# Understanding how the effects of conditional conservatism measurement bias vary with the research context

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ABSTRACT We re-examine previous seminal studies on conditional conservatism (CC) that apply the asymmetric timeliness (AT) measure of Basu (1997) and compare the outcomes with those based on the modified AT (MAT) measure of Badia, Duro, Penalva, & Ryan (2021) and the spread in conditional variances (SCV) measure of Dutta & Patatoukas (2017). Our conclusions are threefold. First, all three measures yield similar inferences in interrupted time-series settings that examine the change in CC following a policy mandate. Second, the inferences drawn from the AT measure in studies that model the determinants of CC based on cross-sectional settings are more sensitive to test specifications and research designs. Third, across the three measures, MAT shows the best empirical performance in terms of aligning with existing theories while being less affected by AT bias.

**Keywords:** Conditional conservatism; Asymmetric timeliness; Measurement bias; Type 1 error.

## 1. Introduction

We examine the extent to which the inferences of prior empirical research on conditional conservatism (CC) based on the Basu (1997) asymmetric timeliness (AT) measure could be affected by the concerns regarding its underlying bias and the potential for Type I error (Dietrich, Muller, & Riedl, 2007; Dutta & Patatoukas, 2017; Gigler & Hemmer, 2001; Patatoukas & Thomas, 2011, 2016). Building on the highly cited contribution of Basu (1997), there is a large and growing literature on CC.[1] Many studies claim to confirm theoretical propositions about CC using AT as their empirical proxy for CC (Dechow, Ge, & Schrand, 2010; Kothari & Wasley, 2019). Examples include the reduction of lender-shareholder conflict (Ahmed, Billings, Morton, & Stanford-Harris, 2002), the strengthening of debt contracting efficiency (Ball, Robin, & Sadka, 2008; Zhang, 2008), the moderation of agency problems (LaFond & Roychowdhury, 2008), and the reduction of information asymmetry (LaFond & Watts, 2008).[2]

However, while studies based on AT confirm the underlying theories associated with the determinants and consequences of CC, there is a parallel debate in the CC literature on the validity of this measure. On the one hand, some studies highlight various sources of bias in AT that could induce Type I error (Dietrich et al., 2007; Gigler & Hemmer, 2001; Patatoukas & Thomas, 2011, 2016) and propose an alternative measure of CC based on the spread in conditional variances (SCV) of earnings or accruals (Dutta & Patatoukas, 2017). On the other hand, other studies suggest adjustments to mitigate such concerns over the AT construct (Ball, Kothari, & Nikolaev, 2013b, 2013a; Collins, Hribar, & Tian, 2014) and develop a modified AT measure (henceforth MAT) that

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[1] For useful reviews on the CC literature, see Beyer, Cohen, Lys, & Beverly (2010), García Lara, García Osma, & Penalva (2014), Mora & Walker (2015), Penalva & Wagenhofer (2019), Ruch & Taylor (2015), Wang, Hógartaigh, & Zijl (2009), and Watts (2003b, 2003a).

[2] To help motivate our study, we conducted a search of all original research articles, since Basu (1997) to date, which apply the AT construct and are published in five leading accounting journals. We counted a total of 148 articles, with the time trend as well as topic and journal distribution provided in Appendix A.

incorporates major improvements and controls for documented sources of bias (Badia et al., 2021). In response to this controversy, we heed the call of Ball (2016b) and Hail, Lang, & Leuz (2020) to revisit previous empirical studies and verify their conclusions using new methodologies and measures.

We reevaluate findings based on the AT measure by employing the MAT measure of Badia et al. (2021) and the SCV measure of Dutta & Patatoukas (2017). The MAT measure controls for (i) firm-persistent characteristics unrelated to CC, (ii) the expected components of earnings and returns, (iii) unconditional conservatism and prior conditional conservatism, and (iv) the return variance effect of Patatoukas & Thomas (2011). Alternatively, the SCV measure captures CC as the spread between the variance of bad news earnings or accruals and the variance of good news earnings or accruals. Dutta & Patatoukas (2017) observe that AT depends not only on CC but also on non-accounting and firm-specific economic factors, such as expected returns, cash-flow persistence, and asymmetric returns distributions. In contrast, Dutta & Patatoukas (2017) show that their SCV measure is more associated with empirically observable accounting conservatism proxies such as asset write-downs and impairments. Nevertheless, Badia et al. (2021) employ their MAT measure in regression analyses and show that it is immune to the sources of bias documented in Patatoukas & Thomas (2011) and Dutta & Patatoukas (2017), among others.

For our empirical analyses, we re-examine the inferences of two sets of previous studies based on the AT construct. To the extent that the interaction between earnings and returns is affected by the choice of the cross-sectional versus (inter-temporal) time-series settings (Beaver, McNichols, & Wang, 2018), our first set of analyses reconsiders previous studies that exploit an interrupted time-series setting to examine how exogenous changes in accounting policy affect CC. Within this context, the Sarbanes-Oxley Act (SOX) and the widespread adoption of International

Financial Reporting Standards (IFRS) are widely recognized in the accounting literature as two of the most far reaching regulatory changes in recent times (Leuz & Wysocki, 2016). As such, the two studies that we re-assess are Lobo & Zhou (2006), who examine the impact of SOX on CC in a U.S. sample, and Ahmed, Neel, & Wang (2013), who evaluate the effect of IFRS on CC in an international sample. Our second set of analyses reevaluates previous studies that model the determinants of CC in cross-sectional settings based on variations in firm characteristics. Within this context, outsider equity investors and debt-holders (particularly public debt-holders) are commonly assumed to be the two most important sources of demand for CC (Watts, 2003a, 2003b). Therefore, the two studies we reconsider are LaFond & Watts (2008), who study the effect of the variation of information asymmetry on CC among firms in a U.S. sample, and Ball et al. (2008) who compare the role of the debt market orientation as a driver of CC against the equity market based on an international cross-section. All four studies that we re-examine have significantly impacted the accounting literature.[3]

Our primary research objective is to evaluate whether the reliance of previous empirical studies based on the AT measure renders their findings sensitive to potential confounding effects introduced through the research settings and test specifications even if the overall inference is consistent with the underlying well-established theoretical rationale. We argue that among the studies based on interrupted time-series settings, the AT measure is more likely to capture CC variations because these studies examine changes in CC among the same group of firms following changes in regulations and standards that are exogenously introduced. In contrast, among studies based on cross-sectional settings, the AT measure is more likely to be affected by confounding

[3] In Appendix B, we provide a selected list of research publications that cite each of these four previous studies we re-examine, showing their importance and relevance to the literature through citation counts based on Google Scholar.

effects associated with economic rather than accounting factors due to cross-sectional heterogeneity across different groups of firms.

Specifically, LaFond and Watts (2008) use the probability of informed trading (*PIN*) as a proxy for information asymmetry, and Ball et al. (2008) use the country-level market value of debt as a proxy for debt market orientation. Although information asymmetry and debt contracting are determinants of CC with well-established theoretical foundations (Watts, 2003b, 2003a), their empirical proxies based on *PIN* (Duarte & Young, 2009; Easley, Hvidkjaer, & O'Hara, 2002) and country-level market value of debt (Barth, Hodder, & Stubben, 2008; Dhaliwal, Heitzman, & Zhen Li, 2006), respectively, could also capture confounding effects associated with economic factors. Existing literature suggests that such economic factors cause the association between negative earnings and returns to be greater than that between positive earnings and returns, i.e., the concave earnings-return relationship (Beaver & Ryan, 2009; Breuer & Windisch, 2019), which could be a potential source of bias embedded in the AT measure (Patatoukas & Thomas, 2011). In other words, the findings of empirical studies based on the AT measure that indicate CC is stronger among firms associated with higher *PIN* and country-level market value of debt could be due to biased AT estimates rather than differences in their CC accounting practices.

In line with the view that cross-sectional heterogeneity in economic factors is a significant driver of cross-sectional variation in AT estimates, we observe a marked contrast in the effects of AT measurement bias between the interrupted time-series and the cross-sectional research settings. In both of our interrupted time-series analyses, we observe that the AT, MAT, and SCV measures generate similar inferences. In contrast, in cross-sectional settings, we find that the main results of the two re-examined studies no longer hold when using MAT or SCV in place of AT. Nevertheless, when we apply modifications related to test specifications and research designs, we find results consistent with the theoretical motivations of the original studies. Specifically, in the reexamination of LaFond & Watts (2008), the MAT and SCV measures do not show that CC increases in the *PIN* score. However, observing that the *PIN* score can be confounded by effects associated with stock illiquidity (Duarte & Young, 2009) and is driven by expected returns (Easley et al., 2002), we replace the *PIN* score with a more direct proxy of information asymmetry, i.e., the bid-ask spread. We then find that the three measures of CC yield consistent evidence that aligns with the theory suggesting that information asymmetry drives CC.

Moving to the other cross-sectional study of Ball et al. (2008), when replacing the AT measure with either MAT or SCV, the results from the country-level regressions do not support the conclusion that the importance of debt markets drives CC. Nevertheless, when we employ a more sophisticated research design, the results show that the AT and MAT measures increase in debt markets with high importance, consistent with theory. Further investigation reveals that the SCV measure does not yield significantly different values between high- and low-importance debt markets. We argue in Section 4 that the SCV measure is not suitable for cross-sectional international comparisons. In sum, we show that the inferences drawn from CC studies based on cross-sectional settings are sensitive to both the choice of the CC measure and other test specification and research design choices.

Our study contributes to the CC literature in three ways. First, unlike previous studies involved in the debate either for (Ball et al., 2013b, 2013a; Collins et al., 2014) or against (Dietrich et al., 2007; Gigler & Hemmer, 2001; Patatoukas & Thomas, 2011, 2016) the AT construct, we provide more direct empirical evidence by revisiting the settings and replicating the analyses applied in previous CC studies. Second, also unlike previous related studies, our paper draws a multifaceted rather than a single-sided conclusion to the issue. We reveal that the impact of the measurement bias associated with AT, and any potential concerns over the inferences drawn from previous studies, depend on the research setting and test specifications involved. To the extent that AT captures both CC and bias effects (Dietrich et al., 2007), it appears to be more effective in capturing the CC effect in studies of exogenous and temporal changes in accounting policy, and more likely to pick up the confounding effect in the studies of cross-sectional variations in firm characteristics. Third, we find that MAT is more successful in mitigating the weaknesses of AT and upholds inferences of prior research that align with the relevant theoretical rationale. On the other hand, despite being theoretically sound, the SCV measure does not seem to perform well in international data where cross-sectional heterogeneity is high. Specifically, the SCV measure is a spread and not a ratio, which renders the SCV measure unreliable when comparing two groups with different statistical attributes of earnings and returns.[4] In addition, the SCV measure is highly sensitive to outliers as it is based on the variance of earnings or accruals.

Overall, our study confirms that the theoretical rationale associated with the prior studies we replicate holds well, but prior studies based on cross-sectional settings are more likely to incur findings that are sensitive to test specification and research design choices. For future CC studies, our findings imply that the MAT measure can be adopted as a primary measure while the SCV measure may be used as a robustness check in some cross-sectional contexts. The rest of this paper is organized as follows. Section 2 reviews the relevant literature and develops our testable hypotheses. Section 3 discusses sample selection and presents the findings from our empirical analyses. Section 4 presents the study's conclusions.

[4] We discuss the issues that may arise when using the ratio instead of the spread of conditional variances in Section 4.

## 2. Literature Review and Hypothesis Development

### *2.1. Conditional conservatism and its measurement*

Conservatism is one of the oldest concepts in accounting (Bliss, 1924; Sterling, 1976). Theoretically, CC is an accounting practice that recognizes economic losses before being realized and recognizes economic gains only when realized (Beaver & Ryan, 2005; Watts, 2003b). Empirical studies define conservatism as the requirement of a lower degree of verification to recognize economic losses compared to the degree of verification required to recognize economic gains (Basu, 1997; Pope & Walker, 1999). Accounting conservatism can be either unconditional or conditional, with the former being associated with the understatement of the book value of net assets regardless of news, while the latter referring to the timelier recognition of bad news than good news in earnings (Ryan, 2006). Examples of unconditional conservatism include the expensing of investment in intangible assets and setting depreciation rates for property plant and equipment above the expected economic rate of depreciation (Beaver & Ryan, 2005), whereas an example of conditional conservatism is achieved through asset impairments in response to bad news about the value of assets in place (Dutta & Patatoukas, 2017).

Basu (1997) developed a construct for measuring the level of CC by regressing earnings on stock returns. This construct has become the most widely adopted CC construct in the literature (Mora & Walker, 2015; Ruch & Taylor, 2015; Wang et al., 2009; Watts, 2003a). The AT construct of Basu (1997) is depicted in the earnings-returns piecewise linear regression below:

$$X_{it} = \beta_0 + \beta_1 RD_{it} + \beta_2 RET_{it} + \beta_3 RD_{it} \times RET_{it} + \varepsilon_{it} \quad (1)$$

where for firm $i$ in year $t$, $X_{it}$ is current year earnings per share deflated by price per share at the end of previous year, $RD$ is a dummy variable that equals 1 if $RET$ is negative, and 0 otherwise, and $RET$ is the abnormal stock returns over the fiscal year, calculated as monthly compounded buy-and-hold observed returns (including distributions) minus expected returns.[5] The coefficient on the interaction term ($\beta_3$) captures the incremental timeliness with which reported earnings reflect bad news relative to good news. We refer to this coefficient as the AT measure of CC.

Appendix A provides a summary of CC studies, published in five leading accounting journals, that apply the AT construct. The summary includes original research articles that utilize the AT measure either for empirical tests or theoretical model development, and excludes review articles. We also include studies that apply the C_Score developed by Khan & Watts (2009) to implement the AT construct on a firm-specific basis. A total of 148 articles were published over the period 1997 to end of 2021, which amounts to nearly six articles per year (and above nine articles per year post 2010). Panel A provides a time trend analysis that indicates a large and steadily growing literature. Panel B provides a topic analysis, which indicates that most of the papers are related to equity and debt markets along with their financial reporting aspects. Measurement issues relating to CC also rank high, which indicates that there is a substantial debate in the accounting literature on the topic we seek to examine in this paper. Panel C provides a journal distribution analysis, showing that the largest number of articles were published in the *Journal of Accounting and Economics*, where Basu (1997) was originally published. Overall, Appendix A helps motivate our study by confirming the importance of the AT construct in the accounting literature on CC.

### *2.2. Debate on the validity of the AT measure*

Despite the widespread adoption of the AT construct in the CC literature, there is a parallel on-going debate within this literature on the validity of this construct. On the one hand, various

[5] In our empirical analyses, we follow recent CC studies and calculate expected returns as the average returns of 5×5 portfolios formed by sorting firms first based on the beginning of year market value of equity and then based on beginning of year book-to-market equity ratio (Badia et al., 2021; Ball et al., 2013b; Dutta & Patatoukas, 2017).

studies highlight potential sources of bias that increase the likelihood of Type I error. Gigler & Hemmer (2001) develop a theoretical model showing that the AT measure might be significantly positive in the absence of CC because researchers fail to control for firms' voluntary disclosure, which jointly affects stock returns and accounting conservatism. Dietrich et al. (2007) demonstrate that the AT measure may indicate CC even in the absence of CC due to sample-variance-ratio bias and truncation bias. These biases arise from the fact that earnings cause returns and not vice versa and that the returns variable is also determined by other news unrelated to earnings. Patatoukas & Thomas (2011) identify two sources of bias driven by the use of stock price as the deflator of the dependent variable in the Basu (1997) piecewise linear regression to estimate the AT measure. First, firms with a low stock price tend to report losses more frequently, and this leads them to have more negative values for the dependent variable in the earnings-returns piecewise regression (i.e., the loss effect). Second, firms with low stock prices have higher fluctuations in their stock prices, which results in a higher variance in stock returns (i.e., the return-variance effect). These two effects *jointly* lead to an upward bias in the AT measure, particularly among firms with low stock prices.[6]

On the other hand, other studies suggest arguments or adjustments to mitigate concerns over the AT measure. For instance, Ryan (2006) questions the severity of the bias identified by Dietrich et al. (2007) by arguing that such concerns have a trivial impact given the low $R^2$ observed empirically from the returns-earnings regressions. Ball et al. (2013a) further argue that the sample-variance-ratio bias suggested by Dietrich et al. (2007) is irrelevant because the covariance between returns and earnings, conditional on returns, is equal for good and bad news when CC is absent.

[6] In line with Patatoukas & Thomas (2011), Breuer & Windisch (2019) provide theoretical insights and empirical evidence suggesting that the earnings-returns concavity is increasing in the volatility of firms' underlying shock processes (i.e., uncertainty). This results in an asymmetric effect in the Basu (1997) piecewise linear regression absent accounting influences.

Furthermore, the price-deflator-related sources of bias in the AT measure highlighted by Patatoukas & Thomas (2011) have motivated other studies to propose modifications to the AT measure to correct the problem. Ball et al. (2013b) argue that the AT bias is driven by the expected components of earnings and returns, which are correlated with firm-specific variables, resulting in an association between the error term and the AT coefficient. They suggest various ways to remove such components and alleviate the AT measure from the bias suggested by Patatoukas & Thomas (2011). Collins et al. (2014) propose that the replacement of the earnings with accruals as the dependent variable in the Basu (1997) regression corrects for the bias raised by Patatoukas & Thomas (2011). Because the persistence of the cash-flow component in earnings increases with the expected component of returns, Collins et al. (2014) suggest that the removal of both components through the use of accruals and unexpected returns in the regression mitigates the spurious asymmetric timeliness in the AT measure.

The debate on the validity, or otherwise, of the AT construct continues. For example, in response to Ball et al. (2013b) and Collins et al. (2014), Patatoukas & Thomas (2016) demonstrate that their revised AT constructs continue to exhibit upward bias in placebo tests based on asymmetry in the conditional relations between the inverse of lagged share price and positive and negative returns news. Furthermore, Dutta & Patatoukas (2017) provide evidence that the revised AT constructs thus far are sensitive to and can be driven by three non-accounting-related and firm-specific economic factors: (i) expected returns (ii) asymmetry in the conditional variances of positive and negative unexpected returns, and (iii) cash-flow persistence. Furthermore, they argue that the revised AT constructs thus far become statistically and economically insignificant in the presence of these non-accounting factors as control variables.

### *2.3. Whether to replace or modify the AT measure*

As an alternative approach to the measurement of CC, Dutta & Patatoukas (2017) propose the SCV measure, which they document to be less affected by the sources of bias that drive the AT measure. According to Dutta & Patatoukas (2017), CC can be empirically estimated by calculating the spread between the variances of bad news earnings or accruals and the variance of good news earnings accruals. They use the sign of unexpected returns as a proxy for good and bad news, deflate earnings or accruals with the lagged stock price, and estimate the measure as follows:

$$SCV = Variance(X_{it} \mid RET_{it} < 0) - Variance(X_{it} \mid RET_{it} > 0) \tag{2}$$

All variables are defined previously and in Table 1. Dutta & Patatoukas (2017) employ accruals as the main variable in constructing SCV, yet they mention that their measure adapts for using earnings. We use earnings instead of accruals for three reasons: (i) the use of earnings strengthens comparability with analysis from the AT and MAT measures, (ii) Badia et al., (2021) emphasize that earnings is the main accounting variable used in contracting and, unlike accruals, it is observed rather than estimated by researchers, and (iii) the use of accruals causes severe loss of observations in international studies when estimated as the difference between earnings and free cash flow (Collins et al., 2014; Dutta & Patatoukas, 2017), in addition to showing outliers that go beyond the conventional 1% threshold in international data, which renders the SCV measure unreliable.

Theoretically, Dutta & Patatoukas (2017) argue that their measure is driven only by variations in CC. Moreover, they claim that, unlike the AT construct, their proposed measure is unaffected by the asymmetric distribution of returns and does not rely on the market efficiency assumption where investors incorporate all information in a timely and efficient manner in stock prices. They also demonstrate empirically that their measure is not affected by non-accounting and firm-specific economic factors that influence the AT construct, even under the modifications of Ball et al. (2013b) and Collins et al. (2014).

Whilst Dutta & Patatoukas (2017) propose the replacement of AT by a very different construct, Badia et al. (2021) develop a modified version of the AT measure, MAT, that builds on prior improvements to the Basu (1997) construct and also addresses the documented sources of bias. Specifically, MAT controls for biases arising from the association between the expected components of earnings and returns (Dietrich et al., 2007; Gigler & Hemmer, 2001), the loss effect documented in Patatoukas & Thomas (2011), other non-accounting factors that yield a non-linear relationship between earnings and returns (Beaver & Ryan, 2009; Breuer & Windisch, 2019), in addition to the persistence of firm-specific attributes that induce spurious CC yet unrelated to CC (Dutta & Patatoukas, 2017). First, the MAT measure uses the unexpected components of returns and earnings following Ball et al. (2013b). Second, the MAT measure is based on firm-fixed effects regressions instead of conventional OLS regressions, which, arguably, eliminates any time-invariant factors that might bias the AT coefficient (i.e., *RD*×*RET*). Third, the MAT measure controls interactively for the returns variance, which captures time-variant non-accounting factors that might induce spurious evidence of CC. Finally, the MAT measure controls interactively for the beginning of year market-to-book ratio to account for firms' unconditional conservatism and prior application of conditional conservatism. The preceding empirical adjustments result in the MAT model of Badia et al. (2021):

$$\begin{aligned} UX_{it} &= \alpha_i + \beta_1 RD_{it} + \beta_2 RET_{it} + \beta_3 RD_{it} \times RET_{it} \\ &+ \beta_4 RVAR_{it} + \beta_5 RD_{it} \times RVAR_{it} + \beta_6 RET_{it} \times RVAR_{it} + \beta_7 RD_{it} \times RET_{it} \times RVAR_{it} \\ &+ \beta_8 MTB_{it-1} + \beta_9 RD_{it} \times MTB_{it-1} + \beta_{10} RET_{it} \times MTB_{it-1} + \beta_{11} RD_{it} \times RET_{it} \times MTB_{it-1} + \varepsilon_{it} \end{aligned} \tag{3}$$

Where the dependent variable *UX* is the unexpected component of earnings, estimated as the residuals from the earnings expectation model that regresses current earnings on lagged earnings, an indicator variable if lagged earnings are negative, and the interaction of lagged earnings and the indicator variable (Badia et al., 2021; Ball et al., 2013b). The earnings expectation model is regressed for each industry-year separately.[7] *RVAR* is the variance of daily stock returns over the fiscal year, *MTB* is the market-to-book ratio of assets at the beginning of the year, and $\alpha_i$ denotes firm fixed effects. All other variables are as defined previously and in Table 1. The MAT measure is the coefficient $\beta_3$. Notably, Badia et al. (2021) show, and our replication of their work confirms, that the coefficient $\beta_7$ is negative and significant, statistically and economically, suggesting that the interactive inclusion of returns variance absorbs a significant part of the AT bias. In validating their MAT measure, Badia et al. (2021) subject their measure to an extensive set of quality checks to ensure its immunity to all documented sources of AT bias in the literature.

### *2.4. Hypothesis development*

As we discussed in Section 2.2, existing studies on the validity of the AT measure largely focus either on the identification of potential sources of bias (Dietrich et al., 2007; Dutta & Patatoukas, 2017; Gigler & Hemmer, 2001; Patatoukas & Thomas, 2011, 2016) or on proposing reasons and solutions to mitigate such bias (Badia et al., 2021; Ball et al., 2013b, 2013a; Collins et al., 2014). However, despite the well-documented concerns about the AT construct, limited attention has been paid to assessing the consequences of these concerns for prior and future research applications of the AT construct. Nevertheless, this is an important issue, particularly for

[7] Ball et al. (2013b) use the two-digit SIC code industry classification. We require each industry-year cross-section to have at least 10 observations and thus we use the Fama and French twelve industry classification to maximize the number of observations and limit data attrition in international datasets where the earnings expectation model is performed at the country-industry-year level. Nevertheless, our results remain unchanged when using the two-digit SIC industry classification.

accounting researchers seeking to rely on past empirical evidence or to develop research designs for future empirical studies.

In the accounting literature, the behavior of the interaction between earnings and returns is affected by the choice of the cross-sectional versus (inter-temporal) time-series settings (Beaver et al., 2018). Accordingly, there are two commonly-adopted research settings in which the AT construct is utilized to empirically examine the determinants of CC. In the first setting, researchers adopt an interrupted time-series design to observe whether and how exogenous changes in accounting policies affect CC for the same group of firms. In the second setting, researchers examine cross-sectional variations to observe whether and how differences in firm characteristics influence CC across different groups of firms. We expect the impact of the measurement error associated with the AT construct to differ across these two research settings for the following reason. As Dietrich et al. (2007) show in their Equations 1.7a and 1.7b, the AT measure is biased for good and bad news, respectively, where each equation comprises two components: the CC component and the bias component. The CC component is affected by accounting decisions related to timely loss recognition, while the bias component is driven by economic factors possibly related to economic factors rather than accounting (Dutta & Patatoukas, 2017; Patatoukas & Thomas, 2011, 2016). In an interrupted time-series setting, the exogenous change in accounting policy is expected to influence the CC component but not the bias component. As such, changes in the AT coefficient estimated for the same sample over a short period of time are, arguably, more likely to be driven by systematic changes in the underlying CC component, while changes in the bias component are likely to offset and wash out. In this case, we would expect inferences based on the AT, MAT, and SCV measures to be broadly consistent. In a cross-sectional setting, however, choices of empirical proxies of firm characteristics assumed to affect the CC component and/or other research design choices could simultaneously capture the bias component. For instance, Dutta & Patatoukas (2017) perform an additional analysis in which they find that the AT measure increases in financial leverage and decreases in firm size and market-to-book ratio due to the association between those firm characteristics and the bias component. As such, differences in the AT coefficients estimated on a cross-sectional basis are likely to be driven by *both* the CC and bias components, and particularly the latter when cross-sectional heterogeneity in economic factors outweighs that of CC. To the extent that the MAT and the SCV measures are evidently less prone to the bias embedded in the AT measure, the scope for conflict between results based on the AT measure and results based on the MAT and SCV measures is expected to be greater in cross-sectional settings. Given these arguments, we formulate the following testable hypotheses:

*H1: In interrupted time-series settings that examine exogenous changes in CC, the AT, MAT, and SCV measures are likely to lead to similar inferences.*

*H2: In cross-sectional settings that examine determinants of CC, the inferences from the AT, MAT, and SCV measures are more likely to be dissimilar.*

To ensure that our findings supporting hypothesis H2 are driven by the sensitivity of the AT measure to test specifications and research designs rather than the underlying theoretical rationale applied in the re-examined studies, we carry out further analysis based on alternative specifications and designs. Specifically, the *PIN* score applied in LaFond & Watts (2008) is likely to be a noisy proxy for information asymmetry, and the high cross-sectional heterogeneity arising from the country-level research design adopted by Ball et al. (2008) could induce confounding effects associated with economic factors that may drive the AT measure even in the absence of CC (Beaver & Ryan, 2009; Breuer & Windisch, 2019). Duarte & Young (2009) show that the *PIN* measure captures stock illiquidity instead of capturing information asymmetry as LaFond & Watts (2008) assume, while Easley et al. (2002) document that *PIN* is associated with expected returns, i.e., driven by an economic (risk) factor. In the same vein, Aktas, de Bodt, Declerck, & Van Oppens (2007) find that *PIN* does not increase when the level of information asymmetry is allegedly higher as in the cases of M&A deals. Similarly, country-level institutional factors applied in Ball et al. (2008) could pick up variations in the rule of law and level of corruption, which could induce economic risk factors (La Porta, Lopez-De-Silanes, Shleifer, & Vishny, 1997, 1998) that bias the AT measure in an unknown direction. In other words, assuming that the underlying theoretical rationales of LaFond & Watts (2008) and Ball et al. (2008) are correct, then the fact that the AT measure yields inconsistent findings relative to other CC measures opens up the possibility that the use of the *PIN* score and country-level market value of debt in these studies drives the findings.

## 3. Empirical Analyses

Our empirical analysis is based on tests of hypotheses H1 and H2. To test hypothesis H1, we re-examine the inferences of two previous studies associated with interrupted time-series settings, namely Lobo & Zhou (2006) for the effect of SOX on CC in the U.S. and Ahmed et al. (2013) for the effect of IFRS adoption on CC across an international sample. To test hypothesis H2, we re-assess the inferences of two previous studies associated with cross-sectional settings, specifically LaFond & Watts (2008) for the effect of information asymmetry on CC in the U.S. and Ball et al. (2008) for the effect of debt market prominence on CC in an international sample. These four studies are chosen primarily due to the importance of their inferences for the CC literature, as evident in Appendix B that provides a list of selected research articles citing these four studies along with their citation counts on Google Scholar.

In each case, we first re-examine the original findings and inferences based on the AT construct, following the sample construction of that study. We then re-examine the replicated findings using the MAT and SCV measures to observe whether the inferences implied by the AT measure still hold.[8] When using the MAT measure, we only report the AT coefficients for brevity and better exposition; however, all MAT regressions interactively include *RVAR* and *MTB* as depicted in Equation (3).We try to maintain the highest possible level of comparability across replicated studies without jeopardizing their research designs. For example, we apply similar data management practices, such as trimming the upper and lower percentiles of the earnings and returns distributions, and use the same definitions of earnings and abnormal returns throughout. Table 1 provides detailed definitions of all the variables used to re-examine the four previous studies. Since each set of re-examinations requires a different sample in accordance with the original study, we provide the description of the sample construction for each case separately in the corresponding sub-sections.

[Insert Table 1 here]

***3.1. Lobo & Zhou (2006) (interrupted time-series setting)***

Lobo & Zhou (2006) examine the change in the level of CC following the Sarbanes-Oxley (SOX) Act in 2002 for U.S. firms. They document that the SOX enactment leads to an increase in the degree of CC estimated by the AT measure, and they argue that this finding provides important empirical evidence that policy makers were able to achieve one of their main goals by improving

[8] Since the SCV is a non-linear parameter, we test the statistical significance of its differences across sub-samples through a non-linear combination of estimators using the delta method (Casella & Berger, 2002; Feiveson, 1999).

timely loss recognition by firms. Furthermore, their paper has informed the debate within a large accounting literature over the economic consequences of SOX (Coates & Srinivasan, 2014).

Consistent with Lobo & Zhou (2006), we use the Compustat fundamentals annual file for accounting data and CRSP for stock returns data. We also follow their sample construction procedure to retrieve accounting and returns data between 2000 and 2004, and exclude firms with penny stocks and observations with negative book value of equity. We then delete the top and bottom percentiles of the earnings and returns distributions, and we require an equal number of observations per firm pre- and post-SOX. The final sample comprises 5,622 (5,622) firm-year observations in the pre-SOX (post-SOX) period. Table 1 defines all the variables applied in this set of analyses.

Table 2 reports our replication and re-examination of Lobo & Zhou (2006). Table 2 Panel A provides summary statistics for the main variables and is comparable with Table 4 in their study. The first column in Table 2 Panel B reports the replication results of model (6b) in Table 4 of Lobo & Zhou (2006). The coefficient on the interaction term *SOX*×*RD*×*RET* is 0.0436 and is significant at the 1% level, which indicates an increase in the degree of CC following the passage of the SOX Act and confirms the finding of Lobo & Zhou (2006). Table 2 Panel C reports the re-examination of the Lobo & Zhou (2006) results using the MAT measure. Consistent with the results of Badia et al. (2021), the coefficient on *RD*×*RET* is positive and significant yet smaller in magnitude, indicating a less-biased estimation of CC. In addition, the unreported coefficients on *RVAR*×*RD*×*RET* and *MTB*×*RD*×*RET* are significantly positive and negative, respectively. More importantly, the coefficient on the interaction term *SOX*×*RD*×*RET* is 0.0672 and is also significant at the 1% level. Finally, Table 2 Panel D reports the change in SCV around SOX, showing a statistically significant increase from 0.0037 in the pre-SOX period to 0.0068 in the post-SOX period. In other words, the MAT and SCV measures corroborate the inferences based on the AT measure generated by Lobo & Zhou (2006) and show an increase in CC after SOX. As such, we provide a set of findings that is consistent with hypothesis H1, which predicts that all examined measures are likely to generate similar inferences for CC under interrupted time-series settings involving exogenous accounting policy changes.

[Insert Table 2 here]

***3.2. Ahmed, Neel, & Wang (2013) (interrupted time-series setting)***

Ahmed et al. (2013) examine the change in CC following the mandatory adoption of IFRS in 2005 across an international sample. They compare the change in the level of CC in 20 countries that mandated the adoption of IFRS in 2005 (i.e., the treatment group) with the change in CC for a control group consisting of 15 countries that did not adopt IFRS in 2005. Following a difference-in-differences research design, Ahmed et al. (2013) document a significant reduction in the level of CC among the treatment group (i.e., the IFRS adoption sample) as compared to the control group (i.e., benchmark sample). This is an important finding for the International Accounting Standards Board (IASB), which is engaged in a policy debate on whether to keep conservatism in its conceptual framework in light of investors' demands for conservative financial reporting (Ball, 2016a; Cooper, 2015).

Following the sample selection criteria in Ahmed et al. (2013), we download accounting data from the Compustat Global fundamental annual file and stock returns data from the Compustat Global security file for all countries. We first keep firms that have non-missing variables between 2002 and 2007, we then delete firms that report using IFRS before 2005 or firms that do not report under IFRS in 2005 and beyond. Next, we delete observations in 2005 (i.e., transition year), observations in the top and bottom 1% of the earnings and returns distributions, and observations with negative book value of equity. Finally, we match each firm in the treatment sample to a firm in the control sample that operates in the same industry. The matching procedure is based on market value of equity, book to market ratio, and net income in the years that precede IFRS adoption. The final sample comprises 6,238 treatment firm-year observations and 6,238 control firm-year observations. We identify the pre- and post-IFRS periods using the indicator variable *Post* that takes the value 1 if the financial year is later than 2005, and zero otherwise. Similarly, we identify treatment firms using the indicator variable *Treat* that takes the value 1 if the firm is listed in a country that mandated the adoption of IFRS in 2005, and zero otherwise. Our re-examination of Ahmed et al (2013) replicates their Equation (6) in which they adapt the Basu (1997) piecewise linear regression to a difference-in-differences design while controlling for book-to-market ratio, firm size, and leverage, along with their interactions with *RD*, *RET*, and *RD*×*RET*. The variables of interest are *RET*×*RD*×*Post* that captures any changes in CC due to a general trend and *RET*×*RD*×*Post*×*Treat* that captures the change in CC among the treatment group due to the IFRS mandate.[9] Table 1 defines all the variables applied in this set of analyses.

Table 3 presents our replication and re-examination of Ahmed et al. (2013). Table 3 Panel A reports the summary statistics for the variables used in our replication of their main findings for the treatment and control groups separately. Table 3 Panel B reports the regression that replicates column 1 of Table 6 in their study.[10] The coefficient on *RET*×*RD*×*Post* is positive, which indicates an increase in the level of CC among firms in the control group, although this increase is

[9] To the extent that the effect of IFRS adoption is found to be more prominent in Europe due to better legal enforcement (Christensen et al., 2013), we have also replicated André, Filip, & Paugam (2015) who examine the change in CC following the mandatory IFRS adoption in 16 European countries. Andre et al. (2015) employ the C_Score measure of CC (Khan & Watts, 2009) and find a significant reduction in C_Score post-2005. Our re-examination of their inference confirms their findings and supports the alignment of AT with the MAT and SCV measures in interrupted time-series settings.

[10] As mentioned earlier, for brevity and exposition purposes, we only report the coefficients on the variables of interest. We state the full regression equation in the notes of Table 3.

statistically insignificant. More interestingly, the significantly negative coefficient on *RET×RD×Post×Treat* (–0.1564 and significant at 1% level) suggests a significant post-IFRS reduction in the degree of CC among firms in the treatment group compared to that among the control group, and this result confirms the finding of Ahmed et al. (2013). Table 3 Panel C reports the results from the re-examination of Ahmed et al. (2013) using the MAT measure.[11] The MAT coefficient estimate (*RET×RD*) is positive and significant yet with a smaller magnitude compared to AT. More importantly, the coefficient on *RET×RD×Post×Treat* continues to be negative and significant, suggesting a reduction in CC among treated firms following IFRS adoption. Similarly, in Table 3 Panel D, the SCV measure for the treatment group further corroborates this inference by showing a statistically significant decrease from 0.0059 in the pre-IFRS period to −0.0013 in the post-IFRS period, i.e., the treatment group does not exhibit CC in financial reporting following the IFRS mandate. On the other hand, the control group shows an insignificant change in the SCV measure around IFRS adoption. In other words, the results reported in Panels C and D are consistent with those from the AT measure reported in Panel B. Collectively, our findings in Table 3 confirm the conclusions of Ahmed et al. (2013) and are consistent with hypothesis H1 suggesting that all examined CC measures are likely to draw similar inferences under interrupted time-series settings that examine how CC is affected by exogenous changes in accounting policies.

[Insert Table 3 here]

### *3.3. LaFond & Watts (2008) (cross-sectional setting)*

LaFond & Watts (2008) examine whether information asymmetry in the equity market generates a demand for CC in financial reporting. They measure information asymmetry through

[11] Given that Ahmed et al. (2013) controls interactively for the book-to-market ratio, we exclude the market-to-book ratio (i.e., the reciprocal of *BTM*) and its interactions from the MAT measure to avoid multi-collinearity.

the probability of information-based trading (*PIN*) developed by Easley & O'Hara (1992), and show a positive relationship between *PIN* and CC. This inference is important because it demonstrates that, in addition to debt-holders (Ahmed et al., 2002; Ball et al., 2008; Zhang, 2008), equity investors also demand timely loss recognition. Their evidence has important implications for standard setters that question the value of accounting conservatism (e.g., FASB, 1980, 2005).

LaFond & Watts (2008) conduct their study based on a U.S. sample that includes firms listed on the NYSE and AMEX exchanges over the period of 1983 to 2001. To construct their sample, they require firms to have data for the *PIN* score and sufficient CRSP and Compustat data for the empirical analyses. We follow their sampling procedure and acquire data for the *PIN* score from the same source.[12] Our final sample comprises 19,640 firm-year observations over the same sample period after excluding the top and bottom percentiles of the earnings and returns distributions, penny stocks, and observations with negative book value of equity. Table 1 defines all variables used in this set of analyses.

Table 4 presents our replication and re-examination of LaFond & Watts (2008). Table 4 Panel A reports summary statistics for the variables we employ. Note that the number of observations and the relevant statistics are generally close to those reported in Table 1 Panel A of their study. Table 4 Panel B reports our replication of their main analyses on the effect of *PIN* on CC measured through the AT measure. Note that the coefficients estimated for $RD \times RET$ (0.0895) and $PIN \times RD \times RET$ (0.5986) are both positive and statistically significant at the 1% level. This yields an inference consistent with Table 2 Panel C of their study and confirms that firms with larger *PIN* scores (i.e., presumably greater information asymmetry) are associated with higher

[12] We are grateful to Soeren Hvidkjaer for providing the *PIN* measure dataset on his website (https://sites.google.com/site/hvidkjaer/). LaFond & Watts' (2008) footnote 10 indicates that they also acquired the *PIN* score from his earlier website (http://www.smith.umd.edu/faculty/hvidkjaer).

levels of CC. Turning to Table 4 Panel C, we re-examine the *PIN* effect on CC through the MAT measure. The significance of the coefficient on *PIN*×*RD*×*RET* disappears when employing the MAT measure. This insignificant coefficient suggests that the *PIN* score is less likely to capture information asymmetry (Duarte & Young, 2009) while being associated with the sources of bias in the AT measure (Easley et al., 2002), yielding spurious evidence of CC. Furthermore, we apply the SCV measure in Table 4 Panel D in which we split the sample into high and low *PIN* score using a dummy variable (*HiPIN*) that takes the value 1 if the *PIN* score is higher than the industry-year median value. The SCV values estimated for the high and low *PIN* groups are 0.0012 and 0.0017 with a *p*-value of 0.176 for the statistical significance of the difference, indicating a positive yet insignificant difference in CC between the high and low *PIN* groups. In other words, the MAT and SCV measures suggest that firms with higher *PIN* scores are not associated with higher levels of CC, which negates the findings of LaFond & Watts (2008) based on the AT measure.

[Insert Table 4 here]

To the extent that the re-examination of LaFond & Watts (2008) based on the MAT measure (and SCV measure) contradicts the theory of CC (Watts, 2003b), we further investigate the possibility that the issue resides in the choice of *PIN* as a noisy proxy that does not capture information asymmetry (Aktas et al., 2007; Duarte & Young, 2009). As such, we move forward and replace the *PIN* score with a more direct measure of information asymmetry, the bid-ask spread (*BIDASK*). Table 5 Panel A reports the summary statistics for the *BIDASK* proxy that we employ in the same sample used to re-examine LaFond & Watts (2008). Table 5 Panel B reports the results of the AT measure while interactively including *BIDASK*, which shows a positive and highly significant association with CC as indicated by the coefficient on *BIDASK*×*RD*×*RET*. Performing the same analysis for the MAT measure, as shown in Table 5 Panel C, yields similar inference as the coefficient on *BIDASK*×*RD*×*RET* is positive and significant, suggesting that CC increases in information asymmetry when captured using the bid-ask spread. It is worth mentioning that the coefficient on *BIDASK*×*RD*×*RET* in the AT construct is 4.0723 while dropping to 1.1050 in the MAT construct (both being significant at the 1% level), which indicates that the MAT measure is successful in eliminating a large part of the bias induced by the AT measure. Finally, in Table 5 Panel D we split the sample into high and low *BIDASK* as we did with the *PIN* score in Table 4 Panel D. The SCV measure for the high *BIDASK* group (0.0021) is almost double that for the low *BIDASK* group (0.0011), with a statistically significant difference at the 5% level.

[Insert Table 5 here]

Overall, the findings across Table 4 and Table 5 are consistent with hypothesis H2 that the MAT and SCV measures might conflict with inferences drawn from the AT measure in cross-sectional settings. Nevertheless, our evidence suggests that the issue is empirical rather than theoretical, and that with the right choice of proxy variables and CC measures, we are able to confirm that higher values of information asymmetry in equity markets drive the demand for more conservative financial reporting.

***3.4. Ball, Robin, & Sadka (2008) (cross-sectional setting)***

Ball et al. (2008) examine whether debt market or equity market orientation constitutes the primary source of demand for timely loss recognition. The empirical findings of Ball et al. (2008) are important because they support the view that CC primarily caters for the information demands of debt investors rather than that of equity investors, and this in turn addresses fundamental issues in the accounting literature regarding the desirable properties of financial statement information (Beyer et al., 2010; Kothari, 2001).

Ball et al. (2008) generate country-level AT coefficient estimates by running the Basu (1997) piecewise linear regression separately for 22 countries over the years 1992-2003. They then run cross-sectional regressions of their 22 AT estimates on country-level proxies for the importance of the debt market and the equity market as their main explanatory variables, along with controls for other country-level characteristics. To re-examine their analyses, we use accounting data from the Compustat Global fundamental annual file and stock returns data from the Compustat Global security file for the selected 22 countries.[13] We delete firms that are cross-listed or belong to the financial and utility sectors, we drop observations in the top and bottom 1% of the earnings and returns distributions, and we delete observations with negative book value of equity. The final sample comprises 95,347 firm-year observations.

Table 6 reports summary statistics for this set of analyses, with Panel A reporting statistics for all variables used to compute the CC measures across the full international sample and Panel B reporting country-level variables, including the AT (*B3*), MAT (*Modified B3*), and SCV measures as well as the control variables. All variables are defined in Table 1.

[Insert Table 6 here]

Table 7 presents the replication and re-examination results of Table 5 in Ball et al. (2008). In Table 7 Panel A, we apply the estimates from the AT, MAT, and SCV measures as the dependent variables, i.e., the country-level estimated values of *B3*, *Modified B3*, and SCV. The first three columns use *B3*, the middle three columns use *Modified B3*, and the last three columns use SCV. Note that for the first three columns where *B3* is the dependent variable, the coefficient on *DEBT* is significantly positive while the coefficient on *EQUITY* is negative and mostly significant. These findings are consistent with those reported in Table 5 of Ball et al. (2008), which

[13] Ball et al. (2008) use the Global Vantage database, which has been succeeded by Compustat Global.

they interpret to imply that the demand for CC is driven more by the needs of debt rather than equity investors. On the contrary, when replacing the dependent variable *B3* either by the *Modified B3* or *SCV*, the significant coefficient on *DEBT* disappears. Notably, the regression results using the MAT measure suggest that the rule of law (*LAW*) is the main determinant of CC at the country level, which is in the spirit of Ball et al. (2000).

To examine whether the results of using a country-level AT regression hold when using a firm-level regression, we first split the sample into high and low importance of debt markets using the indicator variable *HiDebt* that takes the value 1 if the country's importance of debt is higher than the median value of the sample (i.e., the eleven countries with higher importance of debt take the value 1), and then run the AT regression while including *HiDebt* interactively. Table 7 Panel B reports the results of the AT piecewise linear regression while conditioning on *HiDebt* and shows that the coefficient on the triple interaction *HiDebt*×*RD*×*RET* is 0.1058, significant at the 1% level. Moving to Table 7 Panel C, we employ the MAT measure while conditioning on *HiDebt* and find statistical significance on the triple interaction *HiDebt*×*RD*×*RET* yet with almost half the economic magnitude exhibited in the AT regression. This finding is consistent with the contracting theory, yet it shows that the AT measure is sensitive to the choice of research design in cross-sectional settings. Specifically, when using the AT measure in a country-level regression with equal weights, the results contradict those from the MAT measure. However, when allowing for firm-level heterogeneity, the results from the AT and MAT measures coincide yet with different economic significance due to the AT bias. Finally, we examine the difference in the SCV measure between high and low importance of debt countries at the firm-level in Table 7 Panel D. The SCV values of both groups do not show a statistically significant difference, and that is probably due to the lack of power and reliability for the SCV measure when comparing two groups with different statistical attributes of earnings and returns, an issue that we discuss in the Section 4.

[Insert Table 7 here]

The findings presented in Table 7 are consistent with our prediction in hypothesis H2 that the AT measure on one hand, and the MAT and SCV measures on the other hand, are likely to yield dissimilar inferences in cross-sectional settings. The consistency in inferences obtained from our re-examination of Ball et al. (2008) and LaFond & Watts (2008) provides a powerful mutual robustness check across two independent studies.

**4. Discussion and Conclusions**

Given the importance of CC as a research topic in the accounting literature, and the on-going debate over the validity of the widely adopted AT measure of Basu (1997), we believe that it is necessary to heed the calls of Ball (2016) and Hail et al. (2020) to revisit inferences drawn from previous studies using more recently developed research methodologies. As such, unlike previous studies involved in the debate over the validity of the AT measure, we contribute to the CC literature by providing direct evidence on whether and how bias in the AT measure varies across frequently adopted research settings. On the one hand, we provide evidence through the re-examination of Lobo & Zhou (2006) and Ahmed et al. (2013) that inferences based on the AT measure are likely to hold for interrupted time-series settings that examine the impact of exogenous changes in accounting policies. On the other hand, we provide evidence through the re-assessment of LaFond & Watts (2008) and Ball et al. (2008) that the AT measure is sensitive to test specifications and research designs in cross-sectional settings, where the AT measure is likely to generate biased inferences when identifying determinants of CC.

Based on our findings, the MAT measure of Badia et al. (2021) yields robust empirical performance in line with theory. On the one hand, the MAT measure preserves the favorable empirical attributes associated with the AT construct. On the other hand, it is less susceptible to the sources of bias that affects the AT measure as documented in the literature. Regarding the SCV measure of Dutta & Patatoukas (2017), despite its theoretical merit, we believe that there are certain empirical limitations that need to be addressed or at least acknowledged. First, the SCV measure is inefficient when comparing CC between two groups of firms with different statistical attributes of earnings and returns since SCV is a spread and not a ratio. Using the ratio of conditional variances (RCV) instead of the spread also has its downsides and does not completely solve the issue.[14] Second, the SCV measure shows high sensitivity to outliers to the extent that it might be considerably affected by 0.1% of the sample distribution, which makes it yield negative values in some cases. As such, accounting for extreme (influential) observations using regular data management procedures such as winorization or truncation might not be suitable with the SCV measure; instead, more sophisticated procedures may have to be applied, such as the influence diagnostics or robust regression (Leone, Minutti-Meza, & Wasley, 2019). As a result, it is likely the case that the SCV measure lacks power and is prone to Type 2 error, especially when comparing CC across groups with high cross-sectional heterogeneity. Third, the SCV measure cannot be incorporated in a regression analysis that examines the effect of CC on other outcome

[14] To illustrate, assume that a researcher is interested in comparing the level of CC exhibited by publicly listed firms in two different countries with different economic environments and institutional settings. Country A has *Variance (X | RET<0)* = 0.09 and *Variance (X | RET>0)* = 0.06, while Country B has *Variance (X | RET<0)* = 0.02 and *Variance (X | RET>0)* = 0.01; the SCV measure for Country A is 0.03 while that of Country B is 0.01, suggesting that Country A exhibits a higher level of CC. However, that is an unreliable conclusion given that the conditional variances of earnings in both countries are incomparable in magnitude due to high cross-sectional heterogeneity, and thus the ratio of conditional variances (RCV) should be used. When computing the RCV measure for Country A and Country B, the values are respectively 1.5 and 2, resulting in a contradicting conclusion to that inferred from using SCV.

variables. In the same sense, the SCV measure cannot be estimated at the firm-level, and thus cannot incorporate relevant variables that control for CC in a regression analysis.

Our study has important implications for both past and future applications of the AT construct. For past applications, we suggest that prior claims based on cross-sectional applications that rely on the AT measure need to be reworked using more developed measures of CC, like the MAT measure of Badia et al. (2021), that are less likely to suffer from the documented biases associated with the AT measure. For future studies, we highlight the importance of ensuring that the choice of CC measure, along with other research design choices, need to carefully control for potential cross-sectional correlation between the economic characteristics of firms (especially those related to economic risk and option values) and the true (unobservable) level of CC. Furthermore, whilst improved CC measures may improve inferences, like MAT and SCV, it is likely that such measures will need to be integrated into causal research designs (such as difference-in-differences and natural experiments) in order to justify claims about the causes of CC.

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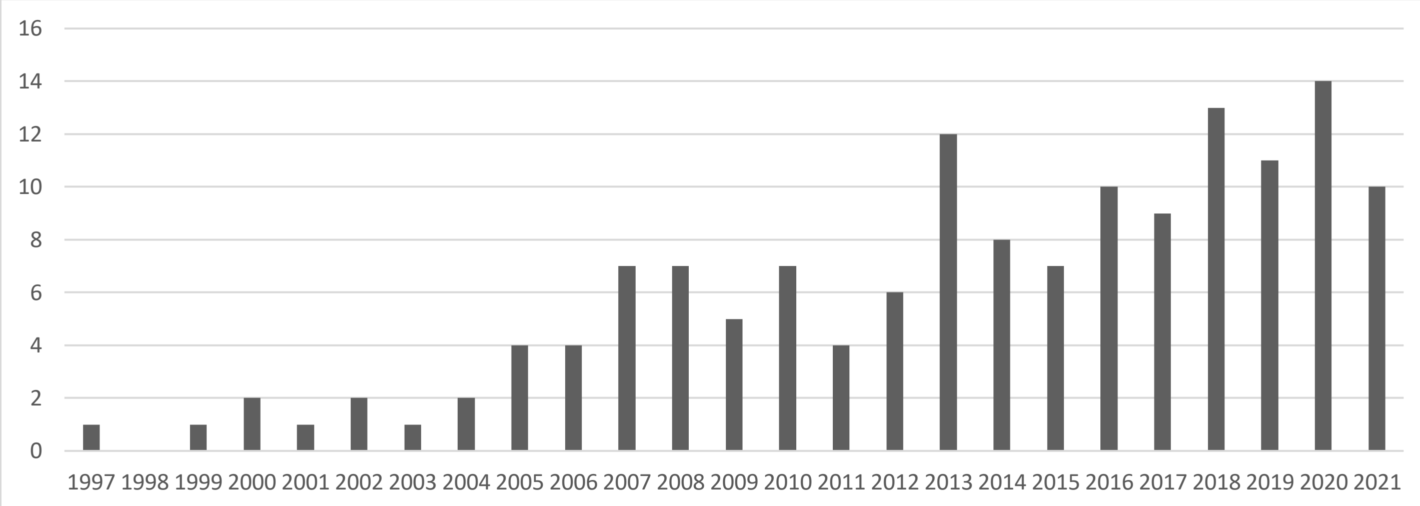
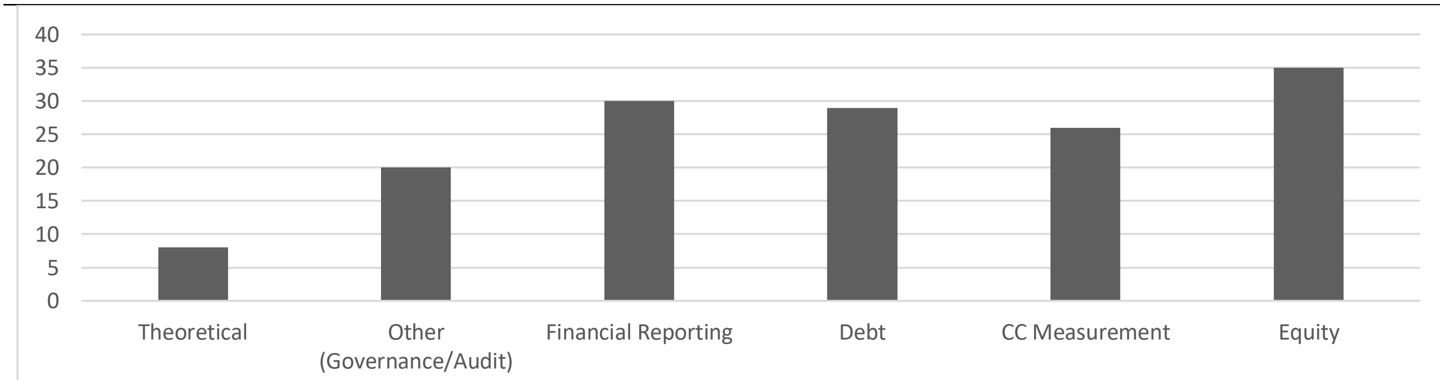
## Appendices

**Appendix A:** Literature analysis

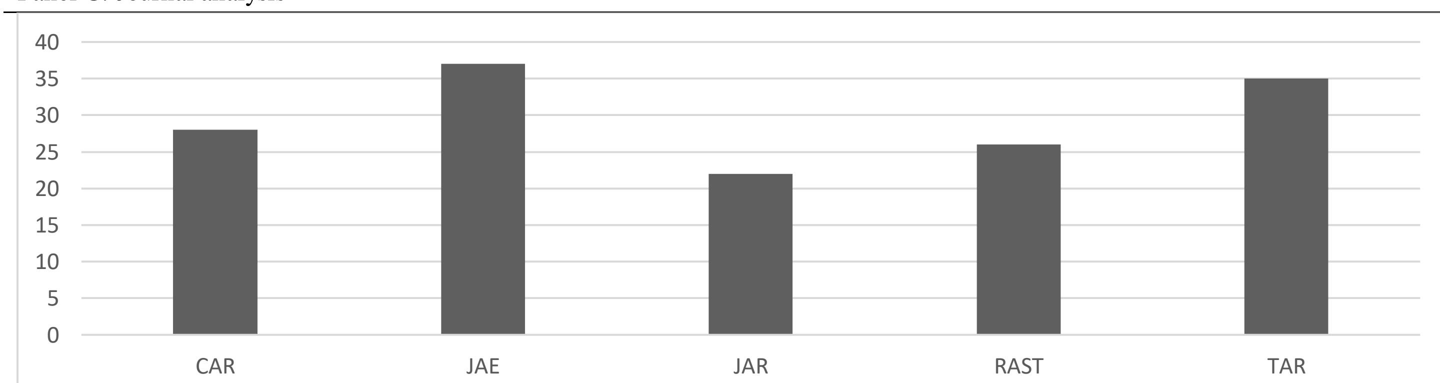
Panel A: Time trend analysis




Panel B: Topic analysis




Panel C: Journal analysis




Notes: This appendix presents the literature analysis of conditional conservatism (CC) studies that applied the Basu (1997) asymmetric timeliness (AT) approach (including the C_Score of Khan & Watts, 2009) either in their empirical analyses or theoretical models. We cover studies published over the period of 1997 to 2021 across five leading accounting journals, including (in alphabetical order) *Contemporary Accounting Research* (CAR), *Journal of Accounting and Economics* (JAE), *Journal of Accounting Research* (JAR), *Review of Accounting Studies* (RAST), and *The Accounting Review* (TAR). We exclude discussion papers associated with some of the studies. The vertical axis in each panel indicates the number of publications. Panels A, B, and C depict the number of publications by year, topic, and journal respectively.

**Appendix B:** Top 10 papers that cited each of the reexamined studies

| Paper: | Citations |
|---|---|
| Panel A: Lobo & Zhou (2006) | 693 |
| Cohen, Dey, & Lys, (2008), *The Accounting Review.* | 3435 |
| Leuz & Wysocki (2016), *Journal of Accounting Research.* | 1145 |
| Krishnan & Visvanathan (2008), *Contemporary Accounting Research.* | 765 |
| García Lara, García Osma, & Penalva (2009), *Review of Accounting Studies.* | 736 |
| Francis, Hasan, Park, & Wu (2015), *Contemporary Accounting Research.* | 479 |
| Lennox & Pittman (2010), *Journal of Accounting and Economics.* | 345 |
| DeFond (2010), *Journal of Accounting and Economics.* | 298 |
| Doukakis (2014), *Journal of Accounting and Public Policy.* | 260 |
| García Lara, García Osma, & Penalva (2007), *European Accounting Review.* | 258 |
| Barua, Lin, & Sbaraglia (2010), *The Accounting Review.* | 209 |
| Panel B: Ahmed, Neel, & Wang (2013) | 904 |
| Leuz & Wysocki (2016), *Journal of Accounting Research.* | 1145 |
| Christensen, Hail, & Leuz (2013), *Journal of Accounting and Economics.* | 913 |
| Horton, Serafeim, & Serafeim (2013), *Contemporary Accounting Research.* | 834 |
| Brüggemann, Hitz, & Sellhorn (2013), *European Accounting Review.* | 549 |
| De George, Li, & Shivakumar (2016), *Review of Accounting Studies.* | 371 |
| DeFond, Hung, Li, & Li (2015), *The Accounting Review.* | 353 |
| Lang & Stice-Lawrence (2015), *Journal of Accounting & Economics.* | 339 |
| Cascino & Gassen (2015), *Review of Accounting Studies.* | 285 |
| Doukakis (2014), *Journal of Accounting and Public Policy.* | 260 |
| Ball, Li, & Shivakumar (2015), *Journal of Accounting Research.* | 236 |

*Continued next page*

| Panel C: LaFond & Watts (2008) | 1539 |
|---|---|
| DeFond & Zhang (2014), *Journal of Accounting and Economics*. | 2232 |
| Khan & Watts (2009), *Journal of Accounting and Economics*. | 1522 |
| Armstrong, Guay, & Weber (2010), *Journal of Accounting and Economics*. | 1477 |
| Francis & Wang (2008), *Contemporary Accounting Research*. | 1268 |
| Ahmed & Duellman (2007), *Journal of Accounting and Economics*. | 1166 |
| LaFond & Roychowdhury (2008), *Journal of Accounting Research*. | 964 |
| Kothari, Ramanna, & Skinner (2010), *Journal of Accounting and Economics*. | 764 |
| Ahmed & Duellman (2013), *Journal of Accounting Research*. | 698 |
| Beatty, Weber, & Yu (2008), *Journal of Accounting and Economics*. | 640 |
| Panel D: Ball, Robin, & Sadka (2008) | 674 |
| Armstrong, Guay, & Weber (2010), *Journal of Accounting and Economics*. | 1477 |
| Nikolaev (2010), *Journal of Accounting Research*. | 533 |
| Francis & Martin (2010), *Journal of Accounting and Economics*. | 464 |
| Kim, Song, & Zhang (2011), *The Accounting Review*. | 448 |
| Bushman, Piotroski, & Smith (2011), *Journal of Business Finance & Accounting*. | 421 |
| Ball & Shivakumar (2008), *Journal of Accounting Research*. | 420 |
| Gigler, Kanodia, Sapra, & Vengopalan (2009), *Journal of Accounting Research*. | 420 |
| Ball, Kothari, & Nikolaev (2013), *Journal of Accounting Research*. | 331 |
| García Lara, García Osma, & Penalva (2011), *Review of Accounting Studies*. | 273 |

Notes: This appendix lists the top 10 papers that cited each of the four re-examined studies based on Google Scholar citation counts as of 31/03/2022. The order of listing in this appendix is based on the descending order of citations.

## Tables and Figures

**Table 1:** Variable definitions (sorted alphabetically)

| Variable | Definition |
|---|---|
| Common variables across the four re-examined studies | |
| *MTB* | Market-to-book ratio of assets at the beginning of the year. |
| *RD* | Bad news dummy variable that equals 1 if *RET* is negative, and 0 otherwise. |
| *RET* | Abnormal stock returns over the fiscal year, calculated as monthly compounded buy-and-hold observed returns (including distributions) minus expected returns. Expected returns are the average returns of 5×5 portfolios formed by sorting firms first based on the beginning of year market value of equity and then based on beginning of year book-to-market equity ratio. |
| *RVAR* | Variance of daily stock returns during the fiscal year. |
| SCV | SCV measure of CC estimated as the variance of bad news earnings minus the variance of good news earnings. SCV = $(X_{it} \mid RD=1) - (X_{it} \mid RD=0)$. |
| | |
| *UX* | Unexpected earnings, computed as the residuals from the following regression model (estimated by industry-year, based on the Fama and French twelve industry classification): $X_{it} = \alpha_0 + \alpha_1 X_{it-1} + \alpha_2 I(X_{it-1}<0) + \alpha_3 X_{it-1} \times I(X_{it-1}<0) + \varepsilon_{it}$, where $I(X_{it-1}<0)$ equals 1 if $X_{it-1}<0$, and 0 otherwise. |
| *X* | Income before extraordinary items deflated by lagged market value of equity. |
| Variables specific to Lobo & Zhou (2006) | |
| *SOX* | SOX indicator that equals 1 for fiscal years ending in August 2002 or beyond, and 0 otherwise. |
| Variables specific to Ahmed, Neel, & Wang (2013) | |
| *BTM* | Ratio of book value of equity to market value of equity. |
| *LEV* | Total liabilities divided by book value of equity. |
| *Post* | Dummy variable that takes the value 1 if the year is greater than 2005, and zero otherwise. |
| *SIZE* | Natural log of market capitalization. |
| *Treat* | Dummy variable that takes the value 1 if the firm is listed in a country that mandated the adoption of IFRS in 2005, and zero otherwise. |

*Continued next page*

**Table 1:** *(continued from previous page)*

| | |
|---|---|
| Variables specific to LaFond & Watts (2008) | |
| *BIDASK* | Median of daily percentage bid-ask spread, calculated the ask price minus the bid price, divided by the average of bid and ask prices. |
| *HiBIDASK* | High *BIDASK* indicator that equals 1 for observations with a *BIDASK* value above the industry-year median value, based on the Fama and French twelve industry classification, and 0 otherwise |
| *HiPIN* | High *PIN* indicator that equals 1 for observations with *PIN* score above the industry-year median value, based on the Fama and French twelve industry classification, and 0 otherwise. |
| *PIN* | the probability of an information-based trade derived from a structural market microstructure model (see Easley, Hvidkjaer, & O'Hara, 2002). |
| Variables specific to Ball, Robin, & Sadka (2008) | |
| *B3* | Country-level coefficient estimate on *RD*×*RET* in the Basu (1997) AT regression. |
| *BTM* | Country-level book-to-market calculated as the median for all firm and years in each country. |
| *CORRP* | Country-level corruption index based on La Porta, Lopez-De-Silanes, Shleifer, and Vishny (1997, 1998). |
| *CRED* | Country-level creditor rights index based on La Porta, Lopez-De-Silanes, Shleifer, and Vishny (1997, 1998). |
| *DEBT* | Country-level debt market importance, calculated as the sum of bank debt of the private sector and outstanding non-financial bonds divided by GNP in 1994, based on La Porta, Lopez-De-Silanes, Shleifer, and Vishny (1997, 1998). |
| *ENGLISH* | English legal origin indicator that equals 1 for such countries, and 0 otherwise. |
| *EQUITY* | Country-level equity market importance, calculated as stock market capitalization held by minorities divided by GNP in 1994, based on La Porta, Lopez-De-Silanes, Shleifer, and Vishny (1997, 1998). |
| *FRENCH* | French legal origin indicator that equals 1 for such countries, and 0 otherwise. |
| *HiDebt* | High debt market importance indicator that equals 1 for countries with *DEBT* value above sample median, and 0 otherwise. |
| *LAW* | Country-level law and order index based on La Porta, Lopez-De-Silanes, Shleifer, and Vishny (1997, 1998). |
| *Modified B3* | Country-level coefficient estimate on *RD*×*RET* in the Badia (2021) MAT regression. |
| *SCAND* | Scandinavian legal origin indicator that equals 1 for such countries, and 0 otherwise. |

Notes: This table presents the definition of all variables applied in our analyses. We make adjustments whenever necessary and appropriate in order to maximize comparability across studies and CC measures.

**Table 2: Lobo & Zhou (2006) inference re-examination (test of hypothesis H1)**

Panel A: Summary Statistics

| | N | Mean | S.D. | Q1 | Median | Q3 |
|---|---|---|---|---|---|---|
| *X* | 11,244 | 0.0150 | 0.1602 | −0.0054 | 0.0472 | 0.0856 |
| *UX* | 11,244 | 0.0049 | 0.1116 | −0.0213 | 0.0089 | 0.0443 |
| *RET* | 11,244 | 0.0128 | 0.6170 | −0.3398 | −0.0738 | 0.2119 |
| *MTB* | 11,244 | 1.7919 | 1.6585 | 1.0297 | 1.2264 | 1.8583 |
| *RVAR* | 11,244 | 0.0427 | 0.3489 | 0.0005 | 0.0010 | 0.0024 |

Panel B: The SOX effect on CC through the AT construct using an OLS regression

| *X* | *RD* | *RET* | *RD*×*RET* | *SOX* | *RD*×*SOX* | *RET*×*SOX* | *RD*×*RET*×*SOX* | |
|---|---|---|---|---|---|---|---|---|
| | 0.0176*** | −0.0043 | 0.2289*** | −0.0115** | −0.0014 | 0.0028 | 0.0436*** | Adj. $R^2$ = 12.93% |
| | (2.94) | (−0.94) | (20.75) | (−2.11) | (−0.17) | (0.41) | (2.63) | N = 11,244 |

Panel C: The SOX effect on CC through the MAT construct using a firm-fixed-effects regression

| *UX* | *RD* | *RET* | *RD*×*RET* | *SOX* | *RD*×*SOX* | *RET*×*SOX* | *RD*×*RET*×*SOX* | |
|---|---|---|---|---|---|---|---|---|
| | 0.0026 | 0.0802*** | 0.0433*** | 0.0130*** | 0.0075 | −0.0217*** | 0.0672*** | Within $R^2$ = 8.16% |
| | (0.32) | (12.00) | (2.65) | (2.59) | (0.94) | (−3.08) | (4.05) | N = 11,244 |

additional variables included are *MTB* and *RVAR*, and their interactions with *RD*, *RET*, and *RD*×*RET*

Panel D: Comparison of the SCV measure values in pre- and post-SOX periods

| Groups | Observations | Average *X* | Std. Dev. |
|---|---|---|---|
| *SOX* = 0, *RD* = 0 | 2,380 | 0.0582 | 0.1509 |
| *SOX* = 0, *RD* = 1 | 3,242 | −0.0050 | 0.1628 |
| *SOX* = 1, *RD* = 0 | 2,459 | 0.0483 | 0.1424 |
| *SOX* = 1, *RD* = 1 | 3,163 | −0.0230 | 0.1647 |

SCV | (*SOX* = 0) = 0.0037 ΔSCV = 0.0031**

SCV | (*SOX* = 1) = 0.0068 $Chi^2$ = 5.73; *p*-value = 0.016

Notes: This table presents replication and re-examination of the inference from Lobo & Zhou (2006) for our analysis to test hypothesis H1 based on an interrupted time-series setting. Panel A reports summary statistics for the variables used in the analysis based on a U.S. sample over the period of 2000-2004. Panel B reports the conditioning effect of SOX through the Basu (1997) AT regression with *X* as the dependent variable. Panel C reports the conditioning effect of SOX through the Badia et al. (2021) MAT regression with *UX* as the dependent variable. Panel D reports the summary statistics of *X* across four observation groups sorted on *SOX* and *RD,* the SCV measures in the pre- and post-SOX periods, and statistical significance of the difference in SCV between both periods. All variables are defined in Table 1. Reported *t*-statistics in parentheses are calculated based on clustered standard errors at the firm level. *, **, *** denote significance at the 10%, 5%, and 1% levels, respectively.

**Table 3: Ahmed, Neel, and Wang (2013) inference re-examination (test of hypothesis H1)**

Panel A: Summary statistics

| | Treatment group | | | | | | | Control group | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | Obs. | Mean | Std. Dev. | Q1 | Median | Q3 | | Obs. | Mean | Std. Dev. | Q1 | Median | Q3 |
| *X* | 6,238 | 0.0059 | 0.1967 | −0.0134 | 0.0412 | 0.0750 | *X* | 6,238 | 0.0166 | 0.1581 | −0.0173 | 0.0452 | 0.0879 |
| *UX* | 6,238 | 0.0007 | 0.1422 | −0.0304 | 0.0051 | 0.0480 | *UX* | 6,238 | 0.0001 | 0.1939 | −0.0322 | 0.0071 | 0.0597 |
| *RET* | 6,238 | 0.0376 | 0.6153 | −0.3243 | −0.0756 | 0.2368 | *RET* | 6,238 | 0.0427 | 0.4179 | −0.1891 | 0.0287 | 0.2583 |
| *BTM* | 6,238 | 0.6632 | 0.6048 | 0.3377 | 0.5463 | 0.8376 | *BTM* | 6,238 | 0.6633 | 0.6047 | 0.3383 | 0.5472 | 0.8369 |
| *RVAR* | 6,238 | 0.0046 | 0.0189 | 0.0005 | 0.0010 | 0.0020 | *RVAR* | 6,238 | 0.0153 | 0.3936 | 0.0003 | 0.0007 | 0.0017 |

Panel B: IFRS effect on CC through the AT construct based on a difference-in-differences research design and using an OLS regression

| *X* | *RET* | *RD* | *RET×RD* | *RET×RD×Post* | *RET×RD×Treat* | *RET×RD×Post×Treat* | |
|---|---|---|---|---|---|---|---|
| | −0.0743$^{*}$ | −0.0851$^{***}$ | 0.1790$^{***}$ | 0.0234 | 0.1670$^{***}$ | −0.1564$^{***}$ | Adj. $R^2$ = 13.26% |
| | (−1.75) | (−2.73) | (2.86) | (0.62) | (3.58) | (−2.77) | N = 12,476 |

Control variables: *SIZE, LEV, BTM*, and their interactions with *RET, RD, and RET×RD.*

Panel C: IFRS effect on CC through the MAT construct based on a difference-in-differences research design and using a firm-fixed-effects regression

| *UX* | *RET* | *RD* | *RET×RD* | *RET×RD×Post* | *RET×RD×Treat* | *RET×RD×Post×Treat* | |
|---|---|---|---|---|---|---|---|
| | −0.0343 | −0.1314$^{***}$ | 0.1207$^{***}$ | 0.0308 | 0.1138$^{***}$ | −0.1202$^{**}$ | Within $R^2$ = 7.11% |
| | (−1.56) | (−5.84) | (2.87) | (0.88) | (3.60) | (−2.12) | N = 12,476 |

Control variables: *SIZE, LEV, BTM, RVAR* and their interactions with *RET, RD, and RET×RD.*

Panel D: Comparison of the SCV values in pre- and post-IFRS periods for the Treatment and Control groups

| Treatment group | | | | Control group | | | |
|---|---|---|---|---|---|---|---|
| Group | Obs. | Average *X* | Std. Dev. | Group | Obs. | Average *X* | Std. Dev. |
| *Post = 0, RD = 0* | 2,553 | 0.0216 | 0.2279 | *Post = 0, RD = 0* | 2,291 | 0.0428 | 0.153 |
| *Post = 0, RD = 1* | 1,333 | −0.0558 | 0.2406 | *Post = 0, RD = 1* | 1,473 | −0.0490 | 0.1734 |
| *Post = 1, RD = 0* | 1,268 | 0.0455 | 0.1036 | *Post = 1, RD = 0* | 1,585 | 0.0619 | 0.1288 |
| *Post = 1, RD = 1* | 1,084 | −0.0017 | 0.0969 | *Post = 1, RD = 1* | 889 | −0.0247 | 0.1482 |
| SCV \| (*Post* = 0) = 0.0059 | | ΔSCV = −0.0072$^{***}$ | | SCV \| (*Post* = 0) = 0.0067 | | ΔSCV = −0.0013 | |
| SCV \| (*Post* = 1) = −0.0013 | | $Chi^2$ = 7.04; *p*-value = 0.008 | | SCV \| (*Post* = 1) = 0.0054 | | $Chi^2$ = 0.52; *p*-value = 0.470 | |

Notes: This table presents replication and re-examination of the inference from Ahmed, Neel, & Wang (2013) for our analysis to test hypothesis H1 based on an interrupted time-series setting. Panel A reports summary statistics for the variables used in the analysis based on an international sample over the period of 2002-2007 (excluding 2005). Panel B reports the conditioning effect of IFRS adoption through the Basu (1997) AT regression with *X* as the dependent variable. Panel C reports the conditioning effect of IFRS adoption through the Badia et al. (2021) MAT regression with *UX* as the dependent variable. For brevity and exposition purposes, we only report the coefficients on the variables of interest. The full regression is based on Equation (6) in Ahmed et al. (2013) is: $X = \alpha_0 + RET + RD + RET\times RD + Post + RET\times Post + RD\times Post + RET\times RD\times Post + Treat + RET\times Treat + RD\times Treat + RET\times RD\times Treat + Post\times Treat + RET\times Post\times Treat + RD\times Post\times Treat + RET\times RD\times Post\times Treat + BTM + RET\times BTM + RD\times BTM + RET\times RD\times BTM + LEV + RET\times LEV + RD\times LEV + RET\times RD\times LEV + SIZE + RET\times SIZE + RD\times SIZE + RET\times RD\times SIZE + \varepsilon$, where $\varepsilon$ is a random error and all other variables are defined in Table 1. Panel D reports the summary statistics in terms of *X* across four observation groups sorted on *Post* and *RD*, the SCV measure in the pre- and post-IFRS periods, for the treatment and control groups separately, and statistical significance of the difference in SCV between both periods. All variables are defined in Table 1. Reported *t*-statistics in parentheses are calculated based on clustered standard errors at the firm level. $^{*}$, $^{**}$, $^{***}$ denote significance at the 10%, 5%, and 1% levels, respectively.

**Table 4: LaFond & Watts (2006) inference re-examination using the *PIN* score (test of hypothesis H2)**

Panel A: Summary Statistics

| | N | Mean | S.D. | Q1 | Median | Q3 |
|---|---|---|---|---|---|---|
| *X* | 19,640 | 0.0518 | 0.1049 | 0.0238 | 0.0616 | 0.0969 |
| *UX* | 19,640 | 0.0018 | 0.1126 | -0.0178 | 0.0072 | 0.0340 |
| *RET* | 19,640 | 0.0000 | 0.4024 | -0.2361 | -0.0432 | 0.1658 |
| *MTB* | 19,640 | 1.5802 | 0.9670 | 1.0525 | 1.2536 | 1.7106 |
| *RVAR* | 19,640 | 0.0294 | 0.4001 | 0.0002 | 0.0004 | 0.0009 |
| *PIN* | 19,640 | 0.1951 | 0.0767 | 0.1391 | 0.1818 | 0.2346 |

Panel B: The *PIN* effect on CC through the AT construct using an OLS regression

| *X* | *RD* | *RET* | *RD*×*RET* | *PIN* | *PIN*×*RD* | *PIN*×*RET* | *PIN*×*RD*×*RET* | |
|---|---|---|---|---|---|---|---|---|
| | 0.0110* | 0.0103 | 0.0895*** | 0.0531** | −0.0005 | 0.0839 | 0.5986*** | Adj. $R^2$ = 18.93% |
| | (1.85) | (0.64) | (3.98) | (2.26) | (−0.02) | (1.17) | (5.61) | N = 19,640 |

Panel C: The *PIN* effect on CC through the MAT construct using a firm-fixed-effects regression

| *UX* | *RD* | *RET* | *RD*×*RET* | *PIN* | *PIN*×*RD* | *PIN*×*RET* | *PIN*×*RD*×*RET* | |
|---|---|---|---|---|---|---|---|---|
| | 0.0001 | −0.0052 | 0.2069*** | 0.0324 | 0.0232 | 0.2087*** | −0.0712 | Within $R^2$ = 9.30% |
| | (0.01) | (−0.32) | (6.52) | (0.96) | (0.55) | (3.59) | (−0.59) | N = 19,640 |
| | additional variables included are *MTB* and *RVAR*, and their interactions with *RD*, *RET*, and *RD*×*RET* | | | | | | | |

Panel D: Comaprison of SCV across high and low *PIN* groups

| Group | Obs. | Average *X* | Std. dev. |
|---|---|---|---|
| *HiPIN* = 0, *RD* = 0 | 4,368 | 0.0795 | 0.0752 |
| *HiPIN* = 0, *RD* = 1 | 5,457 | 0.0416 | 0.0828 |
| *HiPIN* = 1, *RD* = 0 | 4,261 | 0.0877 | 0.1134 |
| *HiPIN* = 1, *RD* = 1 | 5,554 | 0.0125 | 0.1211 |
| | SCV \| (*HiPIN* = 0) = 0.0012 | ΔSCV = 0.0005 | |
| | SCV \| (*HiPIN* = 1) = 0.0017 | $Chi^2$ = 1.83 ; *p*-value = 0.176 | |

Notes: This table presents replication and re-examination of the inference from LaFond & Watts (2008) for our analysis to test hypothesis H2 based on a cross-sectional setting. Panel A reports summary statistics for the variables used in the analysis based on a U.S. sample over the period of 1983-2001. Panel B reports the conditioning effect of *PIN* through the Basu (1997) AT regression with *X* as the dependent variable. Panel C reports the conditioning effect of *PIN* through the Badia et al. (2021) MAT regression with *UX* as the dependent variable. Panel D reports the summary statistics in terms of *X* across four observation groups sorted on *HiPIN* and *RD*, the SCV measure across high and low *PIN* groups, and statistical significance of the difference in SCV between both groups. All variables are defined in Table 1. Reported *t*-statistics in parentheses are calculated based on clustered standard errors at the firm level. *, **, *** denote significance at the 10%, 5%, and 1% levels, respectively

**Table 5: LaFond & Watts (2006) inference re-examination using the bid-ask spread (test of hypothesis H2)**

Panel A: Summary Statistics

| | N | Mean | S.D. | Q1 | Median | Q3 |
|---|---|---|---|---|---|---|
| *BIDASK* | 19,640 | 0.0282 | 0.0219 | 0.0170 | 0.0230 | 0.0327 |

Panel B: The bid-ask spread effect on CC through the AT construct using an OLS regression

| *X* | *RD* | *RET* | *RD*×*RET* | *BIDASK* | *BIDASK*×*RD* | *BIDASK*×*RET* | *BIDASK*×*RD*×*RET* | |
|---|---|---|---|---|---|---|---|---|
| | 0.0229*** | 0.0787*** | 0.0256 | −1.0680*** | −0.7341*** | −1.0870*** | 4.0723*** | Adj. $R^2$ = 15.35% |
| | (2.80) | (6.41) | (1.11) | (−6.05) | (−2.96) | (−4.74) | (9.05) | N = 19,640 |

Panel C: The bid-ask spread effect on CC through the MAT construct using a firm-fixed-effects regression

| *UX* | *RD* | *RET* | *RD*×*RET* | *BIDASK* | *BIDASK*×*RD* | *BIDASK*×*RET* | *BIDASK*×*RD*×*RET* | |
|---|---|---|---|---|---|---|---|---|
| | 0.0127* | 0.0814*** | 0.1267*** | 0.0491 | −0.2217 | −0.6867*** | 1.1050*** | Within $R^2$ = 9.44% |
| | (1.81) | (7.87) | (6.27) | (0.39) | (−1.41) | (−4.86) | (3.76) | N = 19,640 |

additional variables included are *MTB* and *RVAR*, and their interactions with *RD*, *RET*, and *RD*×*RET*

Panel D: Comaprison of SCV across high and low *BIDASK* groups

| Group | Obs. | Average *X* | Std. dev. |
|---|---|---|---|
| *HiBIDASK* = 0, *RD* = 0 | 4,190 | 0.0755 | 0.1100 |
| *HiBIDASK* = 0, *RD* = 1 | 5,843 | 0.0114 | 0.1149 |
| *HiBIDASK* = 1, *RD* = 0 | 4,439 | 0.0910 | 0.0768 |
| *HiBIDASK* = 1, *RD* = 1 | 5,168 | 0.0443 | 0.0896 |

| | |
|---|---|
| SCV \| (*HiBIDASK* = 0) = 0.0011 | ΔSCV = 0.0010** |
| SCV \| (*HiBIDASK* = 1) = 0.0021 | $Chi^2$ = 6.01 ; *p*-value = 0.014 |

Notes: This table presents the re-examination of the inference from LaFond & Watts (2008) for our analysis to test hypothesis H2 based on a cross-sectional setting while replacing the *PIN* score with the bid-ask spread (*BIDASK*) as a proxy for information asymmetry. Panel A reports summary statistics for the bid-ask spread (*BIDASK*) used in the analysis based on a U.S. sample over the period of 1983-2001. Panel B reports the conditioning effect of *BIDASK* through the Basu (1997) AT regression with *X* as the dependent variable. Panel C reports the conditioning effect of *BIDASK* through the Badia et al. (2021) MAT regression with *UX* as the dependent variable. Panel D reports the summary statistics in terms of *X* across four observation groups sorted on *HiBIDASK* and *RD*, the SCV measure across high and low *BIDASK* groups, and statistical significance of the difference in SCV between both groups. All variables are defined in Table 1. Reported *t*-statistics in parentheses are calculated based on clustered standard errors at the firm level. *, **, *** denote significance at the 10%, 5%, and 1% levels, respectively

**Table 6: Summary statistics for Ball, Robin, & Sadka (2008) inference re-examination (test of hypothesis H2)**

Panel A: Summary Statistics

| | N | Mean | S.D. | Q1 | Median | Q3 |
|---|---|---|---|---|---|---|
| *X* | 95,347 | 0.0058 | 0.2450 | −0.0059 | 0.0368 | 0.0789 |
| *UX* | 95,347 | 0.0012 | 0.1613 | −0.0224 | 0.0095 | 0.0476 |
| *MTB* | 95,347 | 2.3506 | 11.2779 | 0.9500 | 1.2096 | 1.7515 |
| *RVAR* | 95,347 | 0.0515 | 0.1563 | 0.0008 | 0.0027 | 0.0420 |
| *RET* | 95,347 | −0.0057 | 0.4911 | −0.2788 | −0.0224 | 0.2211 |

Panel B: Extract from the Full Dataset Used to Replicate Ball et al. (2008)

| Country | *B3* | *Modified B3* | SCV | *DEBT* | *EQUITY* | *LAW* | *CORRUPT* | *CREDIT* | *BTM* |
|---|---|---|---|---|---|---|---|---|---|
| Australia | 0.272 | 0.201 | 0.056 | 0.76 | 0.49 | 10.00 | 8.52 | 1 | 0.646 |
| Canada | 0.293 | 0.160 | 0.038 | 0.72 | 0.39 | 10.00 | 10.00 | 1 | 0.657 |
| Malaysia | 0.160 | 0.109 | 0.028 | 0.84 | 1.48 | 6.78 | 7.38 | 4 | 0.831 |
| Singapore | 0.013 | 0.057 | 0.000 | 0.60 | 1.18 | 8.57 | 8.22 | 3 | 0.861 |
| South Africa | −0.017 | −0.104 | −0.019 | 0.93 | 1.45 | 4.42 | 8.92 | 4 | 0.722 |
| Thailand | 0.365 | 0.070 | −0.018 | 0.93 | 0.56 | 6.25 | 5.18 | 3 | 1.003 |
| UK | 0.193 | 0.148 | 0.020 | 1.13 | 1.00 | 8.57 | 9.10 | 4 | 0.515 |
| USA | 0.203 | 0.146 | 0.007 | 0.81 | 0.58 | 10.00 | 8.63 | 1 | 0.461 |
| Brazil | 0.027 | 0.029 | −0.058 | 0.39 | 0.18 | 6.32 | 6.32 | 1 | 0.003 |
| Chile | 0.116 | 0.073 | 0.004 | 0.63 | 0.80 | 7.02 | 5.30 | 2 | 0.848 |
| France | 0.216 | 0.090 | 0.056 | 0.96 | 0.23 | 8.98 | 9.05 | 0 | 0.690 |
| Indonesia | −0.025 | −0.040 | 0.001 | 0.42 | 0.15 | 3.98 | 2.15 | 4 | 0.767 |
| Italy | 0.129 | 0.062 | 0.004 | 0.55 | 0.08 | 8.33 | 6.13 | 2 | 0.990 |
| Netherlands | 0.221 | 0.142 | 0.009 | 1.08 | 0.52 | 10.00 | 10.00 | 2 | 0.565 |
| Spain | 0.132 | 0.080 | −0.041 | 0.75 | 0.17 | 7.80 | 7.38 | 2 | 0.769 |
| Germany | 0.212 | 0.133 | 0.074 | 1.12 | 0.13 | 9.23 | 8.93 | 3 | 0.693 |
| Japan | 0.081 | 0.005 | 0.011 | 1.22 | 0.62 | 8.98 | 8.52 | 2 | 0.844 |
| South Korea | 0.032 | 0.006 | 0.097 | 0.74 | 0.44 | 5.35 | 5.30 | 3 | 2.000 |
| Denmark | 0.127 | 0.098 | 0.007 | 0.34 | 0.21 | 10.00 | 10.00 | 3 | 0.853 |
| Finland | 0.071 | 0.048 | −0.010 | 0.75 | 0.25 | 10.00 | 10.00 | 1 | 0.829 |
| Norway | 0.230 | 0.151 | 0.023 | 0.64 | 0.22 | 10.00 | 10.00 | 2 | 0.651 |
| Sweden | 0.270 | 0.243 | 0.147 | 0.55 | 0.51 | 10.00 | 10.00 | 2 | 0.672 |
| Mean | 0.168 | 0.087 | 0.02 | 0.766 | 0.529 | 8.208 | 7.956 | 2.273 | 0.767 |
| Median | 0.177 | 0.085 | 0.008 | 0.750 | 0.465 | 8.775 | 8.575 | 2.000 | 0.744 |
| S.D. | 0.096 | 0.078 | 0.045 | 0.245 | 0.416 | 1.947 | 2.102 | 1.162 | 0.346 |

Notes: This table presents the summary statistics for the replication of Ball, Robin, and Sadka (2008) for our analysis to test hypothesis H2 based on a cross-sectional setting. Panel A reports summary statistics for the firm-level variables used in the analysis based on an international sample of 22 countries over the period of 1992-2003. Panel B reports the country-level variables, including the Basu (1997) AT measure (*B3*) coefficient estimates, the Badia et al. (2021) MAT measure (*Modified B3*) coefficient estimates, the SVC measure values, debt and equity market importance, and control variables. All variables are defined in Table 1.

**Table 7: Ball, Robin, and Sadka (2008) inference re-examination (test of hypothesis H2)**

Panel A: The Effect of Debt and Equity Markets on CC using the AT, MAT, and SCV measures

| Dep. Var.: | *B3* | *B3* | *B3* | *Mod. B3* | *Mod. B3* | *Mod. B3* | SCV | SCV | SCV |
|---|---|---|---|---|---|---|---|---|---|
| *DEBT* | 0.2652$^{**}$ | 0.2013$^{*}$ | 0.2571$^{*}$ | 0.1081 | −0.0236 | 0.0056 | 0.0029 | −0.0254 | −0.0153 |
| | (2.79) | (1.87) | (2.06) | (1.27) | (−0.32) | (0.07) | (0.06) | (−0.44) | (−0.22) |
| *EQUITY* | −0.1837$^{***}$ | −0.1451$^{*}$ | −0.1198 | −0.0928 | −0.0134 | −0.004 | −0.0014 | 0.0157 | 0.0152 |
| | (−3.00) | (−2.13) | (−1.45) | (−1.69) | (−0.29) | (−0.07) | (−0.04) | (0.42) | (0.34) |
| *ENGLISH* | 0.2173$^{***}$ | 0.1828$^{**}$ | 0.2047$^{**}$ | 0.1163$^{*}$ | 0.0453 | 0.0479 | −0.0455 | −0.0607 | −0.0397 |
| | (3.18) | (2.50) | (2.38) | (1.89) | (0.90) | (0.82) | (−1.24) | (−1.53) | (−0.84) |
| *FRENCH* | 0.0825 | 0.0697 | 0.0922 | 0.0426 | 0.0163 | 0.0204 | −0.0634$^{*}$ | −0.0690$^{*}$ | −0.0448 |
| | (1.24) | (1.05) | (1.09) | (0.71) | (0.36) | (0.35) | (−1.79) | (−1.92) | (−0.97) |
| *SCAND* | 0.1691$^{**}$ | 0.1101 | 0.1656 | 0.1267$^{*}$ | 0.0054 | 0.0276 | −0.0177 | −0.0438 | −0.026 |
| | (2.19) | (1.22) | (1.55) | (1.83) | (0.09) | (0.38) | (−0.43) | (−0.89) | (−0.44) |
| *LAW* | | 0.0156 | 0.0314 | | 0.0322$^{***}$ | 0.0455$^{***}$ | | 0.0069 | 0.0088 |
| | | (1.21) | (1.50) | | (3.65) | (3.20) | | (0.99) | (0.77) |
| *CORRP* | | | −0.0203 | | | −0.0144 | | | −0.0002 |
| | | | (−0.94) | | | (−0.98) | | | (−0.02) |
| *CRED* | | | −0.001 | | | 0.0099 | | | −0.0012 |
| | | | (−0.04) | | | (0.58) | | | (−0.09) |
| *BTM* | | | 0.0301 | | | −0.0112 | | | 0.0412 |
| | | | (0.41) | | | (−0.22) | | | (1.03) |
| $R^2$ | 52.66% | 56.90% | 61.52% | 30.80% | 63.38% | 67.62% | 25.41% | 29.98% | 36.26% |
| N | 22 | 22 | 22 | | 22 | 22 | 22 | 22 | 22 |

Panel B: The effect of high and low importance of debt markets through the AT measure using an OLS regression

| *X* | *RD* | *RET* | *RD*×*RET* | *HiDebt* | *HiDebt*×*RD* | *HiDebt*×*RET* | *HiDebt*×*RD*×*RET* | |
|---|---|---|---|---|---|---|---|---|
| | −0.0653$^{***}$ | 0.0339$^{***}$ | 0.0926$^{***}$ | −0.0270$^{***}$ | 0.0606$^{***}$ | −0.0271$^{***}$ | 0.1058$^{***}$ | Adj. $R^2$ = 4.4% |
| | (−6.78) | (3.90) | (4.14) | (−5.36) | (6.13) | (−3.51) | (4.56) | N = 95,347 |

Panel C: The effect of high and low importance of debt markets through the AT measure using a firm-fixed-effect regression

| *UX* | *RD* | *RET* | *RD*×*RET* | *HiDebt* | *HiDebt*×*RD* | *HiDebt*×*RET* | *HiDebt*×*RD*×*RET* | |
|---|---|---|---|---|---|---|---|---|
| | −0.0454$^{***}$ | 0.0363$^{***}$ | 0.0972$^{***}$ | . | 0.0381$^{***}$ | −0.0248$^{***}$ | 0.0553$^{***}$ | Within $R^2$ = 3.05% |
| | (−7.43) | (4.21) | (8.05) | . | (5.91) | (−2.77) | (4.32) | N = 95,347 |
| | additional variables included are *MTB* and *RVAR*, and their interactions with *RD, RET*, and *RD*×*RET* | | | | | | | |

Panel D: Comparison of SCV across high and low importance of debt markets

| Groups | N | Average *X* | Std. dev. |
|---|---|---|---|
| *HiDebt* = 0, *RD* = 0 | 6,456 | 0.0696 | 0.2928 |
| *HiDebt* = 0, *RD* = 1 | 6,396 | −0.0450 | 0.3029 |
| *HiDebt* = 1, *RD* = 0 | 40,805 | 0.0428 | 0.2283 |
| *HiDebt* = 1, *RD* = 1 | 41,690 | −0.0325 | 0.2348 |
| SCV \| (*HiDebt* = 0) = 0.0060 | | ΔSCV = 0.003 | |
| SCV \| (*HiDebt* = 1) = 0.0030 | | $Chi^2$ = 1.76; *p*-value = 0.184 | |

Notes: This table presents replication and re-examination of the inference from Ball, Robin, and Sadka (2008) for our analysis to test hypothesis H2 based on a cross-sectional setting. Panels A reports three sets of regression results the country-level values of the AT measure (*B3*), the MAT measure (*Modified B3*), and the SCV measure as dependent variables, with *t*-statistics in parentheses. Panel B reports the conditioning effect of *HiDebt* on CC through the Basu (1997) AT regression with *X* as the dependent variable. Panel C reports the conditioning effect of *HiDebt* through the Badia et al. (2021) MAT regression with *UX* as the dependent variable. Panel D reports the summary statistics in terms of *X* across four observation groups sorted on *HiDebt* and *RD*, the SCV measure across high and low *DEBT* groups, and statistical significance of the difference in SCV between both groups. All variables are defined in Table 1. Reported *t*-statistics in parentheses in Panels B and C are calculated based on clustered standard errors at the firm level. $^{*}$, $^{**}$, $^{***}$ denote significance at the 10%, 5%, and 1% levels, respectively.