Identifying Conditional Conservatism in Accounting Data: Theory and Evidence*

Sunil Dutta
dutta@haas.berkeley.edu

Panos N. Patatoukas
panos@haas.berkeley.edu

Haas School of Business
University of California, Berkeley

This Version: April 6, 2016

* We thank Mary Barth, Jeremy Bertomeu, Peter Easton, Stuart McLeay (discussant), Fernando Peñalva, Jim Ohlson, Stephen Penman, Richard Sloan, Charles Wasley, Ari Yezegel, Paul Zarowin, Xiao-Jun Zhang, two anonymous reviewers, and seminar participants at Berkeley-Haas, Athens University of Economics and Business, Columbia Business School, the University of Rochester, and the 2016 "Methodological and Empirical Advances in Financial Analysis" Research Meeting at the University of Sydney for helpful comments. Panos gratefully acknowledges financial support from the Center for Financial Reporting & Management and the Hellman Fellows Fund.

**Identifying Conditional Conservatism in Accounting Data:
Theory and Evidence**

ABSTRACT

Using a financial reporting and valuation model, we investigate the construct validity of Basu's (1997) asymmetric timeliness (AT) regression coefficient as a measure of asymmetrically timely loss recognition or "conditional conservatism" in corporate financial reporting. Within the context of our model, we predict that the AT coefficient will be positive *even* in the absence of conditional conservatism and it will vary with non-accounting factors *even* if the degree of conditional conservatism is held constant. Our empirical analysis shows that AT coefficient estimates vary in directions predicted by our theory. Specifically, we find that AT coefficient estimates increase with expected returns and asymmetry in the distribution of returns, and decrease with cash flow persistence. Importantly, we identify the spread between the variances of bad news and good news accruals as an alternative measure of conditional conservatism that is free of the effects confounding the AT coefficient. Consistent with a key implication of conditional conservatism, we find evidence that the variance of bad news accruals is significantly higher than the variance of good news accruals primarily due to "conditionally conservative accruals" related to inventory write-downs, long-term asset write-downs, and goodwill impairments. A series of "placebo" tests provides additional support for the construct validity of our alternative measure of conditional conservatism.

Keywords: Conditional conservatism; asymmetric timeliness; loss recognition; corporate financial reporting; valuation; construct validity; placebo tests.

Data Availability: Data are publicly available from the sources indicated in the text.

I. INTRODUCTION

A large body of accounting research relies on Basu's (1997) piecewise linear regression of earnings on positive and negative stock returns to identify asymmetrically timely loss recognition or "conditional conservatism" in corporate financial reporting. By and large, this body of research has intuitively assumed that the incremental coefficient on negative returns, also known as the asymmetric timeliness (AT) coefficient, is a valid measure of the degree of conditional conservatism in accounting. In this paper, we have two objectives. Our first objective is to investigate, both theoretically and empirically, the construct validity of the AT coefficient. Our second objective is to derive and empirically validate an alternative measure of conditional conservatism that is primarily based on the distributional properties of accounting data.

With these two objectives in mind, we develop a financial reporting and valuation model of the relations among fundamental news, accounting earnings, and stock returns. Our model distinguishes between the accounting recognition of news about future cash flows, which we refer to as fundamental news or cash flow news, and the capitalized value of such news in stock returns, which we refer to as return news. The distinction between fundamental news and return news is a key ingredient of our model. Another key feature of our model is the assumption of asymmetry in the distribution of fundamental news, and therefore return news, which is motivated by long-standing evidence of asymmetry in the returns distribution (e.g., Fama 1965).

Researchers in accounting often rely on intuition and "verbal theory" to motivate empirical measures of theoretical constructs, which may introduce confusion with respect to construct validity and inhibit the identification of causal effects (e.g., Bertomeu et al. 2015). In this paper, we adopt a modelling approach that allows us to formalize the assumptions needed to identify conditional conservatism in accounting data and derive an alternative measure of asymmetrically timely loss recognition. Our analytical model also highlights sources of variation in the AT coefficient that have not been considered in prior research.[1]

[1] Prior studies have identified a variety of non-accounting reasons for asymmetric timeliness. For example, Hayn (1995) considers real options (abandonment option), Givoly et al. (2007) consider information environment characteristics (aggregation effect), Patatoukas and Thomas (2011) consider scale-related regularities, and Banker et al. (2016) consider operational reasons (cost stickiness). Patatoukas and Thomas (2016) propose that spurious evidence of asymmetric timeliness could arise for any variable, even if it is unconditionally unrelated to returns, as long as there is a relation between the first moment of that variable and the second or higher moments of the distribution of returns.

Within the context of our model, we find that the AT coefficient depends not only on the degree of conditional conservatism, but also on non-accounting factors. The reason is that the good and bad news slope coefficients in the Basu regression, and hence the AT coefficient, i.e., the difference between the conditional slope coefficients, increase with the inverse of the rate at which the market capitalizes fundamental news into stock returns. Since the capitalization rate *decreases* in expected returns and *increases* in cash flow persistence, the AT coefficient *increases* in expected returns and *decreases* in cash flow persistence even in the absence of variation in the degree of conditional conservatism.

Our analysis also shows that asymmetry in the returns distribution has the potential to yield spurious evidence of asymmetric timeliness *even* if accounting is entirely symmetric. Indeed, our theory predicts a positive AT coefficient for the cash flow component of earnings as long as the variance of returns conditional on good news is higher than that conditional on bad news. This prediction holds despite the fact that cash flows are independent of accounting recognition rules and is consistent with spurious evidence of asymmetry dating back to Basu's (1995) dissertation (e.g., Basu 1997; Ryan and Zarowin 2003; Dietrich et al. 2007; Hsu et al. 2012; Collins et al. 2014). Our theory further predicts that the spuriously positive AT coefficient for cash flows increases with asymmetry in the returns distribution.

Overall, we identify three non-accounting factors—expected returns, cash flow persistence, and asymmetry in the returns distribution—that confound inferences about variation in the degree of conditional conservatism from variation in the AT coefficient. Our analysis predicts that the AT coefficient increases in expected returns and asymmetry in the returns distribution, and decreases in cash flow persistence *even* if the degree of conditional conservatism is held constant. These confounding effects are in play regardless of whether the dependent variable in the AT regression is total earnings, as in Basu (1997), or the unexpected component of earnings, as recommended by Ball et al. (2013a), or the accrual component of earnings, as recommended by Collins et al. (2014).

To test our theory, we compile a comprehensive sample and search for variation in AT coefficient estimates with the three non-accounting factors identified. Consistent with our predictions, we find that AT coefficient estimates increase with expected returns and asymmetry in the conditional variances of positive and negative unexpected returns, and decrease with cash flow persistence. Our efforts to control for variation in the non-accounting factors identified, however, result in substantially weakened AT coefficient estimates. The evidence suggests that it may not be feasible to disentangle the confounding effects of non-accounting factors from the effect of conditional conservatism within the AT regression.

As a way forward, we derive and empirically validate an alternative measure of conditional conservatism based on asymmetry in the distribution of the accrual component of reported earnings. Our alternative measure is motivated by the intuition that in the presence of asymmetrically timely loss recognition, fundamental news is likely to be recognized in accruals to a greater extent when the news is bad than when the news is good. Under conditional conservatism, therefore, one would expect that accruals should be more variable for bad news relative to good news partitions.

Consistent with this intuition, our theoretical analysis shows that the spread between the conditional variances of bad news and good news accruals will be positive if accounting is conditionally conservative, and will be zero if accounting is symmetric. The spread between the conditional variances of bad news and good news accruals is independent of the market capitalization factor and is unaffected by asymmetry in the conditional variances of returns. Thus, unlike the AT coefficient, our alternative measure of conditional conservatism is free of the confounding effects of non-accounting factors, including expected returns, cash flow persistence, and asymmetry in the returns distribution. Moreover, our alternative approach to measure conditional conservatism is primarily based on the distributional properties of accounting data and does not rely on the notion of market efficiency beyond that the sign of unexpected returns is a reasonable proxy for the sign of fundamental news. In contrast, however, the AT regression approach to measure conditional conservatism explicitly relies on the assumption that the market incorporates all news in a timely and efficient manner, regardless of whether or not such news is yet reflected in accounting earnings; i.e., stock market values lead accounting earnings (e.g., Basu 1997; Pope and Walker 1999; Ball et. al. 2013b).

Turning to the empirical results, we find evidence that the variance of bad news accruals is significantly higher than the variance of good news accruals, which is consistent with a key implication of conditional conservatism. Our evidence of asymmetry in the conditional variances of accruals is reassuring given that asymmetrically timely loss recognition is ingrained in GAAP practices of lower-of-cost-or-market accounting for inventories (ASC 330), goodwill impairments (ASC 350), and asset write-downs (ASC 360). Indeed, we also find evidence that asymmetry in the distribution of total accruals is mostly attributed to the subset of "conditionally conservative accruals" related to inventory write-downs, long-term asset write-downs, and goodwill impairments.

A series of construct validity tests provides additional evidence that the spread in the conditional variances of accruals offers a viable alternative measure of conditional conservatism. In contrast to prior spurious evidence of a positive AT coefficient for the cash flow component of earnings, we do not find evidence of a positive spread in the conditional variances of bad news and good news cash flows. In addition, we find *no* evidence of asymmetry in the conditional variances of placebo test variables, including the accrual and cash flow components of lagged earnings, across partitions based on the sign of one-year-ahead unexpected returns consistent with the idea that *lagged* accounting data cannot be conditionally conservative with respect to *future* news.

Prior studies have raised questions about the validity of the AT coefficient as a measure of accounting conservatism (e.g., Dietrich et al. 2007; Givoly et al. 2007; Patatoukas and Thomas 2011, 2016). Ball et al. (2013b) argue that the AT coefficient provides a valid measure of conditional conservatism, identifying conditional conservatism only when it exists. Their analytical model, however, assumes symmetry in the returns distribution and abstracts away from non-accounting sources of variation in the AT coefficient. Pope and Walker (1999) also study the AT coefficient in a valuation framework and point out the confounding effect of variation in expected returns. Pope and Walker (1999), however, assume symmetry in the returns distribution and they do not consider the confounding effect of variation in cash flow persistence.[2] Our theory and evidence corroborate the conclusion of Patatoukas and Thomas (2016) that AT estimates of conditional conservatism are biased upward because they reflect factors unrelated to asymmetrically timely loss recognition in corporate financial reporting.

[2] Other papers with formal analyses of the AT regression include Ball et al. (2000), Ryan and Zarowin (2003), Beaver and Ryan (2005, 2009), Dietrich et al (2007), Callen et al. (2010). These papers, however, do not consider the confounding effects of non-accounting factors. More recently, Armstrong et al. (2015) show that asymmetry in financial reporting surfaces endogenously and generates asymmetry in systematic risk.

The rest of the paper is organized as follows. Section II provides the theoretical analysis. Section III provides the empirical analysis. Section IV concludes.

II. THEORETICAL ANALYSIS

In this section, we develop a simple yet realistic financial reporting and valuation framework to model the relations among fundamental news, accounting earnings, and stock returns. We consider an indefinitely lived and publicly traded firm. Before we discuss accounting measurement and recognition rules used to translate fundamentals into accounting reports, we describe the time-series evolution of the firm's economic performance.

Suppose that the firm's free cash flows $\{CF_t\}$ evolve according to the following autoregressive process:

$$CF_t = wCF_{t-1} + (1-w)m + S\phi_t \tag{1}$$

where w is a persistence parameter between zero and one and $m > 0$ denotes the unconditional mean. The last term of the above equation, $S\phi_t$, is an innovation term, where S represents the scale of the firm's operations and ϕ_t is a random term with mean zero. We assume that the random terms ϕ_t are serially uncorrelated. The process in equation (1) implies that, all else equal, free cash flows are more volatile for larger firms since $Var(CF_t) = S^2 Var(\phi_t)$. This is consistent with the evidence in Figure 1, which shows that the variance of unscaled free cash flows (as well as that of earnings and accruals) increases in the beginning of period market value of equity.

Though we purposefully choose a simple model to illustrate the key insights of our analysis, our results can be extended to more general settings. For instance, one can introduce deterministic growth to the cash flow dynamics in equation (1) by making the scale parameter S grow over time. Our framework can also be extended to accommodate separate dynamics for different free cash flow components (e.g., cash flow from operations and capital expenditures). For expositional simplicity, however, we do not consider such extensions because they do not affect the qualitative nature of our results.

Financial reporting model

Before describing the specifics of the financial reporting model, it is helpful to review how information about firm performance is incorporated in accounting reports and market values. As a concrete example, suppose there is an unfavorable shock to demand for the firm's output. We refer to such information about future cash flows as *fundamental news* or *cash flow news*. This news will clearly affect the value of the firm's inventories, and hence near-term sales and cash flows. However, such unfavorable demand news might also have a persistent component that would negatively impact *all future cash flows* including those from anticipated production and investment choices yet to be undertaken (i.e., future growth opportunities).

An efficient market will rationally reflect the total *capitalized* impact of this news in the current price. We refer to the fully capitalized effect of fundamental news on firm value as *returns news*. In contrast, most financial reporting regimes would require current earnings to reflect, at most, changes to the values of the firm's *assets-in-place* (e.g., inventories) due to this unfavorable demand news. Put differently, current earnings would typically capture only a portion of the total *return news*. Importantly, the relation between fundamental news and return news will vary across firms depending on characteristics such as the discount rate and cash flow persistence. Such sources of cross-sectional variation in the relation between fundamental news and return news play a central role in our analysis of non-accounting sources of variation in the AT coefficient.[3]

Depending on the specific income recognition rules, accounting earnings in period t will reflect not only historical information about the cash flows realized to date, but also additional information about future cash flows. To model the firm's information about future cash flows, we assume that the (scaled) innovation term in equation (1), ϕ_t, is given by

$$\phi_t = f_t + g_t + h_t,$$

where the random terms f_t, g_t, and h_t are independently distributed with zero means and variances of F, G, and H, respectively. We also assume that these three random terms are serially uncorrelated. The firm has perfect advance information about components f_{t+1} and g_{t+1} of its one-period-ahead cash flows at the end of the current period t, but learns about the third component, h_{t+1}, only when the associated cash flows are realized in period t+1.[4] With respect to the recognition

[3] In contrast to our approach, Ball et al. (2013b) abstract away from non-accounting sources of variation in the AT coefficient by assuming that the relation between fundamental news and returns news is identical for all firms. Using a permanent earnings representation of accounting income, Pope and Walker (1999) allow for the relation between fundamental news and return news to vary with expected returns, but not with cash flow persistence.

[4] Our results readily generalize to the case when fundamental news relates not only to one-period-ahead cash flows, but also to information about cash flows in periods beyond t+1.

of fundamental news into current earnings, we assume that g_{t+1} is unconditionally recognized in period t earnings one-period-ahead of its realization as cash flows. Component h_{t+1} is unconditionally recognized in period t +1 earnings concurrently with its realization as cash flows. In contrast, we assume that recognition of f_{t+1} in period t earnings is subject to conditional conservatism.

To model conditional conservatism, let $d_t \in [0,1]$ denote the fraction of cash flow news f_{t+1} recognized in period t accounting earnings. Conditional conservatism implies that d_t is higher when news is bad than when news is good. To model this, let $d_t = \delta \in [0,1]$ denote the fraction of f_{t+1} recognized in period t earnings when news is bad (i.e., $f_{t+1} < 0$). For notational brevity, we assume that $d_t = 0$ when news is good (i.e., $f_{t+1} \geq 0$).[5] The polar case of $\delta = 1$ represents the extreme case of conservatism when *all* bad news, but no good news, is recognized in earnings. At the other extreme, accounting recognition is completely *symmetric* when $\delta = 0$. In this case, revisions in expectations about future cash flows are recognized with delay, regardless of whether such revisions reflect good news or bad news.[6]

Accounting earnings in period t can then be written as

$$Y_t = wCF_{t-1} + (1-w)m + S[d_t f_{t+1} + (1-d_{t-1}) f_t + g_{t+1} + h_t + \mu_t - \mu_{t-1}],$$

where μ_t denotes an accrual estimation "error" that reverses in the next period. The above expression reflects that h_t is always recognized in earnings concurrently with cash flows. Accounting income in period t recognizes cash flow news g_{t+1} irrespective of whether it is positive or negative news, but only fraction d_t of cash flow news f_{t+1}, where d_t depends on the sign of f_{t+1}. We note that the above expression for accounting earnings is consistent with the notion of clean surplus accounting.

---

[5] Our results extend to a setting where a non-zero fraction of good news is recognized in earnings as long as the fraction of news recognized is higher for bad news relative to good news. We also note that, as in Pope and Walker (1999), conditional conservatism is a continuous variable in our model. In contrast, Ball et al. (2013b) examine a binary setting where conservatism is either present or not.

[6] Prior research attributes asymmetric timeliness partially to non-discretionary GAAP accounting practices and partially to managers exercising discretion over accounting practices. For our analysis, it does not matter whether conditional conservatism arises from discretionary or mandatory reporting practices.

Since CF_{t-1}, f_t, and μ_{t-1} are known at the beginning of period t, unexpected earnings in period t are given by $y_t = S[d_t f_{t+1} + g_{t+1} + \mu_t]$. Therefore, scaled unexpected earnings in period t, $x_t = y_t / S$, are simply given by

$$x_t = d_t \cdot f_{t+1} + g_{t+1} + \mu_t. \tag{2}$$

Using the expression for free cash flows CF_t, accounting earnings can also be expressed as $Y_t = CF_t + AC_t$, where $AC_t = S[d_t f_{t+1} - d_{t-1} f_t + g_{t+1} - g_t + \mu_t - \mu_{t-1}]$ denotes the amount of total accruals in period t. Hence, scaled accruals in period t are given by

$$ac_t = d_t f_{t+1} - d_{t-1} f_t + g_{t+1} - g_t + \mu_t - \mu_{t-1}. \tag{3}$$

Valuation model

For brevity, we normalize the number of shares outstanding to one and assume that the firm does not retain any cash. Hence, CF_t is distributed as dividends to the shareholders at the end of period t.[7] The stock market value at the end of period t is then given by

$$P_t = \sum_{\tau=1}^{\infty} \frac{E_t[CF_{t+\tau}]}{(1+R^E)^\tau}$$

where R^E is the firm's expected return (i.e., cost of capital) and $E_t[CF_{t+\tau}]$ denotes the market's expectation of free cash flows in period $t + \tau$ conditional on the information available at date t.

The AT regression approach to measure conditional conservatism is predicated on the assumption that the market incorporates all news in a timely and efficient manner, regardless of whether or not such news is yet reflected in accounting earnings; i.e., stock market values lead accounting earnings (e.g., Basu 1997; Pope and Walker 1999; Ball et. al. 2013b). Consistent with this approach, we assume that period t stock market value P_t incorporates the capitalized values of both fundamental news components g_{t+1} and f_{t+1}, regardless of the extent to which f_{t+1} is reflected in period t accounting earnings. Our analysis will show that the AT coefficient is confounded by variation in non-accounting factors *even* if this strong market efficiency assumption were to hold. In our subsequent analysis, we propose an alternative measure of conditional conservatism that

[7] The assumption that the firm does not carry any cash is without loss of generality, since dividend policy is irrelevant in our symmetric information setting.

does not rely on the notion of market efficiency beyond that the sign of unexpected returns is a reasonable proxy for the sign of fundamental news.

Given the autoregressive process in (1), the market's expectation of free cash flows in period $t+\tau$ conditional on date t information can be expressed as

$$E_t[CF_{t+\tau}] = w^{\tau} CF_t + (1-w^{\tau})m + w^{\tau-1} S(f_{t+1} + g_{t+1}).$$

Substituting the above expression into the valuation formula and simplifying (see Appendix 1 for a detailed derivation) yield the following expression for the stock market value at date t:

$$P_t = \frac{1}{R^E + 1 - w}\left[wCF_t \; + S(f_{t+1} + g_{t+1}) + \frac{(1-w)(1+R^E)}{R^E} m \right]. \tag{4}$$

Here, $(R^E + 1 - w)^{-1}$ represents the rate at which fundamental news (i.e., f_{t+1} and g_{t+1}) is capitalized into the firm's stock market value. The capitalization factor decreases in expected returns R^E and increases in cash flow persistence w. For positive values of the persistence parameter, news about cash flows in the next period affects the market's expectation of cash flows not only in the next period, but also in all subsequent periods. For the extreme case when cash flows follow a random walk (i.e., w = 1), the capitalization factor simplifies to $1/R^E$.

The unexpected return in period t is defined as

$$r_t = \frac{P_t + CF_t - (1+R^E)P_{t-1}}{P_{t-1}}.$$

Substituting for P_{t-1} and P_t from equation (4) into the numerator of the above expression and simplifying give

$$r_t = \frac{S[f_{t+1} + g_{t+1} + (1+R^E)h_t]}{P_{t-1}(R^E + 1 - w)}.$$

Using the beginning-of-period market value of equity as the scale measure (i.e., $S = P_{t-1}$), the unexpected return in period t simplifies to[8]

[8] The qualitative nature of our results would remain unchanged under alternative measures of scale of operations such as book value of equity, book value of total assets, and sales.

$$r_t = \frac{f_{t+1} + g_{t+1} + (1+R^E)h_t}{R^E + 1 - w}. \tag{5}$$

Equation (5) shows that the unexpected return in period t is the capitalized sum of three news components: f_{t+1}, g_{t+1}, and h_t. Component h_t represents news about current cash flows that is contemporaneously reflected in accounting earnings and returns. In contrast, f_{t+1} and g_{t+1} reflect news about cash flows in the next period. While g_{t+1} is always reflected in accounting earnings and returns in period t, f_{t+1} represents the portion of the unexpected return r_t that is subject to conditional conservatism.

Equation (5) makes clear that the unexpected return in period t reflects the *capitalized* impact of current fundamental news, i.e., $f_{t+1} + g_{t+1} + (1+R^E)h_t$, on firm value. We refer to the fully capitalized value of fundamental news, i.e., $r_t = (R^E + 1 - w)^{-1}[f_{t+1} + g_{t+1} + (1+R^E)h_t]$, as returns news. Equations (2) and (5) imply that the regression coefficient of unexpected returns on unexpected earnings (also known as the earnings response coefficient) is proportional to the capitalization factor $(R^E + 1 - w)^{-1}$. Thus, as one would expect from the standard valuation theory (e.g., Kormendi and Lipe 1987; Easton and Zmijewski 1989; Collins and Kothari 1989), the relation between unexpected returns and unexpected earnings depends on firm characteristics such as expected returns R^E and the cash flow persistence parameter w.

AT estimates of conditional conservatism

Basu's (1997) AT measure of conditional conservatism is the incremental slope coefficient on negative returns in a piecewise linear regression of earnings on returns. To mitigate confounding effects due to the expected portions of earnings and returns, Ball et al. (2013a) recommend the use of unexpected earnings and unexpected returns. Following this suggestion, we consider the following regression model for estimating the AT coefficient β_1:

$$x_t = \alpha_0 + \alpha_1 I_t + \beta_0 r_t + \beta_1 I_t \cdot r_t + \varepsilon_t, \tag{6}$$

where I_t is an indicator variable taking the value of one when news is bad and zero otherwise. Since conditionally conservative accounting recognizes bad news sooner than it reflects good news, the AT coefficient β_1 is predicted to be positive and its magnitude is taken as a measure of the degree of conditional conservatism.

For our theoretical analysis, we focus on a scenario in which the researcher can directly observe the sign of fundamental news f_{t+1}. In regression model (6), therefore, the indicator variable I_t takes the value of one for $f_{t+1} < 0$ and zero for $f_{t+1} \geq 0$. For empirical applications, the sign of unexpected returns is used as a proxy for the sign of news subject to conditional conservatism because the researcher cannot directly observe the sign of such news. Nevertheless, the confounding effects we identify apply to actual empirical settings in which the Basu regression is conditioned on the sign of unexpected returns.

Let $\hat{\beta}_1$ denote the OLS estimate of β_1 from equation (6). It follows from the properties of OLS regression that

$$\operatorname{plim}\hat{\beta}_1 = \frac{Cov(x_t, r_t \mid f_{t+1} < 0)}{Var(r_t \mid f_{t+1} < 0)} - \frac{Cov(x_t, r_t \mid f_{t+1} \geq 0)}{Var(r_t \mid f_{t+1} \geq 0)}.$$

To avoid cumbersome notation, hereafter we simply use β_1 to denote $\operatorname{plim}\hat{\beta}_1$. Using expressions (2) and (5) for unexpected earnings x_t and unexpected returns r_t, respectively, we show in Appendix 1 that

$$\beta_1 = (R^E + 1 - w) \cdot \left[\frac{\delta F_b + G + (1+R^E)H}{F_b + G + (1+R^E)^2 H} - \frac{G + (1+R^E)H}{F_g + G + (1+R^E)^2 H} \right], \tag{7}$$

where $F_b \equiv Var(f_{t+1} \mid f_{t+1} < 0)$ and $F_g \equiv Var(f_{t+1} \mid f_{t+1} \geq 0)$.

To investigate how asymmetry in the returns distribution affects the AT coefficient, suppose that the distribution of fundamental news f_{t+1} is asymmetric in the sense that $F_g = F_b + \lambda$, where $\lambda > 0$ measures the degree of distributional asymmetry.[9] It is straightforward to see from equation (5) that this form of asymmetry implies asymmetry in the returns distribution in the sense that $Var(r_t \mid f_{t+1} \geq 0) > Var(r_t \mid f_{t+1} < 0)$. Since the sign of unexpected return r_t is a noisy proxy for the sign of fundamental news f_{t+1}, one would expect that $Var(r_t \mid r_t \geq 0) > Var(r_t \mid r_t < 0)$ for most reasonable distributions. Indeed, Figure 2 provides evidence of asymmetry in the distribution of unexpected returns

[9] The distributions of fundamental news components g_{t+1} and h_t can be similarly asymmetric across positive and negative realizations. Our results hold regardless of whether there is such asymmetry in the distribution of these cash flow news components.

for our sample. The conditional variance of positive unexpected returns is almost six times the conditional variance of negative unexpected returns.[10]

We note that our analysis effectively assumes that asymmetry in the distribution of unexpected returns arises due to asymmetry in the distribution of underlying fundamentals. This assumption is consistent with the theory and evidence in Del Viva et al. (2015) who show that asymmetry in the returns distribution is at least partly due to asymmetry in the distribution of cash flow news, which, in turn, is induced by the firm's active management of its abandonment and expansion options. Nevertheless, our results do not depend on whether asymmetry in the returns distribution is caused by asymmetry in the distribution of fundamental news or some other reason (e.g., leverage, limited liability).

AT coefficient with symmetric accounting

We first consider first a hypothetical scenario when accounting is entirely symmetric, i.e., $\delta = 0$. Let β_1^s denote the corresponding AT coefficient in such a symmetric accounting setting. We obtain the following result:

Proposition 1: *The AT coefficient β_1^s is positive even if accounting is entirely symmetric (i.e., even if $\delta = 0$).*

Proof: All proofs are in Appendix 1.

We show in the proof of Proposition 1 that when $\delta = 0$, the AT coefficient is given by

$$\beta_1^s = (R^E + 1 - w) \cdot \frac{\lambda \cdot \left[G + H(1 + R^E) \right]}{\left[F_b + G + (1 + R^E)^2 H \right] \left[F_b + \lambda + G + (1 + R^E)^2 H \right]}. \tag{8}$$

Equation (8) shows that if the distribution of returns is asymmetric, i.e., when $\lambda > 0$, the AT coefficient will be positive even if there is no accounting asymmetry, i.e., even if $\delta = 0$.

As an immediate implication of Proposition 1, our analysis predicts a *positive* AT coefficient for the cash flow component of earnings as long as the variance of returns conditional

[10] We find a similar spread in the conditional variances of observed returns and market-adjusted returns. In additional analysis, we find that although a log transformation mitigates skewness in the returns distribution, it does not induce symmetry in the conditional variances of log returns.

on good news is higher than that conditional on bad news. This prediction holds despite the fact that cash flows are independent of accounting recognition rules and is consistent with empirical findings dating back to Basu's (1995) dissertation. It can also be easily checked from equation (8) that the AT coefficient for cash flows increases with asymmetry in the returns distribution. Collins et al. (2014) find that the positive AT coefficient for cash flows is higher for early life-cycle firms (i.e., small, young, high growth firms) relative to mature firms. In additional analysis, we find that early life-cycle firms have higher asymmetry in the returns distribution and, therefore, Collins' et al. (2014) evidence of variation in the AT coefficient for cash flows is consistent with our theoretical prediction.

AT coefficient with asymmetric accounting

Consider now the case when accounting is asymmetric, i.e., $\delta > 0$. Then, equation (7) yields

$$\beta_1 = \beta_1^s + B \cdot \delta, \tag{9}$$

where $B \equiv \dfrac{(R^E + 1 - w)F_b}{F_b + G + (1 + R^E)^2 H}$ and β_1^s is as given by (8).

Equation (9) highlights that empirical AT estimates are subject to two spurious effects as captured by β_1^s and *B*, respectively. Proposition 1 shows that $\beta_1^s > 0$ as long as the distribution of returns is asymmetric. For the remainder of the analysis, we assume that $F_b + G \geq 3(1 + R^E)^2 H$. This condition is sufficient to ensure that the AT coefficient increases in the firm's cost of capital R^E. To interpret this condition, note that $F_b + G + (1 + R^E)^2 H$ measures the total return volatility. Condition $F_b + G \geq 3(1 + R^E)^2 H$ requires that the return volatility is driven more by the variability in the expectations of future cash flows than by the variability in current cash flows. This condition is quite reasonable for most ongoing firms (e.g., Liu and Thomas 2000).

The proof of Proposition 2 shows that β_1^s increases in the expected return R^E and asymmetry in the returns distribution λ, and decreases in the cash flow persistence parameter *w*. In addition, *B* is always positive and increases in R^E, while it decreases in *w*. The following result summarizes our findings:

Proposition 2: *Holding δ constant, the AT coefficient β_1 (i) increases in the degree of asymmetry in the returns distribution λ, (ii) increases in the expected return R^E, and (iii) decreases in the cash flow persistence parameter w.*

Researchers are often interested in inferring whether and how financial reporting choices vary with a particular firm characteristic (e.g., firm size, book-to-market, leverage) from the observed relation between the AT coefficient and that characteristic. A key implication of the above result is that evidence of variation in the AT coefficient cannot be necessarily interpreted as evidence of variation in the degree of conditional conservatism.[11]

To see more explicitly how inferences can be confounded by variation in non-accounting factors, we differentiate equation (9) with respect to R^E to yield

$$\frac{d\beta_1}{dR^E} = \frac{d\beta_1^s}{dR^E} + \delta\frac{dB}{dR^E} + B\frac{d\delta}{dR^E}.$$

The above equation makes clear why it is problematic to infer the relation between δ and R^E from the empirical association between β_1 and R^E, since β_1 is also predicted to vary with R^E for reasons unrelated to the firms' degree of conditional conservatism. Therefore, variation in the empirical estimates of β_1 cannot be necessarily interpreted as evidence of variation in the degree of conditional conservatism because of the confounding sources of variation represented by the terms $d\beta_1^s / dR^E$ and dB / dR^E. A similar argument applies for the comparative statics with respect to the cash flow persistence parameter w and asymmetry in the returns distribution λ.

It should be noted that a way to control for the confounding effects of the expected return R^E and the cash flow persistence parameter w might be to use the ratio of the conditional slope coefficients for bad news and good news or the ratio of the conditional R^2's estimated separately

[11] Consider, for instance, prior evidence that AT coefficient estimates increase with book-to-market and leverage, and decrease with firm size (e.g., Pae et al. 2005; Roychowdhury and Watts 2007; Khan and Watts 2009; Ball et al. 2013a; Lawrence et al. 2013; Collins et al. 2014). In additional analysis, we find that book-to-market and leverage are positively related to expected returns and asymmetry in the returns distribution, and are negatively related to cash flow persistence. We also find that firm size is negatively related to expected returns and asymmetry in the returns distribution, and is positively related to cash flow persistence. Given these findings, our theory predicts that AT coefficients will increase with book-to-market and leverage and decrease with firm size even in the absence of any direct association between the degree of conditional conservatism and these firm characteristics.

for bad news and good news subsamples.[12] This is because the term $\left(R^E+1-w\right)$ cancels out when taking these ratios. These alternative measures of conditional conservatism, however, remain upwardly biased due to asymmetry in the returns distribution. To see this, it can be easily checked from our analytical model that even if accounting is entirely symmetric (i.e., δ= 0), the ratio of the conditional slope coefficients as well as the ratio of the conditional R^2's is equal to the ratio of the conditional variances of unexpected returns, i.e., $Var(r_t \mid f_{t+1} \geq 0)/Var(r_t \mid f_{t+1} < 0)$, which will exceed one due to asymmetry in the returns distribution.

AT coefficient for accruals verus AT coefficient for unexpected earnings

In a recent paper, Collins et al. (2014) argue for the use of accruals-based asymmetric timeliness measures because the cash flow component of earnings is not subject to conditional conservatism. To investigate whether accruals-based estimates of conditional conservatism can eliminate the confounding effects identified above, consider the following regression model:

$$ac_t = \alpha_0^a + \alpha_1^a I_t + \beta_0^a r_t + \beta_1^a I_t \times r_t + \varepsilon_t,$$

where I_t is an indicator variable that takes the value of one when economic news is bad (i.e., $f_{t+1} < 0$) and zero otherwise, and β_1^a is the AT coefficient for accruals. Using expressions (3) and (5) for ac_t and r_t, respectively, we get

$$\beta_1^a = k \cdot \beta_1^s + B \cdot \delta, \tag{10}$$

where $k \equiv G/[G+(1+R^E)^2 H]$ is a constant less than one. A comparison with equation (9) reveals that the accounting component of the AT coefficient for accruals is the same as the corresponding component of the AT coefficient for unexpected earnings, but the spurious effect due to asymmetry in the distribution of economic news is smaller for accruals because k < 1. Importantly, equation (10) makes clear that replacing unexpected earnings with the accrual component of earnings as the dependent variable in the Basu regression does *not* eliminate the confounding effects identified by our theory. More specifically, in the presence of asymmetry in the distribution of fundamental

[12] These alternative measures of conditional conservatism were used by Pope and Walker (1999) to draw cross-country comparisons but were criticized by Basu (1999) on the basis of their asymptotic properties.

news, the accruals-based estimate of asymmetric timeliness β_1^a will also (i) be positive *even* in the absence of conditional conservatism, and (ii) increase in expected returns R^E and asymmetry in the distribution of economic news λ, and decrease in the persistence of economic news w *even* if conditional conservatism is held constant.

Overall, our analysis suggests that there is a need for identifying an alternative measure of conditional conservatism which, unlike the AT coefficient, is independent of the market capitalization factor, while it remains unaffected by asymmetry in the returns distribution.[13]

An alternative measure of conditional conservatism

Under conditional conservatism, fundamental news is likely to be recognized in the accrual component of earnings to a greater extent when the news is bad than when the news is good. Therefore, in the presence of conditional conservatism one would expect that accruals should be more variable for bad news relative to good news partitions. This intuition opens the possibility of using the spread in the conditional variances of accruals as a measure of conditional conservatism in corporate financial reporting.[14] To investigate this possibility, let

$$scv(z_t) = Var(z_t \mid f_{t+1} < 0) - Var(z_t \mid f_{t+1} \geq 0),$$

denote the spread in the variances of variable z_t in period t conditional on bad news and good news in that period. We note from equation (3) that scaled accruals in period t are given by

$$ac_t = d_t \cdot f_{t+1} - d_{t-1} \cdot f_t + g_{t+1} - g_t + \mu_t - \mu_{t-1},$$

where $d_t = \delta$ when news is bad (i.e., $f_{t+1} < 0$), and $d_t = 0$ when news is good (i.e., $f_{t+1} \geq 0$).

---

[13] Easton et al. (2016) argue for an enterprise-level approach to estimating the AT coefficient based on a piecewise linear regression of enterprise-level earnings on positive and negative enterprise-level returns. Our model abstracts away from variation in the firm's capital stricture and adopts the commonly used equity-level approach to estimating the AT coefficient. Our theoretical analysis, however, extends to the enterprise level in a straightforward manner. The AT coefficient at the enterprise level is predicted to increase in asymmetry in the distribution of enterprise returns, increase in the weighted-average cost of capital (WACC), and decrease in enterprise free cash flow persistence. Our alternative measure of conditional conservatism is effectively derived at the enterprise level since it focuses on asymmetry in the distribution of operating accruals, while financing accruals (i.e., changes in short-term and long-term debt) should be void of conditional conservatism.

[14] Following this intuition, an alternative measure of conditional conservatism might be the spread across news partitions of the conditional variance of accruals divided by the conditional variance of returns (e.g., Basu 1995). Such a measure, however, would remain upwardly biased due to asymmetry in the returns distribution.

Consistent with the intuition discussed above, this expression shows that accruals are subject to a greater fraction of the variability associated with the contemporaneous fundamental news f_{t+1} when news is good than when news is bad. Specifically, the spread in the conditional variances of bad news and good news accruals is given by the following expression (see the proof of Proposition 3 for details):

$$scv(ac_t) = \delta^2 \cdot F_b. \tag{11}$$

As a measure of conditional conservatism, $scv(ac_t)$ has a number of desirable properties. Specifically, the spread between the conditional variances of bad news and good news accruals (i) it is positive if, and only if, accounting is conditionally conservative (i.e., $\delta > 0$), (ii) it increases in the degree of conditional conservatism as measured by δ, (iii) it is independent of the market capitalization factor $(1+R^E\text{-}w)^{-1}$, and (iv) it is unaffected by asymmetry in the returns distribution.

In addition, our alternative measure of conditional conservatism is primarily based on accounting data and does not rely on the notion of market efficiency beyond that the sign of unexpected returns is a reasonable proxy for the sign of fundamental news. Even the need for relying on the sign of unexpected returns, however, could be done away with by using accounting-based partitioning variables such as the sign of earnings or cash flows. In contrast, the AT regression approach to measure conditional conservatism explicitly relies on the assumption that the market incorporates all information about the firm's future cash flows in a timely and efficient manner, regardless of whether or not such information is yet reflected in accounting earnings; i.e., stock market values lead accounting earnings (e.g., Basu 1997; Pope and Walker 1999; Ball et. al. 2013b).

Admittedly, using the sign of return news offers only a noisy proxy for the (otherwise unobservable) sign of fundamental news subject to conditional conservatism. Noise in our setting would result in the misclassification of some bad news firm-years as good news firm-years and vice versa. Such misclassification, however, is unlikely to be systematic and therefore would only attenuate the spread in the conditional variances of accruals towards zero. Clearly, if fundamental news subject to conditionally conservatism was directly observable there would be no need for a separate measure of conditional conservatism to begin with; researchers would be able to directly identify asymmetry in accounting by comparing the fraction of positive and negative news recognized in earnings.

We note that the spread between the conditional variances of bad news and good news accruals captures the combined effect of δ, which measures the degree to which a *given* piece of bad news is recognized in earnings, and F_b, which can be thought of as the *amount* of news subject to conditionally conservative recognition. As a concrete example, consider a firm's decision to write-down inventories in response to an unfavorable shock to their values. The firm can be more conservative if it either (i) reports a *greater* fraction of its inventories based on the conditionally conservative rule, or (ii) writes down a given fraction of its inventories to *lower* values. Since it seems neither plausible nor desirable to distinguish between these two modes of conservatism, we believe that $scv(ac_t)$ provides a viable measure for the *overall* degree of conditional conservatism.

We hasten to note that the spread in the variances of *unscaled* accruals, $scv(AC_t)$, cannot be used as a measure of conditional conservatism because such a measure would be confounded by cross-sectional differences in the scale of operations. Specifically, all else equal, conditional as well as unconditional variances of accruals would be expected to be larger for bigger firms. Consistent with this intuition, we find that the unconditional variance of accruals increases in the beginning-of-period market value of equity (see Figure 1).

As an example of a differential prediction based on our alternative measure of conditional conservatism, we recall our prediction that the Basu regression will yield a spuriously *positive* AT coefficient for the cash flow component of earnings (see Proposition 1). In contrast, our theory predicts that regardless of the degree of conditional conservatism, the conditional variances of cash flows, $cf_t = CF_t/P_{t-1}$, will be symmetric across bad news and good news partitions. We summarize these findings as follows:

Proposition 3: *If accounting is conditionally conservative, the spread between the conditional variances of bad news and good news accruals is positive (i.e., $scv(ac_t) > 0$). On the other hand, the spread in the conditional variances of cash flows is zero (i.e., $scv(cf_t) = 0$) for all δ.*

Although we model asymmetry in the distribution of returns based on asymmetry in the distribution of the fundamental news component subject to conditional conservatism f_{t+1}, we could introduce asymmetry in the distribution of other components of fundamental news. For example, in the presence of real options such as expansion and contraction options, one would also expect asymmetry in the distribution of fundamental news component that is unconditionally recognized in earnings concurrently with its realization as cash flows (h_t). i.e., $Var(h_t \mid h_t \geq 0) > Var(h_t \mid h_t < 0)$

. Such asymmetry would imply asymmetry in the distribution of cash flows, but in the *opposite* direction from asymmetry in the distribution accruals due to conditional conservatism. Put differently, one would expect that $Var(cf_t \mid h_t \geq 0) > Var(cf_t \mid h_t < 0)$, and therefore $Var(cf_t \mid r_t \geq 0) > Var(cf_t \mid r_t < 0)$.

III. EMPIRICAL ANALYSIS

In our theoretical analysis, we show that the AT coefficient depends not only on the degree of conditional conservatism but also on non-accounting factors, including expected returns, cash flow persistence, and asymmetry in the returns distribution. We identify the spread between the variances of bad news and good news accruals as an alternative measure of conditional conservatism that is free of the effects confounding the AT coefficient. In this section, we empirically test our predictions with respect to variation in the AT coefficient with the non-accounting factors identified by our theory. We then proceed to validate our alternative measure of conditional conservatism.

Sample and descriptive statistics

We obtain accounting data from the Compustat fundamental annual file and stock return data from the CRSP monthly returns file. We measure earnings as income before extraordinary items. We decompose earnings into accruals and cash flows using the cash flow statement approach, which becomes feasible post-1988. We measure free cash flows as cash flows from operating activities minus capital expenditures. We measure accruals as earnings minus free cash flows.[15] We measure "conditionally conservative accruals" using negative Compustat special items and set missing and positive values to zero. We scale earnings, accruals, and free cash flows using the beginning of year market value of equity.

Following Ball et al. (2013a), we measure unexpected earnings (x_{it}) using the residual values from the following autoregressive model estimated by two-digit SIC industry *j* and year *t*:

$$y_{it} = \gamma_{0jt} + \gamma_{1jt} I(y_{it-1} < 0) + \gamma_{2jt} y_{it-1} + \gamma_{3jt} I(y_{it-1} < 0) \cdot y_{it-1} + \varepsilon_{it},$$

[15] Our results are not sensitive when we measure accruals as the difference of earnings and cash flows from operating activities as in Collins et al. (2014) or as the difference of earnings and cash flows from operating and investing activities as in Dechow and Ge (2006).

where y_{it} is current earnings scaled by the beginning of year market value of equity, y_{it-1} is lagged earnings scaled by the beginning of year market value of equity, and $I(y_{it-1} < 0)$ is an indicator variable for negative lagged earnings.

We measure fiscal year stock returns as the buy-and-hold stock return (including distributions) cumulated over the twelve months leading to the fiscal year-end. Our results are not sensitive when we use inter-announcement stock returns cumulated over the twelve months leading to three months after the fiscal year-end. Following Ball et al. (2013a), we measure expected returns using the average returns on 5×5 portfolios constructed each year by sorting firms first based on the beginning of year market value of equity and then based on the beginning of year book-to-market ratio. We measure unexpected returns (r_{it}) as the difference of observed stock returns and expected stock returns.

To reduce the effects of outliers, we trim firm-year observations falling in the top and bottom one percent of the annual cross-sectional distributions of the current and lagged values of scaled accounting data as well as stock returns.[16] Following Ball et al. (2013a), we require a minimum of ten observations per two-digit SIC industry-year. Following Collins et al. (2014), we exclude financial institutions and utilities as well as firms with negative beginning of year book value of equity.

Our sample includes 88,852 firm-year observations over the 26-year period from 1989 to 2014. Appendix 2 describes the variables in detail. Table 1 reports descriptive statistics. The empirical distributions of the variables are similar to those reported in prior research. The average values of the unexpected component of earnings and returns are zero, while the average value of the indicator variable for negative unexpected returns shows that 57.7 percent of the sample is classified as bad news. Special items are strongly positively correlated with earnings and total accruals but are only weakly correlated with free cash flows, which is consistent with the fact that special items are mostly non-cash accrual adjustments (e.g., Dechow and Ge 2006).

[16] Ohlson and Kim (2015) propose the Theil-Sen estimator as an alternative to the OLS estimator that mitigates the unduly effects of outliers. In our empirical analysis, we trim extreme observations to mitigate the effect of outliers.

AT coefficient estimates: Replication of prior evidence

Table 2 reports results from annual cross-sectional AT regression specifications of earnings and earnings components:

$$z_{it} = \alpha_{0t} + \alpha_{1t} I(r_{it} < 0) + \beta_{0t} r_{it} + \beta_{1t} I(r_{it} < 0) \cdot r_{it} + \varepsilon_{it},$$

where z_{it} is reported earnings or earnings components, including expected and unexpected earnings as well as free cash flows and total accruals, r_{it} is unexpected returns, and $I(r_{it}<0)$ is the indicator variable for negative unexpected returns.[17]

Consistent with evidence dating back to Basu (1997), we find that reported earnings are more strongly related with negative unexpected returns (i.e., bad news) relative to positive unexpected returns (i.e., good news). The AT coefficient for total earnings is 0.314. In line with prior research, we also find spurious evidence of asymmetric timeliness in expected earnings and cash flows.[18] The AT coefficients for expected earnings and unexpected earnings are 0.166 and 0.148, respectively, while the AT coefficients for cash flows and accruals are 0.146 and 0.169, respectively. By virtue of OLS regression properties, the sum of the AT coefficients for earnings components are equal to the AT coefficient for total earnings. Therefore, comparing the magnitude of the coefficient estimates shows that approximately 53 percent and 46 percent of evidence of asymmetric timeliness in total earnings is attributed to spurious evidence of asymmetry in expected earnings and cash flows, respectively.

AT coefficient estimates: Variation with non-accounting factors

Our theory predicts that the AT coefficient will vary not only with the degree of conditional conservatism but also with non-accounting factors. Next, we search for variation in AT coefficient estimates after we remove spurious evidence of asymmetry due to either the expected component of earnings, as recommended by Ball et al. (2013a), or the cash flow component of earnings, as recommended by Collins et al. (2014). We refer to the AT regression specification for unexpected

[17] Basu (1997) and Collins et al. (2014) report results using market-adjusted stock returns and observed stock returns. Following Ball et al. (2013a), we use unexpected returns when estimating the AT regressions. Our results are not sensitive to the choice of total, market-adjusted, or unexpected returns.

[18] Spurious evidence of asymmetric timeliness due to the expected component of earnings is traced back to Patatoukas and Thomas' (2011) finding of a positive AT coefficient for lagged earnings, while spurious evidence of asymmetric timeliness due to the cash flow component of earnings was first reported by Basu (1995).

earnings as the Ball et al. or BKN model, and to the AT regression specification for accruals as the Collins et al. or CHT model.

Measurement of non-accounting factors

Consistent with Ball et al. (2013a), we measure expected returns (R^E) as the average returns on 5×5 portfolios formed at the beginning of each year by sorting firms first based on market value of equity and then based on book-to-market. The return expectation model follows Fama and French (1992). We measure unexpected returns or "return news" as the difference between observed returns and expected returns. Across the same 5×5 portfolios, we calculate the conditional variances of unexpected returns separately for "good news" and "bad news" partitions. We then measure asymmetry in the returns distribution (λ) as the difference between the conditional variance of positive unexpected returns and the conditional variance of negative unexpected returns, i.e., $Var(r_{it} \mid r_{it} \geq 0) - Var(r_{it} \mid r_{it} < 0)$. Finally, we measure cash flow persistence (w) as the slope coefficient from a first-order autoregressive model of free cash flows estimated by two-digit SIC industry-year.[19] For brevity, we use the same notation to denote the empirical measure of each non-accounting factor as used to denote the corresponding theoretical construct.

Evidence of variation in AT coefficient estimates with non-accounting factors

Table 3 reports time-series average values of AT coefficient estimates across quintile portfolios of (i) expected returns in Panel A, (ii) asymmetry in the conditional variances of positive and negative unexpected returns in Panel B, and (iii) cash flow persistence in Panel C. First, we observe that the non-accounting factors identified by our theory exhibit significant variation across quintile portfolios.[20] Importantly, consistent with our theoretical predictions, we find that AT

[19] The portfolio-level estimation allow us to measure all three non-accounting factors for our full sample without imposing additional data requirements such as a minimum number of annual observations per firm. The mean (median) number of annual observations per firm in our sample is 8.1 (6), and the maximum (minimum) is 26 (1). Thus, firm-level estimation would likely be noisy and unreliable.

[20] Across quintile portfolios, our measure of expected returns ranges from -2.4 percent to 27 percent with a median value of 11.6 percent, the spread between the conditional variances of positive and negative unexpected returns ranges from 0.092 to 0.435 with a median value of 0.254, and cash flow persistence ranges from 0.108 to 0.727 with a median value of 0.46. We note that the measure of expected returns is based on realized returns, and therefore negative values can be attributed to the ex post return expectation model. In our analysis, we are interested in variation in the level of expected returns, rather than the magnitude of expected returns.

coefficient estimates increase with expected returns and asymmetry in the returns distribution, while they decrease with cash flow persistence.

Focusing on the BKN model, we find that AT coefficient estimates for unexpected earnings (i) increase from 0.094 to 0.176 across portfolios of expected returns, (ii) increase from 0.052 to 0.176 across portfolios of asymmetry in the returns distribution, and (iii) decrease from 0.163 to 0.104 across portfolios of cash flow persistence. Turning to the CHT model, we find that AT coefficient estimates for accruals (i) increase from 0.100 to 0.191 across portfolios of expected returns, (ii) increase from 0.057 to 0.173 across portfolios of asymmetry in the returns distribution, and (iii) decrease from 0.215 to 0.117 across portfolios of cash flow persistence. The differences in AT coefficient estimates across the top and bottom portfolios are statistically significant for both the BKN and the CHT model.

It should be noted that our evidence of a positive association between expected returns and AT coefficient estimates runs contrary to predictions in prior research of a negative relation between the degree of conditional conservatism and expected returns (e.g., Guay and Verrechia 2006; Suijs 2008; García Lara et al. 2011; Li 2015).[21] We acknowledge that even if it is really the case that the "true" degree of conditional conservatism and expected returns are negatively related, the empirical association between AT coefficient estimates and expected returns will depend on the strength of this negative relation relative to the opposite confounding effects identified by our theory. Therefore, one interpretation of our evidence is that these opposite confounding effects prevail in our sample.

AT coefficient estimates: Controlling for variation with non-accounting factors

The analysis so far provides evidence that AT coefficient estimates vary predictably with non-accounting factors, which implies that the Basu regression is misspecified due to omitted correlated variables. Next, we expand the right-hand-side of the Basu regression by including

[21] In a model with strategic reporting behavior on the part of the manager, Guay and Verrechia (2006) show that conditional conservatism results in lower information uncertainty and lower discount rates. Similarly, Suijs (2008) finds that conditional conservatism improves risk sharing across overlapping generations of investors, and hence results in lower discount rates. Consistent with these predictions García Lara et al. (2011) find evidence of a negative association between expected returns and Callen's et al. (2010) alternative measure of asymmetric timeliness, while Li (2015) extends the evidence at the international cross-country setting.

expected returns (R^E), asymmetry in the returns distribution (λ), and cash flow persistence (w), along with their interactions with positive and negative unexpected returns, as follows:

$$\begin{aligned} z_{it} = {} & \alpha_{0t} + \alpha_{1t} I(r_{it} < 0) + \alpha_{2t} R_{it}^E + \alpha_{3t}\lambda_{it} + \alpha_{4t} w_{it} \\ & + \alpha_{5t} R_{it}^E \cdot I(r_{it} < 0) + \alpha_{6t}\lambda_{it} \cdot I(r_{it} < 0) + \alpha_{7t} w_{it} \cdot I(r_{it} < 0) \\ & + \beta_{01t} r_{it} + \beta_{02t} R_{it}^E \cdot r_{it} + \beta_{03t}\lambda_{it} \cdot r_{it} + \beta_{04t} w_{it} \cdot r_{it} \\ & + \beta_{11t} I(r_{it} < 0) \cdot r_{it} + \beta_{12t} R_{it}^E \cdot I(r_{it} < 0) \cdot r_{it} + \beta_{13t}\lambda_{it} \cdot I(r_{it} < 0) \cdot r_{it} + \beta_{14t} w_{it} \cdot I(r_{it} < 0) \cdot r_{it} + \varepsilon_{it}, \end{aligned}$$

where z_{it} is either unexpected earnings or accruals. The slope coefficient β_{11} captures asymmetric timeliness in the data after controlling for variation with non-accounting factors.

Table 4 reports results from annual cross-sectional regressions based on the model described above. To ease the interpretation of the coefficient estimates, we replace the raw values of R^E, λ, and w, with the corresponding annual quintile ranks, *Rank* (·), scaled to vary from zero (lowest quintile) to one (highest quintile). A comparison of the coefficient estimates in Table 4 with those in Table 2 shows that evidence of asymmetric timeliness in accounting data becomes substantially weakened (in terms of both magnitude and statistical significance) after controlling for variation with non-accounting factors. The estimated coefficients on the triple-interactions imply that all three non-accounting factors identified are incrementally relevant for explaining evidence of asymmetric timeliness in unexpected earnings and accruals. Consistent with the portfolio results in Table 3, the multiple regression results in Table 4 show that AT coefficient estimates increase with expected returns and asymmetry in the returns distribution, while they decrease with cash flow persistence.

To be clear, our evidence of substantially weakened AT coefficient estimates in Table 4 cannot be interpreted as suggestive of the absence of conditional conservatism in accounting data. The reason is that the "true" degree of conditional conservatism may itself vary with expected returns, asymmetry in the returns distribution, and cash flow persistence. If this is the case, however, our evidence suggests that there only limited residual variation in the true degree of conditional conservatism that is orthogonal to the three factors identified.

Towards an alternative measure of conditional conservatism

A key implication of conditional conservatism is asymmetry in the distribution of the accrual component of reported earnings. Our theoretical analysis predicts that in the presence of conditional conservatism, the variance of accruals will be higher for bad news relative to good news partitions. Consistent with this prediction, Table 5 provides evidence of asymmetry in the distribution of accruals across partitions based on the sign of contemporaneous unexpected returns. The variance of bad news accruals is 1.35 times the variance of good news accruals and the spread in the conditional variances of accruals is significantly positive, which is consistent with our theoretical prediction about the effect of conditional conservatism.[22] While evidence of asymmetry extends to the distribution of total earnings, we argue that focusing on asymmetry in the distribution of the accrual component of earnings provides a "cleaner" path towards identifying the effect of conditional conservatism in accounting data. This is because the cash flow component of earnings should be free of the effect of conditional conservatism.

Construct validity tests using placebo test variables

Our evidence of a significantly positive spread between the conditional variances of bad news and good news accruals is reassuring given that conservatism is ingrained in GAAP accounting practices. Next, we introduce construct validity tests using placebo test variables that should be free of the effect of conditional conservatism.

For our first test, we start with the observation that while asymmetry in accounting due to conditional conservatism implies that the variance of bad news accruals is higher than the variance of good news accruals, conditional conservatism does not imply asymmetry in the distribution of cash flows. In the presence of real options such as expansion and contraction options, however, one would expect asymmetry in the conditional variances of cash flows but in the *opposite* direction from that of asymmetry in the conditional variances of accruals due to conditional conservatism. While asymmetrically timely loss recognition implies a *positive* spread between the conditional variances of bad news and good news accruals, real options would imply a *negative* spread between the conditional variances of bad news and good news cash flows.

For our second test, we search for asymmetry in lagged accounting data reported in year *t*-1, including (i) the accrual component of lagged earnings, (ii) the cash flow component of lagged earnings, as well as (iii) lagged earnings, across partitions based on the sign of unexpected returns in year *t*. The idea is simple: *lagged* accounting data cannot be conditionally conservative with

[22] Levene's (1960) test for unequal variances yields an F-statistic of 80.73, which indicates that the variances of bad news and good news total accruals are significantly different from each other at the one percent level.

respect to *future* news.[23] Therefore, construct validity would imply no evidence of asymmetry in the conditional variances of lagged accounting data, across partitions based on the sign of future unexpected returns.

Our measure of conditional conservatism is based on asymmetry in the conditional variances of accruals scaled by the beginning of year market value of equity. As an additional construct validity test, we scrub accounting off the variable of interest and focus on asymmetry in the conditional variances of the inverse of lagged market value of equity, denoted $1/P_{it\text{-}1}$. Again, construct validity would imply no evidence of asymmetry in the distribution of the placebo test variable $1/P_{it\text{-}1}$, across partitions based on the sign of future unexpected returns.

Turning to the empirical results, Table 5 reports evidence of asymmetry in the distribution of cash flows but in the opposite direction from that of asymmetry in the distribution of accruals. Although we find a *positive* spread between the conditional variances of bad news and good news accruals, which is consistent with a key implication of asymmetrically timely loss recognition in corporate financial reporting, we find a *negative* spread between the conditional variances of bad news and good news cash flows, which cannot be attributed to conditional conservatism but is consistent with the effect of real options. Specifically, the variance of bad news cash flows is only 0.83 times the variance of good news cash flows.[24]

Focusing on our set of lagged placebo test variables, Panel A of Table 6 reports evidence consistent with construct validity. Specifically, we find no evidence of asymmetry in the conditional variances of lagged earnings components, including lagged accruals and cash flows, as well as lagged total earnings, which is consistent with the idea that accounting data cannot be conditionally conservative with respect to *future* news. In addition, we find no evidence of asymmetry in the conditional variances of $1/P_{it\text{-}1}$, which implies that asymmetry becomes non-existent when accounting is being scrubbed off.

Even though our alternative measure of conditional conservatism passes a series of construct validity tests, we highlight that prior research finds spurious evidence of asymmetric

[23] Using expression (3), we observe that $Var(ac_{it\text{-}1}|f_{t+1}\geq 0) = Var(ac_{it\text{-}1}|f_{t+1}<0)$. This is because period t-1 accruals (i.e., $ac_{it\text{-}1}$) cannot depend on period t cash flow new news (i.e., f_{t+1}).

[24] Levene's (1960) test for unequal variances yields an F-statistic of 65.31, which indicates that the variances of bad news and good news cash flows are significantly different from each other at the one percent level.

timeliness for the same set of placebo test variables. Going all the way back to Basu's (1995) dissertation, several studies report spurious evidence of a positive AT coefficient for the cash flow component of earnings (e.g., Basu 1995, 1997; Ryan and Zarowin 2003; Dietrich et al. 2007; Hsu et al. 2012; Collins et al. 2014). Indeed, our evidence in Table 2 confirms that approximately 46 percent of evidence of asymmetric timeliness in earnings is attributed to spurious evidence of asymmetric timeliness in cash flows. Starting with Patatoukas and Thomas (2011), other studies report spurious evidence of asymmetric timeliness for lagged total earnings, lagged earnings components, and the inverse of lagged scale variables (e.g., Hsu et al. 2012; Ball et al. 2013a; Collins et al. 2014; Patatoukas and Thomas 2016).

Panel B of Table 6 replicates spurious evidence of asymmetric timeliness for lagged placebo test variables and reports positive AT coefficients for the accrual and cash flow components of lagged earnings, as well as for lagged earnings, which cannot be attributed to conditional conservatism. In addition, we find evidence of a significantly negative AT coefficient for $1/P_{it\text{-}1}$. Given that the placebo test variable $1/P_{it\text{-}1}$ excludes accounting information, evidence of asymmetry in the conditional relations of $1/P_{it\text{-}1}$ with positive and negative unexpected returns does not reflect the effect of conditional conservatism. The sharp divergence of the construct validity test results using placebo test variables highlights the relevance of our efforts to introduce and validate our alternative measure of conditional conservatism based on asymmetry in the distribution of accruals.

Construct validity tests using "conditionally conservative accruals"

In practice, asymmetrically timely loss recognition is ingrained in "conditionally conservative accruals" related to lower-of-cost-or-market accounting for inventories (ASC 330), goodwill impairments (ASC 350), and asset write-downs (ASC 360). Our theoretical analysis predicts that asymmetry in the conditional variances of total accruals should only be attributed to the subset of conditionally conservative accruals. To test this prediction, we obtain data on negative special items from the Compustat fundamental annual file. Compustat special items include non-cash, accrual adjustments through which conditional conservatism is primarily applied in practice, including inventory write-downs, long-term asset write-downs, and goodwill impairments, along with other unusual and nonrecurring items.[25]

Even though negative special items are bound to measure conditionally conservative accruals with noise, Panel A of Table 7 reports evidence consistent with our prediction that asymmetry in the distribution of total accruals should be traced back to the subset of accrual items subject to asymmetrically timely loss recognition. Indeed, we find that asymmetry in the conditional variances of total accruals is mostly attributed to asymmetry in the conditional variances of special items. While the variance of bad news special items is 2.43 times the variance of good news special items, we find no evidence of asymmetry in the conditional variances of accruals after excluding special items.[26]

In contrast, Panel B of Table 7 breaks down the AT coefficient for total accruals (0.169) and shows that as much as 57 percent of evidence of asymmetric timeliness is due to all other accruals excluding special items (divide the AT coefficient for accruals excluding special items by the AT coefficient for total accruals or 0.097/0.169). Evidence that only 43 percent of the AT coefficient for total accruals is attributed to conditionally conservative accruals embedded in negative special items is consistent with upward bias in AT estimates of conditional conservatism. Overall, our more granular tests using negative special items provide additional support for the construct validity of our alternative measure of conditional conservatism.

IV. CONCLUSION

Using a financial reporting and valuation model, we investigate the construct validity of the AT coefficient as a measure of conditional conservatism. Within the context of our model, we predict that the AT coefficient will be positive *even* in the absence of conditional conservatism and it will vary with non-accounting factors *even* if the degree of conditional conservatism is held constant. Our empirical analysis shows that AT coefficient estimates vary with non-accounting

[25] As a robustness check, we construct an alternative measure of conditionally conservative accruals using information about funds from operations-other (*FOPO*) from the statement of cash flows. *FOPO* includes asset write-downs along with items not subject to asymmetrically timely loss recognition, most notably excess tax benefits from stock-based compensation (*TXBCO*) and stock-based compensation expense (*STKCO*). Our alternative measure of conditionally conservative accruals, defined as - *FOPO* + *TXBCO* + *STKCO*, is 82 percent correlated with Compustat special items. Our inferences are unchanged using this alternative measure.

[26] Levene's (1960) test for unequal variances yields an F-statistic of 205.08, which indicates that the variances of bad news and good news special items are significantly different from each other at the one percent level.

factors in directions predicted by our theory. Indeed, AT coefficient estimates increase with expected returns and asymmetry in the returns distribution, and decrease with cash flow persistence. Our efforts to control for variation in the non-accounting factors identified, however, result in substantially weakened AT coefficient estimates. The evidence suggests that it may not be feasible to disentangle the confounding effects of non-accounting factors from the effect of conditional conservatism within the AT regression.

We derive and validate the spread between the variances of bad news and good news accruals as an alternative measure of conditional conservatism. Unlike the AT coefficient, our theoretical analysis shows that asymmetry in the distribution of accruals is not subject to the confounding effects of non-accounting factors such as expected returns, asymmetry in the returns distribution, and cash flow persistence. Consistent with a key implication of conditional conservatism, we find evidence that the variance of bad news accruals is significantly higher than the variance of good news accruals primarily due to conditionally conservative accruals related to inventory write-downs, long-term asset write-downs, and goodwill impairments. In addition, we do not find spurious evidence of asymmetry for placebo test variables that should be void of conditional conservatism.

Overall, we conclude that the spread in the conditional variances of accruals offers a viable measure of conditional conservatism. Although our alternative measure can be used to explore variation in conditional conservatism across firms our study is silent with respect to the cross-sectional determinants of financial reporting choices. Future research should attempt to simultaneously model (i) financial reporting choices as a function of firm characteristics, and (ii) stock market value as a function of financial reporting choices. We believe that the time is ripe to incorporate this two-way interplay between financial reporting choices and non-accounting factors for a better understanding of the properties of financial accounting data.

REFERENCES

Armstrong, C S., D. J. Taylor, and R. E. Verrecchia, 2015. Asymmetric reporting. *Journal of Financial Reporting* (forthcoming).

Ball, R., S. P. Kothari, and A. Robin, 2000. The effect of international institutional factors on properties of accounting earnings. *Journal of Accounting and Economics* 29 (1): 1-51.

Ball, R., and L. Shivakumar, 2006. The role of accruals in asymmetrically timely gain and loss recognition. *Journal of Accounting Research* 44 (2): 207-242.

Ball, R., S. P. Kothari, and V. A. Nikolaev, 2013a. On estimating conditional conservatism. *The Accounting Review* 88 (3): 755-787.

Ball, R., S. P. Kothari, and V. A. Nikolaev, 2013b. Econometrics of the Basu asymmetric timeliness coefficient and accounting conservatism. *Journal of Accounting Research* 51 (5): 1071-1097.

Banker, R. D., S. Basu, D. Byzalov, and J. Y. Chen, 2016. The confounding effect of cost stickiness on conservatism estimates. *Journal of Accounting and Economics* 61 (1): 203-220.

Basu, S., 1995. Conservatism and asymmetric timeliness of earnings. Ph.D. Dissertation, University of Rochester.

Basu, S., 1997. The conservatism principle and the asymmetric timeliness of earnings. *Journal of Accounting and Economics* 24 (1): 3-37.

Basu, S., 1999. Discussion of international differences in the timeliness, conservatism, and classification of earnings. *Journal of Accounting Research* 37 (Supplement): 89-99.

Beaver, W. H., and S. G. Ryan, 2005. Conditional and unconditional conservatism: Concepts and modeling. *Review of Accounting Studies* 10 (2-3): 269-309.

Beaver, W. H. and S. G. Ryan, 2009. Mixed-attribute accounting, and the identification of conditional conservatism. SSRN Working Paper: http://dx.doi.org/10.2139/ssrn.1477528.

Bertomeu, J., A. Beyer, and D. J. Taylor, 2015. From casual to causal inference in accounting research: The need for theoretical foundations. SSRN Working Paper: http://dx.doi.org/10.2139/ssrn.2694105.

Callen, J. L., D. Segal, and O. K. Hope, 2010. The pricing of conservative accounting and the measurement of conservatism at the firm-year level. *Review of Accounting Studies* 15 (1): 145-178.

Collins, D. W., and S. P. Kothari, 1989. An analysis of intertemporal and cross-sectional determinants of earnings response coefficients. *Journal of Accounting and Economics* 11 (2): 143-181.

Collins, D. W., P. Hribar, and X. S. Tian, 2014. Cash flow asymmetry: Causes and implications for conditional conservatism research. *Journal of Accounting and Economics* 58 (2): 173-200.

Dechow, P. M., and W. Ge, 2006. The persistence of earnings and cash flows and the role of special items: Implications for the accrual anomaly. *Review of Accounting Studies* 11 (2-3): 253-296.

Del Viva, L., E. Kasanen, and L. Trigeorgis, 2015. Real options and idiosyncratic skewness. *Journal of Financial And Quantitative Analysis* (forthcoming).

Dietrich, D., K. Muller, and E. Riedl, 2007. Asymmetric timeliness tests of accounting conservatism. *Review of Accounting Studies* 12 (1): 95-124.

Easton, P. D., and M. E. Zmijewski, 1989. Cross-sectional variation in the stock market response to accounting earnings announcements. *Journal of Accounting and Economics* 11 (2): 117-141.

Easton, P. D., P. B. Vassallo, and E. H. Weisbrod, 2016. Accounting earnings, change in market value and cash flows. SSRN Working Paper: http://ssrn.com/abstract=2674588.

Fama, E. F., 1965. The behavior of stock-market prices. *Journal of Business* 38 (1): 34-105.

Fama, E. F., and K. R. French, 1992. The cross-section of expected stock returns. *Journal of Finance* 47 (2): 427-465.

Financial Accounting Standards Board (FASB), 2009a. *Inventory*. Accounting Standards Codification (ASC) 330. Norwalk, CT: FASB.

Financial Accounting Standards Board (FASB), 2009b. *Intangibles—Goodwill and Other*. Accounting Standards Codification (ASC) 350. Norwalk, CT: FASB.

Financial Accounting Standards Board (FASB), 2009c. *Property, Plant, and Equipment*. Accounting Standards Codification (ASC) 360. Norwalk, CT: FASB.

García Lara, J. M., B. G. Osma, and F. Penalva, 2011. Conditional conservatism and cost of capital. *Review of Accounting Studies* 16 (2): 247-271.

Givoly, D., C. Hayn, and A. Natarajan, 2007. Measuring reporting conservatism. *The Accounting Review* 82 (1): 65-106.

Guay, W. R., and R. E. Verrecchia, 2007. Conservative disclosure. SSRN Working Paper: http://dx.doi.org/10.2139/ssrn.995562.

Hayn, C., 1995. The information content of losses. *Journal of Accounting and Economics* 20 (2): 125-153.

Hsu, A., J. O' Hanlon, and K. Peasnell, 2012. The Basu measure as an indicator of conditional conservatism: Evidence from U.K. earnings components. *European Accounting Review* 21 (1): 87-113.

Khan, M., and R. L. Watts, 2009. Estimation and empirical properties of a firm-year measure of accounting conservatism. *Journal of Accounting and Economics* 48 (2): 132-150.

Ohlson, J. A. and S. Kim, 2015. On the conditional conservatism measure: A robust estimation approach. SSRN Working Paper: http://dx.doi.org/10.2139/ssrn.2644423.

Kormendi, R., and R. Lipe, 1987. Earnings innovations, earnings persistence, and stock returns. *Journal of Business* 60 (3): 323-345.

Lawrence, A., R. Sloan, and Y. Sun, 2013. Non-discretionary conservatism: Evidence and implications. *Journal of Accounting and Economics* 56 (2): 112-133.

Levene, H., 1960. Robust tests for equality of variances. In *Contributions to Probability and Statistics: Essays in Honor of Harold Hotelling*, edited by I. Olkin, 278-292. Stanford, CA: Stanford University.

Li, X., 2015. Accounting conservatism and cost of capital: International analysis. *Journal of Business Finance and Accounting* 42 (5-6): 555-582.

Pae, J., D. B. Thornton, and M. Welker, 2005. The link between earnings conservatism and the price-to-book ratio. *Contemporary Accounting Research* 22 (3): 693-717.

Patatoukas, P. N., and J. K. Thomas, 2011. More evidence of bias in differential timeliness estimates of conditional conservatism. *The Accounting Review* 86 (5): 1765-1793.

Patatoukas, P. N., and J. K. Thomas, 2016. Placebo tests of conditional conservatism. *The Accounting Review* 91 (2): 625-648.

Pope, P. F., and M. Walker, 1999. International differences in the timeliness, conservatism, and classification of earnings. *Journal of Accounting Research* 37 (Supplement): 53-87.

Roychowdhury, S., R. L. Watts, 2007. Asymmetric timeliness of earnings, market-to-book and conservatism in financial reporting. *Journal of Accounting and Economics* 44 (1): 2-31.

Ryan, S. G., and P. A. Zarowin, 2003. Why has the contemporaneous linear returns-earnings relation declined? *The Accounting Review* 78 (2): 523-553.

Suijs, J., 2008. On the value relevance of asymmetric financial reporting policies. *Journal of Accounting Research* 46 (5): 1297-1321.

APPENDIX 1
Derivations and proofs

Derivation of equation (4):

Substituting $E_t[CF_{t+\tau}] = w^\tau CF_t + (1-w^\tau)m + w^{\tau-1}S(f_{t+1}+g_{t+1})$ into the formula for the stock market value and collecting terms, we get

$$P_t = \left(CF_t - m + \frac{S(f_{t+1}+g_{t+1})}{w}\right)\sum_{\tau=1}^{\infty}\left(\frac{w}{1+R^E}\right)^\tau + \sum_{\tau=1}^{\infty}\frac{m}{(1+R^E)^\tau}.$$

Equation (4) follows after substituting $\sum_{\tau=1}^{\infty}\left(\frac{w}{1+R^E}\right)^\tau = \frac{w}{R^E+1-w}$ and $\sum_{\tau=1}^{\infty}\frac{1}{(1+R^E)^\tau} = \frac{1}{R^E}$ into the above expression and simplifying.

Derivation of equation (7):

Since f_{t+1}, g_{t+1}, and h_t are mutually independent, it follows that $Var(h_t \mid f_{t+1}) = Var(h_t)$ and $Var(g_{t+1} \mid f_{t+1}) = Var(g_{t+1})$. Using equation (5) for the unexpected return then yields

$$Var(r_t \mid f_{t+1} < 0) = k^2[F_b + G + (1+R^E)^2 H]$$

and

$$Var(r_t \mid f_{t+1} \geq 0) = k^2[F_g + G + (1+R^E)^2 H],$$

where $k \equiv (R^E+1-w)^{-1}$ denotes the capitalization factor, $F_b \equiv Var(f_{t+1} \mid f_{t+1} < 0)$, and $F_g \equiv Var(f_{t+1} \mid f_{t+1} \geq 0)$. Similarly, equations (2) and (5) imply that

$$Cov(r_t, x_t \mid f_{t+1} < 0) = k[\delta F_b + G + (1+R^E)H]$$

and

$$Cov(r_t, x_t \mid f_{t+1} \geq 0) = k[G + (1+R^E)H].$$

Equation (7) then follows by substituting the above variance and covariance expressions into the definition of β_1.

Proof of Proposition 1:

When accounting is symmetric, $\delta = 0$. Substituting this into equation (7) and simplifying yield the following expression for the AT coefficient:

$$\beta_1^s = (R^E + 1 - w) \cdot \frac{\lambda \cdot \left[G + H(1+R^E) \right]}{\left[F_b + G + (1+R^E)^2 H \right] \left[F_b + \lambda + G + (1+R^E)^2 H \right]},$$

which is positive as long as $\lambda > 0$.

Proof of Proposition 2:

For the general case when δ and λ are both non-zero, equation (7) can be simplified as

$$\beta_1 = \beta_1^s + B \cdot \delta,$$

where β_1^s is as given above in the proof of Proposition 1 and

$$B = \frac{(R^E + 1 - w) F_b}{F_b + G + (1+R^E)^2 H}.$$

To prove the result, we will show that (i) B increases in R^E and decreases in w, and (ii) β_1^s increases in R^E, decreases in w, and increases in λ. To show part (i), we note that w appears (with a negative sign) only in the numerator of the above expression for B, and hence $dB/dw < 0$. Differentiating B with respect to R^E yield

$$\frac{dB}{dR^E} = F_b \cdot \frac{F_b + G - (1+R^E)^2 H + 2w(1+R^E) H}{[F_b + G + (1+R^E)^2 H]^2}.$$

Since the denominator is positive, the sign of dB/dR^E depends on the sign of the numerator, which is positive since $F_b + G \geq (1+R^E)^2 H$ by assumption. It thus follows that $dB/dR^E > 0$.

To prove part (ii), it can be seen from equation (8) that $d\beta_1^s/dw < 0$. Differentiating (8) with respect to λ and collecting terms reveal

$$\frac{d\beta_1^s}{d\lambda} == \frac{\beta_1^s [F_b + G + \lambda + (1+R_E)^2 H}{\lambda [F_b + G + \lambda + (1+R_E)^2 H]} > 0.$$

Differentiating equation (8) with respect to R^E and simplifying yield

$$\frac{d\beta_1^s}{dR^E} = \lambda \cdot \frac{\Delta_G \cdot G + \Delta_H \cdot H}{\left(F_b + G + (1+R^E)^2 H\right)^2 \left(F_b + G + \lambda + (1+R^E)^2 H\right)^2},$$

where

$$\Delta_H \equiv [2(1+R^E) - w](F_b + G)(F_b + G + \lambda) + [3w - 2(1+R^E)](1+R^E)^3 H^2 + w(1+R^E)^2 H(2F_b + 2G + \lambda)$$

and

$$\Delta_G \equiv (F_b + G)(F_b + G + \lambda) + [4w - 3(1+R^E)](1+R^E)^3 H^2 + (1+R^E)H[2w - (1+R^E)](2F_b + 2G + \lambda).$$

Since $F_b + G > (1+R^E)^2 H$ by assumption, it follows that:

$$(F_b + G)(F_b + G + \lambda) > (1+R^E)^4 H^2,$$

$$2F_b + 2G + \lambda > 2(1+R^E)^2 H.$$

Substituting these into the expression for Δ_H and simplifying yield

$$\Delta_H \geq \left[(2R^E + w)(1+R^E) + 3w\right](1+R^E)^3 H^2 > 0.$$

For brevity, define $a \equiv F_b + G$ and $b \equiv (1+R^E)^2 H$. Since $0 < w < 1$, it follows from the above expression for Δ_G that

$$\Delta_G > a(a+\lambda) - 3b^2 - (2a+\lambda)b$$

$$= (a^2 - 3b^2 - 2ab) + (a-b)\lambda.$$

Since $b < a/3$ by assumption, we observe that $(a^2 - 3b^2 - 2ab)$ and $(a-b)$ are both positive, which implies that $\Delta_G > 0$. This proves that $d\beta_1^s / dR^E < 0$.

Proof of Proposition 3:

Since f_t, g_t, and μ_t are serially and mutually independent, equation (3) yields

$$Var(ac_t \mid f_{t+1} < 0) = \delta^2 F_b + Var(g_{t+1} - g_t + \mu_t - \mu_{t-1} - d_{t-1} f_t)$$

and

$$Var(ac_t \mid f_{t+1} \geq 0) = Var(g_{t+1} - g_t + \mu_t - \mu_{t-1} - d_{t-1} f_t).$$

It thus follows that $scv(ac_t) = \delta^2 F_b,$ which is positive as long as $\delta > 0$.

From the free cash flow process in equation (1) and the fact that $\phi_t = f_t + g_t + h_t$, it follows that CF_t and f_{t+1} (and hence cf_t and f_{t+1}) are uncorrelated. This implies that $Var(cf_t \mid f_{t+1}) = Var(cf_t)$, and therefore $scv(cf_t) = 0$.

APPENDIX 2
Variables definitions

Variable	Definition
y_{it}	Earnings in year *t* measured as income before extraordinary items scaled by the beginning of year market value of equity.
ac_{it}	Accruals in year *t* measured as income before extraordinary items minus free cash flows scaled by the beginning of year market value of equity.
cf_{it}	Free cash flows in year *t* measured as cash flows from operating activities minus capital expenditures scaled by the beginning of year market value of equity.
y_{it}^{E}	Expected earnings for year *t* scaled by the beginning of year market value of equity measured as the fitted values from the autoregressive model of Ball et al. (2013a) estimated by two-digit SIC industry and year.
x_{it}	Unexpected earnings for year *t* scaled by the beginning of year market value of equity measured as the residual values from the autoregressive model of Ball et al. (2013a) estimated by two-digit SIC industry and year.
si_{it}	Compustat special items in year *t* scaled by the beginning of year market value of equity. Missing and positive values are set to zero.
R_{it}	Fiscal year stock returns measured as the buy-and-hold stock return (including distributions) cumulated over the twelve-month window leading to the fiscal year-end.
R_{it}^{E}	Expected stock returns measured as the average returns of 5×5 portfolios formed each year by sorting firms first based on the beginning of year market value of equity and then based on the beginning of year book-to-market ratio.
r_{it}	Unexpected stock returns measured as the difference of observed stock returns and expected stock returns.
$I(r_{it} < 0)$	Indicator variable = 1 if $r_{it} < 0$; and = 0 otherwise.
$Var(r_{it})$	Unconditional variance of unexpected returns measured across 5×5 portfolios formed each year by sorting firms first based on the beginning of year market value of equity and then based on the beginning of year book-to-market ratio.
$Var(r_{it} \mid sign(r_{it}))$	Conditional variances of positive and negative unexpected returns measured across 5×5 portfolios formed each year by sorting firms first based on the beginning of year market value of equity and then based on the beginning of year book-to-market ratio.
λ_{it}	Spread in the conditional variances of unexpected returns measured as the difference of $Var(r_{it} \mid r_{it} \geq 0)$ and $Var(r_{it} \mid r_{it} < 0)$.
w_{it}	Cash flow persistence measured as the slope coefficient from a first-order autoregressive model of free cash flows estimated by two-digit SIC industry and year.

TABLE 1
Descriptive statistics

Panel A: Empirical distributions.

	Mean	Std. Dev.	Skew	Q_1	Median	Q_3
y_{it}	-0.021	0.200	-3.665	-0.050	0.035	0.072
ac_{it}	-0.013	0.204	-2.530	-0.049	0.001	0.052
cf_{it}	-0.008	0.173	-0.713	-0.068	0.014	0.068
si_{it}	-0.025	0.081	-6.812	-0.011	0.000	0.000
x_{it}	0.000	0.153	-2.847	-0.024	0.015	0.052
r_{it}	0.000	0.596	2.097	-0.359	-0.081	0.223
$I(r_{it} < 0)$	0.577	0.494	-0.313	0.000	1.000	1.000

Panel B: Pairwise correlations.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
(1) y_{it}		0.63	0.41	0.59	0.76	0.20	-0.20
(2) ac_{it}	0.42		-0.45	0.53	0.51	0.11	-0.09
(3) cf_{it}	0.48	-0.44		0.05	0.28	0.11	-0.12
(4) si_{it}	0.33	0.31	0.01		0.54	0.11	-0.10
(5) x_{it}	0.59	0.27	0.28	0.30		0.22	-0.19
(6) r_{it}	0.40	0.13	0.22	0.14	0.32		-0.70
(7) $I(r_{it} < 0)$	-0.31	-0.09	-0.18	-0.11	-0.26	-0.86	

Note: All pairwise correlations are significant at the one percent level using two-tailed tests.

This table reports pooled descriptive statistics for the following variables: earnings scaled by lagged market value of equity (y_{it}), accruals scaled by lagged market value of equity (ac_{it}), free cash flows scaled by lagged market value of equity (cf_{it}), special items scaled by lagged market value of equity (si_{it}), unexpected earnings scaled by lagged market value of equity (x_{it}), unexpected returns (r_{it}), and the indicator variable for negative unexpected returns ($I(r_{it}<0)$). Panel A reports the empirical distributions. Panel B reports Pearson (Spearman) correlations above (below) the main diagonal. The sample includes 88,852 firm-year observations from 1989 to 2014. Appendix 2 describes the variables in detail.

TABLE 2
AT estimates of conditional conservatism

	Dependent variable =				
	(1)	(2)	(3)	(4)	(5)
	y_{it}	y_{it}^{E}	x_{it}	cf_{it}	ac_{it}
Intercept	0.030***	0.007*	0.023***	0.021***	0.009
	8.34	*1.72*	*10.66*	*4.66*	*1.34*
$I(r_{it} < 0)$	0.020***	0.008***	0.012***	0.002	0.018***
	5.34	*4.98*	*4.19*	*0.85*	*4.92*
r_{it}	-0.006	-0.030***	0.025***	-0.003	-0.003
	-0.96	*-8.10*	*7.17*	*-0.54*	*-0.59*
$I(r_{it} < 0) \times r_{it}$	0.314***	0.166***	0.148***	0.146***	0.169***
	16.36	*14.31*	*14.16*	*16.56*	*8.40*
Adj. R^2	13.33%	6.18%	8.87%	4.76%	3.59%

This table reports time-series average values of coefficient estimates from annual cross-sectional AT regressions of earnings scaled by lagged market value of equity (y_{it}), expected earnings scaled by lagged market value of equity (y_{it}^{E}), unexpected earnings scaled by lagged market value of equity (x_{it}), free cash flows scaled by lagged market value of equity (cf_{it}), and accruals scaled by lagged market value of equity (ac_{it}). Fama-MacBeth *t*-statistics are reported in italics below the coefficient estimates. *** and * indicate statistical significance at the one and ten percent level, respectively, using two-tailed tests. The sample includes 88,852 firm-year observations from 1989 to 2014.

TABLE 3
AT estimates of conditional conservatism:
Variation with non-accounting factors

Panel A: Variation with expected returns (R^E).

Quintiles of R^E	R^E	AT estimates	
		BKN model	CHT model
Q_1 (Low)	-0.024	0.094***	0.100***
		5.13	*4.73*
Q_2	0.058	0.114***	0.126***
		6.36	*5.57*
Q_3	0.116	0.142***	0.148***
		7.35	*5.12*
Q_4	0.169	0.158***	0.154***
		10.00	*6.58*
Q_5 (High)	0.270	0.176***	0.191***
		8.61	*6.08*
$Q_5 - Q_1$		0.082***	0.091**
		2.82	*2.22*

Panel B: Variation with asymmetry in the returns distribution (λ).

Quintiles of λ	λ	AT estimates	
		BKN model	CHT model
Q_1 (Low)	0.092	0.052***	0.057***
		5.35	*3.83*
Q_2	0.158	0.106***	0.109***
		7.11	*6.60*
Q_3	0.254	0.129***	0.145***
		9.94	*7.15*
Q_4	0.343	0.164***	0.190***
		10.62	*7.54*
Q_5 (High)	0.435	0.176***	0.173***
		11.51	*7.62*
$Q_5 - Q_1$		0.124***	0.115***
		9.85	*6.49*

Panel C: Variation with cash flow persistence (*w*).

Quintiles of w	w	AT estimates	
		BKN model	CHT model
Q_1 (Low)	0.108	0.163***	0.215***
		12.22	*8.07*
Q_2	0.356	0.183***	0.177***
		11.57	*7.58*
Q_3	0.460	0.146***	0.161***
		9.95	*6.11*
Q_4	0.556	0.146***	0.162***
		10.34	*7.38*
Q_5 (High)	0.727	0.104***	0.117***
		11.33	*6.75*
$Q_5 - Q_1$		-0.059***	-0.099***
		-5.06	*-5.27*

Panels A, B, and C report time-series average values of coefficient estimates from annual cross-sectional AT regressions based on Ball's et al. (2013a) or BKN model of unexpected earnings scaled by lagged market value of equity (x_{it}) on positive and negative unexpected returns and Collins' et al. (2014) or CHT model of accruals scaled by lagged market value of equity (ac_{it}) on positive and negative unexpected returns, across quintile portfolios formed independently based on (i) expected returns (R^E), (ii) the spread in the conditional variances of unexpected returns (λ), and (iii) cash flow persistence (w), respectively. Fama-MacBeth *t*-statistics are reported in italics below the coefficient estimates. *** and ** indicate statistical significance at the one and five percent level, respectively, using two-tailed tests. The sample includes 88,852 firm-year observations from 1989 to 2014.

TABLE 4
AT estimates of conditional conservatism:
Controlling for variation with non-accounting factors

	Dependent variable =	
	x_{it}	ac_{it}
Intercept	0.022***	0.024***
	5.45	*3.80*
$I(r_{it} < 0)$	0.003	0.010
	0.79	*1.54*
$Rank(R_{it}^{E})$	0.004	-0.019*
	1.06	*-1.69*
$Rank(\lambda_{it})$	0.001	-0.017**
	0.21	*-2.27*
$Rank(w_{it})$	-0.007**	-0.002
	-2.35	*-0.26*
$Rank(R_{it}^{E}) \times I(r_{it} < 0)$	0.023***	0.020**
	4.26	*2.23*
$Rank(\lambda_{it}) \times I(r_{it} < 0)$	-0.006	0.001
	-1.09	*0.24*
$Rank(w_{it}) \times I(r_{it} < 0)$	-0.007*	-0.014**
	-1.80	*-2.45*
r_{it}	0.034***	0.011
	6.30	*1.36*
$Rank(R_{it}^{E}) \times r_{it}$	0.020***	-0.007
	2.87	*-0.65*
$Rank(\lambda_{it}) \times r_{it}$	-0.015	0.007
	-1.45	*0.73*
$Rank(w_{it}) \times r_{it}$	-0.014***	-0.012
	-2.82	*-1.32*
$I(r_{it} < 0) \times r_{it}$	0.040**	0.043*
	2.02	*1.73*
$Rank(R_{it}^{E}) \times I(r_{it} < 0) \times r_{it}$	0.108***	0.127***
	3.38	*2.82*
$Rank(\lambda_{it}) \times I(r_{it} < 0) \times r_{it}$	0.115***	0.112***
	5.23	*4.49*
$Rank(w_{it}) \times I(r_{it} < 0) \times r_{it}$	-0.055***	-0.066***
	-4.95	*-4.18*
Adj. R^2	11.29%	6.60%

This table reports time-series average values of coefficient estimates from annual cross-sectional AT regressions of unexpected earnings scaled by lagged market value of equity (x_{it}) and accruals scaled by lagged market value of equity (ac_{it}) on positive and negative unexpected returns (r_{it}). The right-hand-side of the AT regression model is expanded to include as additional regressors the expected return (R^E), the spread in the conditional variances of unexpected returns (λ), and cash flow persistence (w), along with their interactions with positive and negative unexpected returns. To ease the interpretation of the coefficient estimates, we replace the raw values of R^E, λ, and w with the corresponding annual quintile ranks scaled to range from zero (lowest quintile) to one (highest quintile). Fama-MacBeth *t*-statistics are reported in italics below the coefficient estimates. ***, **, and * indicate statistical significance at the one, five, and ten percent level, respectively, using two-tailed tests. The sample includes 88,852 firm-year observations from 1989 to 2014.

TABLE 5
Towards an alternative measure of conditional conservatism:
Spread in the conditional variances of accruals

	Test variable z_{it} =		
	ac_{it}	cf_{it}	y_{it}
$Var(z_{it})$	4.01%***	2.98%***	3.96%***
	7.46	*11.49*	*7.81*
$Var(z_{it} \mid r_{it} \geq 0)$	3.30%***	3.26%***	2.90%***
	7.83	*9.54*	*6.97*
$Var(z_{it} \mid r_{it} < 0)$	4.45%***	2.70%***	4.43%***
	7.06	*12.44*	*7.59*
$Var(z_{it} \mid r_{it} < 0) - Var(z_{it} \mid r_{it} \geq 0)$	1.14%***	-0.56%***	1.53%***
	3.41	*-2.89*	*4.11*

This table reports time-series average values of the unconditional and conditional annual cross-sectional variances along with the spread in the annual cross-sectional conditional variances for accruals scaled by lagged market value of equity (ac_{it}), free cash flows scaled by lagged market value of equity (cf_{it}), and earnings scaled by lagged market value of equity (y_{it}), across partitions based on the sign of contemporaneous unexpected returns (r_{it}). Fama-MacBeth t-statistics are reported in italics below our estimates. *** indicates statistical significance at the one percent level using two-tailed tests. The sample includes 88,852 firm-year observations from 1989 to 2014.

TABLE 6
Alternative measure of conditional conservatism:
Construct validity tests using lagged placebo test variables

Panel A: Asymmetry in the conditional variances of lagged placebo test variables.

	Placebo test variable z_{it-1} =			
	ac_{it-1}	cf_{it-1}	y_{it-1}	$1/P_{it-1}$
$Var(z_{it-1})$	5.28%***	3.20%***	5.83%***	0.63%***
	4.57	*7.88*	*4.38*	*3.21*
$Var(z_{it-1} \mid r_{it} \geq 0)$	5.46%***	3.21%***	6.08%***	0.67%***
	3.69	*6.53*	*3.56*	*3.25*
$Var(z_{it-1} \mid r_{it} < 0)$	5.18%***	3.18%***	5.67%***	0.60%***
	5.09	*8.57*	*4.73*	*3.16*
$Var(z_{it-1} \mid r_{it} < 0) - Var(z_{it-1} \mid r_{it} \geq 0)$	-0.29%	-0.04%	-0.41%	-0.07%
	-0.35	*-0.14*	*-0.39*	*-1.54*

Panel B: AT coefficient estimates for lagged placebo test variables.

	Dependent variable =			
	ac_{it-1}	cf_{it-1}	y_{it-1}	$1/P_{it-1}$
Intercept	-0.004	0.016***	0.012**	0.021***
	-0.70	*3.56*	*2.55*	*5.42*
$I(r_{it} < 0)$	0.008***	0.006**	0.014***	0.000
	2.64	*2.29*	*3.83*	*0.36*
r_{it}	-0.030***	-0.032***	-0.062***	0.017***
	-5.34	*-6.52*	*-6.97*	*5.28*
$I(r_{it} < 0) \times r_{it}$	0.064***	0.176***	0.240***	-0.043***
	3.08	*16.56*	*8.39*	*-5.06*
Adj. R^2	0.71%	3.98%	3.97%	1.67%

Panel A reports time-series average values of the unconditional and conditional annual cross-sectional variances along with the spread in the annual cross-sectional conditional variances for a set of placebo test variables, including the accrual component of lagged earnings scaled by lagged market value of equity (ac_{it-1}), lagged free cash flows scaled by lagged market value of equity (cf_{it-1}), lagged earnings scaled by lagged market value of equity (y_{it-1}), and the inverse of lagged market value of equity ($1/P_{it-1}$), across partitions based on the sign of unexpected returns in t (r_{it}). Panel B reports time-series average values of coefficient estimates from annual cross-sectional AT regressions of the placebo test variables on positive and negative unexpected returns in t (r_{it}). Fama-MacBeth t-statistics are reported in italics below our estimates. ***, **, and * indicate statistical significance at the one, five, and ten percent level, respectively, using two-tailed tests. The sample includes 88,852 firm-year observations from 1989 to 2014.

TABLE 7
Alternative measure of conditional conservatism:
Construct validity tests using “conditionally conservative accruals”

Panel A: Asymmetry in the conditional variances of accruals and special items.

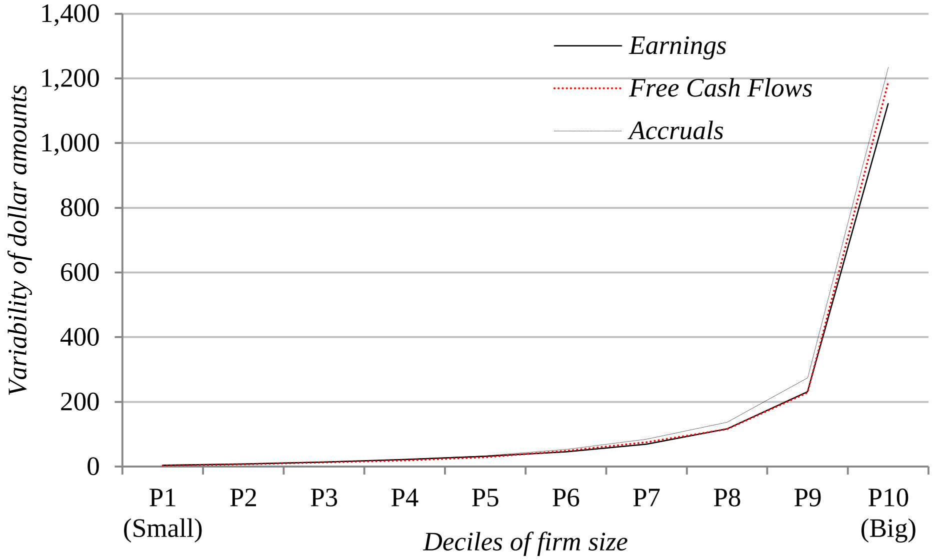
	Test variable Z_{it} =		
	ac_{it}	si_{it}	$ac_{it} - si_{it}$
$Var(Z_{it})$	4.01%***	0.65%***	2.97%***
	7.46	*5.31*	*8.38*
$Var(Z_{it} \mid r_{it} \geq 0)$	3.30%***	0.35%***	2.80%***
	7.83	*5.29*	*8.65*
$Var(Z_{it} \mid r_{it} < 0)$	4.45%***	0.85%***	3.06%***
	7.06	*5.31*	*7.92*
$Var(Z_{it} \mid r_{it} < 0) - Var(Z_{it} \mid r_{it} \geq 0)$	1.14%***	0.50%***	0.26%
	3.41	*5.05*	*1.39*

Panel B: AT coefficient estimates for accruals and special items.

	Dependent variable =		
	ac_{it}	si_{it}	$ac_{it} - si_{it}$
Intercept	0.009	-0.014***	0.023***
	1.34	*-11.88*	*3.96*
$I(r_{it} < 0)$	0.018***	0.007***	0.011***
	4.92	*3.96*	*4.78*
r_{it}	-0.003	-0.002	-0.001
	-0.59	*-1.25*	*-0.17*
$I(r_{it} < 0) \times r_{it}$	0.169***	0.072***	0.097***
	8.40	*7.55*	*7.42*
Adj. R^2	3.59%	3.77%	1.86%

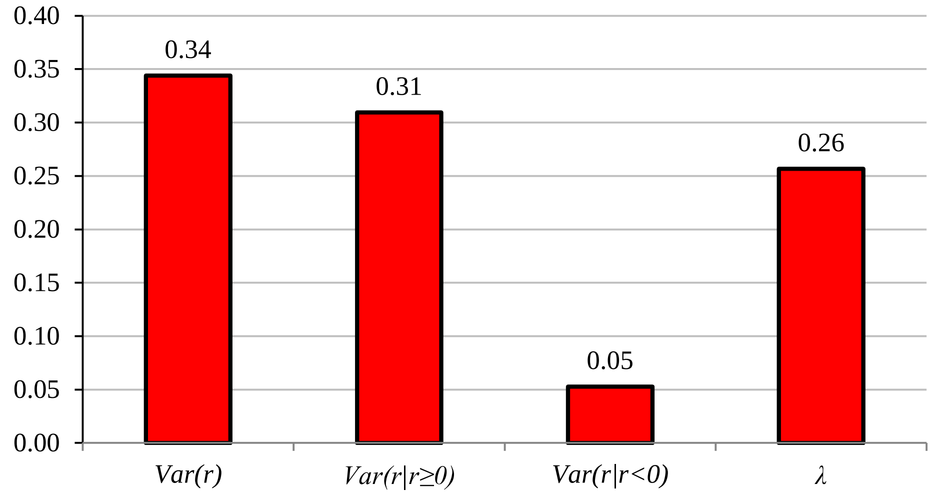
Panel A reports time-series average values of the unconditional and conditional annual cross-sectional variances along with the spread in the annual cross-sectional conditional variances for accruals scaled by lagged market value of equity (ac_{it}), special items scaled by lagged market value of equity (si_{it}), and accruals excluding special items scaled by lagged market value of equity (ac_{it}-s_{it}), across partitions based on the sign of unexpected returns in t (r_{it}). Panel B reports time-series average values of coefficient estimates from annual cross-sectional AT regressions of the same set of test variables on positive and negative unexpected returns in t (r_{it}). Fama-MacBeth t-statistics are reported in italics below our estimates. *** indicates statistical significance at the one percent level using two-tailed tests. The sample includes 88,852 firm-year observations from 1989 to 2014.

FIGURE 1
Evidence of scale effects in raw accounting data



This figure plots the time-series average values of the standard deviation of the dollar values of earnings, the accrual component of earnings, and free cash flows across decile portfolios formed each year based on the beginning of year market value of equity. The sample includes 88,852 from 1989 to 2014.

FIGURE 2
Asymmetry in the returns distribution



Note: The spread between the conditional variances of positive and negative unexpected returns (λ) is significantly different from zero at the one percent level using two-tailed tests.

This figure plots the time-series average values of the unconditional variance of unexpected returns, the conditional variances of unexpected returns, and the spread between the conditional variance of positive unexpected returns and the conditional variance of negative unexpected returns (λ). The sample includes 88,852 from 1989 to 2014.