Conservatism, Disclosure and the Cost of Equity Capital

Tracy C. Artiach*

UQ Business School, The University of Queensland, Brisbane Australia 4072

Peter M. Clarkson

UQ Business School, The University of Queensland, Brisbane Australia 4072
and
Faculty of Business Administration, Simon Fraser University, Burnaby Canada V5A 1S6

Current version: July 2010

Key Words: conservatism, disclosure, the cost of equity capital, signaling, information risk.

JEL classification: M41

*Corresponding author: Tracy Artiach: UQ Business School, The University of Queensland, Brisbane, QLD 4072
t.artiach@business.uq.edu.au; Phone: +61 7 3346 8005

The authors would like to thank Gordon Richardson, Keitha Dunstan, Steven Taylor, Zoltan Matolcsy, Philip Gray, Julie Cotter, Anne Wyatt, Irene Tutticci, Karen Benson, and participants of the University of Southern Queensland and University Technology of Sydney research seminars, and The University of Queensland Business School Research Colloquium for their helpful comments on earlier version of this paper.

Conservatism, Disclosure and the Cost of Equity Capital

Abstract

This paper examines the individual and joint impact of conservatism and disclosure on the cost of equity capital. Motivated by a lack of empirical evidence regarding the economic consequences of conservatism, we find as predicted, evidence of an inverse relation between firm level conservatism and the cost of equity capital for a sample of U.S. listed entities, but that this relation is diminished in environments of high disclosure. Thus, the evidence indicates the existence of positive economic consequences to 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.

Key Words: conservatism, disclosure, the cost of equity capital, signalling, information risk.

JEL classification: M41

1. Introduction

In this study, we seek insights into the economic consequences of accounting conservatism by examining the relationship between conservatism and cost of equity capital. Appealing to both the analytical and empirical literatures, we posit a negative association, with greater accounting conservatism mapping into a reduced cost of equity capital. However, we also posit that the strength of this relation is conditional upon the firm's information environment, with it being the strongest for firms for which the greatest information asymmetry exists and the weakest (and potentially negligible) for firms in the highest information environment. Underlying this prediction is the supposition that conservatism and disclosure represent jointly-determined strategies which form a part of the overall financial reporting strategy of the firm.[1]

As elaborated on in the next section, this study finds its theoretical roots in the models of Gietzmann and Trombetta (2003) and Bagnoli and Watts (2005) who provide a signalling benefit to conservative accounting policy choice. Specifically, Bagnoli and Watts show that in the presence of asymmetric information, conservative accounting policy choice can be used to communicate management's private information, which arguably should then lead to a reduced cost of equity capital. Of more direct relevance, Gietzmann and Trombetta envisage a role for both the adoption of conservative accounting policies and voluntary disclosure within the context of their model. They conclude that the economic implications of either accounting policy choice or voluntary disclosure can not be meaningfully studied in isolation but rather, each has the potential to influence cost of equity capital and additionally, that an investigation the economic benefits of accounting policy choice must "control for the voluntary disclosure strategy of firms."[2]

---

[1] Theoretical support for such a relationship is provided by Gigler and Hemmer (2001) (see footnote 38 for greater detail).

[2] In developing our arguments, the perspective we adopt is that of the impact of conservative accounting policy choice on cost of equity capital, conditional on the firm's information environment. Equally, by symmetry and through the arguments of Gietzmann and Trombetta (2003), we could adopt the converse perspective (i.e., the impact of disclosure on cost of equity capital, conditional on the firm's choice of conservatism). While we recognize both as legitimate perspectives, within the context of this study, our

The study also finds its roots in several related empirical studies. These include Francis, LaFond, Olsson, and Schipper (2004) who examine the relation between seven accounting-based attributes of earnings including conservatism and the cost of equity capital, Francis, Nanda, and Olsson (2008) who examine the effect of voluntary disclosure on the cost of equity capital conditional on the underlying quality of earnings, and Ahmed, Billings, Morton, and Stanford-Harris (2002) who find accounting conservatism to be associated with a lower cost of debt. 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.

Continuing, the predominance of research into conservatism focuses on the existence of conditional conservatism and the contracting incentives of conservatism to mitigate agency conflict (Watts 2003a). Despite theoretical propositions supporting a signalling benefit for conservatism at the firm level (Gietzmann and Trombetta, 2003; Bagnoli and Watts, 2005), the empirical evidence to date has paid little attention to the economic consequences of conservatism from perspectives alternative to agency theory. Finally, conservative reporting practices are optimally selected when they offer greater benefit or impose less cost on the firm relative to alternative financial reporting strategies such as, for example, disclosure.[3] The voluntary disclosure research to date has, for the most part, ignored the impact and potential information content of accounting policy choice, focusing almost exclusively on the association between voluntary disclosure and the cost of equity capital. These gaps in the literature also in part form the motivation for the current study.

Within the context of this study, we view the construct of 'accounting conservatism' as reflecting the firm's overall propensity to undertake on average conservative accounting choices, a view consistent with a traditional definition of conservatism. Conservatism is a function of the firm's

---

interest is the economic impact of conservatism on which there is paucity of evidence within the literature. We therefore adopt the 'conservatism' focus throughout the development of arguments and discussion of results.

[3] For example, the choice of conservative accounting may reduce a firm's ability to meet or beat earnings forecasts and therefore impose a high cost on the firm. Alternatively, the tension for disclosure could include proprietary or litigation costs.

cumulative accounting policies which arise from both discretionary and mandatory policy decisions.[4] Within both mandated and discretionary accounting policy decisions, there are degrees of discretion. For example, although a mandatory requirement, the assessment of net market value required for the application of the lower of cost or net market value involves judgement. This construct also captures aspects of what have been labelled as conditional and unconditional conservatism.[5]

We base our analysis on a sample of 1,782 firm-year observations over the ten-year period 1985 – 1994. Appealing to Botosan and Plumlee (2005) and Botosan, Plumlee, and Wen (2009) for support, we use the Easton (2004) PEG measure as our cost of equity capital estimate. Further, recognizing as described immediately above, the need for a firm-specific measure and also that the extent of conservatism encapsulated within a firm's accounting policies will reflect both mandatory accounting policy adoption and discretionary policy choice, we employ the Givoly and Hayn (2000) negative accruals measure as our proxy for conservatism. Finally, recognizing that the firm's information environment can be influenced by the firm's attitude towards disclosure as well as by external factors such as the extent of analyst following, we conduct our primary analysis using the analyst ratings data published in the Association for Investment Management and Research (AIMR) reports as a proxy for firm disclosure but also conduct sensitivity analysis using analyst following.

Our results provide consistent and unequivocal support for our prediction. Specifically, we find a negative association between our conservatism measure and cost of equity capital estimate, and

---

[4] Mandated conservatism arises from compliance with specific accounting regulation requiring the adoption of a more conservative accounting policy; for example, the expensing rather than capitalizing of R&D or the *ex post* application of the lower of cost or net market value rule for inventory. Discretionary conservatism arises from the firm's decision to select a more conservative accounting policy; for example, the *ex ante* selection of accelerated depreciation over straight-line depreciation or the selection of LIFO over FIFO inventory cost allocation.

[5] Given the mature nature of the literature on conservatism, we only provide a brief overview of the concepts here. Conditional conservatism refers to the manifestation of conservatism as the asymmetry in response of earning to good and bad news measured by the method developed in Basu (1997), arising when the book values are written down under sufficiently adverse circumstances but not written up under favourable circumstances (for example; the application of the lower of cost or market value rule for inventory; the impairment of non-current assets). Conditional conservatism is also referred to as ex post, news-dependent, news-related or income statement conservatism. Conversely, unconditional conservatism refers to the systematic bias in book value relative to market value (Beaver and Ryan, 2005) arising from aspects of the accounting process determined at the inception of assets and liabilities yielding expected unrecorded goodwill (for example, the use of accelerated depreciation or the use of last-in-first-out inventory cost allocation). Unconditional conservatism is also referred to as ex ante, news-independent, news-unrelated or balance sheet conservatism.

further that the marginal impact of conservatism systematically declines as the firm's information environment improves. We find these results to be robust to model form and perspective, and to the choice of proxy for each of our fundamental measures, cost of equity capital, conservatism, and the firm's information environment. Finally, we present preliminary evidence in support of the notion that conservatism and disclosure play a joint role in the financial reporting strategy of the firm.

We argue that this study has the potential to make a contribution from both the regulatory and academic perspectives. 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 conjunction, we also view them as giving credence to the continued observation of conservative reporting practice. Thus, in this fashion, they appear to draw into question the path enunciated in the 2008 IASB/FASB exposure draft away from conservatism as a desirable characteristic of financial reporting. Specifically, the IASB/FASB 'Conceptual Framework for Financial Reporting Exposure Draft' (2008) reveals the Board's intention that conservatism (prudence) not be included as a desirable qualitative characteristic of financial reporting because of its potential to introduce bias and thereby conflict with the desire for neutrality.[6]

In terms of the academic literature, the study contributes to the growing body of literature on accounting policy choice which focuses on the market (economic) benefits of accounting policy decisions. By exploring the interaction between conservatism and disclosure, and their joint

---

[6] In the IASB/FASB Exposure Draft: 'Conceptual Framework for Financial Reporting: The Objective of Financial Reporting and Qualitative Characteristics and Constraints of Decision-Useful Financial Reporting Information' issued in May 2008, the IASB/FASB indicate that: "…Being careful in the presence of uncertainly includes searching for additional information to reduce uncertainly, reflecting the uncertainty of a range of potential amounts in making an estimate, or selecting an amount from the midpoint of a range if a point estimate is required. Going beyond that to reflect conservative estimates of income and equity sometimes has been considered desirable to ensure that financial reports do not reflect excessive optimism or bias on the part of management. 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

influence on the cost of equity capital, it extends our understanding of accounting conservatism and the role that it plays within the firm's overall reporting strategy. In this regard, we argue that it contributes to the accounting policy choice literature. We argue that it also contributes to the literature by adopting a theoretical foundation from the signalling/information risk literature rather than the more commonly employed agency arguments, finding empirical support for arguments emanating from the former.

Finally, our findings are in contrast with the most directly relevant study in the literature, that of Francis, LaFond, Olsson, and Schipper (2004) who examine the relation between seven accounting-based attributes of earnings including conservatism and the cost of equity capital. They find that, on balance, firms with the least favourable values of the accounting-based attributes generally have higher cost of equity capital but fail to find an association between conservatism and the cost of equity capital using a firm-level Basu (1997) measure (conditional conservatism) estimated in time series. This study seeks to extend the insight provided by Francis *et al.* (2004) that accounting-based attributes have an impact on information quality to a more focused investigation into the economic consequences of the adoption of conservative accounting policies. We argue, based on a signalling framework, that the Basu (conditional) conservatism measure adopted by Francis *et al.* (2004) does not adequately reflect the required proxy. Further, Givoly, Hayn and Natarajan (2007) suggest that the Basu measure may be unreliable when estimated in time series as in Francis *et al*. In this sense, by using what is arguably a more appropriate proxy for the underlying construct of accounting conservatism within this particular context, the Givoly and Hayn (2000) negative accruals measure, we also have the opportunity to contribute to the debate surrounding proxy selection as well as confirm the economic benefits of conservatism.

The paper proceeds as follows. Section 2 presents a brief literature review, culminating in 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, a summary and conclusions.

2.0 Literature Review and Hypothesis Development.

2.1 Overview

As outlined in the Introduction, in this study we explore the relationship between the extent of conservatism encapsulated within a firm's accounting policies and its cost of equity capital. As formalized in our hypothesis presented at the end of this section, we posit an inverse relationship between conservatism and cost of equity capital, but only for firms for which information asymmetry is the greatest. For firms for which information asymmetry is low, we alternatively expect conservatism to have a negligible impact on their cost of equity capital. In this regard, we envisage as modelled by Gietzmann and Trombetta (2003) and documented by Artiach (2009), conservatism and information as substitute mechanisms through which firms can influence their cost of equity.

To present the background for the current study and provide it with an underlying theoretical foundation, in the material that follows we first briefly review relevant studies that focus on disclosure. We then briefly review relevant literature on conservatism, followed by a review of studies more directly related to the current study. Finally, we review the directly relevant analytical literature to lay the theoretical foundation for our hypothesis.

2.2 Empirical Literature

Within the literature, there are a number of studies that explore narrowly the relationship between disclosure and cost of equity capital. Of these, two of the most widely cited are Botosan (1997) and Botosan and Plumlee (2002). In motivating these studies and others, authors typically appeal to two sources of theoretical support. The first is the argument that greater disclosure will enhance market liquidity and thereby reduce cost of equity capital through reduced transactions costs and/or increased demand (Botosan and Plumlee, 2002). The second is the argument that disclosure reduces the estimation/information risk associated with the estimation of the parameters of the return distribution (Barry and Brown, 1985).[7] Botosan (1997) explores the relationship between cost of equity capital and disclosure using a sample from a single year and one industry, and using a researcher scored disclosure index. She finds as predicted, an inverse relationship but only for firms with low analyst following. Botosan and Plumlee (2002) extend the analysis to multiple industry sectors and a longer time period, and use AIMR Reports as the source of their disclosure scores. They find that greater annual report disclosure results in a lower cost of equity capital but that greater other publications disclosure is associated with a higher cost of equity capital. Thus, together these results seem to suggest that the benefits of disclosure in terms of a firm's cost of equity capital are dependent on both its information environment and the choice of disclosure outlet.

Conversely, arguably with the exception of Francis *et al.* (2004), to date there have been no studies that narrowly examine the relationship between conservatism and cost of equity capital. However, there have been a number of studies which examine conservatism from a variety of perspectives. For example, there is a significant body of evidence which indicates that financial reporting has become more conservative over time (Basu, 1997; Pope and Walker, 1999; Ball Kothari, and Robin, 2000; Givoly and Hayn, 2000; Holthausen and Watts, 2001).

In summarizing the empirical evidence, Watts (2003a, 2003b) identifies four factors as those commonly advanced as influencing the extent of conservatism: (1) contracting; (2) litigation; (3) income taxation; and (4) accounting regulation. Since the practice of conservatism pre-dates the

---

[7] While displaying modelling differences, subsequent models such as Lambert, Leuz, and Verrecchia (2007) continue to predict an inverse relation between disclosure and cost of equity capital. Notwithstanding these models, as noted by both Botosan (1997) and Botosan and Plumlee (2002), Clarkson, Geudes and Thompson (1996) raise the spectre that estimation/information risk may in fact have a diversifiable element.

litigation, income taxation and regulation explanations, the contracting explanation has received the greatest attention, providing a foundation for both the existence and pervasive nature of conservative accounting practices (Watts 2003a). Existing research provides evidence at the aggregate level of the variation in conservatism across firms, change in conservatism across time periods, the variability of conservatism over quarters, the influence of audit work and auditors' exposure to legal liability on conservatism, how conservatism influences board composition, and its cross-country variation (e.g., Basu, 1997; Pope and Walker, 1999; Ball *et al.*, 2000; Givoly and Hayn, 2000; Ball and Shivakumar, 2005; Roychowdhury and Watts, 2007).[8] Further, Watts (2003a) argues that conservatism, in its verifiability role, continues to be evident in accounting and serves a positive function in contracting efficiencies with evidence, inter alia, of conservatism mitigating agency conflicts and thereby reducing the cost of debt and improving corporate governance (Ahmed *et al.*, 2002; Ahmed and Duellman, 2007). In sum, considering its long-standing history, it is highly likely that conservatism is viewed by management as adding value and/or providing economic benefit. Further, contracting parties (including debt-holders and shareholders) would be unlikely to impose conservative restraints if they perceived that there were no effective benefits. To the counter, aggressive accounting practices and policies are often met with suspicion by market participants. Notwithstanding, as noted at the outset, the economic consequences of conservatism in terms of its impact on the cost of equity remains relatively unexplored to date.

Turning next to literature more directly relevant to the current study, preliminary evidence regarding the economic consequences of conservatism is included in a study by Francis *et al.* (2004) who examine the relationship between the cost of equity capital and seven attributes of earnings: accruals quality, persistence, predictability, smoothness, relevance, timeliness, and conservatism. Using the Basu methodology and a sample of U.S. listed firms for the period 1975 – 2001, Francis *et*

---

8 An extensive review of each of these areas is provided by Watts (2003b) and Givoly, Hayn, and Natarajan (2007).

al. (2004) fail to find support for an association between cost of equity capital and conservatism. Their result is, however, perhaps not surprising given that their measure of conservatism essentially captures an *ex post* response to market information essentially required by mandatory accounting policies. More importantly, Givoly *et al.* (2007) reveal the Basu measure when estimated in time series at the firm level to be unreliable.[9] The contribution of the current study, relative to Francis *et al.* (2004), is the use of a more reliable measure for conservatism which more completely reflects the relevant underlying theoretical construct.

Finally, Francis *et al.* (2008) provide more recent evidence on the effect of voluntary disclosure on the cost of equity capital, and on the effect conditional upon the influence of earnings quality. They find a positive association between earnings quality and voluntary disclosure, indicative of a complementary relationship. They also provide further evidence of a negative association between voluntary disclosure and the cost of equity capital, but find this association to be substantially diminished when controlling for the influence of earnings quality. While their results are robust to alternative measures for the cost of equity capital and earnings quality, Francis *et al.* (2008) find that when using other disclosure proxies (including management forecasts, conference calls, press relations), the results are not consistent.

2.3 Theoretical Foundations

As noted above, historically the greatest attention paid within the literature to the construct of accounting conservatism has been from the contracting perspective, with agency arguments underlying much of the empirical work. For example, LaFond and Watts (2008) argue that conservatism is a rational response to information asymmetry. Specifically, they hypothesize that the demand for conservatism increases as an equilibrium corporate governance response aimed at

---

[9] Gigler and Hemmer (2001) and Dietrich *et al.* (2007) provide further theoretical and empirical evidence which casts significant doubt on the econometric validity of the Basu methodology.

lowering the agency costs associated with managerial overstatement of income and assets. Based on a sample of U.S. listed firms, they find as predicted that conservatism increases in response to increases in information asymmetries in equity markets and conclude that their results highlight conservatism as an equilibrium response to mitigate value reductions arising from these information asymmetries. They do not, however, consider the economic consequences of conservatism.

In contrast with the agency framework underlying studies such as LaFond and Watts (2008), recently several theoretical studies have emerged that envisage an economic role for accounting policy choice from a signalling perspective. Of direct interest, these models envisage an inverse association between accounting conservatism and the cost of equity capital.[10] To begin, along this vein, Easley and O'Hara (2004) propose that in equilibrium, information quantity and quality affect asset prices and show that 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, and through market microstructure. In this sense, 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.

Continuing, the theoretical work most directly relevant to the current study is that of Gietzmann and Trombetta (2003) who provide support for a signalling benefit to the adoption of conservative accounting policies. Specifically, they 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. Here, the voluntary adoption of conservative accounting policies is not a costless signal because it reduces the amount of

[10] In contrast, absent a theory to the contrary, it could be argued that following the arguments of LaFond and Watts to their logic conclusion, their prediction would imply a positive contemporaneous relationship between conservatism and cost of equity capital since increased information asymmetry would lead to higher information risk and therefore an increased cost of equity capital. This is direct opposition to the negative relationship envisaged by the signalling arguments of Gietzmann and Trombetta (2003) and Bagnoli and Watts (2005) on which we base the current study, thereby creating a potentially interesting tension.

financial slack available to the manager which, for example, can reduce their ability manage earnings to meet or beat earnings forecasts. Thus, high quality firms can afford to incur the costs of a more conservative balance sheet whereas their lower quality counterparts cannot and so will not seek to mimic the conservative accounting policy choices. In greater detail, Gietzmann and Trombetta's (2003) theoretical model shows 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. They demonstrate that the optimal disclosure policy for some firms is simply to disclose choice of a conservative accounting policy (p. 187). Thus, firms will use some combination of signals involving conservative accounting choices and/or direct disclosures.[11]

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. In this fashion, there is a parallel with a large number of studies that have investigated the capital market impacts of disclosure policy 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). Whether information risk is diversifiable remains subject to considerable debate in the literature. For example, Lambert, Leuz, and Verrecchia (2008) conclude that the pricing effect in the Easley and O'Hara (2004) model can be diversified away in large economies. However, notwithstanding this conclusion, they also theoretically demonstrate that firm-specific information risk is priced *despite* diversification.[12] This is consistent with the analytical proposition of Clarkson, Guedes and Thompson (1996) that estimation risk has a meaningful and measurable non-

---

[11] In this study, following models such as that of Gietzmann and Trombetta (2003), we propose that firms can use conservative accounting policy choice to signal firm quality and thereby reduce firm-specific information risk. Support for this type of argument is not limited to the theoretical work cited above. It is also supported in concept by the accounting policy choice literature which provides empirical evidence of how firms use accounting policy choice to signal firm type (see, for example, the work of Holthausen (1981), Leftwich (1981), and Holthausen and Leftwich (1983)).

[12] In Lambert *et al.* (2008), the quality of accounting information is directly affected by market participants' perceptions of the distribution of future cash flows and indirectly affected by real decisions that alter the distribution of future cash flows. The direct effect is not diversifiable because the quality of disclosure affects the assessed covariances between the firm's cash flow and other firms' cash flows. In this fashion, they theoretically demonstrate that firm-specific information risk is priced despite diversification.

diversifiable component. Thus, 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.[13]

2.4 Hypothesis

The theoretical literatures reviewed above consistently identify a signalling benefit to accounting conservatism. In this sense, the theory would imply an inverse relation between conservatism and cost of equity capital. However, it also suggests that the relation is somewhat more complex, with the firm's information environment (and thereby its disclosure policy) representing a mitigating factor. For example, Bagnoli and Watts (2005) show within the context of their model that it is only in the presence of asymmetric information that conservative accounting policy choices can be used to communicate management's private information. More directly, Gietzmann and Trombetta (2003) formally model the trade-off between voluntary disclosure and signalling through the firm's choice of conservative accounting policy to reveal management's private information about firm value. They conclude that "the value relevance of either accounting policy adoption or voluntary disclosure can not be meaningfully studied in isolation. Any study of the value relevance of accounting needs to control for the voluntary disclosure strategy of the firm." They then continue on to argue that "Similarly, any study of the link between cost of capital and voluntary disclosure should control for differences in accounting policy adoption."

Thus, as implied by Gietzmann and Trombetta (2003), conservatism and disclosure perspectives are both legitimate perspectives from which to frame arguments and empirical tests regarding their economic consequences. The focus of this study is the economic consequences of conservative accounting policy choice which we investigate by examining the relationship between

[13] These theoretical propositions are consistent with the results and suggestions in Ahmed *et al.* (2002) that the reputation effect from conservative accounting appears to reap real economic benefit in the form of reductions in the cost of debt.

conservatism and cost of equity capital. Thus, we present and test our hypothesis from the perspective of conservatism. Specifically, we posit an inverse relation between the degree of conservatism present in a firm's accounting policy choice and its cost of equity capital. Further, we posit that the strength of this relation is dependent on the firm's information environment, with it being the strongest for firms for which the greatest information asymmetry exists and weakest for firms for which information asymmetry is low. 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 (high disclosure).

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

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

In order to test the relationship between cost of equity capital and conservatism described in H_1, we employ the following econometric model:

$$r_{i,t} = \gamma_0 + \gamma_1\, CONSV_{i,t} + \gamma_2\, RDISC_{i,t} + \gamma_3\, CONSV_{i,t} * RDISC_{i,t} + \gamma_4\, Beta_{i,t} + \gamma_5\, BTM_{i,t} + \gamma_6\, lnSIZE_{i,t} + \varepsilon_i \quad (1)$$

where, in brief, the primary variables, cost of equity capital (r), conservatism ($CONSV$), and disclosure ($RDISC$), are measured as follows:

$r_{i,t}$ = cost of equity capital for firm i for fiscal year t, measured as the Easton (2004) PEG estimate;

$CONSV_{i,t}$ = the accruals-based measure of conservatism (Givoly and Hayn, 2000) for firm i for fiscal year t; and

$RDISC_{i,t}$ = the industry-year ranked total weighted disclosure score from the annual AIMR reports for firm i for fiscal year t (Lang and Lundholm, 1993).

These measures are discussed in detail in dedicated sections below. Based on H_1, we expect the coefficient on *CONSV* to be negative (i.e., $\gamma_1 < 0$) and the coefficient on the interaction term, *CONSV * RDISC*, to be positive (i.e., $\gamma_3 > 0$).[14] Consistent with the expectations of Botosan (1997) and Botosan and Plumlee (2002), we expect the coefficient on *RDISC* to be negative (i.e., $\gamma_2 < 0$).

The choice of control variables included in the model is guided by prior literature, notably Botosan and Plumlee (2002) and Francis *et al.* (2004), and comprises the three risk proxies known to influence the cost of equity capital (Fama and French, 1992, 1993). Specifically, the selected control variables are systematic risk (*Beta*), book-to-market (*BTM*), and firm size (*SIZE*), with the coefficients on *Beta* and *BTM* expected to be positive and the coefficient on *lnSIZE* expected to be negative. These variables are measured as follows:

$Beta_{i,t}$	=	the value-weighted market model beta for firm *i* for fiscal year end *t*, estimated over the preceding 120-month period (minimum 24 months);[15]
$BTM_{i,t}$	=	the book-to-market ratio for firm *i* for fiscal year end *t*, measured as the book value of common equity divided by the market value of common equity; and
$SIZE_{i,t}$	=	market capitalization of firm *i* at fiscal year end *t*.[16]

Finally, as will be discussed in Section 4 below, we conduct our analyses using a pooled cross-sectional sample of U.S. firms for the period 1985 to 1994. Here, the use of pooled cross-sectional time-series (panel) data creates an econometric issue. One of the underlying assumptions of OLS regression is that the regression errors terms are uncorrelated with homogeneous regression variance (Myers, 1989). The independent variables are expected to explain much of the observed differences, with the error term capturing the often-unobserved factors that have not been modelled. The problem arises when the unobserved factors are correlated and the errors are heteroscedastic. If

[14] As previously discussed (see footnote 2 and Section 2.4), our empirical focus is the impact of conservatism on the cost of equity capital conditional on the level of information asymmetry and as such, we interpret the coefficient estimate γ_3 accordingly.

[15] Monthly returns are obtained from the CRSP database. The risk-free rate of interest is the ten-year Treasury Constant Maturity Bond Rate, obtained from the U.S. Federal Reserve.

[16] Data for *BTM* and *Size* are obtained from the Research Insight Database.

OLS regression is used to estimate the model with panel data, the estimates will be biased and inefficient. We therefore estimate a series of generalised linear regression models with both year and firm fixed effects. The fixed effects method overcomes the issue of serial correlation by controlling for unobservable correlated omitted factors (Myers, 1989; Allison, 2008).

3.2 Cost of Equity Capital

Prior literature suggests several approaches to estimating the *ex-ante* firm level cost of equity capital (e.g., Gebhardt, Lee, and Swaminathan, 2001; Easton, 2004). These estimation approaches use price and analysts' earnings forecasts to derive an internal rate of returns as the estimate of the cost of equity capital. Botosan and Plumlee (2005) compare the validity of various proxies for the expected cost of equity, concluding that the Value Line Cost of Equity estimate and Easton (2004)'s PEG estimate outperform the other estimates. Botosan *et al.* (2009) reconfirm that the Value Line Cost of Equity and PEG estimates represent the most reliable proxies for the cost of equity capital because they are generally more consistent with realized returns and firm-specific risk characteristics. The Easton (2004) PEG estimate is, however, far less onerous in terms of data requirements because it only requires price and earnings growth data. We therefore use this estimate as our measure of the cost of equity capital, calculating it as follows:[17]

$$r = \sqrt{(feps_2 - feps_1) / p_0} \tag{2}$$

where

p_0	=	price per share at time 0;
$feps_2$	=	the two-year-ahead median (consensus) forecast of accounting earnings per share; and
$feps_1$	=	a one-year-ahead median (consensus) forecast of accounting earnings per share.

The data required to calculate this measure are obtained from Thomson's I/B/E/S Database.

[17] In untabulated analysis, we undertake a "validation" of this estimate within our sample data by regressing it on the Fama and French (1992, 1993) three risk proxies, Beta, book-to-market, and size. Here we find as expected, our estimate to be positively and significantly related to Beta and book-to-market, and negatively and significantly related to size.

Since as discussed by Francis *et al.* (2004), this type of estimate restricts the sample "in important ways" and further, since one of the methods which Botosan and Plumlee (2005) and Botosan *et al.* (2009) use to identify the most reliable proxies is realized returns, for sensitivity purposes we use one-year-ahead realized monthly returns as an alternative dependent measure (Callen *et al.*, 2010).[18]

3.3 Conservatism

Within the accounting literature, a universally accepted meaning for the construct 'conservatism' has been elusive. Perhaps one of the more commonly recognized is the descriptive definition advanced by Givoly and Hayn (2000) 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". Against this backdrop, it is thereby perhaps not surprising that a number of different proxies for conservatism have developed within the literature, with the various proxies reflecting different perspectives on conservatism. Of these, the most prevalent 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), and Givoly and Hayn's (2000) negative accruals measure. While all have received attention in the literature, the Basu measure has received the greatest (Wang, hOgartaigh, and van Zijl, 2010).

---

[18] One of the assumptions underlying the PEG estimation approach is that changes in abnormal earnings beyond the forecast horizon will be zero. Botosan and Plumlee (2002, 2005) suggest that long-run earnings forecasts can be substituted for the short-run earning forecasts to provide a greater likelihood that the change in abnormal earnings beyond the forecast horizon will be zero. In response, we alternatively estimated r using $feps_4$ and $feps_3$ instead of $feps_2$ and $feps_1$ in Equation (2). Based upon this alternative estimate of r, we no longer find support for H_1. Specifically, while the coefficients on *CONSV* and *RDISCL* remain negative and the coefficient on *CONSV*RDISCL* remains positive, none are statistically significant at conventional levels. However, use of the longer-run earnings forecasts also limits the sample, reducing it from 1,782 to 521 firm-year observation, with descriptive statistics revealing the reduced sample firms to be significantly more conservative (*CONSV*), better disclosers (*RDISCL*), larger, more profitable, and to have higher analyst following as well as lower Betas and lower book-to-market ratios relative to the full sample. Given these fundamental differences, the lack of statistical support for H_1 is perhaps not surprising, especially the lack of significance on the interaction term given that the reduced sample represents only the higher information environment sample firms.

For this study, our cross-sectional tests require a firm-specific measure of conservatism. Further, we require a construct which captures conservatism resulting from both mandatory and discretionary policy choices. In this regard, in general neither the Basu nor the Ball and Shivakumar proxies represent a firm-specific measure.[19] 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 the specific accounting standards such as those relating to R&D and inventory. Since the book-to-market ratio is a known risk factor and thereby already captured within our econometric model, we adopt the Givoly and Hayn (2000) measure as the most reasonable proxy for our purposes.

The intuition underlying the Givoly and Hayn (2000) 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 (Givoly and Hayn, 2000).

To calculate our proxy, we focus on non-operating accruals rather than total accruals because total accruals incorporate operating accruals which arise from the day-to-day business of the firm. Specifically, operating accruals exhibit positive accumulation over time consistent with growth in operations and thereby likely economic characteristics unrelated to conservatism (Givoly and Hayn 2000).[20] Further, when determined using non-operating accruals, the measure has the desired attribute of capturing accruals resulting from both mandated accounting regulation and discretion

[19] As pointed out by Wang *et al.* (2010), 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, the measure performs poorly in terms of its ability to detect conservatism.

[20] Non-operating accruals 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). In this regard, our proxy is consistent with the original proxy specification provided by Givoly and Hayn. In contrast, for their conservatism measure, Ahmed *et al.* (2002) use total accruals rather than non-operating accruals; hence their measure arguably reflects both growth in operations and conservatism.

accounting choice. To reveal their persistence, the accumulation of non-operating accruals must be measured over a sufficiently long period. 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. In sum, our accruals-based conservatism proxy (*CONSV*) is the ratio of non-operating accruals to total assets determined using the indirect method,[21] averaged over a six-year period, and multiplied by -1 to produce an increasing measure of conservatism (i.e., larger values of *CONSV* indicate greater conservatism). Thus, our measure is:

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

where $NOACC_{i,t}$ is non-operating accruals calculated as total accruals minus operating accruals for firm i for fiscal year end t and $TA_{i,t}$ is the total assets of firm i at fiscal year end t.[22]

3.4 Disclosure

We base our disclosure measure on the analyst ratings data published in the AIMR reports during the period 1985 to 1994. In this fashion, we follow a significant sector of the disclosure literature which includes Lang and Lundholm (1993, 1996), Healy, Hutton, and Palepeu (1999), Botosan and Plumlee (2002), and recently Drake, Myers and Myers (2009), among others, who appeal to the AIMR reports as the basis for their disclosure measures.[23]

---

[21] Accruals can be determined using either the indirect or the direct approach (Givoly and Hayn, 2000; Hribar and Collins, 2002; Dechow and Dichev, 2002). The direct approach using the Statement of Cash Flow is not available prior to 1988 when SFAS No. 95 became effective. We adopt the indirect method since our sample covers the period 1985 to 1994, 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 (Hribar and Collins 2002; Francis *et al.* 2004).

[22] Total accruals = Net Income before Extraordinary Items (IB) + Depreciation (DP) - Cash flow from operations. Cash flow from operation = Funds from operation (FOPT) - [ΔCurrent Assets (ACT) + ΔDebt in Current Liabilities (DLC) - ΔCurrent Liabilities (LCT) - ΔCash (CHE)]. Operating accruals = ΔAccounts Receivable (RECT) + ΔInventories (INVT) + ΔPrepayments (XPP) - ΔAccounts Payable (AP) - ΔAccrued Expenses (XACC) - ΔTaxes Payable (TXP). The data are obtained from the Research Insight database, with the relevant data mnemonic provided in parentheses.

[23] Results from these many disclosure studies collectively demonstrate the validity of the AIMR disclosure scores and support the conjecture that the AIMR disclosure rankings capture meaningful variation in the quality of disclosure practices across firms. For a

In brief, the 'Annual Review of Corporate Reporting Practices' was produced annually from report year 1978-1979 until 1995-1996 by the AIMR's Corporate Information Committee (CIC). The reports review the corporate reporting practices of a wide cross-section of industries. For each industry under review (on average nineteen industries annually), a subcommittee of approximately thirteen industry specialist financial analysts assessed the disclosure practices of firms based on three categories: (i) annual and required published information; (ii) quarterly and other published information; and (iii) investor relations. The subcommittees then developed a total company disclosure score as a weighted combination of these three categories.

To standardise the rating process, the CIC provided guidelines for the evaluation of company disclosure within each industry. The subcommittees then developed their assessment checklists from these guidelines. However, notwithstanding the guidelines, the subcommittees typically used discretion to develop checklists appropriately suited to the characteristics of their particular industry. In this regard, the number of points used to determine disclosure scores differs by subcommittee. The CIC also suggested weights to be applied to each of the three disclosure categories in the determination of the total score, although again the ultimate choice of weights was determined by each subcommittee.[24] Thus, on this basis, the subcommittees provided a consensus judgment score for each of the three categories, as well as a total weighted score.

We focus exclusively on the total disclosure in our analysis for the following reasons. First, within the context of our study, the disclosure measure represents a control variable designed to capture the firm's relative information environment. In this sense, we expect the total disclosure score to more fully reflect the information set available to investors. Second, use of the total disclosure score has the additional advantage that it is available for all firms within our sample. As

---

discussion on the relative advantages and disadvantages of the AIMR disclosure rankings specifically and analyst scores of disclosure more generally, see Lang and Lundholm (1996) and Healy *et al.* (1999).

[24] The suggested weights are 40-50% for the annual report, 30-40% for the quarterly and other published information, and 23-30% for investor relations (Botosan and Plumlee, 2002).

noted by Lang and Lundholm (1993) and recognised by Drake *et al.* (2009), for a non-trivial portion of the firms covered, the AIMR reports only include the total disclosure score.[25]

Finally, since each industry is evaluated separately by a subcommittee of analysts and the subcommittees used a degree of judgement in arriving at their total disclosure score, we follow the prior literature by converting the total weighted disclosure scores to industry-year percentile ranks and base our empirical analysis on these percentile rank data (*RDISC*).[26]

4. Sample Data

The preliminary sample consists of all U.S. listed firms with available AIMR disclosure rankings at some point over the ten-year period 1985 to 1994. The choice of time period is driven by our need for access to the AIMR disclosure data.[27] This provides an initial sample of 4,556 firm-year observations. From this, we then delete the following firm-year observations: those with insufficient data to determine the total weighted disclosure score (378 firm-year observations);[28] those where the firm could not be validly matched with a firm file on the Research Insight and/or I/B/E/S databases (145 firm-year observations); those lacking the requisite financial data to calculate the various measures used in the analyses (1,604 firm-year observations);[29] those from the financial institutions sector, given different regulations relating to both accounting and disclosure (712 firm-year

[25] For example, Lang and Lunholm (1993) note that for their sample which covers the 1985 – 1989 AIMR reports, the reports contain only the total disclosure score for "approximately 19% of the industries (38% of the firms)."
[26] The within industry-year percentile ranks are determined as: (rank – 1) / (# firms – 1).
[27] As noted in the Introduction, while the choice of sample period is driven by the use of the AIMR data, the study period selected facilitates two key advantages. First, the period predates regulatory changes which resulted in significant changes in disclosure and reporting requirements. Second, over the study period, conservatism was viewed a desirable characteristic of financial information.
[28] In several industry sectors, the AIMR reports provide the company's final rankings only without the provision of the raw scores required to determine the total weighted disclosure. For example, in the 1994/1995 report, the software and services subcommittee do not provide any scores but rather provide only a qualitative discussion regarding the company rankings and disclosure practices within the industry.
[29] Financial information used to calculate each variable is obtained from a variety of databases including 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 required to estimate the cost of equity capital causes the most significant sample attrition.

observations); and those from the two industry sectors with only one year of disclosure rankings data, given concerns regarding generalizability (11 firm-year observations).

Table 1 presents a frequency distribution by year and GICS industry code for the final sample of 1,782 firm-year observations. As can be seen, while there are slightly more observations over the sub-period 1990 to 1992, on balance the sample is relatively uniformly spread across the study period. The by-year distribution ranges from 7.7 percent of the total sample in 1985 to 12.1 percent in 1990. 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 10.0 and 9.5 percent of the total sample, respectively, while the smallest are in the international pharmaceutical and nonferrous mining sectors with 0.2 and 0.7 percent, respectively. Given this relatively broad-based industry representation, it is unlikely that results will be driven by specific types of industries such as, for example, capital intensive industries.

Table 2 presents descriptive statistics for the various regression model variables, as well as several additional measures, based on the pooled sample of 1,782 firm-year observations. As revealed, there is broad variation in all measures across the sample. The mean and median values of the cost of equity capital estimate (*r*) are 0.121 and 0.106, with a standard deviation of 0.059. These figures are similar to Easton (2004) who reports a median estimate of 0.113 for the sample period 1981-1999 and Easton and Monahan (2005) who report mean and median estimates of 0.110 and 0.106, with a standard deviation of 0.032.[30]

Our conservatism measure, *CON*, ranges from a minimum of -0.452 to a maximum of 0.412, with mean and median values of 0.032 and 0.036, respectively, and a standard deviation of 0.055. Unfortunately, it is difficult to directly compare descriptive statistics with prior studies, given

[30] The descriptive statistics cannot readily be compared with those reported in Botosan and Plumlee (2005). They report descriptive statistics for the cost of capital estimates minus the risk-free rate of interest represented by the Five-Year Treasury Constant Maturity Rate as at the end of the month in which the estimate is made.

differences in the way which conservatism is calculated. Nevertheless by way of preliminary context, for their measure based on total accruals rather than non-operating accruals, Ahmed *et al.* (2002) report mean and median measures of 0.004 and 0.003, respectively, with a standard deviation of 0.031 for a sample of U.S. firms from 1998. Alternatively, Givoly and Hayn (2000), who use non-operating accruals scaled by total assets, report mean and median values of -0.016 and -0.003 for the period 1951-1980, and mean and median values of -0.047 and -0.037 for the period 1981-1998.

Turning to the total weighted disclosure measure, the statistics presented in Table 2 are for the measure determined using the percentage of the total points available in each category (Lang and Lunholm, 1996). As revealed, the mean and median values for our sample are 0.735 and 0.750, with a standard deviation of 0.132. While slightly higher, these figures are similar to the mean and median figures of 0.70 and 0.72 reported by Lang and Lundholm (1996) for a sample of 2,272 firm-year observations for the period 1985-1989 (see their Table 1).[31] As expected, there is consideration variation in the disclosure measure, with a minimum value of 0.129 and a maximum value of 0.967.

Finally, none of the pair-wise correlations (not tabulated) among the measures in our econometric model exceed 0.269. Thus, there is unlikely to be a threat of multicollinearity.

5. Results

5.1 Pooled Sample Regression Results

Table 3 presents the results for variants of the econometric model (equation (1)) used to test our hypothesis, H_1. Recall, H_1 predicts a negative association between the conservatism measure, *CONSV*, and cost of equity capital, *r*. Further, it predicts that the strength of this association is conditional upon the firm's information environment, with *CONSV* exhibiting the strongest marginal

[31] The small difference is likely to be an artefacts of our sample selection process wherein smaller firms are more likely to have been deleted due to missing data required to calculate the cost of equity capital measure.

impact on cost of equity capital for firms with the greatest information asymmetry and a potentially negligible impact for firms in the highest information environment.

Primary results based on the pooled sample of 1,782 firm-year observations are presented in Panel A of Table 3. As previously discussed, the analyses are undertaken using generalized linear regression models with both year and firm fixed effects. All reported *p*-values are two-tailed. Finally, as an aside before considering our primary measures, the results for the control variables indicate that across all specifications, as expected the coefficient on *Beta* is positive and significant while the coefficient on *lnSize* is negative and significant. Alternatively, contrary to expectations, the coefficient on *BTM* is negative but typically insignificant.

Turning to the experimental variables, for the first specification, Model (1), which includes only the conservatism measure in addition to the control variables, we find as predicted, a negative and significant coefficient on *CONSV* of -0.102 (p = 0.002). Thus, the results from this specification provide preliminary support for the notion that firm level conservatism has an unconditional economic benefit in terms of a reduced cost of equity capital for our sample firms.

Second, to provide additional confidence in our measures and to align our results with the arguments and findings of Botosan and Plumlee (2002), we include only the disclosure measure in Model (2). Here, we find a negative but insignificant coefficient on *RDISCL* of -0.005 (p = 0.268), a finding consistent with Botosan and Plumlee who also fail to document an inverse relation using a similar total disclosure measure. Continuing, in Model (3) we include both *CONSV* and *RDISCL* 'unconditionally', finding again the coefficient on *CONSV* to be negative and significant (-0.103; p = 0.002) and the coefficient on *RDISCL* to be negative but insignificant (-0.005; p = 0.240).

Last and most importantly, results for the complete model which provides a direct test of H_1 are provided under Model (4). Here, we find that the coefficients on *CONSV* and *RDISCL* continue to be negative as predicted but now both are highly significant. Specifically, within this model, the coefficient on *CONSV* is -0.167 ($p < 0.001$) while the coefficient on *RDISCL* is -0.011 ($p = 0.024$). Further and also as predicted, the coefficient on the interaction term, *CONSV*RDISCL*, is positive and significant. Its estimated value is 0.167 ($p = 0.016$). Thus, in a fashion consistent with H_1, the results indicate that conservatism has a beneficial influence on cost of equity capital but also that this influence is conditional on the information environment. Specifically, they indicate that the marginal benefits of conservatism are diminished in environments of low information asymmetry (high disclosure).[32] For the firms with the lowest disclosure percentile rankings (*RDISCL* = 0.00), based on its estimated coefficient (γ_1), a 0.1 unit increase in the *CONSV* measure is estimated to result in a 0.0167 reduction in cost of equity capital. However, as a firm's disclosure percentile ranking increases, the marginal impact of *CONSV* systematically declines reaching an effective level of zero ($\gamma_1 + \gamma_3$) for the firms with the highest ranking (*RDISCL* = 1.00). From a statistical perspective, the linear restriction capturing the marginal impact of *CONSV* conditioned on the information environment, $\gamma_1 + \gamma_3$, looses statistical significance for values of *RDISCL* above 0.507.[33]

To further illustrate the reliance of the relationship between conservatism and cost of equity capital on the firm's information (disclosure) environment, we run a reduced form of our econometric model on two subgroups of our sample data. Specifically, we separately consider those firms for which their industry-year disclosure percentile rank (*RDISCL*) is either at least 0.65 or at most 0.35. In this fashion, we separately focus on sample firms with a relatively high information environment and those with a relatively low information environment. The reduced form of the model excludes the interaction term given the analysis partitions the sample on the information (disclosure) measure. Results, presented in Panel B of Table 3, effectively confirm the takeaway

[32] Again, results of sensitivity analysis indicate that the results are robust to the inclusion of the equal-weighted as opposed to the value-weighted beta and also to the inclusion of *LEV* as an additional control variable. For example, with *LEV* included in the model, the coefficient estimates become: *CONSV* (-0.167, $p < 0.001$), *RDISC* (-0.011, $p = 0.023$), and *CONSV*RDISC* (0.165, $p = 0.017$), while the coefficient on *LEV* is -0.024 ($p = 0.062$).

[33] The p-value on the F-test that the linear restriction is equal to zero (i.e., $\gamma_1 + \gamma_3 = 0$) is 0.100 when *RDISCL* = 0.507.

from the pooled analysis discussed above. For the low information partition, the coefficient on *CONSV* is negative and significant (-0.213; p = 0.002) whereas for the high information partition, it is insignificant at conventional levels (-0.002; p = 0.975). Thus, based on this illustration, the documented relation between conservatism and cost of equity capital appears restricted to only those firms within a low information (high information asymmetry) environment.[34] As an aside, the coefficient on *RDISCL* is perhaps not unexpectedly insignificant given the reduced cross-sectional variation in the measure under the partitioning process.

5.2 Annual Regression Results

Table 4 presents mean coefficient values, re-estimating our econometric model (equation (1)) using OLS for each of the 10 years in our sample period (1985–1994) and then averaging across the years. Here, we determine annual t-statistics based on White's consistent covariance matrix. As revealed, the mean coefficient values are similar to those for the pooled analysis presented in Table 3 and as such, the annual results also provide strong support for H_1. Also of note, the coefficients on the three control variables are universally of the sign predicted across the 10 years and significant in the vast majority of years. Thus, these results contrast with the pooled analysis where the coefficient on *BTM*, while insignificant, was of the opposite sign to that predicted. Lastly, the adjusted R^2s of the annual regressions range from 0.136 (in 1993) to 0.327 (in 1991).

In detail, the estimated *CONSV* coefficient is negative in all 10 years with |t-statistics| > 1.65 in 7 of the years while the estimated coefficient on the interaction term, *CONSV* * *RDISCL*, is positive in 9 of the 10 years with |t-statistics| > 1.65 in 8 years. The results for the disclosure measure, *RDISCL*, are more muted, with its coefficients being negative in nine of the 10 years but

[34] We acknowledge that the choice of a partitioning point is arbitrary and that the tabulated results based on partitions of 'at least 0.65' and 'at most 0.35' are at best illustrative. The results are somewhat weaker but the message similar when we partition the sample at the median value of *RDISCL*. Here, the p-value on *CONSV* for the high disclosure partition is 0.731 while the p-value on *CONSV* for the low disclosure partition is 0.094.

only significant at the 10% level in 5. To test for the statistical significance of the mean coefficients, we consider several approaches. First, following Abarbanell and Bernard (2000), we calculate the standard error from the distribution of yearly coefficients and then make an adjustment for serial correlation in the coefficients.[35] To supplement this measure, we also calculate the two Z-statistics (Z_1 and Z_2) employed by Aboody and Lev (1998) and Healy, Kang, and Palepu (1987).[36] Reported significance levels are two-tailed for all measures. Appealing to these measures, we find the mean coefficients on both *CONSV* and the interaction term, *CONSV* * *RDISCL*, to be statistically significant at the 1% level. For example, the Abarbanell and Bernard *t*-statistics for these two coefficients are -4.072 and 3.617, respectively. We also find the mean coefficient on *RDISCL* to be weakly significant. Its *t*-statistic is -1.798.

5.3 Rank Regression Results

One potential alternative explanation for the results presented above is the possibility that *CONSV*, measured as the accumulation of non-operating accruals, is related to capital intensity. However, within our data the correlation between capital intensity and *CONSV* is only -0.072 (p = 0.005). Thus, while the correlation is significant, given its magnitude it is unlikely that capital intensity represents an important omitted variable within our model.

More generally, this challenge raises the much broader issue that, with the exception of the disclosure measure, potential industry effects have been overlooked within our primary analysis. As such, in a manner consistent with Botosan and Plumlee (2002), for sensitivity purposes we rank all measures in our model (equation (1)) within year and industry and then re-run the analysis based on

---

35 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. As noted by Abarbanell and Bernard, this adjustment factor assumes that the serial correlation is first-order autoregressive.

36 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.

the pooled sample of 1,782 firm-years using these percentile rank data. Here, higher percentile rank values imply higher values of the relevant measure. By ranking within year and industry, 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.[37] In this sense, the ranking procedure controls for cross-sectional and contemporaneous variations in our sample.

The results for this 'rank analysis' are presented in Panel B of Table 4. As revealed, the results for the primary measures are broadly consistent with the primary results presented in Table 3. Specifically, the coefficients on the percentile rank conservatism measure, denoted R*CONSV* (-0.076, p = 0.095) and *RDISCL* (-0.113, p = 0.015) remain negative and significant, and the coefficient on R*CONSV**RDISCL* (0.104, p = 0.043) remains positive and significant. Thus, our results and conclusions appear robust to the use of industry and year percentile rank data, with these rank regression results again providing strong support for H_1. Perhaps more importantly, they also serve to directly confront concerns regarding differences across industry or time in both our primary and control measures, indicative that our results are unlikely to have been driven by industry effects. Finally, also of note is the fact that, in a manner consistent with the annual results presented in Panel A, the coefficients on the three control measures have the predicted sign and are significant.

[37] Additionally, this rank analysis should also mitigate concerns regarding the potential influence of outliers. Nevertheless, to directly confront the potential influence of outliers, we also determined the Cook's Distance to identify influential data points in the regression analysis (Myers, 1989). Specifically, for each of the models reported in Panel A of Table 3, the Cook's *D*-statistic was estimated to detect influential observations and then analysis was then repeated after the deletion of these observations. Results based on this reduced sample (not tabulated) as qualitatively similar to those reported in Panel A of Table 3. For example, for the complete model (Model 4), based upon the remaining sample of 1,778 firm-year observations, the coefficients (p-values) on *CONSV*, *RDISCL*, and *CONSV*RDISCL* are -0.164 (p = 0.003), -0.106 (p < 0.001), and 0.166 (p = 0.020), respectively. Similar conclusions follow when the data are alternatively Windsorized at the three standard deviation level.

5.4 Additional Considerations

In this section, we present results relating to several additional considerations. First, we provide preliminary direct evidence regarding the relation between conservatism and disclosure, constructs which the preceding analysis implicitly treats as substitute mechanisms. Following, we then, in turn, explore the sensitivity of our results to the choice of proxy for cost of equity capital, conservatism, and disclosure (the firm's information environment). As will be seen, these additional analyses reveal results and conclusions are robust to our choice of proxy for each measure.

5.3.1 The Relation between Conservatism and Disclosure

The arguments underlying H_1 and the results presented above are consistent with management viewing conservatism and disclosure as substitute mechanisms. Artiach (2009) formally explores this potential substitution effect, finding as implied an inverse relationship.[38] For completeness, we also briefly consider the relationship between disclosure and conservatism within our data of 1,782 firm-years using a disclosure model adapted from Artiach (2009). Specifically, in addition to *CONSV*, this disclosure model includes the following factors identified in the literature as determinants of disclosure (e.g., Lang and Lundholm, 1993, 1996): firm size (*lnSIZE*); performance measured as return on assets (*ROA*); analyst following (*AF*); leverage measured as long-term debt divided by total assets (*LEV*); and growth options measured as book-to-market (*BTM*). Consistent with Artiach, the results, presented in Panel A of Table 5, confirm an inverse association within the

---

[38] Theoretical support follows from, among others, Gigler and Hemmer (2001) who model optimal disclosure policy and timeliness of financial reporting in the presence of conservatism. While the main goal of their model is to establish a theoretical link between the properties of mandatory financial reports and the amount of information management provides through voluntary disclosure, they also demonstrate that firms with relatively more conservative accounting are relatively less likely to make timely voluntary disclosure. Additionally, they demonstrate that to accurately measure the degree of conservatism in an accounting regime (by comparing the earnings response coefficient of good versus bad news consistent with Basu (1997) methodology), it is necessary to control for the omitted correlated influence of the amount of voluntary disclosure across firms. Hence, an association between conservatism and disclosure is implicit in their model.

context of this simple test. Specifically, the coefficient on *CONSV* at -0.326 is negative and significant (p = 0.010). Coefficients on the remaining measures in the model are typically of the predicted sign and significant. Thus, within this preliminary analysis, the results appear consistent with the substitutability of conservatism and disclosure.

5.3.2 Sensitivity to the Selected Proxy for Cost of Equity Capital

As discussed by Francis *et al.* (2004), a limitation associated with the use of earnings-based approaches to estimating the *ex ante* cost of equity capital is that they require either or both positive earnings and increasing earnings. For example, for our selected measure, the Easton (2004) PEG measure, Francis *et al.* note that it requires that earnings forecasts are both positive and increasing. In this way, Francis *et al.* argue that the earnings-based approaches restrict the sample. Nevertheless, such approaches have been identified within the literature as the dominant approaches on the basis of their association with known risk factors and/or with realized returns (Guay *et al.*, 2003; Francis *et al.*, 2004; Botosan and Plumlee, 2005; Botosan *et al.*, 2009).

To examine the sensitivity of our results to the use of the ex ante Easton (2004) PEG estimate, based on this use of realized returns to evaluate different ex ante estimation approaches and following Callen, Khan, and Lu (2010), we use one-year-ahead monthly stock returns as an alternative "proxy" for cost of equity capital. Results reported in Panel B of Table 5 reveal our conclusions to be robust to the use of this alternative measure as the dependent variable in equation (1). Specifically, the coefficients on *CONSV* (-0.056, $p < 0.001$) and *RDISCL* (-0.004, $p = 0.024$) remain negative and significant, and the coefficient on the interaction term, *CONSV * RDISCL*, remains positive and significant (0.056, $p = 0.016$). Further, results for the control variables within the model are predominantly as expected and similar to those reported in Panel A of Table 3 based on the Easton (2004) PEG estimate of the cost of equity capital (r).

5.3.3 Sensitivity to the Selected Proxy for Conservatism

As discussed in Section 3.3, one measure advanced within the literature as a plausible alternative conservatism proxy to the Givoly and Hayn (2000) negative accruals measure employed in our primary analysis is the market-to-book ratio (*MB*). As currently developed, our econometric model includes its inverse, the book-to-market ratio (*BTM*) as a risk parameter. In this regard, we run the possibility that in fact the model includes two (potentially overlapping) measures of conservatism. To explore the sensitivity of our results to the joint inclusion of *CONSV* and *BTM*, we adjust our analysis in two ways, with the results for both presented in Panel C of Table 5.

First, we drop *BTM* from our econometric model so that the model only includes one prospective proxy for conservatism. Here, we find the results to qualitatively identical to those reported in Table 3 for the complete model. Specifically, the coefficients on *CONSV* (-0.165, $p <$ 0.001) and *RDISCL* (-0.011, p = 0.025) remain negative and significant, and the coefficient on *CONSV*RDISCL* (0.163, p = 0.018) remains positive and significant. Thus, our results and conclusions appear robust to the inclusion or exclusion of *BTM*.

Second, we replace *CONSV* with the market-to-book (*MB*) ratio as the measure of conservatism within the model. Note, as with *CONSV*, a higher value of *MB* implies a higher degree of accounting conservatism. Here, we also find that the coefficients on the primary variables of interest have the predicted signs and are significant. The coefficients on *MB* and *RDISCL* are -0.002 (p = 0.002) and -0.014 (p = 0.007), respectively, and the coefficient on *MB*RDISCL* is 0.004 (p = 0.002). Thus, results and conclusions also appear robust to the choice of proxy for conservatism.

5.3.4 Sensitivity to the Selected Proxy for the Firm's Information Environment

H_1 implies that the strength of this association between conservatism and cost of equity capital is conditional upon the firm's information environment. In conducting our primary tests reported above, we employ a measure of the firm's disclosure (*RDISCL*) based on AIMR reports as our proxy for the firm's information environment. We argue that this is the preferred measure given that both conservatism and disclosure represent choices made by management wherein they can explicitly determine the relative trade-off. Here, we consider the sensitivity of our results to the use of *RDISCL* by alternatively using the extent of analysts following (*AF*) to proxy for the firm's information environment. Since analyst following may also depend upon the industry and time period, we again use an industry and year percentile rank measure denoted *RAF*. The results, presented in Panel D of Table 5, indicate that while slightly weaker, our conclusions are robust to the choice of proxy for the firm's information environment. Specifically, the coefficients on *CONSV* (-0.149, p = 0.004) and *AF* (-0.009, p = 0.092) are again negative and significant, and the coefficient on the interaction term, *CONSV * AF*, is again positive and significant (0.105, p = 0.038).

5. Summary and Conclusion

In this study, we seek insights into the economic consequences of accounting conservatism by examining the relationship between conservatism and cost of equity capital. Based on both the analytical and empirical literatures, we posit a negative association, with greater accounting conservatism mapping into a reduced cost of equity capital. However, we also posit that the strength of this relation is conditional upon the firm's information environment, with it being the strongest for firms with the greatest information asymmetry and the weakest (and potentially negligible) for firms in the highest information environment. Underlying this prediction is the supposition that conservatism and disclosure represent jointly-determined strategies which form a part of the overall financial reporting strategy of the firm.

We base our analysis on a sample of 1,782 firm-year observations over the ten-year period 1985 – 1994. In our primary analysis, we use the Easton (2004) PEG measure as our cost of equity capital estimate, the Givoly and Hayn (2000) negative accruals measure as our proxy for conservatism, and analyst ratings data published in the AIMR reports as a proxy for firm disclosure (the firm's information environment). We conduct our analysis from several perspectives. First, we base our analysis on the pooled sample of 1,782 firm-year observations drawn over the study period. We then supplement this analysis by conducting a year-by-year analysis and also a rank regression analysis. Our results provide consistent and unequivocal support for our prediction. Specifically, we find a negative association between our conservatism measure and cost of equity capital estimate, and further 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 our fundamental measures, cost of equity capital, conservatism, and the firm's information environment. Finally, we present preliminary evidence in support of the notion that conservatism and disclosure play a joint role in the financial reporting strategy of the firm.

We argue 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 conjunction, we also view them as giving credence to the continued observation of conservative reporting practice. Thus, in this fashion, they appear to draw into question the path enunciated in the 2008 IASB/FASB exposure draft away from conservatism as a desirable characteristic of financial reporting.

In terms of the academic literature, it contributes to the growing body of literature on accounting policy choice which focuses on the market (economic) benefits of accounting policy decisions. By exploring the interaction between conservatism and disclosure, 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 reporting strategy. In this regard, we argue that it contributes to the accounting policy choice literature. We argue that it also contributes to the literature by adopting a theoretical foundation from signalling / information risk literatures rather than the more commonly employed agency arguments, finding empirical support for arguments emanating from the former.

In sum, we view the evidence presented in this study as making a contribution to academic research, regulators and industry by enhancing our understanding of the positive role and benefits of accounting conservative. In future research, we believe that it would be interesting to identify whether there is an equilibrium setting in which the substitutive effects of conservatism and disclosure can be determined. This hopefully would shed greater light on the costs and benefits associated with each of these reporting strategies. We believe that it would also be interesting to identify a causal relationship between conservatism and disclosure by exploring a multi-period changes relationship.

References

Abarbanell, J. and V. Bernard, 2000. Is the U.S. stock market myopic? *Journal of Accounting Research*, Autumn, 221-242.

Aboody, D. and B. Lev, 1998. The value-relevance of intangibles: The case of software capitalization, *Journal of Accounting Research*, Supplement, 161-191.

Allison, P. 'Fixed Effects Regression Methods for Longitudinal Date Using SAS', Viewed 14 December 2008 <http://support.sas.com/publishing/bbu/companion_site/58348.html>.

Ahmed, A., B. Billings, R. Morton, and M. Stanford-Harris, 2002. The role of accounting conservatism in mitigating bondholder-shareholder conflicts over dividend policy and in reducing debt costs, *The Accounting Review*, 77(4), 867-890.

Ahmed, A. and S. Duellman, 2007. Accounting conservatism and board of director characteristics: An empirical analysis, *Journal of Accounting and Economics*, 43(2/3), 411-437.

Artiach, T., 2009. The effect of financial reporting conservatism and disclosure on the cost of equity capital, PhD Thesis, The University of Queensland.

Bagnoli, M. and S. Watts, 2005, Conservative accounting choices, *Management Science*, 51(5), 786-801.

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

Ball, R. and L. Shivakumar, 2005, Earnings quality in U.K. private firms: Comparative loss recognition timeliness', *Journal of Accounting and Economics*, 39, 83-128.

Barry, C. and S. Brown, 1985. Differential information and security market equilibrium, *Journal of Financial and Quantitative Analysis*, December, 407-422.

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

Beaver, W. and S. Ryan, 2000. Biases and lags in book value and their effects on the ability of the book-to-market ratio to predict book return on equity, *Journal of Accounting Research*, 38(1), 127-148.

Beaver, W. and S. Ryan, 2005. Conditional and unconditional conservatism: concepts and modelling, *Review of Accounting Studies*, 10(2/3), 269-309.

Botosan, C., 1997. Disclosure level and the cost of equity capital. *The Accounting Review*, 72, 323-349.

Botosan, C. and M. Plumlee, 2002. A re-examination of disclosure level and the expected cost of equity capital, *Journal of Accounting Research*, 40(1), 21-40.

Botosan, C. and M. Plumlee, 2005. Assessing alternative proxies for expected risk premium, *The Accounting Review*, 80(1), 21-53.

Botosan, C., M. Plumlee, and H. Wen., 2009. The relation between expected returns, realized returns, and firm risk characteristics, working paper, Univ. of Utah.

Callen, J., M. Khan, and H. Lu, 2010. Accounting quality, stock price delay and future stock returns, working paper, University of Toronto.

Clarkson, P., J. Guedes, and R. Thompson, 1996. On the diversification, observability, and measurement of estimation risk', *Journal of Financial and Quantitative Analysis*, 31(1), 69-84.

Dechow, P. and I. Dichev, 2002. The quality of accurals and earnings: The role of accruals estimation errors', *The Accounting Review*, 77(4), 35-59.

Dietrich, J., K. Muller, and E. Riedl, 2007. Asymmetric timeliness tests of accounting conservatism, *Review of Accounting Studies*, 12(1), 95-142.

Drake, M., J. Myers, and L. Myers, 2009. Disclosure quality and the mispricing of accruals and cash flow, *Journal of Accounting , Auditing & Finance*, 357-384.

Easley, D. and M. O'Hara, 2004. Information and the cost of capital', *Journal of Finance*, 59(4), 1553-1583.

Easton, P., 2004. PE ratios, PEG ratios, and estimating the implied expected rate of return on equity capital, *The Accounting Review*, 79(1), 73-95.

Easton, P. and S. Monahan, 2005. An evaluation of accounting-based measures of expected returns, *The Accounting Review*, 80(2), 501-538.

Fama, E. and K. French, 1992. The cross-section of expected stock returns', *Journal of Financial Economics*, 47, 427-465.

Fama, E. and K. French, 1993, Common risk factors in the returns on stocks and bonds', *Journal of Financial Economics*, 33, 3-56.

Francis, J., R. LaFond, P. Olsson, and K. Schipper, 2004. Costs of equity and earnings attributes, *The Accounting Review*, 79(4), 967-1010.

Francis, J, D. Nanda, and P. Olsson, 2008. Voluntary disclosure, earnings quality, and cost of capital', *Journal of Accounting Research*, 46(1), 53-99.

Gebhardt, W., C. Lee, and B. Swaminathan. 2001. Toward an implied cost of capital. *Journal of Accounting Research*, 39, 135-176.

Gietzmann, M. and M. Trombetta, 2003. Disclosure interactions: Accounting policy choice and voluntary disclosure effects on the cost of raising outside capital, *Accounting and Business Research*, 33(3), 187-205.

Gigler, F. and T. Hemmer, 2001. Conservatism, optimal disclosure policy, and the timeliness of financial reports, *The Accounting Review*, 76(4), 471-493.

Givoly, D. and C. Hayn, 2000. The changing time-series properties of earnings, cash flows and accruals: Has financial reporting become more conservative?, *Journal of Accounting and Economics*, 29(3), 287-320.

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

Guay, W., S.P. Kothari, and S. Shu, 2003. Properties of the implied cost of capital using analysts' forecasts. working paper, University of Pennsylvania.

Healy, P., A. Hutton, and K. Palepu, 1999. Stock performance and intermediation changes surrounding sustained increases in disclosures, *Contemporary Accounting Research*, 16(3), 485-520.

Healy, P., S-H Kang, and K. Palepeu, 1987. The effect of accounting procedure changes on CEO's cash salary and bonus compensation, *Journal of Accounting and Economics*, April, 7-35.

Holthausen, R., 1981. Evidence of the effect of bond covenants and management compensation contracts on the choice of accounting techniques: The case of the depreciation switch-back, *Journal of Accounting and Economics*, 3, 73-109.

Holthausen R. and R. Leftwich, 1983. The economic consequences of accounting choice: Implications of costly contracting and monitoring, *Journal of Accounting and Economics*, 5, 77-117.

Holthausen, R. and R. Watts, 2001. The relevance of the value relevance literature for financial accounting standard setting, *Journal of Accounting and Economics*, 31(1/3), 3-75.

Hribar, P. and D. Collins, 2002. Errors in estimating accruals: Implications for empirical research, *Journal of Accounting Research*, 40(1), 105-134.

LaFond, R. and R. Watts, 2008. 'The information role of conservatism, *The Accounting Review*, 83(2), 447-448.

Lambert, R., C. Leuz, and R. Verrecchia, 2007. Accounting information, disclosure and the cost of capital, *Journal of Accounting Research*, 45(2), 385-420.

Lang, M. and R. Lundholm, 1993. Cross-sectional determinants of analyst ratings of corporate disclosure, *Journal of Accounting Research*, 31, 246-271.

Lang, M. and R. Lundholm, 1996. Corporate disclosure policy and analyst behavior, *The Accounting Review*, 71(4), 467-492.

Leftwich, R., 1981. Evidence of the impact of mandatory changes in accounting principles on corporate loan agreements, *Journal of Accounting and Economics*, 3, 3-36.

Myers, R., 1989. *Classical and Modern Regression with Applications*, 2nd Ed., Duxbury Press, California.

Penman, S. and X. Zhang, 2002. Accounting conservatism, the quality of earnings, and stock returns, *The Accounting Review*, 77(2), 237-264.

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

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

Wang, R., C. hOgartaight, and T. van Zijl, 2010. Measures of accounting conservatism: A construct validity perspective, *Journal of Accounting Literature* (forthcoming).

Watts, R., 2003a. Conservatism in accounting Part I: Explanations and implications, *Accounting Horizons*, 17(3), 207-221.

Watts, R.,2003b, 'Conservatism in accounting Part II: Evidence and research opportunities, *Accounting Horizons*, 17(4), 287-301.

TABLE 1 Frequency distribution by year and GICS Industry Sector for a Sample of U.S. firms from the period 1985 – 1994

GICS Industry	**1985**	**1986**	**1987**	**1988**	**1989**	**1990**	**1991**	**1992**	**1993**	**1994**	***Total***	***%***
Aerospace	7	6	6	7	7	7	-	-	-	-	40	2.2
Airlines	3	4	3	3	-	4	4	3	4	5	33	1.9
Apparel	6	6	7	7	7	9	9	8	9	-	68	3.8
Chemical	-	6	5	7	4	5	7	7	-	8	49	2.7
Construction	-	-	-	8	7	11	-	13	-	-	39	2.2
Container/Packaging	-	-	-	5	5	5	5	5	-	-	25	1.4
Diversified	-	-	3	2	3	5	5	5	-	-	23	1.3
Electrical	9	9	9	9	8	9	9	10	-	11	83	4.7
Environmental Control	3	6	9	10	12	12	8	11	8	7	86	4.8
Financial Services	-	6	8	8	8	7	7	6	9	-	59	3.3
Food Beverages	-	-	-	-	14	15	16	17	16	16	94	5.3
Healthcare Services	13	17	16	13	12	12	12	14	14	13	136	7.6
Insurance	11	12	12	10	9	13	14	15	14	15	125	7.0
Int. Pharmaceutical	-	-	-	-	-	-	1	1	1	-	3	0.2
Machinery	8	9	7	7	7	9	8	6	12	9	82	4.6
Motor Carriers	4	4	4	5	5	5	-	-	-	-	27	1.5
Natural Gas Distributors	6	6	6	6	6	6	6	6	5	6	59	3.3
Natural Gas Pipelines	6	4	5	5	5	5	5	5	5	4	49	2.7
NonFerrous Mining/Metal	-	-	-	-	2	4	6	-	-	-	12	0.7
Paper and Forest	-	-	8	9	3	3	9	10	11	11	64	3.6
Petroleum	20	16	22	22	14	13	20	14	14	14	169	9.5
Precious Metals	-	-	-	-	-	4	4	4	4	4	20	1.1
Publishing Broadcasting	8	8	9	9	12	12	12	12	12	13	107	6.0
Railroad	6	5	5	4	6	5	5	5	5	5	51	2.9
Retail Trade	16	16	16	17	19	21	21	17	17	18	178	10.0
Specialty Chemical	11	11	-	12	12	12	10	9	9	-	86	4.8
Textile	1	1	1	1	2	2	2	3	2	-	15	0.8
Total	**138**	**152**	**161**	**186**	**189**	**215**	**205**	**206**	**171**	**159**	**1782**	**100**
%	**7.7**	**8.5**	**9.0**	**10.4**	**10.6**	**12.1**	**11.5**	**11.6**	**9.6**	**8.9**	**100**	

TABLE 2 Descriptive Statistics for a Pooled Sample of 1,782 firm-year Observations over the Period 1985 – 1994

==

Variable	**Mean**	**Median**	**Std Dev**	**Minimum**	**Maximum**
r	0.1214	0.1062	0.0586	0.0138	0.4967
CONSV	0.0316	0.0363	0.0552	-0.4516	0.4117
DISC	0.7345	0.7500	0.1319	0.1286	0.9670
Beta	1.1302	1.1000	0.3431	0.2700	5.0600
BTM	0.5479	0.5070	0.3069	0.0000	2.6485
Size ($'000)	5,375.612	2,220.120	9,689.910	16.1100	87,004.32
AF	19.6218	19.0000	9.4897	1.0000	48.0000
LEV	0.1906	0.1799	0.1421	0.0000	1.4150
ROA	0.0565	0.0497	0.0632	-0.4302	0.5228

Variable definitions: *r* is the cost of equity capital estimates derived from Equation (2); *CONSV* is the accruals-based conservatism proxy measured as the negative of the average over six years of the ratio of non-operating accruals scaled by total assets calculated using the indirect (balance sheet) method; *DISC* is 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; *BTM* is the book value of common equity divided by the market value of common equity; *Size* is the market value of common equity in $millions; *AF* is the number of analysts following the firm; *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.

TABLE 3 Regression Model Results for a Pooled Sample of 1,782 firm-year Observations over the Period 1985 – 1994

Panel A: Pooled

Model	Intercept	*CONSV* (–)	*RDISCL* (–)	*CONSV * RDISCL* (+)	*Beta* (+)	*BTM* (+)	*lnSIZE* (–)
(1)	0.456 (< 0.001)	***-0.102*** ***(0.002)***	---	---	*0.012* *(0.072)*	-0.008 (0.144)	***-0.047*** ***(<0.001)***
(2)	0.447 (< 0.001)	---	-0.005 (0.268)	---	*0.012* *(0.063)*	-0.008 (0.158)	***-0.046*** ***(<0.001)***
(3)	0.456 (< 0.001)	***-0.103*** ***(0.002)***	-0.005 (0.240)	---	*0.012* *(0.069)*	-0.008 (0.146)	***-0.046*** ***(<0.001)***
(4)	0.456 (< 0.001)	***-0.167*** ***(< 0.001)***	***-0.011*** ***(0.024)***	***0.167*** ***(0.016)***	*0.012* *(0.081)*	-0.009 (0.125)	***-0.046*** ***(<0.001)***

Panel B: Partitioned by Disclosure Rank

Partition	Inter	*CONSV* (–)	*RDISCL* (–)	*Beta* (+)	*BTM* (+)	*lnSIZE* (–)
RDISCL $\geq$ 0.65	0.496 (< 0.001)	-0.002 (0.975)	-0.001 (0.945)	-0.005 (0.715)	*-0.019* *(0.076)*	*-0.049* *(< 0.001)*
RDISCL $\leq$ 0.35	0.302 (< 0.001)	***-0.213*** ***(0.002)***	-0.011 (0.496)	-0.013 (0.370)	***0.022*** ***(0.034)***	***-0.041*** ***(< 0.001)***

Variable definitions: *r* is the cost of equity capital estimates derived from Equation (2); *CONSV* is the accruals-based conservatism proxy measured as the negative of the average over six years of the ratio of non-operating accruals scaled by total assets calculated using the indirect (balance sheet) method; *DISC* is 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; *BTM* is the book value of common equity divided by the market value of common equity; and *Size* is the market value of common equity in $millions.

TABLE 4 Annual Regression Model and Rank Regression Results for a Sample of Observations over the Period 1985 – 1994

===

Panel A: Annual Regression Model Results

	CONSV	***RDISCL***	***CONSV * RDISCL***	***Beta***	***BTM***	***lnSIZE***
	(–)	**(–)**	**(+)**	**(+)**	**(+)**	**(–)**
Mean coefficient	-0.254***	-0.011*	0.288***	0.044***	0.064***	-0.043***
No. of coefficients > 0	0	1	9	10	10	0
No. of |*t*-statistics| > 1.645	7	5	8	9	9	8
Z_1 (Aboody and Lev)	-5.990	-3.584	5.665	11.227	13.875	-6.486
Z_2 (Aboody and Lev)	-6.356	-4.185	4.784	8.416	7.194	-14.372
Abarbanell / Bernard *t*-stat	-4.072	-1.798	3.617	7.902	6.444	5.291

Panel B: Rank Regression Results

Intercept	**R*CONSV***	***RDISCL***	**R*CONSV * RDISCL***	**R*Beta***	**R*BTM***	**R*lnSIZE***
	(–)	**(–)**	**(+)**	**(+)**	**(+)**	**(–)**
0.168 (0.477)	*-0.076* *(0.095)*	***-0.113*** ***(0.015)***	***0.104*** ***(0.043)***	***0.165*** ***(< 0.001)***	***0.111*** ***(< 0.001)***	***-0.278*** ***(< 0.001)***

Variable definitions: *r* is the cost of equity capital estimates derived from Equation (2); *CONSV* is the accruals-based conservatism proxy measured as the negative of the average over six years of the ratio of non-operating accruals scaled by total assets calculated using the indirect (balance sheet) method; *DISC* is 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; *BTM* is the book value of common equity divided by the market value of common equity; and *Size* is the market value of common equity in $millions.

In Panel A, 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. 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 Panel B, all measures are ranked within year and industry, denoted by modifying the notation to include an 'R'.

TABLE 5 Additional Considerations

==

Panel A: The Relation between Conservatism and Disclosure

Intercept	*CONSV* (–)	*lnSIZE* (+)	*ROA* (+)	*AF* (+)	*LEV* (+)	*BTM* (+)
0.678	***-0.326***	***0.123***	*-0.111*	*0.002*	***-0.132***	-0.223
(< 0.001)	***(0.010)***	***(<0.001)***	*(0.028)*	*(0.093)*	***(<0.001)***	(0.366)

Panel B:One-Year-Ahead Monthly Returns as the Dependent Variable

Intercept	*CONSV* (–)	*RDISCL* (–)	*CONSV * RDISCL* (+)	*Beta* (+)	*BTM* (+)	*lnSIZE* (–)
0.192	***-0.056***	***-0.004***	***0.056***	*0.004*	-0.003	***-0.015***
(< 0.001)	***(<0.001)***	***(0.024)***	***(0.016)***	*(0.081)*	(0.125)	***(<0.001)***

Panel C: Sensitivity to the Choice of Conservatism Measure

Conservatism proxy	Intercept	*Conservatism* (–)	*RDISCL* (–)	*Conservatism * RDISCL* (+)	*Beta* (+)	*lnSIZE* (–)
CONSV	0.436	***-0.165***	***-0.011***	***0.163***	*0.011*	***-0.044***
	(< 0.001)	***(<0.001)***	***(0.025)***	***(0.018)***	*(0.097)*	***(<0.001)***
MB	0.439	-0.002	*-0.014*	*0.004*	*0.011*	***-0.044***
	(< 0.001)	(0.002)	*(0.007)*	*(0.002)*	*(0.099)*	***(<0.001)***

Panel D: Analyst Following as the Measure of Information

Intercept	*CONSV* (–)	*RAF* (–)	*CONSV * RAF* (+)	*Beta* (+)	*BTM* (+)	*lnSIZE* (–)
0.458	***-0.149***	*-0.009*	***0.105***	***0.018***	-0.008	***-0.048***
(< 0.001)	***(0.004)***	*(0.092)*	***(0.038)***	***(0.021)***	(0.184)	***(<0.001)***

Variable definitions: *r* is the cost of equity capital estimates derived from Equation (2); *CONSV* is the accruals-based conservatism proxy measured as the negative of the average over six years of the ratio of non-operating accruals scaled by total assets calculated using the indirect (balance sheet) method; *DISC* is 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; *BTM* is the book value of common equity divided by the market value of common equity and *MB*, the market-to-book ratio, is its inverse; *Size* is the market value of common equity in $millions; and *RAF* is the industry-year percentile rank analyst following.