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Conservatism, disclosure and the cost of equity capital

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Abstract

This study seeks insights into the economic consequences of accounting conservatism by examining the relation between conservatism and cost of equity capital. Appealing to the analytical and empirical literatures, we posit an inverse relation. Importantly, we also posit that the strength of the relation is conditional on the firm's information environment, being the strongest for firms with high information asymmetry and the weakest (potentially negligible) for firms with low information asymmetry. Based on a sample of US-listed entities, we find, as predicted, an inverse relation between conservatism and the cost of equity capital, but further, that this relation is diminished for firms with low information asymmetry environments. This evidence indicates that there are economic benefits associated with the adoption of conservative reporting practices and leads us to conclude that conservatism has a positive role in accounting principles and practices, despite its increasing rejection by accounting standard setters.

JEL Classification: **M41**

Keywords

Conservatism, cost of equity capital, disclosure, information risk, signalling

1. Introduction

In this study, we seek insights into the economic consequences of accounting conservatism by examining the relation between conservatism and cost of equity capital. Appealing to both the analytical and empirical literatures, we posit an inverse relation, with greater accounting conservatism mapping into a reduced cost of equity capital. Importantly, we also posit that the strength of the relation is conditional on the firm's information environment, being the strongest for firms with high information asymmetry and the weakest (potentially negligible) for firms with low

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information asymmetry. Underlying this prediction is the supposition that conservatism forms a part of the overall communication strategy of the firm (Gigler and Hemmer, 2001).

We view a study into the economic benefits of conservatism as timely, given the ongoing debate throughout the 2000s regarding the merits of neutrality, as opposed to conservatism (prudence), as a desirable characteristic of accounting numbers. Ultimately, through a series of joint exposure drafts, the IASB and FASB have adopted the view that neutrality is the higher order objective. For example, their May 2008 joint exposure draft contains the following statement:

> However, the Boards concluded that describing prudence or conservatism as a qualitative characteristic or a desirable response to uncertainty could conflict with the quality of neutrality because, even with the proscriptions of deliberate misstatement that appear in the existing frameworks, an admonition to be prudent is likely to lead to a bias in the reported financial position and financial performance. … Accordingly, the proposed framework does not include prudence or conservatism as desirable qualities of financial reporting information. (BC2.21)

Reinforcing this position, the September 2010 FASB exposure draft explicitly states: ‘Chapter 3 does not include prudence or conservatism as an aspect of faithful representation because including either would be inconsistent with neutrality’. (BC3.27)

The IASB/FASB position has not been universally endorsed. For example, 179 comment letters were received in response to the 2006 discussion paper, with the IASB minutes of the board meeting dated 20 February 2007 revealing that ‘[m]any constituents argued that prudence, or conservatism, should be included as a characteristic or component of a characteristic’ (point 72). More explicitly, in response to the 2008 joint exposure draft, the British Accounting Association stated: ‘We approve of the present Framework’s acceptance of prudence, and regret its withdrawal in the Exposure Draft’. Further, in concluding their submission, they argued:

> All of these considerations mean that the reliability of measurements is an important characteristic and that statistical bias, in the form of conservatism, can be a desirable characteristic in many situations. *It would be scandalous if the wisdom of generations of accountants, supported by modern economic analysis, were to be ignored.* (p. 3, emphasis added)

Historically, the predominance of research into conservatism has focused on the contracting incentives for conservatism to mitigate agency conflict (Watts, 2003a). Despite theoretical propositions supporting a signalling benefit (e.g. Bagnoli and Watts, 2005; Gietzmann and Trombetta, 2003), the empirical evidence to date has paid relatively little attention to the economic consequences of conservatism. As signalling theory suggests, conservative reporting practices are optimally selected when they offer greater benefit or impose less cost on the firm relative to alternative communication strategies such as disclosure. The voluntary disclosure research has, for the most part, ignored the potential information content of accounting policy choice, focusing primarily on the association between voluntary disclosure and the cost of equity capital.

The recent argument by Guay and Verrecchia (2007) also provides support for an investigation into the economic consequences of conservatism. They suggest a framework for *ex ante* conservatism in an attempt to shift researchers from a paradigm of considering conservatism as an *ex post* response to information (Basu, 1997) to a notion congruent with an *ex ante* commitment to conservative reporting that is relevant to increasing firm disclosure and reducing uncertainty.

From the theoretical perspective, we draw on Gietzmann and Trombetta (2003) and Bagnoli and Watts (2005). Of direct relevance, Gietzmann and Trombetta (2003: 200) envisage a role for both conservative accounting policies and voluntary disclosure. They conclude that the economic implications of accounting policy choice or voluntary disclosure can not be meaningfully studied in isolation; rather, each has the potential to influence cost of equity capital and, additionally, an investigation of the economic benefits of accounting policy choice must 'control for the voluntary disclosure strategy of firms'.

From the empirical perspective, Francis et al. (2004) examine the relation between accounting-based attributes of earnings including conservatism and cost of equity capital, Gietzmann and Ireland (2005) and Francis et al. (2008) examine the effect of voluntary disclosure on cost of equity capital conditional on the underlying quality of earnings and Ahmed et al. (2002) find accounting conservatism to be associated with a lower cost of debt. Most recently, Garcia Lara et al. (2011a) examine the relation between conditional conservatism and cost of equity capital, Garcia Lara et al. (2011b) examine the information consequences of conservatism for analysts, debtholders and shareholders, and Kim et al. (2012) document a reduced adverse price reaction at the time of the announcement of a seasoned equity offering (SEO) for firms with more conservative financial reporting. We view an investigation into the direct relation between conservatism and cost of equity capital conditional on the firm's information environment as a natural extension of these studies.

We base our analysis on a sample of 3138 firm-year observations from the period 1985–2000, a period selected so as to pre-date both the previously discussed debate regarding the merits of neutrality versus conservatism as a desirable characteristic of accounting numbers and the significant changes in disclosure requirements brought about by regulatory changes such as Sarbannes–Oxley (SOX). We argue that in this earlier experimental window, a firm's decision to adopt conservative accounting policies is more likely to reflect a part of its strategic communication decision in the spirit of Gietzmann and Trombetta (2003), unaffected by either the debate or the regulatory changes.

The primary measures we use to proxy for the three fundamental inputs into our econometric model are as follows. First, appealing to Botosan et al. (2011), we use the Easton (2004) modified price–earnings–growth (MPEG) measure as our proxy for cost of equity capital. Second, recognising the need for both a firm-specific measure to conduct our cross-sectional analysis and a measure that captures conservatism resultant from both mandatory accounting policy adoption and discretionary policy choice, we adopt the Givoly and Hayn (2000) negative accruals measure as our proxy for accounting conservatism. This decision reflects our view of the construct 'accounting conservatism' as reflecting the firm's overall propensity to undertake, on average, conservative accounting choices. Finally, we adopt idiosyncratic volatility, measured as the standard deviation of the residuals from a daily market-model regression, as our proxy for information asymmetry (Dierkens, 1991; Moeller et al., 2007).

Our results provide consistent support for our predictions. We document an inverse association between our conservatism measure and cost of equity capital estimate, and further find that the marginal impact of conservatism systematically declines as the firm's information environment improves. We also find these results to be robust to the choice of proxy for each of the fundamental inputs.

We argue that this study has the potential to make important contributions from both the regulatory and the academic perspectives. From a regulatory perspective, we interpret the findings as indicative that a firm's decision to adopt conservative reporting policies has the potential to provide real economic benefits and, thereby, that accounting conservatism has a beneficial role within accounting principles and practices. We also view them as giving credence to the continued observation of conservative reporting practice. In this fashion, they appear to draw into question the path enunciated in recent joint IASB/FASB exposure drafts, away from conservatism (prudence) and towards neutrality as a desirable qualitative characteristic of financial reporting.

The study also contributes to the growing body of academic literature that focuses on the market (economic) benefits of accounting policy decisions. We argue that it contributes to the conservatism literature by taking its foundation from the signalling literature rather than the more commonly employed agency arguments. By exploring the influence of conservatism conditional on the firm's information environment, it extends our understanding of accounting conservatism and the role that it plays within the firm's overall communication strategy. To date, the majority focus within the literature has been on the cost of equity capital implications of disclosure. While some thought has been given to accounting policy choice, typically it has been as a control. For example, Gietzmann and Ireland (2005) and Francis et al. (2008) introduce an earnings quality measure into their 'cost of equity capital – disclosure' models in a linear additive fashion. More directly, Garcia Lara et al. (2011a, 2011b) introduce conservatism into excess returns models. Our analysis extends this work by conditioning its investigation into the economic consequences of conservative accounting policy choice on the firm's information environment.

The paper proceeds as follows. Section 2 presents a brief literature review, culminating with the study's hypothesis. Section 3 describes the econometric model and Section 4 describes the sample data. Section 5 then presents the empirical results and Section 6 concludes.

2. Literature review and hypothesis development

2.1. Empirical literature

A significant number of studies have explored the relation between disclosure and cost of capital, notably including Botosan (1997) and Botosan and Plumlee (2002). On balance, the results provide support for the conjectured inverse relation, but also suggest that the benefits are dependent on the firm's information environment. Here, authors typically appeal to two arguments. The first is that greater disclosure will enhance market liquidity (Botosan and Plumlee, 2002). The second is that disclosure reduces the information risk associated with the estimation of the parameters of the return distribution (Barry and Brown, 1985; Lambert et al., 2007).[1]

In contrast, with the exceptions of Garcia Lara et al. (2011a, 2011b), Kim et al. (2012) and Francis et al. (2004), to date conservatism has been examined from a variety of perspectives other than its economic benefits. For example, existing research provides evidence of variation in conservatism across firms, quarters and jurisdictions, of increased conservatism over time, about the influence that auditors' exposure to legal liability has on conservatism, and about how conservatism influences board composition. Extensive reviews of this work are provided in Watts (2003b) and Givoly et al. (2007). Notwithstanding the lack of empirical evidence on its economic benefits, given its long-standing history, it is likely that conservatism is viewed by management as value-adding. Countering this, aggressive accounting practices are often met with suspicion by market participants.

Of more direct relevance, Francis et al. (2004) examine the relation between cost of equity capital and seven attributes of earnings including conservatism. Using the Basu (1997) asymmetric timeliness measure, they fail to document an association with cost of equity capital. Their result is, however, perhaps not surprising, given that the Basu measure captures an *ex post* response to market information essentially required by mandatory accounting policies. More importantly, Givoly et al. (2007) reveal the Basu measure to be unreliable when estimated in time series at the firm level. Alternatively, Gietzmann and Ireland (2005) and Francis et al. (2008) provide evidence on the association between disclosure and the cost of equity capital after controlling for earnings quality. Both document a negative association, but also find it to be substantially diminished when controlling for the influence of earnings quality. Gietzmann and Ireland interpret their discretionary accruals measure as distinguishing aggressive versus conservative accounting choice.[2]

Finally, several recent studies directly examine the cost-of-equity-capital implications of conservatism. First, Garcia Lara et al. (2011a) investigate the association between conditional conservatism and cost of equity capital, finding, as predicted, a negative association. They use a firm-specific measure of conditional conservatism based on Callen et al. (2010) and introduce it in a linear fashion into an excess returns model that controls for known risk factors, beta, firm size, book-to-market and momentum (Fama and French, 1993). Alternatively, Garcia Lara et al. (2011b) show that the same conservatism measure is inversely associated with information asymmetry, stock return volatility, analyst forecast error and credit risk. They conclude that their results are 'consistent with conservatism improving the information environment of the firm' (2011: 25). Lastly, Kim et al. (2012) investigate whether the magnitude of the adverse share-price reaction at the time of a seasoned equity offering (SEO) announcement is related to a firm's decision to adopt conservative financial reporting. They find that significantly less underpricing is experienced by issuers with high information asymmetry that have more conservative financial reporting.

2.2. Theoretical foundations

Historically, the greatest amount of attention paid to the construct of accounting conservatism has been from the contrasting perspective, with agency arguments underlying much of the empirical work. For example, LaFond and Watts (2008) hypothesise that the demand for conservatism arises as an equilibrium corporate governance response aimed at lowering agency costs associated with managerial overstatement of income and assets. For a sample of US firms, they find, as predicted, that conservatism increases in response to increased information asymmetries in equity markets.

In contrast, several theoretical studies have emerged recently that envisage an economic role for accounting policy choice from a signalling perspective. These models predict an inverse association between the tone of accounting policy choice (aggressive or conservative) and the cost of equity capital. Easley and O'Hara (2004) show that in equilibrium, firms can influence their cost of capital by affecting the precision and quantity of information available to investors through the selection of accounting standards and corporate disclosure practices. They envisage a role for both disclosure and accounting policy choice. More directly, Bagnoli and Watts (2005) develop a signalling model and show theoretically that in the presence of asymmetric information, management can use conservative accounting policy choices to signal private information.

The most relevant theoretical work is that of Gietzmann and Trombetta (2003), who demonstrate within the context of their model that managers having private information regarding the firm's future earnings can choose conservative accounting policies and/or voluntary disclosure as quality signals. The voluntary adoption of conservative accounting policies is not a costless signal, because it reduces the amount of financial slack available to the manager, which can – for example – reduce their ability to manage earnings to meet earnings forecasts. They show that, in equilibrium, the choice of accounting policy can be used to signal private information about future earnings because the choice interacts with the optimal firm strategy for voluntary disclosure. Thus, firms will use some combination of signals involving conservative accounting choices and/or direct disclosures.

Finally, in a general sense, the arguments underlying these models can equally be couched in terms of the ability of conservatism to reduce information risk. As noted, a number of studies have investigated the economic benefits of disclosure based on the underlying assumption that disclosure can reduce the firm's cost of equity capital through a reduction in non-diversifiable information risk (e.g. Botosan, 1997; Botosan and Plumlee, 2002; Francis et al., 2004). From this perspective, factors argued to impact information risk – such as, for example, accounting policy choice in the form of conservatism and/or corporate disclosure policy – will also impact the cost of equity capital.

2.3. Hypothesis

The theoretical literature reviewed above consistently identifies a signalling benefit to accounting conservatism. However, it also suggests that the relation is somewhat more complex, with the firm's information environment representing a mitigating factor. Bagnoli and Watts (2005) show that it is only in the presence of asymmetric information that conservative accounting policy choices can be used to communicate management's private information. Gietzmann and Trombetta (2003: 200) conclude that accounting policy choice and voluntary disclosure can not be meaningfully studied in isolation, but rather, 'any study of the value relevance of accounting needs to control for the voluntary disclosure strategy of the firm' and 'any study of the link between cost of capital and voluntary disclosure should control for differences in accounting policy adoption'.

As implied by Gietzmann and Trombetta (2003), conservatism and disclosure (information asymmetry) are both legitimate perspectives from which to frame arguments and empirical tests. Our focus is the economic consequences of conservative accounting policy choice. Thus, we present and test our hypothesis from the perspective of conservatism. We initially posit an inverse relation between conservatism and cost of equity capital. We then further posit that the strength of this relation is dependent on the firm's information environment, being the strongest for firms with the greatest information asymmetry and weakest for firms for with low information asymmetry. In this sense, we argue that the signalling benefits of conservatism as manifest in the cost of equity capital are reduced in environments where there is low information asymmetry.

Formally, our hypothesis (expressed in the alternative form) is, then:

> *H_1: There is an inverse relation between the extent to which a firm adopts conservative accounting policies and its cost of equity capital, with the strength of the relation increasing in the extent of information asymmetry surrounding the firm.*

3. Research methodology

3.1. Econometric model

To investigate the relation between cost of equity capital and conservatism described by H_1, we employ the following econometric model:

$$r_{i,t} = \gamma_0 + \gamma_1\, CON_{j,t} + \gamma_2\, IA_{j,t} + \gamma_3\, CON_{j,t} * IA_{j,t} + \gamma_4\, Beta_{j,t} + \gamma_5\, lnSIZE_{j,t} + \varepsilon_j \quad (1)$$

where the primary variables are measured as follows (subscripts j and t denote firm and year):
$r_{j,t}$ = cost of equity capital, based on the Easton (2004) modified PEG measure (*MPEG*);
$CON_{j,t}$ = conservatism, based on the Givoly and Hayn (2000) accruals-based measure of conservatism (CON_{GH}); and
$IA_{j,t}$ = information asymmetry, measured as idiosyncratic volatility (σ_ε) (Dierkens, 1991).

These measures are discussed in detail in dedicated sections below, along with the alternative measures of each employed for sensitivity purposes. Based on H_1, we expect (unconditionally) a negative coefficient on *CON* ($\gamma_1 < 0$) and a negative coefficient on the interaction term,

$CON * IA$ ($\gamma_3 < 0$). Further, consistent with prior literature, we expect a positive coefficient on IA ($\gamma_2 > 0$).

The choice of control variables is guided by prior literature (e.g., Botosan et al., 2011; Francis et al., 2004), and comprises two of the known risk proxies, systematic risk (*Beta*) and firm size (*SIZE*) (Fama and French, 1993).[3] The coefficients are expected to be positive for *Beta* and negative for *lnSIZE*. These variables are measured as follows:

$Beta_{j,t}$ = the value-weighted market-model beta, estimated over the preceding 120-month period (minimum 24 months); and
$SIZE_{j,t}$ = market capitalisation at fiscal year-end.

3.2. Cost of equity capital

Prior literature suggests a number of approaches for estimating the *ex ante* firm-level cost of capital using analyst forecasts of earnings and target prices. Several recent studies have attempted to evaluate the validity of these various proxies against known firm-specific risk factors and future realised returns. Here, arguably, the most rigorous investigation is that by Botosan et al. (2011) who evaluate the validity of 12 proxies based on their associations with both realised returns and known risk factors, finding support for the target price estimate and long-term PEG measures.

Unfortunately, target price forecasts are not generally available given our source of analyst forecast data, I/B/E/S. Further, given the requirement for four- and five-year ahead earnings forecasts, the long-term PEG measure is only available for less than one-quarter of our sample firms. From the remaining measures examined by Botosan et al. (2011), their results suggest that the modified PEG measure (MPEG) is best: they find it to be significantly correlated in the predicted direction with realised returns, as well as with four of the five risk factors they consider. We therefore adopt MPEG estimated using analyst forecast data as our primary cost of capital proxy. Following Easton (2004) and consistent with Botosan et al. (2011), we measure MPEG as follows:

$$MPEG = A + \sqrt{A^2 + (feps_2 - feps_1) / P_0} \tag{2}$$

where $A = dps_1/2P_0$, P_0 is current price per share, dps_1 is current-year dividends, and $feps_1$ and $feps_2$ are one- and two-period ahead analysts' consensus forecasts of accounting earnings per share.

For robustness purposes, we consider two additional approaches to estimating the cost of equity capital. First, we recalculate the modified PEG measure after adjusting for predictable forecast error using the de-biasing approach proposed by Larocque (2012). The results of Botosan et al. (2011) suggest that de-biasing analyst forecasts can increase power (e.g. Guay et al., 2011; Hughes et al., 2008). Following Larocque's methodology, we initially predict errors in analysts' forecasts of earnings and then deduct the predicted error from the consensus analyst forecast to obtain the bias-adjusted forecast. The bias-adjusted forecasts are then used to recalculate the measure which we denote as $MPEG^{adj}$. Second, in light of the data limitations involving analysts' forecast data discussed above, following Hou et al. (2012), we use earnings forecasts generated from a cross-sectional model based on historical data. Since forecasts can be generated for all of our sample firm-year observations, we now calculate both the modified and long-term PEG measures using these earnings forecasts, given the support they receive in Botosan et al. (2011). We denote these alternative measures as $MPEG^{HVZ}$ and $PEGLT^{HVZ}$.[4]

3.3. Conservatism

Within the accounting literature, a universally accepted meaning for the construct 'conservatism' has been elusive. In this regard, Guay and Verrecchia (2007: 6) state:

> Because notions of conservatism have evolved somewhat informally from observations about various accounting policies and empirical regularities in financial data, offering a theory of conservatism is fraught with controversy. In particular, the accounting community has come to associate different empirical regularities with different attributes of conservatism. This has evolved to the point that often it is difficult to find consensus in the literature as to what conservatism means, and what role it serves within an efficient reporting system.

Against this backdrop, it is perhaps not surprising that a number of proxies for conservatism have developed within the literature, with the various proxies reflecting different perspectives on conservatism. Of these, the most prominent are Basu's (1997) asymmetric timeliness measure, Ball and Shivakumar's (2005) asymmetric cash flow to accruals measure, Penman and Zhang's (2002) hidden reserves measure, the market-to-book ratio (Beaver and Ryan, 2000), Givoly and Hayn's (2000) negative accruals measure and, more recently, the firm-specific measures of asymmetric timeliness proposed by Khan and Watts (2009) and Callen et al. (2010).

Ideally, when selecting among alternative proxies, researchers could appeal to evidence regarding their relative merits in much the same fashion that they can seek guidance on the choice of proxy for cost of equity capital by appealing to studies such as Botosan et al. (2011). However, because each proxy has been developed from its own perspective, there is no single or common set of external criteria against which the measures can be judged. Hence, it is neither feasible nor possible to identify a universally dominant proxy for accounting conservatism. This conclusion is underscored by two recent studies which attempt to evaluate the relative merits of the various proxies. Here, Wang et al. (2009: 197) conclude that 'the measures of conservatism employed in the literature may have a low degree of construct validity' while Artiach et al. (2012: 2) conclude 'the alternative proxies exhibit starkly different characteristics and show little consistency between them, suggesting they are capturing unrelated constructs'. This leads us to the position that, consistent with the conclusion drawn by Artiach and Clarkson (2011: 41), 'the choice of proxy can ultimately at best be guided by the researcher's assessment of the advantages and disadvantages of each of the alternative proxies relative to their experimental setting and by the researcher's view or notion of accounting conservatism'.

As presented, tests of H_1 require both a firm-specific measure of conservatism and one that captures conservatism resulting from discretionary, as well as mandatory, policy choices. Neither the Basu (1997) nor the Ball and Shivakumar (2005) proxies are firm-specific measures.[5] Further, in addition to requiring several difficult-to-find data items such as R&D and advertising expenses, the Penman and Zhang measure captures a level of conservatism common to many firms due to the mandatory nature of specific accounting standards. Finally, since the book-to-market ratio is a known risk factor, this effectively rules out the Beaver and Ryan (2000) book-to-market measure within a cost-of-equity-capital model. Thus, at first pass, the Givoly and Hayn (2000), Khan and Watts (2009) and Callen et al. (2010) measures appear to be the most reasonable for our purposes.

Of these, we view the Givoly and Hayn (2000: 292) negative accruals measure as (*ex ante*) the best suited and adopt it as our primary conservatism proxy. In addition to being a firm-specific measure, it captures both discretionary and mandatory dimensions of conservatism. Non-operating accruals, on which the measure is based, consist of those arising from both managerial action resulting from mandated accounting regulations and those arising from managerial discretion in the timing and amount of both accounting policy choices and accounting estimates.

Givoly and Hayn (2000: 292) base their measure on the descriptive definition of conservatism as ‘a selection criterion between accounting principles that leads to the minimization of cumulative reported earnings by slower revenue recognition, faster expense recognition, lower asset valuation, and higher liability valuation’. Thus, the intuition underlying their measure is that conservative accounting results in persistently negative accruals, with more negative average accruals reflecting more conservative accounting. Without intervention, accruals can typically be expected to reverse over time, with operating income converging to cash flows from operations. Thus, persistence in the level of cumulative negative accruals over time should reflect a conservatism bias within the firm’s accounting system rather than the transitory nature of accruals.

Following Givoly and Hayn (2000), we focus on non-operating accruals rather than total accruals. As they note, total accruals incorporate operating accruals and thereby arguably reflect both growth in operations and conservatism. We use a six-year accumulation period, consistent with Ahmed et al. (2002) and Francis et al. (2004), who view this as a sufficiently long period to enable the identification of persistence in accumulated accruals. Thus, our accruals-based conservatism proxy (CON_{GH}) is the ratio of non-operating accruals to total assets determined using the indirect method, averaged over a six-year period, and multiplied by −1 to produce an increasing measure: that is,

$$CON_{GHj,t} = -1 \times \left(\frac{1}{6} \sum_{t=1}^{6} \frac{NOACC_{j,t}}{TA_{j,t}} \right) \tag{3}$$

where, for firm j at fiscal year-end t, $NOACC_{j,t}$ is non-operating accruals and $TA_{j,t}$ is total assets.[6]

For robustness purposes, we consider the Khan and Watts (2009) firm-specific measure of asymmetric timeliness, denoted as *C_Score*. This measure is an extension of the Basu (1997) measure and enables the estimation of a firm year-specific measure of the asymmetric timeliness of earnings.

3.4. Information asymmetry

As our primary proxy for the firm’s information environment, we adopt the indirect measure of information asymmetry frequently employed within the literature, idiosyncratic volatility (σ_ε) (e.g. Dierkens, 1991; Moeller et al., 2007). Consistent with both studies, we measure σ_ε as the standard deviation of the residuals from a market-model regression estimated on daily data over the preceding year. The intuition underlying this measure, as expressed by Dierkens (1991: 186), is that it assumes ‘market fluctuations are the only information shared by the managers of the firm and the market’; thereby, fluctuations in the difference between price and value as reflected in σ_ε ‘capture… the information asymmetry between the managers of the firm and the market’. Higher values of σ_ε imply greater information asymmetry.

For robustness purposes, we appeal to a direct measure of disclosure, analysts’ ratings of the firm’s total disclosure as published in annual AIMR reports from 1978–79 to 1995–96. These ratings have been employed in a significant body of disclosure literature, including Lang and Lundholm (1993, 1996), Healy et al. (1999), Botosan and Plumlee (2002), and Drake et al. (2009), and are described in these studies. As described in Section 5.3.3, these ratings are only available for a subset of our sample. Since each industry is evaluated separately by a subcommittee of analysts and the subcommittees used a degree of judgement in arriving at a total disclosure score, we follow prior literature by converting the total weighted disclosure scores to industry-year percentile ranks and base our analysis on these percentile-rank data (*DISC*). As measured, higher values of *DISC* imply less information asymmetry.[7]

4. Sample data

Our sample consists of 3138 firm-year observations relating to the US-listed firms for which data to measure the variables in our econometric model, equation (1), could be obtained at some point over the period 1985–2000. The choice of study period is driven by our interest in a period when conservatism was viewed as a desirable characteristic of financial information, and also one prior to regulatory changes which resulted in significant changes in disclosure and reporting requirements. Data used to calculate the various measures were obtained from Standard and Poor's Research Insight Database, Thomson's DataStream and I/B/E/S Database, and the Centre for Research in Security Prices (CRSP) Database. The lack of I/B/E/S data for earnings forecasts to estimate the cost of equity capital led to the greatest sample attrition.

Table 1 presents a frequency distribution by year (panel A) and GICS industry code (panel B) for the sample observations. While there are slightly fewer observations over the first five years, on balance the sample is relatively uniformly spread across the study period. From 1990 onward, the sample stabilises at slightly in excess of 200 firms each year, with the differences driven by the requirements for a meaningful calculation of *MPEG*. Table 1 also reveals the sample to be broadly spread across the 27 identified industry sectors, with no apparent industry concentration. The largest representations are in the retail trade and petroleum sectors, with 9.44% and 7.85% of the total sample respectively, while the smallest are in the international pharmaceutical and non-ferrous mining sectors, with 0.31% and 1.00% respectively. As such, it is unlikely that our results are driven by specific types of industries, such as the capital-intensive industries.

Table 2 presents descriptive statistics for the regression model variables, as well as several additional measures, based on the pooled sample of 3138 firm-year observations. As revealed, there is broad variation in all measures. For the cost of equity capital estimates, the mean and median values of *MPEG* are 12.4% and 10.9% respectively, with a standard deviation of 6.0%. These figures are similar to the mean, median and standard deviation figures of 12.4%, 11.3% and 4.6% reported in Botosan et al. (2011). Consistent with Larocque (2012), the mean and median figures for $MPEG^{adj}$ are lower, at 8.8% and 8.4% respectively. Finally, based on the Hou et al. (2012) cross-sectional approach, the mean (median) values for $MPEG^{HVZ}$ and $PEGLT^{HVZ}$ are 13.9% (12.4%) and 14.1% (13.7%) respectively.

For the conservatism proxies, the mean and median values of CON_{GH} are 0.033 and 0.037 respectively, with a standard deviation of 0.051. The value at the first percentile is −0.167 and the value at the 99th percentile is 0.160. Unfortunately, it is difficult to directly compare descriptive statistics with prior studies, given differences in the way conservatism is calculated. Nevertheless, to provide context, for their measure based on total accruals, Ahmed et al. (2002) report mean and median values of 0.004 and 0.003 respectively, with a standard deviation of 0.031. Alternatively, using non-operating accruals scaled by total assets, Givoly and Hayn (2000) report mean (median) values of −0.016 (−0.003) and −0.047 (−0.037) for the periods 1951–80 and 1981–98, respectively. The Khan and Watts (2009) *C_Score* measure has mean and median values of 0.085 and 0.099 respectively, and a standard deviation of 0.170. These figures are relatively similar to the mean and median values of 0.105 and 0.097 that Khan and Watts report for the period 1963–2005.

Lastly, for the measures of information asymmetry, the mean and median values of σ_ε are 0.029 and 0.028 respectively, with a standard deviation of 0.149. These figures are similar to the mean

Table 1. Frequency distribution by year and GICS industry sector for a sample of 3138 firm-year observations from the period 1985–2000.

Panel A: Frequency distribution by year

Year	# Obs	%	Year	# Obs	%
1985	138	4.40%	1993	210	6.69%
1986	152	4.84%	1994	206	6.56%
1987	161	5.13%	1995	208	6.63%
1988	186	5.93%	1996	211	6.72%
1989	189	6.02%	1997	213	6.79%
1990	215	6.85%	1998	206	6.56%
1991	212	6.76%	1999	209	6.66%
1992	211	6.72%	2000	211	6.72%

Panel B: Frequency distribution by industry

GICS Industry	# Obs	%	GICS Industry	# Obs	%
Aerospace	77	2.46%	Machinery	128	4.08%
Airlines	50	1.59%	Motor carriers	52	1.65%
Apparel	116	3.69%	Natural gas distributors	94	2.99%
Chemical	92	2.93%	Natural gas pipelines	78	2.49%
Construction	118	3.76%	Non-ferrous mining/metal	31	1.00%
Container/Packaging	57	1.81%	Paper and forest	126	4.02%
Diversified	55	1.76%	Petroleum	246	7.85%
Electrical	143	4.55%	Precious metals	46	1.45%
Environmental control	152	4.86%	Publishing/broadcasting	178	5.67%
Financial services	94	2.99%	Railroad	80	2.54%
Food/beverages	201	6.42%	Retail trade	296	9.44%
Healthcare services	217	6.93%	Speciality chemical	152	4.86%
Insurance	215	6.84%	Textile	34	1.09%
Int. Pharmaceutical	10	0.31%			

The figures presented in each panel are the number of observations per year or industry sector (# Obs), as appropriate, and the percentage of the total of 3138 firm-year observations found within each cell (%).

and median figures reported by Moeller et al. (2007) of 0.029 and 0.026, and the mean values reported by Dierkens (1991) of between 0.023 and 0.025. Alternatively, for *DISC*, the mean and median values are 0.735 and 0.750, with a standard deviation of 0.132; these figures are similar to the mean and median figures of 0.70 and 0.72 reported by Lang and Lundholm (1996). Analyst following ranges from three at the first percentile to 43 at the 99th percentile.

Finally, none of the pair-wise correlations (not tabulated) across the different measures in our econometric model exceed 0.302. Thus, there is unlikely to be a threat of multicollinearity.

5. Results

5.1. Pooled sample regression results

Table 3 presents results for variants of our econometric model (equation (1)). All analyses employ the primary measures, *MPEG* for cost of equity capital, CON_{GH} for conservatism and σ_ε for information asymmetry. Panel A presents results based on the raw values of the model inputs, panel B

Table 2. Descriptive statistics for a pooled sample of 3138 firm-year observations over the period 1985–2000.

Variable	Mean	Median	Std dev	1st percentile	99th percentile
Cost of equity capital (r)					
MPEG	0.124	0.109	0.060	0.030	0.372
$MPEG^{adj}$	0.088	0.084	0.060	0.028	0.341
$MPEG^{HVZ}$	0.139	0.124	0.074	0.038	0.351
$PEGLT^{HVZ}$	0.141	0.137	0.063	0.044	0.336
Conservatism (CON)					
CON_{GH}	0.033	0.037	0.051	−0.167	0.160
C_Score	0.085	0.099	0.170	−0.213	0.530
Information asymmetry (IA)					
σ_{ε}	0.029	0.028	0.149	0.001	0.493
AF	20.042	20	9.396	3	43
DISC (*n* = 1782)	0.735	0.750	0.132	0.360	0.958
Additional firm characteristics					
Beta	1.119	1.090	0.340	0.410	2.070
Size ($000)	5,649.00	2,302.41	9,994.78	67.16	57,468.15
BTM	0.549	0.509	0.306	0.074	1.454
LEV	0.186	0.174	0.138	0.000	0.601
ROA	0.057	0.051	0.061	−0.106	0.218

Variable definitions: *r* is the cost of equity capital, estimated using the Easton (2004) modified PEG (*MPEG*) measure described in equation (2), the measure alternatively recalculated after adjusting for predictable analyst forecast error using the de-biasing approach proposed by Larocque (2012) (*$MPEG^{adj}$*) and based on earnings forecasts generated by the Hou et al. (2012) cross-sectional approach (*$MPEG^{HVZ}$*), and finally, the Easton (2004) long-term PEG measure as described in footnote 4 also based on earnings forecasts generated by the Hou et al. (2012) cross-sectional approach (*$PEGLT^{HVZ}$*); *CON* is a firm-specific measure of reporting conservatism, estimated alternatively using the Givoly and Hayn (2000) negative accruals measure (*CON_{GH}*) described in equation (3) and the Khan and Watts (2009) asymmetric timeliness measure (*C_Score*); information asymmetry (*IA*) is measured alternatively as σ_{ε} equal to the standard deviation of the residuals from a market-model regression estimated on daily data over the preceding year, *AF* the number of analysts following the firm and *DISC* equal to the total weighted disclosure score expressed as a percentage of total points available from the AIMR report; *Beta* is the value-weighted market-model beta estimated over the preceding 120-month period; *Size* is the market value of common equity in $millions; *BTM* is the book value of common equity divided by the market value of common equity; *LEV* is leverage measured as long-term debt divided by total assets; and *ROA* is return on assets measured as net income before extraordinary items divided by total assets.

results based on industry-year percentile-rank values and panel C mean coefficient values, determined by estimating the model yearly based on raw input values across the 16-year study period and then averaging across years. For Panels A and B, three specifications of the model are reported; in addition to the control variables, the first specification includes only CON_{GH}, the second includes both CON_{GH} and σ_{ε} and the third represents the complete model inclusive of the interaction term, $CON_{GH} * \sigma_{\varepsilon}$. These models are run as pooled cross-sectional regressions with corrections for clustering of standard errors by firm and year (Petersen, 2009). For panel C, which reports results based only on the complete model, the annual regressions are run using OLS.

Finally, across all specifications, the coefficients on *Beta* are uniformly positive and the coefficients on *lnSIZE* uniformly negative, with all significant at better than the 1% level. As such, the discussions below focus on the coefficients for the primary terms, CON_{GH}, σ_{ε} and $CON_{GH} * \sigma_{\varepsilon}$.

Table 3. Regression model results for a pooled sample of 3138 firm-year observations over the period 1985–2000.

Panel A: Results based on raw data

Intercept	CON_{GH} (–)	σ_ε (+)	$CON_{GH} * \sigma_\varepsilon$ (–)	*Beta* (+)	*lnSIZE* (–)	*Adj* R^2
0.140	***−0.098***	—	—	0.042	−0.009	0.216
(< 0.001)	***(<0.001)***			(< 0.001)	(< 0.001)	
0.127	***−0.093***	***0.249***	—	0.042	−0.008	0.222
(< 0.001)	***(< 0.001)***	***(0.001)***		(< 0.001)	(< 0.001)	
0.129	−0.052	***0.209***	***−1.277***	0.042	−0.008	0.234
(< 0.001)	(0.277)	***(0.013)***	***(0.039)***	(< 0.001)	(< 0.001)	

Panel B: Rank regression results

Intercept	*Rank*(CON_{GH}) (–)	*Rank*(σ_ε) (+)	*Rank*(CON_{GH}) **Rank*(σ_ε) (–)	*Rank(Beta)* (+)	*Rank(lnSIZE)* (–)	*Adj* R^2
0.464	***−0.085***	—	—	0.196	−0.157	0.242
(< 0.001)	***(<0.001)***			(< 0.001)	(< 0.001)	
0.456	***−0.079***	***0.022***	—	0.196	−0.161	0.261
(< 0.001)	***(0.001)***	***(0.023)***		(< 0.001)	(< 0.001)	
0.469	−0.048	***0.015***	***−0.079***	0.196	−0.161	0.269
(< 0.001)	(0.230)	***(0.041)***	***(0.014)***	(< 0.001)	(< 0.001)	

Panel C: Annual regression results

	CON_{GH} (–)	σ_ε (+)	$CON_{GH} * \sigma_\varepsilon$ (–)	*Beta* (+)	*lnSIZE* (–)
Mean coefficient	−0.047	0.224**	−1.152***	0.045***	−0.013***
No. of coefficients > 0	4	14	1	16	0
No. of |*t*-statistics| > 1.645	1	11	14	15	16
Z_1 (Aboody and Lev)	−1.415	3.814	−5.101	10.025	−8.631
Z_2 (Aboody and Lev)	−1.823	4.190	−4.138	7.818	−11.390
Abarbanell/Bernard *t*-stat	−1.430	2.091	−3.097	7.221	−6.195

Variable definitions: *r* is the cost of equity capital, estimated using the Easton (2004) modified PEG (*MPEG*) measure described in equation (2); CON_{GH} is a firm-specific measure of reporting conservatism, estimated using the Givoly and Hayn (2000) negative accruals measure described in equation (3); information asymmetry (*IA*) is measured as σ_ε equal to the standard deviation of the residuals from a market-model regression estimated on daily data over the preceding year; *Beta* is the value-weighted market-model beta estimated over the preceding 120-month period; and *Size* is the market value of common equity in $millions.
For panel B, all measures are ranked within year and industry (denoted *Rank*).
In panel C, the Abarbanell/Bernard *t*-statistic adjusts for the estimated first-order autocorrelation in the independent variables over the sample period by adjusting standard errors using the following factor: $\{[(1 + \varphi)/(1 - \varphi)] - [2\varphi(1 - \varphi^n)/n(1 - \varphi)^2)]\}^{1/2}$, where *n* is the number of years and φ is the estimated first-order autocorrelation in the yearly coefficients. As noted by Abarbanell and Bernard, this adjustment factor assumes that the serial correlation is first-order autoregressive.
The Z_1 statistic, which assumes residual independence, is $(1/n)^{1/2} \Sigma^n [t_i/\{ k_i/(k_i - 2)\}^{1/2}]$, where t_i is the White's *t*-statistic for year *n*, k_i are the degrees of freedom, and *n* is the number of years. The Z_2 statistic is: mean *t*-statistic/(standard deviation of *t*-statistics/$\{n - 1\}^{1/2}$).
For Panels A and B, the *p*-values (in parenthesis) are two-tailed, with corrections for clustering of standard errors by year and industry.
For panel C, *** and ** indicate that the mean coefficient value is significantly different from zero at the 1% and 5% levels (two-tailed), respectively, based on the Abarbanell/Bernard *t*-statistic.

5.1.1. Regression results based on raw input values. Results based on raw values of the model inputs are reported in Panel A of Table 3. For the first specification, which includes only CON_{GH}, its coefficient at −0.098 ($p < 0.001$) is negative and significant. Thus, the results from this specification provide preliminary support for the notion that conservatism has an unconditional economic benefit in terms of cost of equity capital. Similarly, for the second specification, which includes both CON_{GH} and σ_ε in a linear fashion, the coefficient on CON_{GH} of −0.093 ($p < 0.001$) remains negative and significant. Additionally, the coefficient on σ_ε of 0.249 ($p = 0.001$) is positive and significant.

For the complete model which provides a direct test of H_1, the coefficient on CON_{GH} at −0.052 remains negative but is now insignificant ($p = 0.277$). Such a finding is, however, not fully unexpected, since this coefficient captures the low information asymmetry situation (the lowest value of σ_ε). Additionally, the coefficient on σ_ε remains positive and significant (0.209, $p = 0.013$). Importantly, the coefficient on the interaction term, $CON_{GH} * \sigma_\varepsilon$, is negative and significant. Its estimated value is −1.227 ($p = 0.039$). Thus, the strength of the association between conservatism and cost of equity capital is increasing in the level of information asymmetry as measured by σ_ε. From a purely statistical perspective, the linear restriction capturing the marginal impact of CON_{GH} conditional on σ_ε, $\gamma_1 + \gamma_3$ becomes statistically significant at the 5% level for values of σ_ε above 0.0339. Within our sample, 37.6% of the observations exceed this cut-off point. Thus, consistent with H_1, these results indicate that conservatism has a beneficial influence on cost of equity capital, but that this influence is conditional on the firm's information environment, with the marginal benefits of conservatism accruing only in environments of higher information asymmetry.[8,9]

Finally, to provide economic context, these results indicate that for a firm ranked at the 75th percentile within our sample in terms of σ_ε ($\sigma_\varepsilon = 0.0388$), a one-standard-deviation increase in CON_{GH} is associated with a 52-basis-point reduction in cost of equity capital (i.e. 0.051 * (−0.052 – 1.277(0.0388)) = 0.0052). Alternatively, for a firm ranked at the 95th percentile ($\sigma_\varepsilon = 0.0554$), a one-standard-deviation increase in CON_{GH} is associated with a 63-basis-point reduction.

5.1.2. Rank regression results. One alternative explanation for the results above is the possibility that $\mathrm{CON}_{\mathrm{GH}}$ is related to capital intensity. However, within our data, the correlation between capital intensity and $\mathrm{CON}_{\mathrm{GH}}$ is only −0.083 (p = 0.007). Thus, while the correlation is significant, given its magnitude, it is unlikely that capital intensity represents an important omitted variable.

More generally, this challenge raises the broader issue that potential industry effects have been overlooked. As such, consistent with Botosan and Plumlee (2002), we rank all measures within year and industry and then re-run the analysis using these percentile-rank data. Here, higher percentile-rank values imply higher values of each measure. By ranking, concerns regarding differences in management's attitude towards conservatism across industries (and, more broadly, differences in any of the measures across either industry or time) should be greatly alleviated.

As revealed in Panel B of Table 3, the results are consistent with those reported above. Focusing on the complete model, the coefficient on the conservatism measure at −0.048 ($p = 0.230$) is again insignificant, while the coefficient on the interaction term at −0.079 ($p = 0.014$) is again significant. Thus, the evidence continues to suggest that conservatism has a beneficial effect on cost of equity capital, but only for high information asymmetry firms. Additionally, the coefficient on the information asymmetry measure is positive (0.015, $p = 0.041$). In sum, results and conclusions appear robust to the use of industry–year percentile-rank data, again providing strong support for H_1. They also suggest that our results are unlikely to have been driven by industry effects.

Table 4. Regression model results for a pooled sample of 3138 firm-year observations over the period 1985 – 2000 partitioned by information asymmetry (σ_ε) quintile.

Quintile	Intercept	CON_{GH}	σ_ε	*Beta*	*lnSIZE*	*Adj* R^2
		(–)	(+)	(+)	(–)	
Q1	0.134	−0.054	0.303	0.022	−0.005	0.032
(lowest σ_ε)	(< 0.001)	(0.183)	(0.571)	(< 0.001)	(0.008)	
Q2	0.141	−0.077	1.091	0.047	−0.007	0.082
	(< 0.001)	(0.100)	(0.289)	(< 0.001)	(< 0.001)	
Q3	0.098	***−0.108***	1.449	0.027	−0.006	0.077
	(0.001)	***(0.017)***	(0.110)	(< 0.001)	(< 0.001)	
Q4	0.144	***−0.154***	0.157	0.044	−0.010	0.163
	(< 0.001)	***(0.001)***	(0.822)	(< 0.001)	(< 0.001)	
Q5	0.112	***−0.096***	0.319	0.078	−0.011	0.149
(highest σ_ε)	(< 0.001)	***(0.039)***	(0.348)	(< 0.001)	(< 0.001)	

Variable definitions: *r* is the cost of equity capital, estimated using the Easton (2004) modified PEG (*MPEG*) measure described in equation (2); CON_{GH} is a firm-specific measure of reporting conservatism, estimated using the Givoly and Hayn (2000) negative accruals measure described in equation (3); information asymmetry (*IA*) is measured as σ_ε equal to the standard deviation of the residuals from a market-model regression estimated on daily data over the preceding year; *Beta* is the value-weighted market-model beta estimated over the preceding 120-month period; and *Size* is the market value of common equity in $millions.
For this analysis, the pooled data set of 3138 firm-year observations has been partitioned into quintiles using the information asymmetry measure σ_ε. The first quintile, Q1, consists of the firm-year observations with the lowest measure of σ_ε (i.e., the lowest level of information asymmetry) and the fifth quintile, Q5, consists of the firm-year observations with the highest measure of σ_ε (i.e., the highest level of information asymmetry).
Reported *p*-values (in parenthesis) are two-tailed, with corrections for clustering of standard errors by year and industry.

5.1.3. Annual regression results. Panel C of Table 4 presents mean coefficient values for the complete model re-estimated annually based on raw values of the model inputs. Here also, the results provide strong support for H_1. While the estimated coefficient on CON_{GH} is negative in 12 of the 16 years, it is only significant in one year. In contrast, the estimated coefficient on the interaction term, $CON_{GH} * \sigma_\varepsilon$, is negative in 15 years, with |*t*-statistics| > 1.65 in 14 of those years. For σ_ε, its coefficient is positive in 14 years and significant in 11 years. To test for the statistical significance of the mean coefficients, we consider several approaches. Following Abarbanell and Bernard (2000), we calculate the standard error from the distribution of yearly coefficients and then make an adjustment for serial correlation. To supplement this measure, we also calculate the two Z-statistics (Z_1 and Z_2) employed by Aboody and Lev (1998) and Healy et al. (1987).[10] Appealing to these measures, we find the mean coefficient on CON_{GH} to be statistically insignificant at conventional levels across all test statistics, but the mean coefficient on the interaction term, $CON_{GH} * \sigma_\varepsilon$, to be consistently significant at the 1% level. For example, the Abarbanell/Bernard *t*-statistics for these two coefficients are −1.430 and −3.097, respectively. We also find the mean coefficient on σ_ε to be significant at the 5% level. Its *t*-statistic is 2.091. For these annual regressions, the adjusted R^2s range from 0.164 (in 1993) to 0.327 (in 1998).

5.2. Partitioned data

To more directly illustrate the reliance of the relation between conservatism and cost of equity capital on the firm's information environment, we partition our pooled sample observations into quintiles on the basis of the information asymmetry measure, σ_ε, and then run a reduced form of our econometric model which excludes the interaction term on each of the partitions using raw input values. The results are presented in Table 4.

As revealed, the results confirm the takeaway from the pooled analysis discussed above. For the first two quintiles (Q1 and Q2) which have the lowest measures of σ_ε, the coefficients on CON_{GH} are statistically insignificant at conventional levels. In contrast, for the remaining three quintiles (Q3, Q4, and Q5), which exhibit the higher values of σ_ε (greater information asymmetry), the coefficient on CON_{GH} is consistently negative and significant. Thus, the documented relation between conservatism and cost of equity capital again appears restricted to only those firms within the higher information asymmetry environments. As an aside, the coefficient on σ_ε is insignificant across all five analyses – perhaps not unexpectedly, given the reduced cross-sectional variation in the measure under the partitioning process. The coefficients on both *Beta* and *lnSIZE* are, however, highly significant in the predicted direction.

5.3. Sensitivity considerations

In Table 5, we present results from analyses designed to explore the sensitivity of our conclusions to our choice of proxies for the primary measures, cost of equity capital, conservatism and information asymmetry. For parsimony, we only report results for the complete model. As will be seen, results and conclusions appear robust to the choice of proxies.

5.3.1. Sensitivity to the cost of equity capital measure. Panel A of Table 5 presents results for our econometric model run with each of the three alternative proxies for cost of equity capital as the dependent variable, the Easton (2004) modified PEG measure recalculated after adjusting for predictable forecast error using the de-biasing approach proposed by Larocque (2012) ($MPEG^{adj}$) and the Easton (2004) modified PEG and long-term PEG measures based on earnings forecasts generated from the Hou et al. (2012) cross-sectional approach ($MPEG^{HVZ}$ and $PEGLT^{HVZ}$, respectively). We continue to use our primary measures for conservatism (CON_{GH}) and information asymmetry (σ_ε).

The results are consistent with those based on the primary measure reported in Panel A of Table 3. The coefficient on CON_{GH} remains insignificant while the coefficient on σ_ε is positive and significant. Of greatest interest, the coefficient on the interaction term is uniformly negative and significant. For example, for the third model using $PEGLT^{HVZ}$, the coefficient on CON_{GH} is −0.053 ($p = 0.267$), the coefficient on σ_ε is 0.209 ($p = 0.014$) and the coefficient on $CON_{GH} * \sigma_\varepsilon$ is −1.233 ($p = 0.036$). Further, based on these coefficient estimates, the linear restriction relating to the conditional coefficient on CON_{GH} becomes significant at the 5% level for values of σ_ε above 0.0308 (representing 56.9% of the sample observations with the highest measures of σ_ε), and for a firm ranked at the 75th percentile in terms of σ_ε, a one-standard-deviation increase in CON_{GH} is associated with a 51-basis-point reduction in cost of equity capital. Thus, our results appear robust to the choice of proxy for cost of equity capital. Note that such a finding is not totally unexpected, given that all pair-wise correlations among the four proxies for cost of equity capital exceed 0.53.

Finally, to confront concerns regarding the possibility of a mechanical relationship between the implied cost of equity capital and conservatism measures, we alternatively use realised returns as the dependent variable.[11] Following the literature, we make adjustments for cash flow news (Botosan et al., 2011). Again, the results (not tabulated) reveal an inverse relation with conservatism, but only in the high information setting. The coefficients (p-values) on CON_{GH} and $CON_{GH} * \sigma_\varepsilon$ are −0.039 ($p = 0.320$) and −1.448 ($p = 0.027$) respectively. We interpret these results as indicating that it is relatively unlikely that our primary results derive exclusively from a

Table 5. Additional considerations.

Panel A: Sensitivity to the choice of cost of equity capital measure (r)

Measure	Intercept	CON_{GH}	σ_ε	$CON_{GH} * \sigma_\varepsilon$	*Beta*	*lnSIZE*	*Adj R*^{2}
		(–)	(+)	(–)	(+)	(–)	
***MPEG**^{adj}*	0.095	−0.046	***0.199***	***−1.471***	0.040	−0.008	0.245
	(< 0.001)	(0.335)	***(0.017)***	***(0.026)***	(<0.001)	(<0.001)	
***MPEG**^{HVZ}*	0.145	−0.052	***0.207***	***−1.263***	0.040	−0.008	0.239
	(< 0.001)	(0.274)	***(0.014)***	***(0.034)***	(<0.001)	(<0.001)	
***PEGLT**^{HVZ}*	0.147	−0.053	***0.209***	***−1.233***	0.040	−0.008	0.250
	(< 0.001)	(0.267)	***(0.014)***	***(0.036)***	(<0.001)	(<0.001)	

Panel B: Sensitivity to the choice of conservatism measure (CON)

Measure	Intercept	*C_Score*	σ_ε	*C_Score* * σ_ε	*Beta*	*lnSIZE*	*Adj R*^{2}
		(–)	(+)	(–)	(+)	(–)	
C_Score	0.129	−0.013	***0.123***	***−0.588***	0.039	−0.010	0.228
	(< 0.001)	(0.325)	***(0.015)***	***(0.025)***	(<0.001)	(<0.001)	

Panel C: Sensitivity to the choice of information asymmetry measure (IA)

Measure	Intercept	CON_{GH}	*(1 – DISC)*	CON_{GH} * *(1 – DISC)*	*Beta*	*lnSIZE*	*Adj R*^{2}
		(–)	(+)	(–)	(+)	(–)	
DISC	0.131	−0.029	***0.010***	***−0.103***	0.038	−0.008	0.273
(n = 1782)	(< 0.001)	(0.347)	***(0.032)***	***(0.023)***	(<0.001)	(<0.001)	

Variable definitions: *r* is the cost of equity capital, alternatively estimated using the Easton (2004) modified PEG measure alternatively calculated after adjusting for predictable analyst forecast error using the de-biasing approach proposed by Larocque (2012) (*MPEG*^{adj}) and based on earnings forecasts generated by the Hou et al. (2012) cross-sectional approach (*MPEG*^{HVZ}), and the Easton (2004) long-term PEG measure as described in footnote 4 also based on earnings forecasts generated by the Hou et al. (2012) cross-sectional approach (*PEGLT*^{HVZ}); *CON* is a firm-specific measure of reporting conservatism, estimated alternatively using the Khan and Watts (2009) asymmetric timeliness measure (*C_Score*); information asymmetry (*IA*) is measured *DISC* equal to the total weighted disclosure score expressed as a percentage of total points available from the AIMR report; *Beta* is the value-weighted market-model beta estimated over the preceding 120-month period; and *Size* is the market value of common equity in $millions.
p-values (in parenthesis) are two-tailed, with corrections for clustering of standard errors by industry and year.

possible mechanical relationship. Notwithstanding, we also acknowledge the very clear and explicit statement by Botosan et al. (2011: 1119) that ‘we caution against the use of realized returns to proxy for cost of equity capital before or after controlling for cash flow news’.

5.3.2. Sensitivity to the selected proxy for conservatism. Panel B of Table 5 presents results for our econometric model run with the Khan and Watts (2009) *C_Score* measure of conservatism. We return to our primary measure of cost of equity capital, *MPEG*, and continue to use σ_ε as the proxy for information asymmetry. As revealed, the coefficient on *C_Score* is negative but insignificant (−0.013, p = 0.325), while the coefficient on the interaction term, *C_Score* * σ_ε, is negative and significant (−0.588, p = 0.025). The coefficient on σ_ε remains positive and significant (0.123, p =

0.015). Based on these estimated coefficients, the linear restriction relating to the conditional coefficient on *C_Score* becomes significant at the 5% level for values of σ_ε above 0.0297 (representing the 48.9% of the sample observations with the highest measures of σ_ε), and for a firm ranked at the 75th percentile in terms of σ_ε, a one-standard-deviation increase in *C_Score* is estimated to result in a 61-basis-point reduction in cost of equity capital. Thus, these results continue to provide strong support for H_1 and indicate that our results are robust to the choice between CON_{GH} and *C_Score* as the proxy for conservatism.

5.3.3. Sensitivity to the selected proxy for the firm's information environment. Finally, we alternatively proxy for information asymmetry using the measure of the firm's disclosure activity based on the AIMR ratings (*DISC*). This analysis is restricted to the 1782 firm-year observations covering the period 1985–94 for which the ratings can be obtained. We view *DISC* as a reasonable alternative measure because it is a direct measure of the firm's information environment, and because conservatism and disclosure represent choices made by management wherein they can explicitly determine the relative trade-off. Since information asymmetry is decreasing in *DISC*, we run the model using ($1 - DISC$) to allow for a consistent interpretation of the coefficient estimates.

The results, presented in Panel C of Table 5, indicate that conclusions are also robust to the choice of proxy for information asymmetry. The coefficients (*p*-values) on CON_{GH} and ($1 - DISC$) are −0.029 ($p = 0.347$) and 0.010 ($p = 0.032$), respectively, and the coefficient on CON_{GH} * ($1 - DISC$), is −0.103 ($p = 0.023$). These results confirm a negative and significant relation between cost of equity capital and conservatism for the high information (low *DISC*) firms. The linear restriction capturing the conditional coefficient on CON_{GH} becomes significant at the 5% level for values of ($1 - DISC$) above 0.513 (representing the 48.7% of the sample observations with the lowest levels of disclosure). For firms with the lowest disclosure percentile rankings ($DISC = 0.000$), a one-standard-deviation increase in CON_{GH} is associated with a 67-basis-point reduction in cost of equity capital.

6. Summary and conclusion

In this study, we seek insights into the economic consequences of accounting conservatism by examining the relation between conservatism and cost of equity capital. Based on both the analytical and the empirical literatures, we posit an inverse association, with greater accounting conservatism mapping into a reduced cost of equity capital. Importantly, we also posit that the strength of the relation is conditional on the firm's information environment, being the strongest for firms with high information asymmetry and the weakest (and potentially negligible) for firms with low information asymmetry. Underlying this prediction is the supposition that conservatism represents a part of the overall communication-reporting strategy of the firm.

We base our analysis on a sample of 3138 firm-year observations from the period 1985–2000. This period has been selected to pre-date the current debate regarding the merits of neutrality as opposed to conservatism (prudence) as a desirable characteristic of accounting numbers (FASB, 2010; IASB, 2010) and to pre-date significant changes in disclosure requirements brought about by regulatory change such as the adoption of SOX in 2002. We argue that by considering a pre-2000 experimental window, a firm's decision to adopt conservative accounting policies reflects a strategic decision unaffected by the debate or by regulatory change.

Our results provide unequivocal support for our predictions. We find an inverse association between conservatism and cost of equity capital and, further, that the marginal impact of conservatism systematically declines as the firm's information environment improves. The results of sensitivity analyses reveal our conclusions to be robust to the choice of proxy for each of the three main variables: cost of equity capital, conservatism and information asymmetry.

As argued, we believe that this study has the potential to make a contribution from the perspective of both the regulator and the academic literature. From a regulatory perspective, we interpret the findings as indicative that a firm's decision to adopt conservative reporting practices has the potential to provide real economic benefits and, thereby, that accounting conservatism has a potentially beneficial role within accounting principles and practices. In this fashion, the findings appear to draw into question the path enunciated in the 2010 IASB and 2010 FASB exposure drafts away from conservatism as a desirable characteristic of financial reporting.

In terms of the academic literature, this study contributes to the growing body of literature on accounting policy choice which focuses on market (economic) benefits. By exploring the interaction between conservatism and the firm's information environment, and their joint influence on the cost of equity capital, it extends our understanding of accounting conservatism and the role that it plays in the firm's overall communication strategy.

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Notes

1. Clarkson et al. (1996) raise the spectre that information risk may in fact have a diversifiable element. Here, for example, Lambert and Verrecchia (2010) suggest that market competition may play a mitigating role.
2. Gietzmann and Ireland (2005: 602) find that 'the effect of disclosure on cost of capital is only significant for those firms that make aggressive accounting choices'. Francis et al. (2008: 60) suggest that this finding is consistent with Gietzmann and Trombetta (2003), whose work they interpret as implying that 'voluntary disclosure is predicted to affect the cost of equity capital only for firms with aggressive accounting policies'.
3. Since, as discussed in Section 3.3, Beaver and Ryan (2000) propose the market-to-book ratio as a measure of accounting conservatism, inclusion of the third Fama and French (1993) risk fact, book-to-market, is likely to be problematic. Nevertheless, when included in the model, results (untabulated) are qualitatively identical to those based on the primary equation (1).
4. The long-term *PEG* measure is estimated as $\sqrt{(feps_5 - feps_4)/p_0}$. When based on analyst forecast data, this measure is only available for 703 firm-year observations, with descriptive statistics revealing this subsample to be more conservative, larger and more profitable, and to have lower information asymmetry, higher analyst following and lower betas relative to the full sample.
5. As pointed out by Wang et al. (2009), while in principle the Basu measure can be estimated from time-series data, Givoly et al. (2007) reveal that when estimated in the time series, it performs poorly in terms of its ability to detect conservatism.

6. Non-operating accruals (*NOACC*) are calculated as total accruals minus operating accruals where total accruals is equal to net income before extraordinary items plus depreciation minus cash flow from operations, and consist primarily of items such as bad and doubtful debt allowances, restructuring charges, the effect of change in estimates, gains or losses on disposal of assets, asset write-downs, accrual and capitalisation of expenses and deferral and recognition of revenue (Givoly and Hayn, 2000). Accruals can be determined using either the indirect or the direct approach (Dechow and Dichev, 2002; Givoly and Hayn, 2000; Hribar and Collins, 2002). Since the direct approach using the Statement of Cash Flow was not available prior to 1988, when SFAS No. 95 became effective, we adopt the indirect method, using the balance sheet to estimate accruals. Although this approach suffers from measurement error, especially for firms with discontinued operations or subject to merger and acquisition activity, it provides a consistent measure over our study period (1985–2000) (Francis et al., 2004; Hribar and Collins, 2002).
7. As an additional information asymmetry measure, we also considered the extent of analyst following (*AF*) (Botosan, 1997). We find all results (untabulated) and conclusions to be robust to the use of this measure as a proxy for information asymmetry.
8. To address the potential influence of outliers, we alternatively determined the Cook's Distance to identify influential data points in the regression analysis (Myers, 1989), 'windsorised' at the three-standard-deviation level, and truncated all variables at the 1% and 99% levels annually. In each instance, results remain qualitatively similar to those reported in Table 3. Thus, results and conclusions do not appear to have been driven by outliers within the data. Further support follows from the rank analysis reported in panel B, where the ranking process mitigates the potential influence of outliers.
9. Additional analyses (untabulated) reveal that the results are robust to the inclusion of *LEV* as an additional control variable.
10. The Abarbanell/Bernard *t*-statistic adjusts for the estimated first-order autocorrelation in the independent variables over the sample period by adjusting standard errors using following factor: $\{[(1 + \varphi) / (1 - \varphi)] - [2\varphi(1 - \varphi^n) / n(1 - \varphi)^2)]\}^{1/2}$, where n is the number of years and φ is the estimated first-order autocorrelation in the yearly coefficients. This adjustment factor assumes that the serial correlation is first-order autoregressive. The Z_1 statistic, which assumes residual independence, is $(1/n)^{1/2}\, \Sigma^n\, [t_i / \{ k_i / (k_i - 2)\}^{1/2}]$, where t_i is the White's *t*-statistic for year n, k_i are the degrees of freedom and n is the number of years. The Z_2 statistic is: mean *t*-statistic / (standard deviation of *t*-statistics / $\{n - 1\}^{1/2}$). See White (1984) for further support.
11. *MPEG* relies on the earnings–price relation, while CON_{GH} is indirectly related to the level of earnings. As such, the possibility of a mechanical relationship exists. We do not, however, view such a possibility as overly likely, since *MPEG* is based on the forecasted year-over-year change in earnings while CON_{GH} is based on the six-year average non-operating accruals to total asset ratio. Notwithstanding, by alternatively adopting realised returns, we consider the relationship where the dependent variable has not been mechanically calculated using forecasts of earnings.

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