values of separable assets.[11] Given our assumption that investors are risk-neutral and expected returns are therefore zero, the independence of λ_t and r_t does not appear unreasonable. On the other hand, in observed markets, where risk-averse investors presumably price securities, there is a well documented relation between returns and market-to-book ratios (Fama and French, 1992). This suggests that λ_t has a relation with expected returns, but perhaps not unexpected returns, in observed markets. It is possible, however, to somewhat control for this relation by adding explanatory variables to the regression that proxy for factors that determine expected returns.

Another assumption is that returns r_t are independent of the accounting recognition $\delta_t f_{t-1} / v_{t-1}$ and $\omega_t g_{t-1} / v_{t-1}$ of value previously impounded into the market price. We expect this to approximately hold because neither $\delta_t f_{t-1}$ nor $\omega_t g_{t-1}$ reflect changes in the market value of the firm. They both represent the accounting recognition of value that has been previously impounded in market prices. We therefore expect a correlation between returns and $(\delta_t f_{t-1} + \omega_t g_{t-1}) / v_{t-1}$ only to the extent that share prices fail to reflect the value of the firm.

Finally, we assume that the firm's accounting parameter θ_t for recognizing shocks to separable assets is independent of returns conditional on the sign of returns. The above assumptions imply that earnings recognize a greater amount of bad news about separable assets than good news if and only if $\mathrm{E}[\theta_t \mid r_t < 0] > \mathrm{E}[\theta_t \mid r_t > 0]$.

[11] A potential concern is the capitalization of software under SFAS 86 (FASB, 1985) and of purchased R&D under SFAS 141(R) (FASB, 2007); however, these activities are, in principle, separable assets rather than future rents.

Regressions of Earnings on Returns Using Price-scaled Earnings

The typical test of asymmetric timeliness estimates a regression of the following form, where 1_A denotes an indicator variable that equals one when A is true and equals zero otherwise:

$$\frac{x_t}{v_{t-1}} = \beta_0 + \beta_1 1_{r_t<0} + \beta_2 r_t + \beta_3 1_{r_t<0} r_t + e_t \tag{5}$$

The coefficient of interest in the tests is β_3. Ordinary least squares gives:

$$\operatorname{plim}\hat{\beta}_3 = \frac{\operatorname{cov}(r_t, x_t / v_{t-1} \mid r_t < 0)}{\operatorname{var}(r_t \mid r_t < 0)} - \frac{\operatorname{cov}(r_t, x_t / v_{t-1} \mid r_t \geq 0)}{\operatorname{var}(r_t \mid r_t \geq 0)} \tag{6}$$

The independence of $(\delta_t f_{t-1} + \omega_t g_{t-1}) / v_{t-1}$ and r_t imply that $\operatorname{cov}(r_t, x_t / v_{t-1} \mid r_t < 0)$ equals $\operatorname{cov}(r_t, \theta_t \lambda_t r_t \mid r_t < 0)$. The assumptions on the relation between r_t, θ_t and λ_t then imply:

$$\operatorname{cov}(r_t, x_t / v_{t-1} \mid r_t < 0) = \mathrm{E}[\theta_t \mid r_t < 0]\,\mathrm{E}[\lambda_t]\operatorname{var}(r_t \mid r_t < 0) \tag{7}$$

A similar analysis applies to $\operatorname{cov}(r_t, x_t / v_{t-1} \mid r_t \geq 0)$ giving:

$$\operatorname{plim}\hat{\beta}_3 = (\mathrm{E}[\theta_t \mid r_t < 0] - \mathrm{E}[\theta_t \mid r_t \geq 0])\,\mathrm{E}[\lambda_t] \tag{8}$$

The sign of $\hat{\beta}_3$ correctly identifies asymmetric timeliness in large samples; however, the magnitude is dampened by the expected portion $\mathrm{E}[\lambda_t]$ of returns attributable to separable assets, on account of our assumption that $\lambda_t \in [0,1]$. This can be viewed as a form of multiplicative measurement error (Hwang, 1986) and differs from the classical errors-in-variables problem but nonetheless has a similar attenuation effect in our setting. This poses no problem if the researcher only wishes to test whether or not there is asymmetric timeliness; however, this creates a bias in cross-sectional comparisons of asymmetric timeliness.

Researchers can remove the bias by examining a ratio of regression coefficients that can be tested using the delta method:

$$\operatorname{plim} \frac{\hat{\beta}_3}{\hat{\beta}_2 + \hat{\beta}_3} = \frac{\mathrm{E}[\theta_t \mid r_t < 0] - \mathrm{E}[\theta_t \mid r_t \geq 0]}{\mathrm{E}[\theta_t \mid r_t < 0]} \tag{9}$$

The use of the delta method test statistic requires that the ratio be a continuous function of the parameters at the true parameter values, which in this case requires that the true value of $\beta_2 + \beta_3 \neq 0$. This assumption is plausible on both the theoretical ground that earnings recognize some current period bad news and the empirical ground that estimated asymmetric timeliness coefficients are typically positive and significant. Because of the need to ensure continuity of the ratio function, researchers should define the ratio so that the denominator includes the more precisely measured coefficient rather than the coefficient on negative returns, per se.[12]

In order to address concerns with the applicability of the delta method, we present confidence intervals for the ratio based on Fieller's theorem (Fieller, 1954) that translates the ratio into a linear hypothesis as described in the Appendix. This method is more robust than the delta method, particularly in small samples (Staiger, Stock and Watson, 1997). We illustrate both delta method and Fieller confidence intervals in order to illustrate the extent to which the delta method statistics appear reasonable. When making cross-sectional comparisons of the ratio, it is not possible to translate the comparison of ratios into a linear hypothesis and we therefore rely on the delta method. Similarity between the delta method and Fieller confidence intervals supports the use of delta method tests when making cross-sectional comparisons.

Cross-sectional Comparisons of Asymmetric Timeliness Coefficients

Several studies test cross-sectional differences in asymmetric timeliness. For example,

[12] Empirically, the coefficient on negative returns is measured more precisely. See, for example, Ryan and Zarowin (2003, Table 2).

Givoly, Hayn and Natarajan (2007) find a negative relation between asymmetric timeliness and the market-to-book ratio, another measure of accounting conservatism. The finding is puzzling because a high market-to-book ratio signifies firm value that is not recorded in book value and therefore suggests conservative accounting. Roychowdhury and Watts (2007) hypothesize that this negative relation is due to a high proportion of returns being attributable to revisions of expected future rents. In the context of our model, this is indeed the case because (8) implies that estimates of asymmetric timeliness will be lower for firms with a low proportion of returns attributable to separable assets (low $\mathrm{E}[\lambda_t]$), which are likely to have a high market-to-book ratio.

In order to analyze the impact of such bias on cross-sectional comparisons, we evaluate regressions of the following form based on a single sample partition Z_{t-1} that depends on data available prior to the return accumulation period:

$$\frac{x_t}{v_{t-1}} = \beta_0 + \beta_1 1_{r_t<0} + \beta_2 r_t + \beta_3 1_{r_t<0} r_t + 1_{Z_{t-1}}\left(\beta_4 + \beta_5 1_{r_t<0} + \beta_6 r_t + \beta_7 1_{r_t<0} r_t\right) + e_t \tag{10}$$

The coefficient β_7 measures the difference in the asymmetric timeliness of earnings for members of the partition vis-à-vis non-members. Similar computations as before give:

$$\operatorname{plim}\hat{\beta}_3 = \left(\mathrm{E}[\theta_t \mid r_t < 0, \neg Z_{t-1}] - \mathrm{E}[\theta_t \mid r_t \geq 0, \neg Z_{t-1}]\right)\mathrm{E}[\lambda_t \mid \neg Z_{t-1}] \tag{11}$$

$$\begin{aligned}\operatorname{plim}\hat{\beta}_7 = {} & \left(\mathrm{E}[\theta_t \mid r_t < 0, Z_{t-1}] - \mathrm{E}[\theta_t \mid r_t \geq 0, Z_{t-1}]\right)\mathrm{E}[\lambda_t \mid Z_{t-1}] \\ & -\left(\mathrm{E}[\theta_t \mid r_t < 0, \neg Z_{t-1}] - \mathrm{E}[\theta_t \mid r_t \geq 0, \neg Z_{t-1}]\right)\mathrm{E}[\lambda_t \mid \neg Z_{t-1}]\end{aligned} \tag{12}$$

If the portion λ_t of returns attributable to separable assets is associated with the partition variable Z_{t-1}, then the estimated difference in asymmetric timeliness is biased because $\mathrm{E}[\lambda_t \mid Z_{t-1}] \neq \mathrm{E}[\lambda_t \mid \neg Z_{t-1}]$. For example, if the partition variable is market-to-book, firms with high market-to-book likely have relatively low values of λ_t.

The ratio statistic (9) may be employed as a means to address cross-sectional differences in λ_t. In the context of regression (10), researchers can test the relative degree of asymmetric timeliness using ratios of regression coefficients:[13]

$$\begin{aligned}&\operatorname{plim}\frac{\hat\beta_3+\hat\beta_7}{\hat\beta_2+\hat\beta_3+\hat\beta_6+\hat\beta_7}-\frac{\hat\beta_3}{\hat\beta_2+\hat\beta_3}\\ &=\frac{\mathrm{E}[\theta_t \mid r_t<0, Z_{t-1}]-\mathrm{E}[\theta_t \mid r_t\geq 0, Z_{t-1}]}{\mathrm{E}[\theta_t \mid r_t<0, Z_{t-1}]}-\frac{\mathrm{E}[\theta_t \mid r_t<0, \neg Z_{t-1}]-\mathrm{E}[\theta_t \mid r_t\geq 0, \neg Z_{t-1}]}{\mathrm{E}[\theta_t \mid r_t<0, \neg Z_{t-1}]}\end{aligned} \tag{13}$$

Regressions that Include Pre-Returns Information

Basu's (1997) regression tests of asymmetric timeliness implicitly assume a relation between unexpected earnings and unexpected price changes (Dietrich, Muller and Riedl, 2007). Our econometric model explicitly makes this assumption, which we can use to refine the regression test by including a set of variables Y_{t-1} from investors' pre-returns information set:

$$\frac{x_t}{v_{t-1}}=\beta_0+\beta_1 1_{r_t<0}+\beta_2 r_t+\beta_3 1_{r_t<0} r_t+\beta_y' Y_{t-1}+e_t \tag{14}$$

The regression (14) is equivalent to:

$$x_t^e=\gamma_0+\gamma_1 1_{r_t<0}^e+\gamma_2 r_t^e+\gamma_3 (1_{r_t<0} r_t)^e+u_t \tag{15}$$

where the *e* superscript denotes the residuals from regressing the variables in (14) on Y_{t-1}. Under the assumption that returns are independent of $t-1$ information, $1_{r_t<0}^e=1_{r_t<0}$ and $r_t^e=r_t$.

Furthermore, this implies that:

[13] If the setting is such that researchers can reasonably impose the restriction that groups have comparable recognition of good news about separable assets, $\mathrm{E}[\theta_t \mid r_t\geq 0, Z_{t-1}]=\mathrm{E}[\theta_t \mid r_t\geq 0, \neg Z_{t-1}]$, the ratio test statistic (13) is positive and statistically significant if and only if the partition group recognizes a greater portion of negative shocks to separable assets, $\mathrm{E}[\theta_t \mid r_t<0, Z_{t-1}]-\mathrm{E}[\theta_t \mid r_t<0, \neg Z_{t-1}]$. The restriction on good news recognition does not imply that the two sets of firms have identical regression coefficients for good news; rather, it implies that any variation in the coefficients on good news stems from differing proportions of returns attributable to rents.

$$\operatorname{plim}\hat{\gamma}_2 = \operatorname{plim}\hat{\beta}_2 = \mathrm{E}[\theta_t \mid r_t \geq 0]\,\mathrm{E}[\lambda_t]$$
$$\operatorname{plim}\hat{\gamma}_3 = \operatorname{plim}\hat{\beta}_3 = (\mathrm{E}[\theta_t \mid r_t < 0] - \mathrm{E}[\theta_t \mid r_t \geq 0])\,\mathrm{E}[\lambda_t] \tag{16}$$

The coefficients are the same as predicted in the regression without Y_{t-1}, but the residual x_t^e represents the unexpected portion of earnings, which are associated with returns. In a typical OLS model, this reduction in 'noise' in the dependent variable would improve the precision of regression estimates. If λ_t is constant, this intuition applies to our setting and the estimates will be more precise; otherwise, it is possible that the estimates will be less precise.[14] This approach to including forecasts or lagged earnings as a regressor is preferable to a regression that uses forecast errors or earnings changes as the dependent variable because it allows for the fact that expected earnings may differ from the forecast.

The specification (14) is also useful because prior studies suggest the assumption that returns are independent of $t-1$ information is invalid. The inclusion of $t-1$ variables can serve to control for both expected earnings and expected returns. Several attributes such as the price-earnings ratio (Basu, 1977), size and book-to-market ratio also affect expected returns (Fama and French, 1992). This is not particularly problematic to the extent that the $t-1$ information reflects risk factors because the residual from regressing returns on the information can be viewed as a risk-adjusted return. However, it does suggest that including a forecast in Y_{t-1} may affect the coefficients on returns. Furthermore, we later show that the coefficient on positive returns tends to be positive and significant when including controls such as lagged earnings. Prior literature shows that the coefficient β_2 on positive returns in specification (5) is often

[14] See section 4.4 of Wooldridge (2002) for a description of the former case in which λ_t is constant.

measured imprecisely or even negative.[15]

Empirical Predictions

The econometric model implies several characteristics of asymmetric timeliness tests. We develop and later test these predictions in order to assess the extent to which our derived results are robust to violations of the assumptions in our econometric model. First, (8) suggests a negative relation between market-to-book ratios and asymmetric timeliness tests, as in Givoly, Hayn and Natarajan (2007) and Roychowdhury and Watts (2007):

P1 The estimated asymmetric timeliness coefficient β_3 is lower, but positive, for firms with a high amount of value based on rents.

Beatty (2007) suggests using observed market-to-book ratios to control for the impact of rents on the earnings-returns relation. In our setting, including market-to-book as a regressor will control for expected earnings and returns, but will have little effect on the bias in estimating asymmetric timeliness.[16] An alternative approach uses book-to-market to proxy for the portion λ_t of returns attributable to separable assets. This gives the following regression specification:

$$\frac{x_t}{v_{t-1}} = \delta_0 + \delta_1 1_{r_t<0} + \delta_2 \frac{b_{t-1}}{v_{t-1}} r_t + \delta_3 1_{r_t<0} \frac{b_{t-1}}{v_{t-1}} r_t + e_t \tag{17}$$

Under the assumptions of our econometric model, plus the 'best-case' assumption that the book-to-market ratio is a perfect proxy, $\lambda_t = b_{t-1} / v_{t-1}$, we have the following:[17]

[15] See, for example, Givoly, Hayn and Natarajan (2007, Table 5).
[16] See the section 'Regressions that Include Pre-Returns Information'.
[17] The derivation is available from the authors upon request.

$$\text{plim}\,\hat{\delta}_3 = \text{E}[\theta_t \mid r_t < 0] - \text{E}[\theta_t \mid r_t \geq 0]$$
$$+ \underbrace{\left(\frac{\text{E}[r_t \mid r_t < 0]}{\text{var}(r_t b_{t-1} / v_{t-1} \mid r_t < 0)} - \frac{\text{E}[r_t \mid r_t \geq 0]}{\text{var}(r_t b_{t-1} / v_{t-1} \mid r_t \geq 0)} \right)}_{<0} \times \text{cov}\left(\frac{b_{t-1}}{v_{t-1}}, \frac{\delta_t f_{t-1} + \omega_t g_{t-1}}{v_{t-1}} \right) \quad (18)$$

The covariance term in (18) is likely negative because book value $b_{t-1} = v_{t-1} - f_{t-1} - g_{t-1}$ by definition, so that regression (17) yields a positively biased measure of asymmetric timeliness.[18] Because we predict that the standard specification (5) yields a dampened measure, we predict the following:

P2 A regression of the form (17) that interacts returns with the book-to-market ratio yields larger asymmetric timeliness coefficients than in the standard specification (5).

Taking as given that the covariance term in (18) is negative, a ratio test $\hat{\delta}_3 / (\hat{\delta}_2 + \hat{\delta}_3)$ is greater than in the standard specification if and only if:

$$1 - \frac{\text{E}[r_t \mid r_t \geq 0]}{\text{E}[r_t \mid r_t < 0]} \frac{\text{var}(r_t b_{t-1} / v_{t-1} \mid r_t < 0)}{\text{var}(r_t b_{t-1} / v_{t-1} \mid r_t \geq 0)} > \frac{\text{E}[\theta_t \mid r_t < 0] - \text{E}[\theta_t \mid r_t \geq 0]}{\text{E}[\theta_t \mid r_t < 0]} \quad (19)$$

The left-hand-side of (19) is greater than one since $\text{E}[r_t \mid r_t < 0] < 0$. Assuming that some good news is recognized, $\text{E}[\theta_t \mid r_t \geq 0] > 0$, the right-hand-side of (19) is less than one, giving the following prediction:

P3 A ratio test $\hat{\delta}_3 / (\hat{\delta}_2 + \hat{\delta}_3)$ based on (17) yields a larger, biased ratio than the test $\hat{\beta}_3 / (\hat{\beta}_2 + \hat{\beta}_3)$ based on (5).

We can also derive the expected coefficient in a regression of earnings per share on dollar returns to evaluate Dietrich, Muller, and Riedl's (2007) evidence of aggressive accounting when

[18] The term preceding it is clearly negative because variances are positive, $\text{E}[r_t \mid r_t < 0] < 0$ and $\text{E}[r_t \mid r_t \geq 0] \geq 0$.

using per share figures. An unscaled regression estimates the following:[19]

$$x_t = \alpha_0 + \alpha_1 1_{r_t<0} + \alpha_2 \Delta v_t + \alpha_3 1_{r_t<0} \Delta v_t + e_t \tag{20}$$

Similar computations as before give:

$$\begin{aligned} \operatorname{plim} \hat{\alpha}_3 &= (\mathrm{E}[\theta_t \mid r_t < 0] - \mathrm{E}[\theta_t \mid r_t \geq 0])\, \mathrm{E}[\lambda_t] \\ &+ \left(\frac{\mathrm{E}[r_t \mid r_t < 0]}{\operatorname{var}(\Delta v_t \mid r_t < 0)} - \frac{\mathrm{E}[r_t \mid r_t \geq 0]}{\operatorname{var}(\Delta v_t \mid r_t \geq 0)} \right) \operatorname{cov}(v_{t-1}, \delta_t f_{t-1} + \omega_t g_{t-1}) \end{aligned} \tag{21}$$

Expression (21) suggests a negative bias in $\hat{\alpha}_3$ for purposes of conservatism tests. The term in parentheses is clearly negative. Because of scale effects, we also expect a positive covariance between firm size v_{t-1} and $\delta_t f_{t-1} + \omega_t g_{t-1}$, which roughly corresponds to the predictable portion of earnings. If the conjecture on the covariance term is true, then $\hat{\alpha}_3$ is negatively biased, giving the following prediction:

P4 The estimated asymmetric timeliness coefficient α_3 in an unscaled regression is negatively biased and possibly negative, even if $\mathrm{E}[\theta_t \mid r_t < 0] > \mathrm{E}[\theta_t \mid r_t \geq 0]$.

The inclusion of goodwill violates our assumption that book values reflect separable assets (Roychowdhury and Watts, 2007). We can examine the impact of goodwill on the returns-earnings relation under SFAS 141(R) (FASB, 2007) where goodwill is subject to impairment. In this case, earnings include an additional component $\psi_t (1-\lambda_t) r_t v_{t-1}$ that reflects the shock to the prior year value f_{t-1} of rents. Under impairment accounting, $\mathrm{E}[\psi_t \mid r_t \geq 0] = 0$ since goodwill impairments reflect only bad news. In this case, the asymmetric timeliness coefficient includes an additional term from the effect of goodwill impairments on earnings:

[19] We continue to use negative returns to partition good and bad news. In our setting, dollar returns are negative if and only if returns are negative. In settings that allow for payouts and equity issuance, the change in equity depends on both news and transactions with shareholders so that returns remains the appropriate means to partition good and bad news.

$$\text{plim}\,\hat{\beta}_3 = \left(\text{E}[\theta_t \mid r_t < 0] - \text{E}[\theta_t \mid r_t \geq 0]\right)\text{E}[\lambda_t] + \text{E}[\psi_t \mid r_t < 0]\,\text{E}[1-\lambda_t] \tag{22}$$

Researchers can eliminate the effect of goodwill impairments on earnings because they are reported separately. This allows us to test the following prediction:[20]

P5 The asymmetric timeliness coefficient is lower in a regression where earnings exclude the effect of goodwill impairments.

Given (22) and $\text{plim}\,\hat{\beta}_2 + \hat{\beta}_3 = \text{E}[\theta_t \mid r_t < 0]\,\text{E}[\lambda_t] + \text{E}[\psi_t \mid r_t < 0]\,\text{E}[1-\lambda_t]$, a ratio test when earnings include goodwill impairments is less than one if and only if the ratio $\left(\text{E}[\theta_t \mid r_t < 0] - \text{E}[\theta_t \mid r_t \geq 0]\right) / \text{E}[\theta_t \mid r_t < 0]$ is less than one. This leads to our final prediction:

P6 If the ratio test $\hat{\beta}_3 / (\hat{\beta}_2 + \hat{\beta}_3)$ is less than one with unadjusted earnings, it is lower, yet, when earnings are adjusted to remove goodwill impairments.

Impact of Relaxing Assumptions of the Econometric Model

We conclude this section by discussing our assumption that the portion λ_t of returns due to net assets is independent of returns r_t. This assumption ensures that regression estimates are unaffected by the sample truncation bias predicted by Dietrich, Muller and Riedl (2007). If positive returns have a greater tendency to affect the expected value of abnormal profits, then this assumption is violated. For example, Givoly, Hayn and Natarajan (2007) claim that events such as the announcement of new contracts and mergers and acquisitions tend to have a large positive impact on returns but not on earnings. If λ_t is correlated with returns, the coefficient β_3 on asymmetric timeliness and the ratio for regression (5) are:

[20] SFAS 141 applies to business combinations initiated after June 30, 2001 so that this prediction only applies to fiscal years beginning later than the June 30, 2001.

$$\operatorname{plim}\hat{\beta}_3 = \mathrm{E}[\theta_t \mid r_t < 0]\frac{\operatorname{cov}(\lambda_t r_t, r_t \mid r_t < 0)}{\operatorname{var}(r_t \mid r_t < 0)} - \mathrm{E}[\theta_t \mid r_t \geq 0]\frac{\operatorname{cov}(\lambda_t r_t, r_t \mid r_t \geq 0)}{\operatorname{var}(r_t \mid r_t \geq 0)} \tag{23}$$

$$\operatorname{plim}\frac{\hat{\beta}_3}{\hat{\beta}_2 + \hat{\beta}_3} = 1 - \frac{\mathrm{E}[\theta_t \mid r_t \geq 0]}{\mathrm{E}[\theta_t \mid r_t < 0]}\frac{\operatorname{cov}(\lambda_t r_t, r_t \mid r_t \geq 0)}{\operatorname{var}(r_t \mid r_t \geq 0)}\frac{\operatorname{var}(r_t \mid r_t < 0)}{\operatorname{cov}(\lambda_t r_t, r_t \mid r_t < 0)} \tag{24}$$

The ratio $\operatorname{cov}(\lambda_t r_t, r_t) / \operatorname{var}(r_t)$ is the fraction of returns attributable to net assets. If positive returns relate more to rents than separable assets, then the term in (24) that is multiplied by $\mathrm{E}[\theta_t \mid r_t \geq 0] / \mathrm{E}[\theta_t \mid r_t < 0]$ is greater than one so that both the coefficient $\hat{\beta}_3$ and the ratio test $\hat{\beta}_3 / (\hat{\beta}_2 + \hat{\beta}_3)$ are biased in favor of identifying accounting conservatism. This need not be the case. For example, it is impossible to have the lowest possible return of -1 unless both the values of net assets and expected abnormal profits decline to zero. On the other hand, it is possible that expected abnormal profits cause extreme positive returns absent a large change in the value of net assets in place.

The ratio measure provides valid cross-sectional comparisons of asymmetric timeliness so long as the two cross-sectional groups do not have a large difference in the degree to which positive returns tend to be attributed to future abnormal profits. If λ_t is correlated with returns, we can rewrite the cross-sectional ratio comparison (13) as:

$$\begin{aligned}
&\operatorname{plim}\frac{\hat{\beta}_3 + \hat{\beta}_7}{\hat{\beta}_2 + \hat{\beta}_3 + \hat{\beta}_6 + \hat{\beta}_7} - \frac{\hat{\beta}_3}{\hat{\beta}_2 + \hat{\beta}_3} \\
&\quad = \left(\frac{\mathrm{E}[\theta_t \mid r_t \geq 0, \neg Z_{t-1}]}{\mathrm{E}[\theta_t \mid r_t < 0, \neg Z_{t-1}]} - \frac{\mathrm{E}[\theta_t \mid r_t \geq 0, Z_{t-1}]}{\mathrm{E}[\theta_t \mid r_t < 0, Z_{t-1}]}\right)\frac{\operatorname{cov}(\lambda_t r_t, r_t | r_t \geq 0, Z_t) / \operatorname{var}(r_t | r_t < 0, Z_t)}{\operatorname{cov}(\lambda_t r_t, r_t | r_t < 0, Z_t) / \operatorname{var}(r_t | r_t < 0, Z_t)} \\
&\quad + \frac{\mathrm{E}[\theta_t \mid r_t \geq 0, \neg Z_{t-1}]}{\mathrm{E}[\theta_t \mid r_t < 0, \neg Z_{t-1}]}\left(\frac{\operatorname{cov}(\lambda_t r_t, r_t | r_t \geq 0, \neg Z_t) / \operatorname{var}(r_t | r_t < 0, \neg Z_t)}{\operatorname{cov}(\lambda_t r_t, r_t | r_t < 0, \neg Z_t) / \operatorname{var}(r_t | r_t < 0, \neg Z_t)} - \frac{\operatorname{cov}(\lambda_t r_t, r_t | r_t \geq 0, Z_t) / \operatorname{var}(r_t | r_t < 0, Z_t)}{\operatorname{cov}(\lambda_t r_t, r_t | r_t < 0, Z_t) / \operatorname{var}(r_t | r_t < 0, Z_t)}\right)
\end{aligned} \tag{25}$$

If the two cross-sectional groups have similar proportions of positive and negative returns attributable to net assets, then the last term is approximately zero and the ratio test provides a valid measure of whether the group represented by Z_t exhibits greater asymmetric timeliness than the group $\neg Z_t$.

III. DATA AND SAMPLE SELECTION

We conduct our empirical analysis on the sample of firm-years reported in Table 1 Panel A. We begin with the 254,127 firm-years in the *Compustat* Fundamental Annual table and restrict the sample to firm-years with matches to *CRSP* identifiers in the *CRSP*/*Compustat* merged database (205,448 firm-years). Missing variables from *Compustat* or *CRSP* reduces the sample to 179,885 firm-year observations. The final sample includes 18,681 unique firms covering the years 1963-2007. Some of our cross-sectional regressions use a smaller sample of firms with market-to-book ratios either below the 35th percentile or above the 65th percentile of our full forecast sample. This sub-sample includes 125,918 firm-years and 17,854 firms.

Panel B of Table 1 reports descriptive statistics for the sample firms. All variables are winsorized at the 1st and 99th percentiles within *Compustat* fiscal years. Assets (*Compustat* Annual Xpressfeed item *AT*), Market value (*PRCC_F* × *CSHO*), Book value (*CEQ*), and Market-to-book (Market value/Book value) are calculated at the beginning of the period. Earnings is earnings per share before extraordinary items (*EPSPX*) scaled by the beginning-of-fiscal-year price (*PRCC_F*), which we adjust for stock splits by multiplying by the ratio of the current-to-prior year price adjustment factors (*AJEX*). We obtain annual buy-and-hold returns from the *CRSP* monthly stock file beginning at the start of the fourth month of the fiscal year. The descriptive statistics show the mean assets, market and book values of equity to be $1.88 billion, $787.8 million, and $456.4 million for our full sample of firms, respectively. We also note the mean market-to-book for our sample is 2.52. In one of our tests, we include the forecast of current earnings as a control in the asymmetric timeliness regression. We obtain these forecasts from the *I/B/E/S* Summary Unadjusted file. We use the most recent mean analyst forecast of earnings per share just after the release of *t* - 1 earnings to measure the market's expectation of earnings prior to the earnings and returns period. We adjust the forecasts for stock splits by multiplying them by the ratio of the *CRSP* share adjustment factor (*CFACSHR*) at the earnings announcement date to the factor at the forecast date. As with the other variables, we scale the forecast by the split adjusted, beginning-of-fiscal-year stock price.

IV. EMPIRICAL RESULTS

Effect of Including Prior Information

Table 2 presents estimates of asymmetric timeliness using both the specification (5) as in Basu (1997) and a specification of the form (14) that includes proxies for expected earnings and expected returns. We condition on past earnings as a component of both expected earnings and expected returns, where the latter derives from the empirical regularity that firms with high price/earnings ratios tend to have lower future returns (Basu, 1977). We also condition on book-to-market to control for factors noted in Fama and French (1992).

The results in Table 2 show that estimation of specification (5) (column (1)) generates a positive and significant coefficient on negative returns consistent with asymmetric timeliness. The coefficient on positive returns is negative (-0.022) and significant (p-value 0.000), which is consistent with prior studies. Columns (2) and (3) add controls for expected earnings using both a forecast of current EPS and lagged EPS. The coefficient on positive returns is no longer negative when using lagged EPS or a forecast of EPS, but is only positive and significant when using lagged EPS. Although the coefficient on negative returns remains significantly positive, it is diminished (0.411 vs. 0.240 and 0.253) relative to the original specification, suggesting that some information in returns proxies for expected earnings. Finally, the inclusion of book-to-market in column (4) has little effect on the asymmetric timeliness coefficients relative to the original specification with similar coefficients on both negative returns (0.425 vs. 0.411) and positive returns (-0.020 vs. -0.022). Because the control for lagged scaled EPS provides a better specified regression for asymmetric timeliness and produces a positive coefficient on positive returns we include this additional control in our subsequent tests.

Effect of Rents on Returns

Table 3 presents the regression that examine cross-sectional differences in asymmetric timeliness by separating firms with high and low market-to-book ratios at the start of the returns period. The regression includes firm-years with beginning-of-period market-to-book ratios either below the 35th percentile (low market-to-book) or above the 65th percentile (high market-to-book) of our sample firm-years. The results demonstrate that firms with high market-to-book ratios appear to have less conservative earnings as indicated by a coefficient of -0.379 (significant at 1%) on the interaction between the high market-to-book indicator and the asymmetric timeliness coefficient. This result is consistent with prior literature and our prediction P1. On the other hand, the difference ratio test statistic as in equation (13) is positive 0.066 (significant at 1%) when we partition by high market-to-book, suggesting that firms with high market-to-book ratios have higher asymmetric timeliness than low market-to-book firms. These results are consistent with our conjecture that the lower asymmetric timeliness coefficient for high market-to-book firms is due to bias that can be corrected using a ratio of the coefficients. The delta method confidence intervals are nearly identical to those based on the Fieller method, which supports the validity of the delta method test statistic for the difference in ratios.[21]

In Table 4, we explicitly test the effects of including book-to-market in the asymmetric timeliness regression as a correction for the bias caused by returns including information about rents. We do this by testing our predictions P2 and P3 that a regression including the interaction of returns and book-to-market as in (17) produces positively biased asymmetric timeliness coefficient and ratio estimates. In Panel A, the table reports coefficient estimates of the original Basu (1997) specification and the estimates based on (17). Panel B reports the tests of differences in asymmetric timeliness coefficient and ratio estimates. Specifically, we test the restriction that the estimated coefficient or ratio from the interaction regression is equal to the point estimate from the original specification. Consistent with our predictions, the estimates using the interaction of returns and book-to-market are higher than the original specification for the coefficient (0.306 vs. 0.253, statistically different at 1%). Although the ratio is also higher using this specification (0.940 vs. 0.933), the difference is not statistically significant. These results suggest that including book-to-market in the regression does not control for the bias caused by the impact of future rents on returns.[22]

Table 5 provides additional support for the ratio measure by examining cross-sectional differences in asymmetric timeliness for REITs (Real Estate Investment Trusts) relative to other firms. We conjecture that REITs have returns more associated with changes in net assets than other firms, and test whether the ratio measure controls for this higher expected λ. We do not have any expectations that accounting rules related to REITs cause more asymmetric timeliness relative to other firms. We use the same specification as in Table 3, but use an indicator equal to

[21] The confidence intervals differ only at the fourth decimal place and therefore appear to be identical in the table.

[22] We also note that, as in Table 3, the confidence intervals for the ratio differ between the delta method and Fieller method only at the fourth decimal place.

one if the firm is a REIT and zero otherwise.[23] The interaction coefficient for REIT and negative returns is 0.180 (significant at 1%), suggesting that REITs tend to exhibit more asymmetric timeliness. However, our ratio test statistic that controls for cross-sectional differences in the amount of returns due to net assets does not indicate that REITs have greater asymmetric timeliness (test statistic of -0.019, p-value of 0.752). To the extent that REITs exhibit no more or less asymmetric loss recognition than other firms, this suggests that the positive interaction coefficient is due to the high proportion of returns that are attributable to net assets, a bias remedied by using comparisons based on the ratio measure.[24]

Effects of Scale

We next examine the relation between asymmetric timeliness regressions and alternative scaling variables. Easton and Sommers (2003) argue that market value of equity is the appropriate scalar for regressions that measure the relation between market-based and accounting-based variables. Prediction P4 states that estimates of the asymmetric timeliness coefficient are negatively biased in regressions that use per share amounts. We first present regressions as in Basu (1997) that scale earnings by beginning-of-period market value of equity and use returns as the independent variable. We then estimate asymmetric timeliness using alternative scaling variables. We exclude lagged EPS as a control in these tests because including it controls for the scaling effect we are demonstrating. Our primary interest is in the regression that scales by shares outstanding, which indicate *aggressive* accounting in Dietrich, Muller and Riedl (2007).

[23] A firm is classified as a REIT if the SIC is equal to 6798. A broader definition of real estate firms (including firms with SIC between 6500 and 6599 or equal to 6798 or 6799) yields very similar estimates.

[24] We again note that the delta method confidence intervals for the ratios provide a very close approximation of the Fieller method confidence intervals.

Column (1a) of Table 6 presents the standard regression of price-scaled earnings on returns consistent with Basu (1997). We find a positive coefficient on the interaction between a negative-returns indicator and returns, suggesting that earnings have a greater tendency to recognize economic losses. Column (1b) presents a regression that includes the beginning-of-period price for the purpose of comparison to the other regressions in Table 6. Here we note that the coefficients on returns differ between columns (1a) and (1b), which suggests a violation of the assumption that returns are independent of price levels. The difference in the coefficients is significant at the 1% level and perhaps reflects a size effect on returns.

Column (2a) of Table 6 regresses earnings per share on dollar returns as in of Dietrich, Muller and Riedl (2007, Table 3). Consistent with their results, we find a negative coefficient estimate (coefficient -0.079, significant at 1%) for asymmetric timeliness that suggests aggressive accounting. The coefficient for returns is 0.074 and significant at the 1% level. However, because we predict that scale effects bias the coefficient estimates in this specification, in column (2b) we add the beginning of period price to the regression to determine whether price is acting as a correlated omitted variable in the specification of column (2a). In column (2b), the estimate of asymmetric timeliness is 0.122 and significant at the 1% level, which suggests that scaling is indeed responsible for the negative coefficient in Dietrich, Muller and Riedl (2007).

To conclude this section, we present regressions that scale earnings and dollar returns by assets and the book value of equity in order to show that the inclusion of price does not induce a spurious relation between earnings and negative returns. Specifications (3) and (4) in Table 6 present these regressions both with and without beginning-of-period price as a regressor. The coefficient on price is significant in both specifications and has a statistical effect on the asymmetric timeliness coefficients, but only reduces the coefficients by 13.5 percent and 8.5

percent between specifications (3a) and (3b) and (4a) and (4b), respectively.

Effects of Goodwill

Finally, we examine our predictions (P5 and P6) that the inclusion of goodwill write-downs in earnings biases estimates of conservatism. As mentioned previously, the inclusion of these changes in the value of rents in earnings is problematic when we are only interested in describing conservatism relative to net assets. Table 7 presents the results of estimation using the original Basu (1997) specification and excluding the effects of goodwill write-downs on earnings. Since we are only interested in the period where SFAS 141 is in effect (which requires goodwill write-downs), we restrict the sample in Table 7 to firm-years 2001 to 2007 (FASB, 2007). The table shows that, consistent with P5, the asymmetric timeliness coefficient is lower when excluding the effects of goodwill write-downs (0.299 v. 0.268). The difference is statistically significant at 1 percent. However, the ratio appears to be unaffected by whether goodwill write-downs are included or excluded from EPS (0.960 v. 0.962, p-value 0.841). These results suggest that the inclusion of goodwill write-downs does bias the coefficient estimate of asymmetric timeliness upwards. Contrary to our prediction, goodwill write-downs do not appear to affect the estimation of conservatism when using the ratio statistic.

V. EMPRICAL PROPERTIES OF THE RATIO MEASURE

In this section we provide evidence that the ratio measure exhibits empirical properties that suggest it is a well-behaved and valid measure of accounting conservatism. As documented in our prior empirical tests, estimates of asymmetric timeliness coefficients may be biased relative to the ratio statistic in cross-sectional comparisons. Prior research has shown that asymmetric timeliness coefficients are negatively associated with other conservatism measures, which suggests that they fail to accurately represent the underlying construct of accounting conservatism (Givoly, Hayn and Natarajan, 2007). We therefore test how well the ratio statistic correlates with other conservatism measures when compared across industries. Furthermore, Givoly and Hayn (2000) and Givoly, Hayn, and Natarajan (2007) document that the asymmetric timeliness coefficient is not stable over time, which inconsistent with the notion that accounting rules are fairly stable. We therefore test the time-series stability of the ratio statistic.

The Ratio Measure and Asymmetric Timeliness Coefficient Across Industries

Our prior results on cross-sectional estimates of conservatism suggest that the asymmetric timeliness coefficient can be biased relative to the ratio measure. However, we expect this bias to differ across industries due to differences in the proportion of news relative to rents and net assets. We examine how the coefficient and ratio measure behave across industries to examine this conjecture. Figure 1 shows the asymmetric timeliness coefficient and ratio statistic estimated by 48 Fama and French (1997) industry groups. We partition the graph into four quadrants based on the mean values of the coefficient and ratio to help visually identify when the ratio and coefficient estimates differ. Industries in the upper left and lower right quadrants represent industries where the bias appears to affect cross-sectional comparisons of conservatism by altering the ranking of firms based on measures of asymmetric timeliness.

We note a few particular industries where the ranking appears to differ. Pharmaceutical Products, Tobacco, and Medical Equipment industries appear to have low coefficients relative to the ratio. For pharmaceuticals and medical equipment, the low coefficient measure is likely biased downward because a large portion of news for these firms is related to growth opportunities (including new drug and other research opportunities) and not net assets. The presence of tobacco companies in this quadrant could be due to the nature of their legal liabilities in which they recorded current obligations but also faced significant future payouts that reduce the value of future rents.[25] In the opposite quadrant, Shipbuilding and Railroads appear to have a higher coefficient relative to the ratio statistic. Unlike the previously mentioned industries, this industry has relatively little news related to growth opportunities (high $\mathrm{E}[\lambda_t]$), which dampens the coefficient less relative to other industries. These descriptive results are consistent with our prior predictions and suggest that researchers should consider using the ratio measure when examining conservatism across industries.

We also compare different measures of conservatism across the 48 industries. It is important to distinguish measures that reflect *ex post* conservatism in the form of recognizing bad news from measures of *ex ante* conservatism. Both the market-to-book ratio and the cumulative discretionary accrual (CDA) measure from Givoly and Hayn (2000) provide measures that reflect *ex post* conservatism because the recognition of bad news reduces both book values and cumulative discretionary accruals. We therefore expect a positive relation between the ratio and market-to-book and between negative relation between the ratio and CDA.

Easton and Pae (2004) provide two measures of conservatism based on an empirical implementation of Feltham and Ohlson (1996). We expect no relation between the ratio and the measure of conservatism due to positive NPV projects, a form of *ex ante* conservatism, because positive NPV investments are rents that we explicitly assume are excluded from book values while the ratio captures the recognition of news about separable assets but not rents. Easton and Pae's (2004) other measure of conservatism reflects the use of high depreciation rates, which is

[25] For example, Phillip Morris recognized a \$3.1 billion charge in 1998 for litigation and also experienced poor stock returns during the year. The settlements also require varying amounts of payments in the future. See Phillip Morris' 1998 10-K for details.

also a form of *ex ante* conservatism. We expect a weak, if any, relation between this measure and the ratio because consistently low depreciation indicates that the firm avoids recognizing bad news about assets while consistently high depreciation indicates that the firm's carrying values are so low that they fall below any negative news about their market values.

Lastly, we consider Penman and Zhang's (2002) 'hidden reserves' measure comprised of the LIFO reserve and hypothetical advertising and R&D assets. This measure represents a form of *ex ante* conservatism. We therefore expect no relation between the ratio and hidden reserves.

We test the relation between measures of conservatism by computing them for each of the Fama and French (1997) 48 industry groups. Table 8 compares Pearson and Spearman rank correlation measures across these groups. Consistent with Givoly, Hayn and Natarajan (2007), the asymmetric timeliness coefficient is negatively associated with market-to-book and conservatism due to NPV projects.[26] Moreover, the asymmetric timeliness coefficient is negatively associated with hidden reserves and positively associated with cumulative discretionary accruals.[27] All four of these correlations are inconsistent with the coefficient measuring the underlying construct of asymmetric recognition of bad news.

Table 8 shows that the ratio statistic exhibits the predicted correlations for all of the reported measures of conservatism. As predicted, we do not find a statistically significant correlation between the ratio measure and the measures of *ex ante* conservatism. We do find, however, that the ratio measure is positively associated with market-to-book and negatively associated with cumulative discretionary accruals, as predicted given how asymmetric recognition of bad news affects these measures.

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[26] The correlation between the Basu (1997) coefficient and conservatism due to NPV is statistically significant only for the rank correlation.

[27] The correlation with cumulative discretionary accruals is statistically significant only for the rank correlation.

Time Series Properties of the Ratio Measure

Figure 2 presents annual estimates of the asymmetric timeliness coefficient and the ratio measure from 1963 to 2007. The figure shows that, starting in the mid-1970s, the asymmetric timeliness ratio statistic has been relatively stable compared to the asymmetric timeliness coefficient. In untabulated analysis, the coefficients of variation (standard deviation divided by the mean) for the coefficient and ratio measures are 0.674 and 0.310, respectively, and 0.428 and 0.124 for years after 1975. The ratio statistic has less variability over time than the asymmetric timeliness coefficient, a desirable property since asymmetric timeliness as required by accounting standards does not change dramatically from year to year.

It appears from Figure 2 that there was a fundamental shift in both the asymmetric timeliness coefficient and the ratio in the mid-1970s. This shift could be explained by the emergence of the Financial Accounting Standards Board (FASB) in 1973 and the early standards written by the board. In 1975, the FASB issue Statement No. 5, *Accounting for Contingencies*, which required the recognition of loss contingencies but not gain contingencies (FASB, 1975). Statement No. 15, *Accounting by Debtors and Creditors for Troubled Debt Restructuring*, was issued in 1977, requiring creditors to reduce yields for adjustments of payment terms of troubled debt if the aggregate payments were not less than the carrying amount; however, if the aggregate payments were less than the carrying amount, the creditor must reduce the carrying amount and record a loss. Therefore, smaller losses adjusted future gains, while larger losses required current period recognition consistent with asymmetric timeliness.[28]

[28] In 1974, the FASB issued Statement No. 2, *Accounting for Research and Development Costs*, requiring the immediate expensing of all research and development costs (FASB, 1974). While this standard did not contribute to asymmetric timeliness in accounting, it shows the FASB's early use of the general principle of conservatism.

VI. CONCLUSION

Many empirical studies on accounting conservatism employ regression-based measures of earnings' asymmetric recognition of bad news as in Basu (1997). Recent literature has shown that such tests may produce invalid test statistics and, in particular, may be unsuitable for comparisons of conservatism across groups of firms or across time (Givoly, Hayn and Natarajan, 2007). We develop an econometric model of the earnings and returns process that provides analytical support for these concerns. The model shows that coefficient estimates appear to provide valid indications of the presence of conservatism but biased measures of the degree of conservatism, rendering cross-group comparisons invalid. We show that a simple test using a ratio of regression coefficients allows for valid comparisons of conservatism across groups.

We derive several predictions from our empirical model and test them on a large sample from the *CRSP* and *Compustat* databases often used in tests of accounting conservatism. We find support for our predictions, which suggests that the model provides a reasonable description of actual data. Furthermore, we verify that the ratio tests are well-behaved, in a statistical sense, by demonstrating that delta method confidence intervals for the ratios provide very close approximations to more robust confidence intervals based on Fieller's Theorem (Fieller, 1954).

We conclude by noting that the usefulness of coefficient ratios extends beyond studies of accounting conservatism. The test design in this study compared the incremental response of an event, bad news, across groups of firms. A variety of studies in accounting, such as those that compare the effects of regulation across groups of firms, utilize a similar test design. We conjecture that inferences based on coefficient ratios may be more appropriate than inferences based on the coefficients, themselves, in cases where coefficient biases may be associated with the sample partition.

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APPENDIX

This appendix describes the construction of confidence intervals for ratios of regression coefficients based on Fieller's Theorem (Fieller 1954) as implemented in Zerbe (1978).[29] The confidence intervals are based on transforming a hypothesis for the ratio to an equivalent linear hypothesis. Given a value k, a vector $\boldsymbol{\beta}$ of p estimated coefficients and vectors $\boldsymbol{n}$ and $\boldsymbol{d}$, the ratio $\boldsymbol{n'\beta} / \boldsymbol{d'\beta} > k$ if and only if $\boldsymbol{n'\beta} - k\boldsymbol{d'\beta} > 0$.[30] A t-statistic for this inequality follows, where $\hat{\boldsymbol{\beta}}$ and $\hat{\boldsymbol{\Sigma}}$ are the estimated coefficients and their estimated covariance matrix, respectively:

$$T = (\boldsymbol{n'\hat{\beta}} - k\boldsymbol{d'\hat{\beta}}) / \sqrt{\boldsymbol{n'\hat{\Sigma}N'} - 2k\boldsymbol{n'\hat{\Sigma}d} + k^2\boldsymbol{d'\hat{\Sigma}d}} \qquad \text{(A1)}$$

Given a confidence level α, sample size N and critical t-statistic $t_{\alpha/2}$ with $N-p$ degrees of freedom such that $\mathrm{P}(-t_{\alpha/2} < T < t_{\alpha/2}) = 1-\alpha$, the confidence intervals on the ratio level k solve:

$$a_0 + 2a_1 k + a_2 k^2 \leq 0 \qquad \text{(A2)}$$

$$a_0 = (\boldsymbol{n'\beta})^2 - t_{\alpha/2}^2\boldsymbol{n'\hat{\Sigma}n} \quad a_1 = t_{\alpha/2}^2\boldsymbol{n'\hat{\Sigma}d} - (\boldsymbol{n'\beta})(\boldsymbol{d'\beta}) \quad a_2 = (\boldsymbol{d'\beta})^2 - t_{\alpha/2}^2\boldsymbol{d'\hat{\Sigma}d} \qquad \text{(A3)}$$

If $a_2 > 0$, then the $1-\alpha$ confidence interval for the ratio is $(-a_1 \pm \sqrt{a_1^2 - a_0 a_2}) / a_2$.[31] If $a_2 < 0$ then the confidence interval is unbounded.[32] In this case, if $a_1^2 - a_0 a_2 < 0$, then the confidence interval is the entire real line. If $a_2 < 0$ and $a_1^2 - a_0 a_2 > 0$, then the confidence interval is the complement of $(-a_1 \pm \sqrt{a_1^2 - a_0 a_2}) / a_2$.

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[29] Also see Staiger, Stock and Watson (1997) for an example application of this approach.

[30] If $\boldsymbol{d'\beta} < 0$, the second inequality is reversed, but (A2) still determines the confidence interval.

[31] See Buonaccorsi (1979) and Zerbe's response for a discussion that $a_2 > 0$ implies $a_1^2 - a_0 a_2 > 0$ and of the confidence interval when $a_2 < 0$.

[32] For example, if the ratio's denominator is the j^{th} regressor $\boldsymbol{\beta}_j$, then $a_2 = (\boldsymbol{\beta}_j^2 / \hat{\boldsymbol{\sigma}}_{jj}^2 - t_{\alpha/2}^2)\hat{\boldsymbol{\sigma}}_{jj}^2$ so that $a_2 < 0$ and the confidence interval is unbounded if $\boldsymbol{\beta}_j$ is insignificant based on a two-tailed test with significance α.

Figure 1: Asymmetric Timeliness Coefficient and Ratio Estimates by Industry

This figure displays negative news coefficients and ratios of the positive and negative news coefficient estimates based on Basu's (1997) asymmetric timeliness regression by 48 Fama and French (1997) industry groups. The estimates are obtained from the following regression of price-scaled earnings per share on returns:

$$EPS_{it}/P_{i,t-1} = \beta_0 + \beta_1 1(r_{it} < 0) + \beta_2 r_{it} + \beta_3 1(r_{it} < 0) r_{it} + \beta_4 EPS_{i,t-1}/P_{i,t-2} + e_t$$

where the indicator variable $1(r_{it} < 0)$ equals one when returns are negative and zero otherwise. The AT Coefficient is the estimate β_3 from the above regression. The AT Ratio equals the estimates $\beta_3 / (\beta_2 + \beta_3)$ from the above regression.

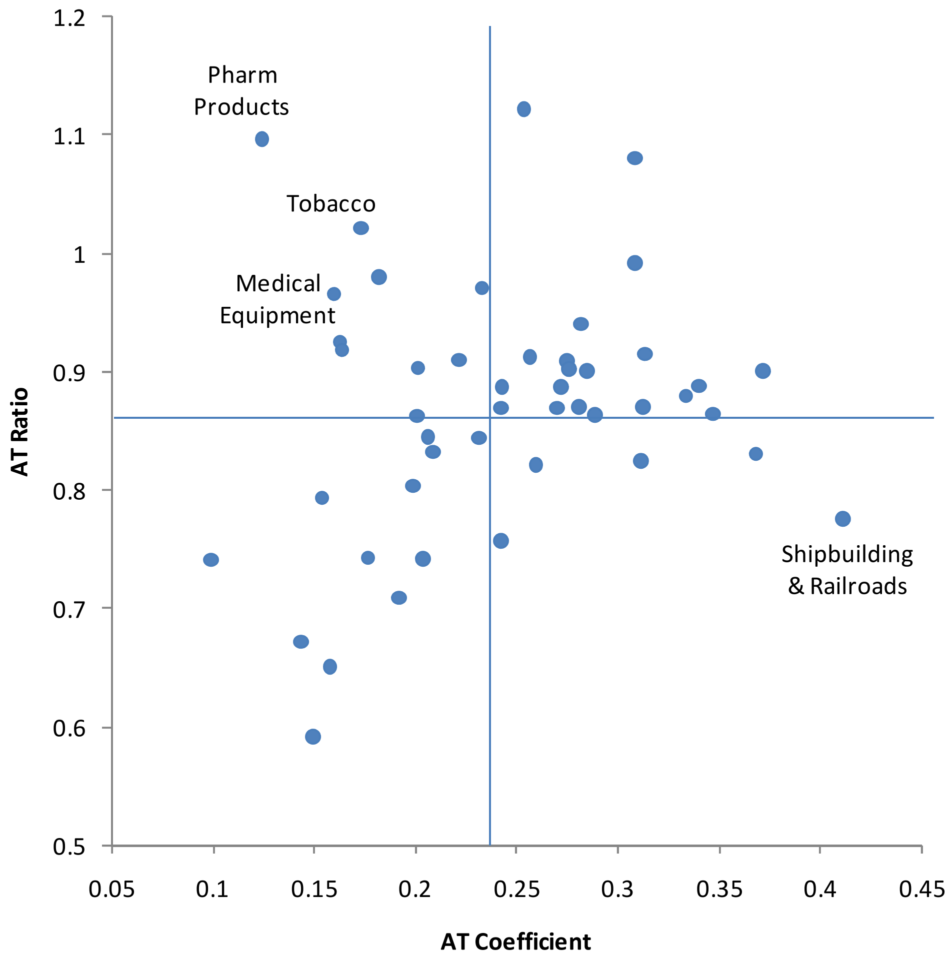

Figure 2: Asymmetric Timeliness Estimates by Year

This figure displays estimates of positive and negative news coefficients and a ratio of those estimates based on Basu's (1997) asymmetric timeliness regression by year from 1963 - 2007. The estimates are obtained from the following regression of price-scaled earnings per share on returns: $EPS_{it}/P_{it-1} = \beta_0 + \beta_1 1(r_{it}<0) + \beta_2 r_{it} + \beta_3 1(r_{it}<0) r_{it} + \beta_4 EPS_{it-1}/P_{it-2} + e_t$ for firm i and year t where the indicator variable $1(r_{it}<0)$ equals one when returns are negative and zero otherwise. The positive and negative news coefficients are the estimates of β_2 and β_3. The AT Ratio equals the estimates $\beta_3/(\beta_2+\beta_3)$ from the above regression.

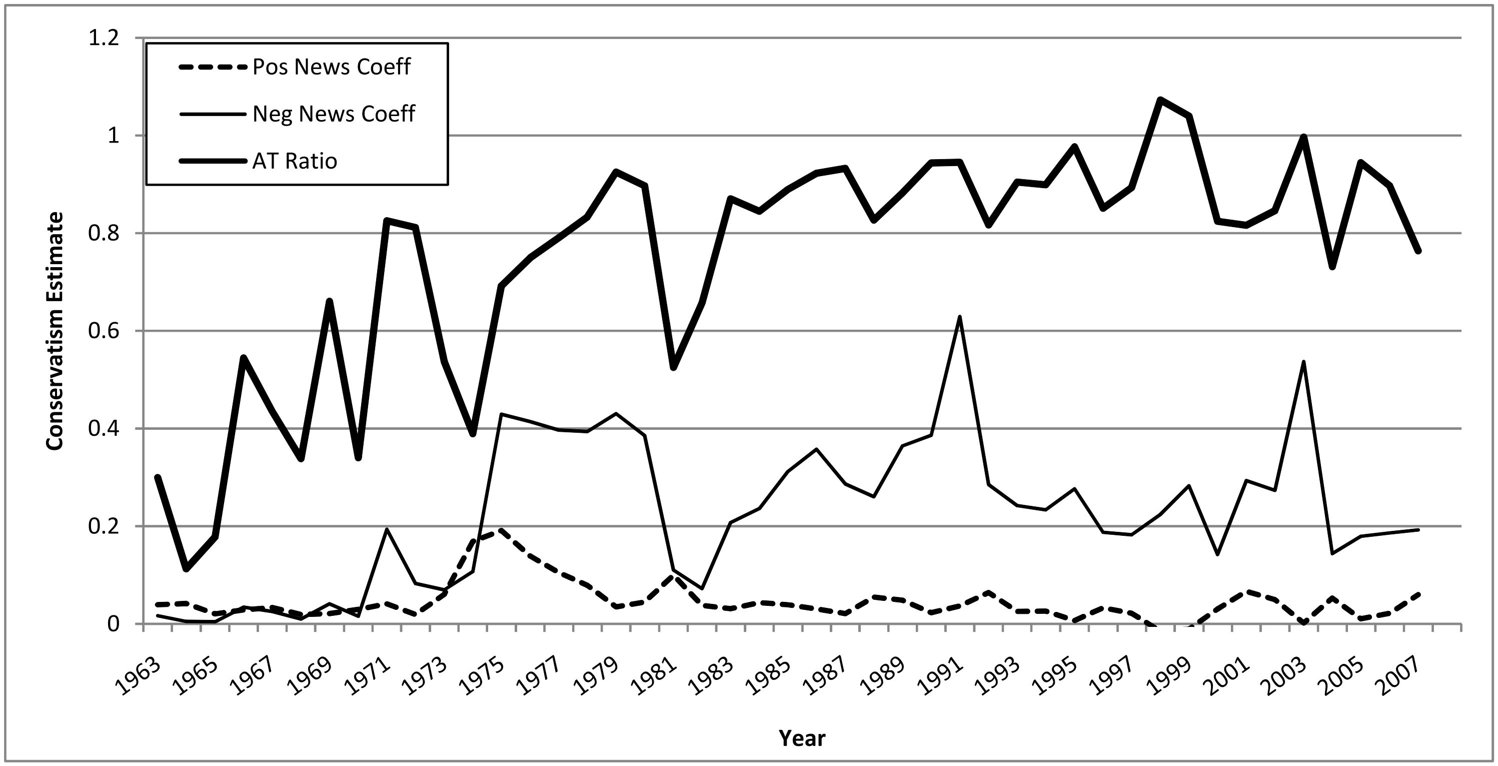

Table 1: Sample Selection and Summary Statistics

This table presents the attrition of the sample and summary statistics. Panel A reports how the sample was obtained after data requirements. The observations cover the years 1963-2007. Panel B reports summary statistics for the sample of firms. Assets (*Compustat* Annual Xpressfeed item *AT*), Book Value of Equity (*Compustat CEQ*), Market Value of Equity (*CRSP PRC* × *SHROUT*), and Book-to-market (Market Value of Equity/Book Value of Equity) are calculated as of the beginning of the fiscal year. Returns are annual buy-and-hold returns obtained from the *CRSP* monthly stock file and begin at the start of the fourth month of the fiscal year. Earnings is earnings per share before extraordinary items (*Compustat EPSPX*) scaled by the stock price (*PRCC_F*) at the beginning of the fiscal year, where we adjust the price for stock splits by multiplying the ratio of the price adjustment factor (*AJEX*) at the earnings announcement date to the factor at the beginning of the fiscal year. The forecast of current earnings is the first mean *I/B/E/S* earnings per share forecast following the announcement of the prior year's earnings. We adjust the forecast for stock splits using the ratio of the *CRSP* share adjustment factor (*CFACSHR*) at the earnings announcement date to the factor at the forecast date. We scale the forecast by the split adjusted beginning-of-fiscal-year share price. All variables are winsorized at 1%/99% within *Compustat* fiscal years.

Panel A: Sample Selection

All U.S. firm-years on *Compustat*	254,127
Less firms with no match on *CRSP*	(48,679)
Less missing *Compustat* /*CRSP* data	(25,563)
Firm-year observations	179,885
Unique firms	18,681

Panel B: Summary Statistics

Variable	N	Mean	Std. dev.	25th pctl.	Median	75th pctl.
Beginning total assets	179,885	1,881.3	8,490.1	28.4	121.3	648.4
Beginning market value of equity	179,885	787.8	3,217.2	18.8	75.1	354.5
Beginning book value of equity	179,885	456.4	1,746.0	12.4	48.6	208.4
Beginning market-to-Book (MTB)	179,885	2.524	3.757	0.922	1.558	2.773
Returns	179,885	0.155	0.654	- 0.211	0.062	0.368
Earnings	179,885	0.011	0.240	- 0.007	0.057	0.104
Forecast of current earnings	98,447	0.069	0.089	0.045	0.073	0.107

Table 2: Asymmetric Timeliness Tests with Prior Information

This table reports coefficient estimates from Basu (1997)-type regressions of earnings on returns with the inclusion of *t* - 1 information to improve the specification. Table 1 describes our sample of firms and variable definitions. Earnings are earnings per share before extraordinary items (*Compustat EPSPX*) scaled by the beginning-of-fiscal-year price. Returns are annual buy-and-hold returns. Forecast is the price-scaled beginning-of-year mean analyst forecast of earnings. Lag EPS is the 1-period lagged Earnings variable (dependent variable). Beginning *BTM* is the beginning-of-fiscal-year book value of equity scaled by market value of equity. All variables except indicator variables are winsorized at the 1st and 99th percentiles within *Compustat* fiscal years. Two-tailed, hetroskedasticity-robust t-statistics are shown below coefficient estimates.

	(1)		(2)		(3)		(4)	
Negative return indicator	- 0.008	***	0.003	*	- 0.003	*	- 0.008	***
	- 4.640		1.920		- 1.880		- 4.820	
Returns	- 0.022	***	0.002		0.018	***	- 0.020	***
	- 12.800		1.120		11.280		- 11.630	
Negative indicator × Returns	0.411	***	0.240	***	0.253	***	0.425	***
	77.970		37.990		54.930		80.180	
Forecast			1.003	***				
			62.170					
Lag EPS					0.406	***		
					67.980			
Beginning BTM							- 0.038	***
							- 18.400	
Intercept	0.072	***	- 0.018	***	0.047	***	0.103	***
	81.690		- 11.150		53.820		57.380	
N	179,885		98,447		179,885		179,885	
F	4,485.3	***	2,991.8	***	5,669.7	***	3,505.2	***
R^2	0.117		0.339		0.399		0.129	

* p<0.10, ** p<0.05, *** p<0.01

Table 3: Results of Cross-Sectional Estimates of Asymmetric Timeliness using Market-to-Book

This table reports coefficient estimates from Basu (1997)-type regressions of earnings on returns with cross-sectional comparisons based on market-to book. Table 1 describes our sample of firms and variable definitions. The high *MTB* indicator is set to one if beginning *MTB* is greater than the 65th percentile of *MTB* for all sample firms and zero if *MTB* is below the 35th percentile (thereby restricting the sample to high and low *MTB* firms). Earnings are earnings per share before extraordinary items scaled by the beginning-of-fiscal-year price. Returns are annual buy-and-hold returns. All variables except indicator variables are winsorized at the 1st and 99th percentiles within *Compustat* fiscal years. Two-tailed, heteroskedasticity-robust t-statistics are shown below coefficient estimates. Significance levels for the ratio tests are computed using the delta method, however we present confidence intervals using both the delta method and the Fieller method as explained in the Appendix (using 5 percent significance). *, **, and *** indicate significance at 0.10, 0.05, and 0.01 percent.

Negative return indicator	0.002	
	0.440	
Returns	0.033	***
	10.210	
Negative indicator × Returns	0.523	***
	40.240	
High MTB indicator	- 0.026	***
	- 13.080	
MTB indicator × Negative indicator	0.002	
	0.400	
MTB indicator × Returns	- 0.034	***
	- 9.150	
MTB indicator × Negative indicator × Returns	- 0.379	***
	- 28.540	
Lag EPS	0.397	***
	61.300	
Intercept	0.059	***
	31.600	
N	125,918	
F	2,586.8	***
R^2	0.429	

Additional Tests:

			Confidence Intervals	
AT coefficient/Recognition coefficient for negative retur	Ratio		Delta	Fieller
High market-to-book	1.006		[0.981, 1.031]	[0.981, 1.031]
Low market-to-book	0.940		[0.928, 0.952]	[0.928, 0.952]
Difference	0.066	***		
F-statistic	21.951			
p-value	0.000			

Table 4: Results of Asymmetric Timeliness Estimates with Book-to-Market Interactions

This table reports coefficient and ratio estimates from Basu (1997)-type regressions of earnings on returns using the original specification and including interactions with book-to-market. Table 1 describes our sample of firms and variable definitions. Earnings are earnings per share before extraordinary items scaled by the beginning-of-fiscal-year price. Returns are annual buy-and-hold returns. All variables except indicator variables are winsorized at the 1st and 99th percentiles within *Compustat* fiscal years. Two-tailed, heteroskedasticity-robust t-statistics are shown below coefficient estimates. Significance levels for the ratio tests are computed using the delta method, however we present confidence intervals using both the delta method and the Fieller method as explained in the Appendix (using 5 percent significance). *, **, and *** indicate significance at 0.10, 0.05, and 0.01 percent.

Panel A: Estimates for the original specification and including BTM as control

$$\frac{x_t}{v_{t-1}} = \beta_0 + \beta_1 1_{r_{t<0}} + \beta_2 r_t + \beta_3 1_{r_{t<0}} r_t + \beta_4 \frac{x_{t-1}}{v_{t-2}} + e_t$$

	β_0	β_1	β_2	β_3	β_4
Coefficient	0.047 ***	-0.003 *	0.018 ***	0.253 ***	0.253 ***
t-statistic	53.82	-1.88	11.28	54.93	54.93

$$\frac{x_t}{v_{t-1}} = \delta_0 + \delta_1 1_{r_{t<0}} + \delta_2 \frac{b_{t-1}}{v_{t-1}} r_t + \delta_3 1_{r_{t<0}} \frac{b_{t-1}}{v_{t-1}} r_t + \delta_4 \frac{x_{t-1}}{v_{t-2}} + e_t$$

	δ_0	δ_1	δ_2	δ_3	δ_4
Coefficient	0.048 ***	-0.022 ***	0.020 ***	0.306 ***	0.306 ***
t-statistic	55.31	-14.75	10.97	41.06	41.06

Panel B: Tests for differences in coefficients and ratios

			Confidence Intervals	
	Basu Coeff.	Ratio	Delta	Fieller
Original Specification	0.253	0.933	[0.922, 0.945]	[0.922, 0.945]
Including BTM as Regressor	0.306	0.940	[0.929, 0.951]	[0.929, 0.951]
Difference	- 0.053 ***	- 0.006		
F-statistic	50.680	1.320		
p-value	0.000	0.250		

Table 5: Results of Cross-Sectional Estimates of Asymmetric Timeliness for REITs and non-REITs

This table reports cross-sectional comparisons of coefficient and ratio estimates from Basu (1997)-type regressions of earnings on returnsfor REITs (Real Estate Investment Trusts) and non-REITs. Table 1 describes our sample of firms and variable definitions. The REIT is set to one if the firm's SIC code is equal to 6798 and zero otherwise. Earnings are earnings per share before extraordinary items scaled by the beginning-of-fiscal-year price. Returns are annual buy-and-hold returns. All variables except indicator variables are winsorized at the 1st and 99th percentiles within Compustat fiscal years. Two-tailed, heteroskedasticity-robust t-statistics are shown below coefficient estimates. Significance levels for the ratio tests are computed using the delta method, however we present confidence intervals using both the delta method and the Fieller method as explained in the Appendix (using 5 percent significance). *, **, and *** indicate significance at 0.10, 0.05, and 0.01 percent.

Negative return indicator	- 0.003	**
	- 2.320	
Returns	0.018	***
	11.010	
Negative indicator × Returns	0.252	***
	54.620	
REIT	- 0.028	***
	- 3.920	
REIT × Negative indicator	0.044	***
	3.610	
REIT × Returns	0.022	
	0.810	
REIT × Negative indicator × Returns	0.180	***
	2.740	
Lag EPS	0.406	***
	67.920	
Intercept	0.048	***
	54.000	
N	179,885	
F	2,842.0	***
R^2	0.399	

Additional Test:

		Confidence Intervals	
AT coefficient/Recognition coefficient for negative returns	Ratio	Delta	Fieller
REITs	0.916	[0.803, 1.030]	[0.793, 1.030]
Non-REITs	0.935	[0.923, 0.946]	[0.923, 0.946]
Difference	- 0.019		
F-statistic	0.100		
p-value	0.752		

Table 6: Asymmetric Timeliness Tests using Alternative Scaling Factors

This table reports coefficient estimates from Basu (1997)-type regressions of earnings on returns using alternative scaling variables. Table 1describes our sample of firms and variable definitions. The beginning price is the split adjusted beginning-of-fiscal-year price. Earnings in regression (1) are price-scaled earnings per share before extraordinary items. Earnings in regression (2) are earnings per share before extraordinary items (*Compustat EPSPX*). Earnings in regressions (3) and (4) are net income before extraordinary items (*Compustat IB*) divided by the beginning-of-fiscal-year scaling variable. Returns in regression (1) are annual buy-and-hold returns. Returns in regression (2) are annual buy-and-hold returns multiplied by the beginning price. Returns in regression (3) and (4) are annual buy-and-hold returns multiplied by the beginning price and divided by the beginning-of-fiscal-year scaling variable. All variables except indicators are winsorized at the 1st and 99th percentiles within *Compustat* fiscal years. Two-tailed, hetroskedasticity-robust t-statistics are shown below coefficient estimates.

	(1a)	(1b)	(2a)	(2b)	(3a)	(3b)	(4a)	(4b)
	Scale by Market value of equity		Scale by Common shares		Scale by Assets		Scale by Book value of equity	
	Orig	w/Price	Orig	w/Price	Orig	w/Assets	Orig	w/Book value
Negative return indicator	- 0.008 ***	- 0.005 ***	- 0.588 ***	- 0.370 ***	- 0.059 ***	- 0.052 ***	- 0.112 ***	- 0.099 ***
	- 4.645	- 3.280	- 48.421	- 39.829	- 59.192	- 55.229	- 39.417	- 35.258
Returns	- 0.022 ***	- 0.008 ***	0.074 ***	0.006 ***	- 0.028 ***	0.004	- 0.037 ***	- 0.007
	- 12.802	- 4.725	55.585	5.859	- 5.862	0.910	- 5.310	- 0.985
Negative indicator × Returns	0.411 ***	0.367 ***	- 0.079 ***	0.122 ***	0.379 ***	0.328 ***	0.542 ***	0.496 ***
	77.973	70.956	- 31.628	54.251	32.403	28.699	35.571	32.875
Beginning price		0.003 ***		0.078 ***		0.003 ***		0.006 ***
		76.171		192.323		91.995		73.130
Intercept	0.072 ***	0.022 ***	0.903 ***	0.012 *	0.038 ***	- 0.014 ***	0.102 ***	- 0.003
	81.690	16.943	106.053	1.795	80.926	- 17.156	66.526	- 1.401
N	179,885	179,885	179,885	179,885	179,885	179,885	179,885	179,885
F	4,485.3 ***	4,402.0 ***	5,749.4 ***	15,247.1 ***	3,079.7 ***	4,544.8 ***	2,095.7 ***	2,967.3 ***
R^2	0.117	0.140	0.128	0.399	0.085	0.133	0.088	0.116

* p<0.10, ** p<0.05, *** p<0.01

Table 7: Results of Asymmetric Timeliness Estimates Including and Excluding Goodwill Writedowns

This table reports coefficient and ratio estimates from Basu (1997)-type regressions of earnings on returns with both earnings as reported and earnings excluding the effect of goodwill write-downs. Table 1describes our sample of firms and variable definitions. The sample of firms in this table only includes firm-years from 2001 to 2007 because this represents the period in which SFAS 141 was in effect. Earnings are earnings per share before extraordinary items scaled by the beginning-of-fiscal-year price. The effect of goodwill on earnings per share is obtained from *Compustsat* (*GDWLIEPS*). Returns are annual buy-and-hold returns. All variables except indicator variables are winsorized at the 1st and 99th percentiles within *Compustat* fiscal years. Two-tailed, heteroskedasticity-robust t-statistics are shown below coefficient estimates. Significance levels for the ratio tests are computed using the delta method, however we present confidence intervals using both the delta method and the Fieller method as explained in the Appendix (using 5 percent significance). *, **, and *** indicate significance at 0.10, 0.05, and 0.01 percent.

	Earnings as Reported		Excluding Goodwill Writedowns	
Negative return indicator	0.010	***	0.005	*
	2.840		1.660	
Returns	0.013	***	0.010	***
	3.180		2.710	
Negative indicator × Returns	0.299	***	0.268	***
	25.200		24.440	
Lag EPS	0.315	***	0.299	***
	32.900		32.800	
Intercept	0.026	***	0.029	***
	12.880		14.870	
N	33,417		33,417	
F	926.3	***	933.0	***
R^2	0.404		0.403	

Additional Tests:

	Basu Coeff.		Ratio	Confidence Intervals: Delta	Confidence Intervals: Fieller
EPS including goodwill writedowns	0.299		0.960	[0.935, 0.985]	[0.934, 0.985]
EPS excluding goodwill writedowns	0.268		0.962	[0.935, 0.990]	[0.935, 0.990]
Difference	0.031	***	- 0.003		
F-statistic	8.020		0.040		
p-value	0.005		0.841		

Table 8: Correlation Between Conservatism Measures

This table reports the correlations between different measures of accounting conservatism based on estimates for each of the 48 Fama and French (1997) industries. The table displays Pearson correlations below the diagonal and Spearman correlations above, with p-values reported in parenthesis below the correlation estimates. The asymmetric timeliness (AT) coefficient is the estimate of β_3 in a regression of price-scaled earnings per share on returns: $EPS_{it}/P_{i,t-1} = \beta_0 + \beta_1 1(r_{it}<0) + \beta_2 r_{it} + \beta_3 1(r_{it}<0)\, r_{it} + e_{it}$ for firm i and year t where the indicator variable $1(r_{it}<0)$ equals one when returns are negative and zero otherwise. Ratio equals the estimates $\beta_3/(\beta_2+\beta_3)$. Conservatism due to the net present value of investments and Conservatism due to accounting rules are the estimates of β_4 and β_5, respectively, from the regression $r_{it} = \beta_0 + \beta_1 x_{it}/p_{i,t-1} + \beta_2 \Delta x_{it}/p_{i,t-1} + \beta_3 d_{i,t-1}/p_{i,t-1} + \beta_4 \Delta c_{it}/p_{it} + \beta_5 \Delta oa_{i,t-1}/p_{i,t-1} + e_{it}$ of returns on price-scaled comprehensive income (x_{it}), cash payout (d_{it}), cash investments (ci_{it}) and operating assets (oa_{it}) as defined in the text (Easton and Pae, 2004). Hidden reserves are the sum of the LIFO reserve, hypothetical capitalized R&D expenditures and hypothetical capitalized advertising expenditures, scaled by net operating assets as defined in the text (Penman and Zhang, 2002). Cumulative discretionary accruals are the sum of total accruals (net income + depreciation - cash from operations) less operating accruals (change in accounts receivable + change in inventory - change in accounts payable - change in taxes payable) scaled by total assets in the firm's first year with available data (Givoly and Hayn, 2000). We list our predictions for correlations with the coefficient and ratio measures only listed as 'pos' for positive, 'neg' for negative, and 'none' for no prediction.

	Prediction (coeff & ratio)	AT coefficient	Ratio	Market-to-book	Conservatism due to NPV	Conservatism due to rules	Hidden reserves	Discretionary accruals
AT coefficient			0.226 (0.123)	- 0.364 (0.011)	- 0.310 (0.032)	- 0.023 (0.878)	- 0.597 (0.000)	0.259 (0.076)
Ratio	pos	0.294 (0.043)		0.455 (0.001)	- 0.042 (0.777)	0.151 (0.307)	0.065 (0.663)	- 0.396 (0.005)
Market-to-book	pos	- 0.380 (0.008)	0.432 (0.002)		0.206 (0.161)	0.056 (0.705)	0.508 (0.000)	- 0.717 (0.000)
Conservatism due to NPV	none	- 0.174 (0.238)	- 0.184 (0.210)	- 0.049 (0.740)		0.268 (0.066)	0.221 (0.132)	- 0.396 (0.005)
Conservatism due to rules	none	0.074 (0.616)	0.168 (0.255)	0.082 (0.580)	- 0.941 (0.000)		- 0.160 (0.276)	- 0.159 (0.281)
Hidden reserves	none	- 0.543 (0.000)	0.205 (0.163)	0.684 (0.000)	0.094 (0.524)	- 0.081 (0.583)		- 0.287 (0.048)
Discretionary accruals	neg	0.222 (0.129)	- 0.410 (0.004)	- 0.781 (0.000)	0.021 (0.886)	- 0.142 (0.336)	- 0.499 (0.000)	