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Management earnings forecast disclosure policy and the cost of equity capital

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Abstract We examine the relation between management earnings forecast disclosure policy and the cost of equity capital in a cross-section of 1,355 firms over a 4-year post-Regulation Fair Disclosure period (2001 through 2004). We find evidence of a negative association between the quality of management earnings forecasting policy and cost of equity capital, and we document that the strength of the association is greater for firms with higher disclosure costs and for firms with more relevant quarterly management earnings forecasts. Our results are robust to the use of multiple methods to address both endogeneity and the measurement error in firm-specific estimates of implied cost of equity capital.

Keywords Cost of equity capital · Voluntary disclosure · Management earnings forecasts

JEL Classification M41

1 Introduction

We examine the relation between management earnings forecast disclosure policy and the cost of equity capital. We develop a firm-specific management forecast policy metric that jointly captures whether a firm is a supplier of quarterly management earnings forecasts over a four-year post-Regulation Fair Disclosure (Reg FD) period (2001 through 2004), the frequency of quarterly management

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earnings forecasts over the period, and the average precision of those forecasts. We then test for a cross-sectional correlation between the disclosure policy metric and various proxies for cost of equity capital. We also examine whether the strength of the cross-sectional disclosure policy/cost of capital relation is increasing in disclosure costs and the value relevance of quarterly management earnings forecasts.

Economic theory predicts a negative association between voluntary disclosure and cost of equity capital (Diamond and Verrecchia 1991; Easley and O'Hara 2004).[1] Many practitioners and policymakers hold beliefs that the negative association exists (AICPA 1994) and that the relation is intuitive (Foster 2003). Consistent with these theories and beliefs, several empirical studies find a negative association between disclosure and both cost of equity capital (e.g., Botosan 1997; Botosan et al. 2004; Hail 2002) and other measures of the information environment with links to cost of equity capital, such as bid/ask spreads and volume (e.g., Coller and Yohn 1997; Healy et al. 1999; Leuz and Verrecchia 2000). However, not all empirical work finds that the negative association exists or that it persists after control for earnings quality. For example, Botosan and Plumlee (2002) show that the relation between disclosure and cost of equity capital switches from a negative to a positive relation when switching from the annual report to more timely types of disclosure. Francis et al. (2008) detect the negative association between voluntary disclosure in annual reports and 10-K filings in 2001 and both ex ante measures of cost of capital and ex post realized returns but show that the association is not incremental to a control for earnings quality in their tests on ex ante cost of capital.

Rather than characterize voluntary disclosure as an amalgamation of disclosures of varying types, precisions, and relations to payoffs of interests, we focus on a single voluntary disclosure type, management forecasts of earnings, which are high profile voluntary disclosures linked most closely with the payoff forecasting task faced by investors. Because of their direct relation to payoffs (e.g., future earnings or future dividends/cash flows via the earnings quality link), management forecasts of earnings likely possess the greatest chance to reduce information risk, information asymmetry or both as envisioned by disclosure theorists and, therefore, provide a powerful opportunity to test the relation between disclosure quality and cost of equity capital. However, empirical work on direct links between management forecasting behavior and cost of equity capital is limited to a supplemental test in Francis et al. (2008), which finds a positive association of this high profile and relatively precise voluntary disclosure with cost of equity capital in a single year.[2]

[1] As noted by Botosan (2006), the effects of voluntary disclosure are to reduce information asymmetry, information risk, or both. We review several analytical models in Section 2 that provide varying predictions about the relations among information risk, information asymmetry, and cost of equity capital.

[2] The difference between single period forecasting behavior and forecasting policy is particularly important for interpretation of evidence on the cost of equity capital effects of management forecasting. Francis et al. (2008) limit their study to 2001, the first year following the passage of Reg FD. Reg FD changed both management forecast disclosure policy and the capital market information environment (Wang 2007). It is highly likely that investors would find it difficult to infer longer-run forecasting behavior from a single year's forecasting activity immediately after Reg FD's passage. In fairness to the authors, their study focuses on voluntary disclosures in annual reports and 10-Ks which are far less likely to vary over time, and given the disclosure mechanism, are unaffected by Reg FD. Their supplemental

We emphasize disclosure *policy* in our measurement of management forecasting behavior by using multiple periods to characterize disclosure rather than a short time-frame. We measure management earnings forecasting behavior over a four-year period (2001 through 2004).[3] We begin our tests after the effective date of Reg FD to avoid the potential effects of a change in disclosure regulation on our tests, to enhance the likelihood that publicly released management forecasts are a less noisy proxy for all forecast disclosures, both public and private, and to transact on a period of time after a major regulatory change during which firms must re-establish management forecasting policy.[4] In supplemental tests, we also investigate definitions of management earnings forecasting policy based on single year disclosures and the patterns of disclosures within our sample period.

Prior research generally does not test for cross-sectional differences in the disclosure/cost of capital relation based on the expected costs and benefits of disclosure. Because disclosure is costly, the relation should be stronger (i.e., the disclosure should be associated with a lower cost of capital) for the set of firms with higher disclosure costs. Further, the capital market benefits of a policy to disclose a particular piece of accounting information are increasing in the usefulness of the information in security pricing. We identify high disclosure cost firms based on several proxies used in the literature (industry concentration, capital intensity, high-tech industry membership, and growth opportunities as measured by book-to-market) and examine whether the strength of the disclosure/cost of capital relation is stronger for high disclosure cost firms. We identify the firms with higher information content of management quarterly earnings forecasts and examine whether the strength of the relation is stronger for these high disclosure benefit firms.

Finally, empirical disclosure studies are criticized for failure to control for the endogeneity of disclosure, which may lead to spurious inferences regarding the economic relation between disclosure and cost of equity capital (Healy and Palepu 2001; Core 2001; Larcker and Rusticus 2010; Nikolaev and Van Lent 2005). Also, prior research has raised reliability concerns about firm-specific cost of capital estimates due to low quality analyst forecasts (Easton and Monahan 2005). We address endogeneity concerns and employ several approaches to deal with potentially unreliable cost of capital estimates in our empirical tests.

We find evidence that the quality of management earnings forecasting policy is negatively associated with cost of equity capital both before and after control for earnings quality, CAPM beta, and additional Fama–French determinants of expected return, firm size, and book-to-market (Fama and French 1992). We document a

Footnote 2 continued
test on management forecasts is based on the same year, and they recognize that management forecasting has greater intertemporal variation.

[3] We use quarterly management earnings forecasts because of both their strong link to security prices (Pownall et al. 1993; Baginski et al. 1993) and the greater tension provided by the fact that quarterly forecasts are more timely, and Botosan and Plumlee (2002) detect the unexpected positive relation between more timely disclosure and cost of equity capital.

[4] Reg FD was implemented on October 23, 2000, in an effort to level the playing field for all investors by eliminating selective disclosure. Prior research supports the argument that Reg FD was successful in reducing the amount of selective disclosure (Gintschel and Markov 2004).

stronger negative association for high disclosure cost firms and firms for which quarterly management earnings forecasts generate greater valuation effects. Our results are stronger when we measure forecasting policy over multiple years rather than proxy it with a single year, although our results for single years are significant when the particular single year forecasting quality could be better predicted by the forecasting quality of other years. We are unable to detect a significant effect of patterns of forecasting within our sample period, likely because, after 2 years, firm forecasting behavior appears to be established and somewhat invariant. Our results are robust with respect to estimation method: two-stage instrumental variables approach to address endogeneity, ordinary least squares estimation, and Heckman (1979)-type treatment effects estimation. The negative association between the quality of management earnings forecast policy and cost of equity capital also exists in ex post realized returns-based tests in a substantially larger, more representative sample. Finally, our analysis using the portfolio-level approach suggested by Easton (2009), which avoids firm-specific cost of capital estimation and does not require assumptions about growth, yields the same finding, that is, that the quality of management earnings forecast policy is negatively associated with cost of equity capital.

The paper proceeds as follows. Section 2 reviews prior research and states our hypotheses. Section 3 presents definitions of our empirical variables. Section 4 presents our primary empirical results. Section 5 provides supplemental robustness tests, and Section 6 concludes.

2 Prior research and hypotheses

2.1 Theory and evidence on the relation between disclosure and cost of equity capital

As noted by Botosan (1997, 2006), prior theoretical work has linked disclosure with cost of equity capital in two ways, through the effect of disclosure on transactions costs/*information asymmetry* and through the effect of disclosure on *information risk*. With respect to the *information asymmetry* effect, Amihud and Mendleson (1986) assert that disclosure reduces the adverse selection component of the bid-ask spread and reduces the firm's cost of equity capital. Diamond and Verrecchia (1991) show that disclosure reduces the adverse price impact of a large trade, causing investors to take a larger position in a firm's stock, increasing demand for the firm's stock, and thus reducing the firm's cost of equity capital. With respect to the *information risk* effect, Barry and Brown (1985), Handa and Linn (1993), and Coles et al. (1995) use a Bayesian framework to analyze cost of capital effects. They argue that investors face uncertainty in predicting the true parameters of the return distribution. They conclude that this estimation risk is nondiversifiable and is not reflected in CAPM beta.

More recently, Easley and O'Hara (2004) also analytically link greater public disclosure and lower information risk to lower costs of equity capital. In their model, cost of equity is higher for firms with a larger proportion of private information because uninformed investors require compensation for transacting with informed investors. That is, cost of equity is higher because investors are asymmetrically informed. Public disclosure mitigates information asymmetry by displacing private information, and cost of equity capital is consequently lower.[5] Also, if, in total, information in public disclosure and information in (widely dispersed) private information revealed in prices is more precise, cost of equity is reduced as well.[6] Lambert et al. (2007) argue that, in Easley and O'Hara's pure competition setting, the effects of reducing information asymmetry on cost of equity capital only occur when accompanied by an increase in the average level of information precision. Bhattacharya et al. (2007) interpret Lambert et al. (2007) as implying the possibility of an indirect link between information precision and cost of equity capital in imperfectly competitive environments that is mediated by information asymmetry.

Botosan (1997), Botosan and Plumlee (2002), and Botosan et al. (2004) provide empirical evidence on *aggregate* disclosure's *direct* link to cost of equity capital. Botosan (1997) documents a negative association between an annual report-based disclosure index in a single industry and a cost of equity capital estimate from an accounting-based valuation formula rooted in early work by Preinreich (1938) and Edwards and Bell (1961). Botosan and Plumlee (2002) examine all firms with Association for Investment Management and Research (AIMR) disclosure scores and document a negative association between cost of equity capital and annual report disclosure level. Additionally, they find a *positive* relation between the cost of equity capital and the ratings of more timely disclosures (i.e., quarterly reports). While this finding contradicted their expectations, they note that it is consistent with managers' claims that a higher volume of timely disclosure increases the cost of equity capital through increased stock price volatility. Botosan et al. (2004) examine the association between disclosure quality (both private and public) and cost of equity capital at the aggregate disclosure level. They capture the underlying quality of investors' public and private information from properties of financial analysts forecasts, which represent an ex post reflection of the consequences of all disclosure decisions. They find that an inverse relation exists between the quality of public disclosure and cost of equity capital, as predicted by Easley and O'Hara (2004) but that this relation is more than offset by the positive relation that exists between the cost of equity capital and private disclosure quality.[7]

[5] This assertion itself is a subject of debate because, although intuitive and (generally) supported by empirical evidence, alternative analytic models specify conditions under which the assertion will not hold. These alternative models can be found in Diamond (1985), Lundholm (1988, 1991), Bushman (1991), Alles and Lundholm (1993), Kim and Verrecchia (1991, 1994), and McNichols and Trueman (1994). In summary, the conditions that call the assertion into question are the correlation of private and public signal errors, the ability of informed investors to create more information precision with their private information, and the predictability of the disclosure event.

[6] Leuz and Verrecchia (2006) analytically examine the link between information quality, which they define as higher reporting precision, and a firm's cost of equity capital. They also find that higher quality leads to a lower cost of equity capital, and they also show that this link does not disappear when diverse portfolios are formed. Hughes et al. (2007) show that the Easley and O'Hara (2004) result is driven by underdiversification in a finite economy.

[7] Other papers indirectly link disclosure to cost of equity capital by linking individual disclosure types to various capital market variables (bid/ask spread, volatility, etc.), which proxy for information risk/ information asymmetry conditions that likely lead to higher cost of equity capital. The results are mixed. Coller and Yohn (1997) document decreases in bid/ask spread pursuant to management forecast release. Piotroski (2002) finds support of managers' claims of increased volatility following disclosure.

These empirical studies do not examine the direct link between a single disclosure type and cost of equity capital. "Disclosure" is an aggregate concept that does not differentiate between mandatory and voluntary types. Typically, disclosure studies construct a proprietary index of aggregate disclosure or rely on others' published assessments of disclosure "quality" such as the AIMR scores or disclosure scores reported by the *Report of the Financial Analysts Federation Corporate Information Committee* (Lang and Lundholm 1996) and Standard & Poor's Transparency and Disclosure scores (Patel and Dallas 2002).

Francis et al. (2008) focus solely on the voluntary dimension of aggregate disclosure by examining whether an index of voluntary disclosure derived from annual reports and 10-Ks for a given firm is related to its cost of equity capital. They construct the voluntary disclosure index for a single year and show that the index is negatively related to cost of capital in that year but not incrementally related after control for earnings quality. However, voluntary disclosures included in an index are not homogenous and may be further divided into management earnings forecasts and other voluntary disclosures, which include cash flow projections, forecasts of future dividends, sales projections, plant closings, strategic business changes, and explanations of cost increases. Disclosures of items more precisely related to payoffs, such as management earnings forecasts, are more likely to reduce information risk relative to disclosures of items that assist forecasts of payoffs but that are further removed from payoffs. Requiring investors to engage in more analysis to understand the implications of these latter disclosures results in higher forecast errors and raises the possibility that investors with greater information processing capabilities use indirect disclosures to gain an informational advantage (Kim and Verrecchia 1994). Francis et al. (2008) present supplemental tests on specific types of voluntary disclosure. They document that management earnings forecast disclosure in 2001 is *positively* related to cost of equity capital both before and after control for earnings quality.

Francis et al. (2008) emphasize that:

> [T]he theoretical research used to motivate our hypotheses is predicated on a firm's commitment to a voluntary disclosure policy. We interpret the notions of commitment and policy to mean a stable set of disclosure practices. Our review indicates that disclosures made in annual reports and 10-K filings are relatively stable from one period to the next; as such they are likely subject to less discretion than is, for example, the decision to issue a management forecast. (p. 11)

Like Francis et al. (2008), we interpret the notion of *policy* to mean a stable set of disclosure practices. By measuring forecast policy over a period of time rather than in a single period, single period behaviors receive less weight in describing the policy. As we discuss later in our empirical tests, our sample of firms moved very quickly to relatively stable forecasting behavior by the end of 2004 as evidenced by the ability to predict 2004 forecasting behavior using most recent prior or subsequent year's forecasting behavior. *Commitment* to a policy, however, is difficult to ascertain. *In a given period*, omission of a forecast can mean either a reneging of the policy, if managers possess private information, or simply that managers have no private information to reveal. If the cause of nondisclosure is not observable by market participants, then whether a firm can credibly pre-commit to a given policy is questionable. However, the planning process within firms suggests that managers possess internal forecasts, and *over a number of periods*, observing frequent and precise forecasts suggests that managers possess private information in the form of forecasts and regularly publicly disseminate it. Thus, one can assume that the market forms beliefs about whether managers possess private information, whether the firm has a policy to disclose private information, what that policy is, whether the policy benefits investors, and whether it will continue. Of course, a commitment to disclosure policy can be reneged, and although estimating if and when the reneging occurs is difficult, prior research argues that it can have significant costs (Leuz and Verrecchia 2000; Healy and Palepu 2001; Einhorn and Ziv 2008).

The value of a high quality management forecasting policy is normally considered in terms of a pre-commitment to reduction of information asymmetry (e.g., King et al. 1990), which, for the reasons given above, is likely to be weakened to some extent by the difficulty in credibly signaling pre-commitment. However, a high quality forecasting policy can also affect cost of capital by increasing average information precision (Botosan 2006; Lambert et al. 2007). Management forecasts are explicit disclosures of direct inputs into equity valuation models, as opposed to indirect voluntary disclosures (e.g., capital expansion forecasts, sales forecasts, etc.). Baginski et al. (1993) document that management forecasts are, on average, more precise (i.e., have a narrower range) than the range of contemporaneous financial analyst forecasts and are associated with a reduction in the range of analysts' forecasts when publicly released. It is important to the interpretation of our study to note that these financial analysts have observed the remaining set of publicly released voluntary disclosures and possess information on current earnings and its components. Yet, management forecasts are more precise. Further, in laboratory settings, Hirst et al. (1999) and Libby et al. (2006) find that analysts are more confident in their forecasts when management forecasts are more precise.

We combine three dimensions of forecasting behavior into a single disclosure quality measure to capture potential beneficial effects of removing information asymmetry, increasing information precision, or both. First, our measure captures whether a firm is a *supplier* of at least one quarterly management earnings forecast over 16 quarters (a 4-year period). The decision not to forecast under any market condition is a powerful dimension of disclosure choice, and not issuing a single forecast over a 4-year period would be unusual given a high probability of at least occasional differences between private and public information. Second, regular revelation of management's private information (forecast *frequency*) is necessary to remove regularly occurring opportunities for private information acquisition. Also, investor forecasting tasks subject to information estimation risk exist over multiple periods, and therefore more frequent management earnings forecasting is necessary to reduce information estimation risk (assuming that management has more precise information). Third, management forecasts are released in alternative forms, including less precise range, open interval, and general impression forecasts. To the extent managers provide precise revelations of their private information (forecast *precision*), information asymmetry, information estimation risk, or both should be reduced.

Our goal is not to separate the effects of a high quality management earning forecast policy into effects on information asymmetry and effects on information precision. Instead, our focus is on discovering whether the quality of management earnings forecasting policy is associated with cost of equity capital, using theory as a basis for the expectation that it is and that the effect derives from the ability of management earnings forecasting policy to reduce information asymmetry, increase information precision, or both.[8] Our hypothesis (stated in the alternative) is:

$H1_A$ The quality of a firm's management earnings forecasting policy is negatively associated with its cost of equity capital.

2.2 Disclosure costs and the strength of the relation

Disclosure costs play an important role in determining which firms will have the highest quality disclosure policies. Disclosure costs are lower when barriers-to-entry are high in product markets. Proprietary information and legal exposure also increase disclosure costs. High quality disclosers believe that disclosure benefits exceed a disclosure cost threshold.

Low disclosure cost firms will have a higher quality disclosure policy if they receive relatively low disclosure benefits or relatively high disclosure benefits because both benefit levels exceed the disclosure cost threshold. On the other hand, high disclosure cost firms that have high disclosure quality must have received relatively high disclosure benefits. Given that high disclosure cost firms that do not disclose receive no disclosure benefits, the association between disclosure and cost of capital disclosure benefits should be larger for high disclosure cost firms relative to low disclosure cost firms.

$H2_A$ The negative association between the quality of a firm's management earnings forecasting policy and its cost of equity capital is increasing in its disclosure costs.

Following Cohen (2006), we use four proxies for disclosure costs—current product market competition, capital intensity, expected litigation costs, and growth opportunities—to examine whether disclosure costs effect the association between forecasting policy and cost of equity capital.

2.3 Quarterly management earnings forecasts and firm valuation

As noted previously, we chose quarterly management earnings forecasts as the voluntary disclosure variable of interest because of its direct conveyance of a payoff used in valuation—earnings. We believe that, in contrast to earnings quality research, most voluntary disclosure research has not necessarily focused on the most

[8] Mapping the precise path from disclosure to cost of equity capital is a complex task and beyond the scope of this paper. See Bhattacharya et al. (2007) for recent work in this area.

value relevant voluntary disclosures. As a result, a bias is introduced in favor of the null hypothesis of no association between voluntary disclosure and cost of capital.

At issue then is whether quarterly management earnings forecasts are sufficiently value relevant so as to have detectable effects of their disclosure policy on cost of equity capital. If not, then our tests are also biased toward the null of no association, and a non-result is interpretable only in the context of a joint hypothesis of sufficient value relevance and the existence of a cost of capital effect.

Reasons exist to expect that quarterly management earnings forecasts are sufficiently value relevant. Security price reaction measures the present value of changes in expected payoffs over the infinite horizon. If a quarterly management forecast conveys information that only changes expectations about the current quarter, price reaction will be relatively small. However, price reactions to quarterly management earnings forecasts are relatively large, implying that the revision in expectations goes beyond the current quarter. For example, for firms with no expected growth in nonzero net present value projects, expected abnormal earnings growth is zero, and price equals next period's capitalized earnings (Ohlson and Juettner-Nauroth 2005). Further, if current earnings captures changes in the firm's investment opportunities with little or no implications for future abnormal earnings growth, uncertainty about current earnings captures total uncertainty, and quarterly earnings disclosures have the potential to reduce estimation risk significantly.[9]

While the sufficiency of value relevance is ultimately an empirical issue, we take advantage of the opportunity to test for cross sectional differences in the disclosure policy/cost of capital relation based on firm-specific management quarterly earnings forecast relevance.

$H3_A$ The negative association between the quality of a firm's management earnings forecasting policy and its cost of equity capital is increasing in the relevance of its quarterly management earnings forecasts.

3 Empirical proxies

3.1 Cost of equity capital

In our primary tests, we measure a firm's cost of equity capital at a given point in time using the PEG method, which is derived from the dividend discount model. The formula below is taken from Easton (2004):

$$r_{PEG} = \sqrt{\frac{E_0(eps_2) - E_0(eps_1)}{P_0}} \tag{1}$$

[9] We thank a referee for pointing out these two examples.

where r_{PEG} is estimated cost of equity capital; E_0 is the expectations operator; P_0 is stock price at end of sample period; and eps_t is earnings per share at time t, where eps_1 (eps_2) represents expected EPS one (two) period(s) ahead.[10]

For our primary tests performed in the cross-section at the end of 2004, the EPS one period ahead reflects fiscal year end 2005 and two periods ahead reflects fiscal year end 2006. The I/B/E/S summary statistics file releases the consensus analyst forecasts on the third Thursday of every month. We use the consensus from the January 2005 release in order to ensure capturing the effect of all management earnings forecasts issued in 2004. We did not calculate cost of equity capital for firms where EPS two periods ahead is less than EPS one period ahead. Approximately 4.5% of the firms had decreasing EPS forecasts.

Prior studies have used a variety of methods to derive an ex ante cost of equity capital (e.g., Claus and Thomas 2001; Easton 2004; Gebhardt et al. 2001; Gode and Mohanram 2003; Ohlson and Juettner-Nauroth 2005). Botosan and Plumlee (2005) evaluate alternative cost of equity capital proxies by assessing their association with known risk proxies, beta, size, book-to-market, as well as leverage and growth. They find that the cost of capital measure based on the PEG method is most highly associated with these known risk factors.[11] However, using the return decomposition approach provided by Vuolteenaho (2002), Easton and Monahan (2005) provide evidence that calls into question the reliability of these ex ante cost of equity capital measures by showing that the measures are not associated with ex post realized returns unless financial analyst forecasts are of high quality. Some past research has averaged several ex ante cost of equity capital measures as a means of offsetting measurement error in individual measures (e.g., Dhaliwal et al. 2005). However, it is unlikely that averaging several highly correlated measures (see Easton and Monahan 2005) all based on potentially low quality financial analyst forecasts addresses the problem, especially given Easton and Monahan's finding

[10] Easton (2004, p. 77) notes that the PEG measure is equal to the price-earnings ratio divided by an earnings growth rate. Using a price-to-forward earnings ratio (PE) for stock recommendations requires that a high (low) PE implies a low (high) expected rate of return, but the earnings of next period may not be indicative of the future stream of earnings. Thus, the PEG captures the comparison of the PE ratio and earnings growth rate as a basis for stock recommendations. Easton (2004) refers readers to http://www.fool.com/School/TheFoolRatio.htm for a description of the PEG ratio. This site notes that the PEG ratio may not work for firms in the financial industry because firms in these industries "have low P/E's that virtually never reach their growth rates, mainly because their companies are valued off assets they hold (like oil deposits and real estate) rather than operating earnings." Accordingly, we exclude financial firms from our primary analyses.

[11] Botosan and Plumlee (2005) also find that another ex ante cost of equity capital estimation approach, the target price method (see Brav et al. 2005), yields estimates of cost of capital that exhibit relatively high correlations with well-known risk factors. We do not employ the target price method in our tests for three reasons. First, the method requires Value Line data, which dramatically reduces sample size. Second, the Value Line sample is dominated by the largest, most followed, less risky, and oldest firms. As argued by Leuz and Verrecchia (2000) it is unlikely that theoretically suggested disclosure effects can be documented in highly developed information environments. Third, if analysts are basing their target prices on the assumption of market inefficiency, then it is not clear that implied rates of return derived from target prices reflect capital market beliefs about cost of equity capital and future cash flows.

that all of the cost of equity capital measures suffer from the low analyst forecast quality problem. Accordingly, in tests described later, we examine the robustness of our results using ex post realized returns tests that do not rely on financial analyst forecast quality and, due to the relaxation of data requirements, permit the use of a much larger sample. Also, we re-examine our main hypothesis using approaches suggested by Easton (2009) that do not require explicit assumptions about growth and that mitigate measurement error associated with firm-specific estimation via large portfolio groupings and reverse regression.

3.2 The quality of management earnings forecast policy

We use the FirstCall database to obtain management earnings forecast data. We define three dimensions, which we combine in multiplicative fashion, to measure management earnings forecast policy. *Supplier* equals one if a firm issued at least one quarterly management earnings forecast over the 2001–2004 period and zero otherwise. *Frequency* equals the number of quarterly management earnings forecasts issued by a firm over the same period. *Precision* equals the average of a given firm's forecast precisions over the same period. Tests of the causes and consequences of management earnings forecast disclosures define forecast precision in terms of alternative forecast forms (Baginski et al. 1993; Baginski and Hassell 1997; Bamber and Cheon 1998). Consistent with this approach, forecast precision equals 0 if no forecast exists, 1 for general impression forecasts, 2 for minimum and maximum forecasts, 3 for range forecasts, and 4 for point forecasts. The quality of management earnings forecast policy for firm *i* (*MFDiscPol*) is then computed as follows:

$$MFDiscPol_i = Supplier_i \times Frequency_i \times Precision_i \tag{2}$$

The natural log of $MFDiscPol_i + 1$ ($lnMFDiscPol_i$) is used in the main empirical analysis.

We use a multiple year determination of management forecasting policy to enhance construct validity. To draw an analogy of the reasoning behind our approach to the reasoning behind a similar approach used in the earnings quality literature, consider the following statement in Francis et al. (2008):

> Note that all of the earnings quality metrics are estimated over a multi-year period…. We believe the use of a multiyear period enhances the construct validity of these earnings quality proxies because the disclosure theories described in Section 2 speak to the underlying earnings quality of the firm, not to earnings quality measured in any individual year (which might be influenced by transitory managerial incentives). (p. 68)

We begin the forecast policy measurement after the passage of Reg FD. We assume that firms choose disclosure policy to maximize firm value. Prior to passage of Reg FD, disclosure policy would include both public disclosures and private disclosures to analysts when doing so maximized firm value (e.g., avoiding public disclosure of proprietary information; King et al. 1990; Wang 2007). Because of the private disclosure avenue, disclosure metrics derived from public disclosure counts are not valid indicators of voluntary disclosure before Reg FD.[12]

4 Primary empirical model

Formulation of management earnings forecast policy is likely endogenous. Several studies identify the source of this endogeneity. For example, Brown and Hilligeist (2007) argue that, if better voluntary disclosure quality leads to less information asymmetry (and by extension, a lower cost of equity capital), then high information asymmetry firms will have greater incentives to choose high quality voluntary disclosure to reduce information asymmetry. Larcker and Rusticus (2010) also note an endogeneity problem caused by the omitted variable bias in studies of disclosure's effect on cost of capital. They cite disclosure costs as an example of an omitted variable that explains lower disclosure and, if present in higher risk firms, higher cost of capital. Nikolaev and Van Lent (2005) couch the problem in general terms as unobserved omitted variables that capture firm-specific heterogeneity, citing management reputation and costs of disclosure as examples.[13] Endogeneity causes OLS parameter estimates to be inconsistent. A common approach to mitigate the problem is to use instrumental variables to replace the endogenous variable with a predicted variable and use a two-stage least squares (2SLS) estimation. Accordingly, we use instrumental variables and two-stage least squares to estimate the effect of disclosure dimensions on cost of equity capital. In the first stage, we obtain fitted values from regressions of the quality of management earnings forecast policy on a set of instruments generally used to describe voluntary disclosure choice:

$$lnMFDiscPol_i = \lambda_0 + \lambda INSTRUMENTS_i + \varepsilon_i \tag{3}$$

where $INSTRUMENTS_i$, a vector of instrumental variables.[14]

To select instruments for Eq. 3, we surveyed a number of studies that address both disclosure choice and the quality of the resulting disclosures (e.g., Baginski and Hassell 1997; Bamber and Cheon 1998; Leuz and Verrecchia 2000; Barton and

[12] One could argue that pre-Reg FD counts are valid measures of *public* disclosure and therefore are relevant if the disclosure/cost of capital relation is driven by information asymmetry reduction. However, one cannot infer whether a given level of public disclosure is accompanied by private disclosure as well, and the disclosure/cost of capital relation may be driven solely through the reduction in information risk.

[13] Their specific example of a disclosure cost heterogeneity is cross-sectional differences in the sophistication of firms' investors, which leads to cross-sectional differences in disclosure practices (Dye 1985, 1998), and cross-sectional differences in required return to compensate for information risk.

[14] When we estimate Eq. 3, we also include the exogenous variables from the second stage (Wooldridge 2002). Brown and Hilligeist (2007) and Larcker and Rusticus (2010) employ 2SLS to address these particular sources of endogeneity, and Nikolaev and Van Lent (2005) cite the use of 2SLS as a potential solution. See Barton and Waymire (2004), Cohen (2006), and Heflin et al. (2005) for additional recent examples of the use of 2SLS. Leuz and Verrecchia (2000) model a similar disclosure choice/disclosure outcome as a self-selection problem and use a treatment effects model (see Heckman 1979) to mitigate the effects of disclosure choice. We replicate our analysis using the Heckman approach in supplemental tests.

Waymire 2004; Cohen 2006). We proxy for the demand for high quality disclosure by using the age of the firm, the number of shareholders, and the amount of financial analyst following. *AGE* equals the number of days a firm has been listed in the CRSP database measured as the difference between the firm's first day on CRSP and December 31, 2000. Barton and Waymire (2004) argue that age partially proxies for the demand for information about newer firms. *Shareholders* equals the log of the number of common shareholders at the end of 2000. *Shareholders* proxies for the information demand created by dispersed ownership (Bushee et al. 2003; Cohen 2006). *Analyst* equals the number of financial analysts following each firm at the end of 2000. Analyst following captures the demand for information by analysts, from either public disclosure or private information search (Lang and Lundholm 1996). We capture incentives to supply forecasts in several ways. *BegROA* equals average return on assets for 2000, where return on assets is measured by dividing income before extraordinary items by total assets for each quarter in 2000 and using the average of these quarterly observations. Return on assets measures performance, which has been found to be associated with disclosure (Miller 2002). We use the sign of *BegROA* in the model ($SIGNROA = 1$ for $BegROA > 0$ and 0 otherwise) to indicate positive and negative performance. Because managers are likely to increase disclosure quantity when accessing the capital markets to reduce information risk or "hype" the stock (Lang and Lundholm 2000), we use two proxies for financing activities. *CapIntensity* equals total assets minus current assets all divided by total assets at the end of 2000. Capital intensity is a proxy for the level of financing needs, and it also proxies for barriers-to-entry, which is associated with lower disclosure costs (Cohen 2006). *Offer* is an actual measure of financing activity. It equals the percent change in common shares outstanding adjusted for stock splits from the end of 2000 to the end of 2004. Finally, to proxy for longer-run disclosure costs and benefits (to coincide with our focus on forecasting policy rather than period-specific forecasts), we include three proxies for industry effects. *HHI* equals the Herfindahl-Hirschman Index at the end of 2000, an industry concentration measure that proxies for barriers-to-entry (Cohen 2006). *REGULATE* equals 1 for firms in regulated industries and 0 otherwise, where regulated firms are in the following SICs: 4811 through 4899, 4922 through 4924, 4931, 4941, 4833, 6021 through 6023, 6035, 6036, 6141, 6311, 6321, and 6331. Regulated industries provide a substantial amount of mandated and other industry-related disclosures, which are substitutes for management forecasts. *HIGHTECH* equals 1 for high-tech firms and 0 otherwise, where high-tech firms are in the following SICs: 2833 through 2836, 3570 through 3577, 3600 through 3674, 7371 through 7379, and 8731 through 8734. Membership in a high-tech industry proxies for litigation risk (Kasznik and Lev 1995).[15]

[15] Age and performance are also related to litigation risk (Kasznik and Lev 1995; Rogers and Stocken 2005). Regulated industries include all communication sectors; electric and other services; water supply; natural gas transmission and distribution sectors; national and state commercial banks; chartered savings institutions; personal credit institutions; life insurance; accident and health insurance; and fire, marine, and casualty insurance. High-tech industries include drug and pharmaceuticals; computers, electronics, computer programming, data processing and other computer related; and research, development, and testing services.

Ideally, the exogenous variables that we identify as instruments for the first-stage regression are uncorrelated with the unobservable error in the second-stage structural equation. As noted by van den Berg (2006):

> Basically, if one is interested in the effect of a "treatment variable" on an outcome variable, and the treatment is not exogenously assigned, then one may perform causal inference by exploiting the presence of variables that causally affect the treatment status but do not have a direct causal effect on the outcome. The latter restriction is called an exclusion restriction. Exclusion restrictions are identifying restrictions, so they cannot be tested. This means that empirical results critically depend on the validity of the exclusion restriction, and that this restriction needs to be justified on a priori grounds. (p. 2)

Our a priori reason for the exclusion restriction on the instruments we identify rests with their lack of inclusion in theoretical models of cost of equity capital. As noted below and in Section 2, these models link cost of equity capital with beta and the effects of disclosure and earnings quality on information asymmetry and information risk.

The second stage structural cross-sectional OLS model is:

$$COC_i = \phi_0 + \phi fitMFDiscPol_i + \phi' X_i + \pi_i \tag{4}$$

where COC_i is cost of equity capital level at the end of 2004 calculated using the PEG method (i.e., r_{PEG} from Eq. 1); $fitMFDiscPol_i$ is a vector of fitted values from the first stage regression (Eq. 3); and X_i is a vector of control variables discussed below.

We estimate several versions of Eq. 4 using statistical approaches that correctly compute second stage standard errors, adding different sets of control variables. The capital asset pricing model (Lintner 1965; Mossin 1966; Sharpe 1964) indicates that market beta (*BETA*) is the theoretical determinant of cost of equity capital. Accordingly, we estimate a version of Eq. 4 that controls for market beta, which we expect to be positively associated with cost of capital. We estimate *BETA* using the market model with a minimum of 30 out of 60 monthly returns and a market index return equal to the value-weighted NYSE/AMEX return.

Fama and French (1992) identify two other determinants of expected returns, firm size and the book-to-market ratio. Accordingly, we estimate an additional version of Eq. 4 that also includes log of market value of equity (*lnEndSize*) and the log of the book-to-market ratio (*lnBM*), measured as the log of the common equity of the firm scaled by the market value of equity, as controls in addition to *BETA*. Fama and French (1992), Gebhardt et al. (2001), and Baginski and Wahlen (2003) find a positive relation between the book-to-market ratio and the cost of equity capital.[16]

[16] In untabulated tests, we also include leverage (*LEV*) as a control variable, measured as long-term debt plus any debt in current liabilities divided by total assets, to proxy for the amount of debt in the firm's capital structure. Botosan and Plumlee (2005) find *LEV* to be positively associated with cost of equity capital. Likewise, although not identified as an additional risk factor/anomaly by Fama and French (1992), we also include long-term growth rates (*LTG*) to control for risk associated with growth

In addition to the aforementioned controls for risk, we control for earnings quality in some of our second-stage regressions. *EarnVar* is the standard deviation of earnings before extraordinary items divided by total assets estimated over the 1992–2004 period (*lnEarnVar* used in empirical analysis). Francis et al. (2004) advance earnings quality as a potential determinant of cost of equity capital. Francis et al. (2008) show that the negative relation between voluntary disclosure and cost of equity capital is not as strong in the presence of a control for earnings quality.

4.1 Sample and descriptive statistics

Our sample consists of 1,355 nonfinancial firms with sufficient data from CRSP, I/B/E/S, and Compustat to estimate our research design variables. Table 1, Panel A presents descriptive statistics for continuous variables. Mean (median) *COC* is just under 11% (just under 10%). *MFDiscPol* is highly skewed (mean over 19, median equal 8), justifying using its log in empirical tests. Approximately 80% of our sample firms issued at least one quarterly management earnings forecast during the four-year period. Among these forecasting firms, the median number of forecasts issued is 10 (over 16 quarters). The median forecast form is a range forecast (not tabulated). Median annual earnings variance (*EarnVar*), calculated as the standard deviation of ROA using up to 13 annual earnings observations, is 0.0459. Mean *EarnVar* is substantially higher due to several larger values. Accordingly, we use the log in empirical analyses. We also log transform firm size and the book-to-market variables, consistent with prior research.

4.2 Results: Association of management earnings forecast policy and (ex ante) cost of equity capital (H1)

Tables 2 through 4 present our primary tests of whether the quality of management earnings forecast policy and cost of equity capital are negatively associated (H1) using two-stage least squares and a structural regression model with control variables. Table 2 presents a correlation matrix of the variables used in the first stage regression. Of particular note are the number of significant intercorrelations among variables and the magnitude of some of the correlation coefficients. Our primary interest is not in interpreting individual regression coefficients but making sure that theoretically relevant instruments are included in the first stage regression.

Table 3 presents an OLS first stage regression for each alternative second stage specification. Because some sample firms do not forecast during the time period, we have a censored sample. Therefore, we present t-statistics from a TOBIT regression for comparison purposes. We continue this sensitivity test in Table 4 by using both first-stage fitted values from OLS and TOBIT in the second stage regression (with appropriate standard error correction).

Footnote 16 continued
opportunities for each firm, estimated using the long-term growth rates from I/B/E/S. Gebhardt et al. (2001) and Botosan and Plumlee (2005) find *LTG* to be positively associated with cost of equity capital. Inclusion of *LEV* and *LTG* does not affect our inferences, and we limit our main controls to well-known risk factors identified in the CAPM and by Fama and French.

Table 1 Variable distributions for 1,355 firms used in cross-sectional cost of equity capital regressions

Variable	Mean	Standard deviation	25th percentile	Median	75th percentile
Panel A: Variable distributions for continuous variables					
Dependent variable					
COC	0.1079	0.0542	0.0788	0.0953	0.1211
Disclosure score					
MFDiscPol	19.3682	24.698	0	8	33
Control variables					
EarnVar	0.1067	0.2456	0.0236	0.0459	0.1063
Beta	0.9298	0.5597	0.5622	0.7915	1.1673
Size ($M)	6,412.21	21,227.32	401.10	1,122.49	3,925.54
BM	0.4247	0.2360	0.2579	0.3865	0.5587
Unique first stage instruments					
AGE	5,941.54	6,204.92	1,655.00	3,465.00	10,106.00
CapIntensity	0.2546	0.2173	0.0849	0.1936	0.3640
Shareholders (millions)	23.4405	101.8190	0.6380	2.6900	9.7930
Analyst	7.6487	6.6093	3.0000	5.0000	11.0000
HHI	0.0640	0.0644	0.0255	0.0397	0.0819
Offer	0.1868	0.6130	0.0067	0.0700	0.2308

Panel B: Variable distributions for categorical variables first stage instruments

	Positive = 1 (%)	Negative = 0 (%)
SIGNROA	1,086 (80.15%)	269 (19.85%)
	Yes = 1 (%)	No = 0 (%)
REGULATE	42 (3.10%)	1,313 (96.90%)
HIGHTECH	328 (24.21%)	1,027 (75.79%)

COC is cost of equity capital at the end of 2004 for each firm calculated using Eq. 1. *MFDiscPol* is a measure of the quality of a firm's management forecast disclosure policy, measured by summing the number of quarterly management earnings forecasts issued by each firm from 2001 to 2004, weighted by the precision of each forecast. The precision of a forecast equals 4, 3, 2, and 1 for point, range, open-ended, and qualitative forecasts, respectively. The natural log of *MFDiscPol* + 1 (*lnMFDiscPol*) is used in the empirical analysis. *EarnVar* is the standard deviation of earnings before extraordinary items divided by total assets estimated over the 1992–2004 period (*lnEarnVar* used in empirical analysis). *BETA* is estimated over the period just prior to the end of 2004 using the market model with a minimum of 30 out of 60 monthly returns and a market index equal to the value weighted NYSE/AMEX return. *Size* is the market value of equity at December 31, 2004 (*lnSize* used in the empirical analysis). *BM* is the book-to-market ratio measured as book value of equity divided by common shares outstanding multiplied by end of 2004 stock price (*lnBM* used in the empirical analysis). *AGE* equals the number of days a firm has been listed in the CRSP database +1, measured as the difference between the firm's first day on CRSP and December 31, 2000 (*lnAGE* used in the empirical analysis). *SIGNROA* equals 1 if the average return on assets for 2000, measured by dividing income before extraordinary items by total assets for each quarter in 2000 and using the average of these quarterly observations, is positive and 0 otherwise. *CapIntensity* is total assets less current assets divided by total assets at the end of 2000 (*lnCapIntensity* used in the empirical analysis). *Shareholders* is the number of common shareholders at the end of 2000 (*lnShareholders* used in the empirical analysis). *Analyst* is the number of analysts following each firm at the end of 2000 (the natural log of *Analyst* + 1, *lnAnalyst*, used in the empirical analysis). *HHI* is the Herfindahl-Hirschman Index at the end of 2000 (*lnHHI* used in the empirical analysis). *Offer* is the percent change in common shares outstanding over the sample period adjusted for stock splits (the natural log of *Offer* + 1, *lnOffer*, used in the empirical analysis). *REGULATE* is 1 if the firm is in a regulated industry and 0 otherwise. *HIGHTECH* is 1 if the firm is in a high tech industry and 0 otherwise

Table 2 Pearson product-moment correlation matrix for variables used in first stage of two-stage least squares estimation ($N = 1,355$ firms)

Variable	*lnMFDiscPol*	*lnAGE*	*SIGNROA*	*lnCapIntensity*	*lnShare Holders*	*lnHHI*	*lnOffer*	*REGULATE*	*HIGHTECH*	*lnAnalyst*	*lnEarn Var*	*Beta*	*lnSize*
lnAGE	−0.0199												
	0.4627												
SIGNROA	0.1467	0.3050											
	<0.0001	<0.0001											
lnCapIntensity	−0.0016	0.1867	0.1353										
	0.9530	<0.0001	<0.0001										
lnShareHolders	−0.0067	0.4391	0.1883	0.1535									
	0.8044	<0.0001	<0.0001	<0.0001									
lnHHI	0.1155	0.0655	0.1495	0.1900	−0.0286								
	<0.0001	0.0159	<0.0001	<0.0001	0.2924								
lnOffer	−0.1365	−0.1805	−0.2607	−0.0727	−0.1799	−0.1047							
	<0.0001	<0.0001	<0.0001	0.0074	<0.0001	<0.0001							
REGULATE	−0.0978	0.0266	−0.0497	0.1466	0.1092	−0.1935	−0.0006						
	0.0003	0.3273	0.0607	<0.0001	<0.0001	<0.0001	0.9813						
HIGHTECH	0.0097	−0.2135	−0.2197	−0.2854	−0.1153	−0.4414	0.1235	−0.1010					
	0.7201	<0.0001	<0.0001	<0.0001	<0.0001	<0.0001	<0.0001	0.0002					
lnAnalyst	0.2848	0.2264	0.1709	0.0235	0.3711	−0.0615	−0.1870	0.0030	0.0194				
	<0.0001	<0.0001	<0.0001	0.3874	<0.0001	0.0234	<0.0001	0.9112	0.4753				
lnEarnVar	0.0651	−0.4058	−0.4365	−0.1678	−0.3600	−0.1522	0.2046	−0.0880	0.3939	−0.1755			
	0.0165	<0.0001	<0.0001	<0.0001	<0.0001	<0.0001	<0.0001	0.0012	<0.0001	<0.0001			
Beta	0.1698	−0.2441	−0.1449	−0.3225	−0.1193	−0.2475	0.0733	−0.0294	0.4865	0.2577	0.4037		
	<0.0001	<0.0001	<0.0001	<0.0001	<0.0001	<0.0001	0.0069	0.2793	<0.0001	<0.0001	<0.0001		
lnSize	0.1001	0.3596	0.2548	0.0725	0.5619	−0.0064	−0.1509	0.0575	−0.0889	0.6703	−0.4221	0.0363	
	0.0002	<0.0001	<0.0001	0.0076	<0.0001	0.8127	<0.0001	0.0343	0.0010	<0.0001	<0.0001	0.1810	
lnBM	0.0017	0.0695	0.1574	0.0567	−0.0135	0.0835	0.0671	0.0890	−0.1452	−0.1460	−0.2446	−0.0782	−0.2550

Table 2 continued

Variable	*lnMFDiscPol*	*lnAGE*	*SIGNROA*	*lnCapIntensity*	*lnShare Holders*	*lnHHI*	*lnOffer*	*REGULATE*	*HIGHTECH*	*lnAnalyst*	*lnEarn Var*	*Beta*	*lnSize*
	0.9475	0.0104	<0.0001	0.0367	0.6195	0.0021	0.0135	0.0010	<0.0001	<0.0001	<0.0001	0.0040	<0.0001

lnMFDiscPol is the natural log of *MFDiscPol* + 1, measured by summing the number of quarterly management earnings forecasts issued by each firm from 2001 to 2004, weighted by the precision of each forecast. The precision of a forecast equals 4, 3, 2, and 1 for point, range, open-ended, and qualitative forecasts, respectively. *lnAGE* is the natural log of the number of days a firm has been listed in the CRSP database +1, measured as the difference between the firm's first day on CRSP and December 31, 2000. *SIGNROA* equals 1 if the average return on assets for 2000, measured by dividing income before extraordinary items by total assets for each quarter in 2000 and using the average of these quarterly observations, is positive and 0 otherwise. *lnCapIntensity* is the natural log of total assets less current assets divided by total assets at the end of 2000. *lnShareholders* is the natural log of the number of common shareholders at the end of 2000. *lnAnalyst* is the natural log of the number of analysts + 1 following each firm at the end of 2000. *lnHHI* is the natural log of the Herfindahl-Hirschman Index at the end of 2000. *lnOffer* is the natural log of the percent change in common shares outstanding over the sample period adjusted for stock splits + 1. *REGULATE* is 1 if the firm is in a regulated industry and 0 otherwise. *HIGHTECH* is 1 if the firm is in a high-tech industry and 0 otherwise. *lnEarnVar* is the log of the standard deviation of earnings before extraordinary items divided by total assets estimated over the 1992–2004 period. *BETA* is estimated over the period just prior to the end of 2004 using the market model with a minimum of 30 out of 60 monthly returns and a market index equal to the value weighted NYSE/AMEX return. *lnSize* is the log of the market value of equity at December 31, 2004. *lnBM* is the log of the book-to-market ratio measured as book value of equity divided by common shares outstanding multiplied by end of 2004 stock price

Table 3 OLS estimation of first stage of two-stage least squares regression (dependent variable is *lnMFDiscPol*; columns correspond to alternative structural model specifications in Table 4; *t*-statistics from using TOBIT in first stage also provided)

Instrument	Coefficient estimate (*t*-statistic) (TOBIT *t*-statistic)	Coefficient estimate (*t*-statistic) (TOBIT *t*-statistic)	Coefficient estimate (*t*-statistic) (TOBIT *t*-statistic)	Coefficient estimate (*t*-statistic) (TOBIT *t*-statistic)
Intercept	3.3687	3.2543	2.9034	3.1066
	(8.42***)	(8.20***)	(6.94***)	(7.26***)
	(4.52***)	(4.92***)	(3.48***)	(3.77***)
lnAGE	−0.1195	−0.0807	−0.0688	−0.0685
	(−2.99***)	(−2.01**)	(−1.70*)	(−1.70*)
	(−2.75***)	(−1.76*)	(−1.54)	(−1.54)
SIGNROA	0.4462	0.6369	0.6160	0.5956
	(3.97***)	(5.44***)	(5.26***)	(5.06***)
	(3.59***)	(4.92***)	(4.77***)	(4.59***)
lnCapIntensity	0.0394	0.0097	0.1677	0.1466
	(0.15)	(0.04)	(0.62)	(0.54)
	(0.19)	(0.12)	(0.56)	(0.52)
lnShareholders	−0.0843	−0.0599	−0.0564	−0.0445
	(−3.51***)	(−2.48**)	(−2.33**)	(−1.72*)
	(−3.28***)	(−2.37**)	(−2.24**)	(−1.85*)
lnAnalyst	0.6902	0.6972	0.6333	0.7092
	(11.38***)	(11.61***)	(9.78***)	(8.97***)
	(11.30***)	(11.53***)	(9.94***)	(8.57***)
lnHHI	0.2515	0.2497	0.2505	0.2472
	(3.90***)	(3.91***)	(3.93***)	(3.88***)
	(4.06***)	(4.14***)	(4.14***)	(4.08***)
lnOffer	−0.5379	−0.5737	−0.5833	−0.6126
	(−2.90***)	(−3.12***)	(−3.18***)	(−3.29***)
	(−2.82***)	(−3.06***)	(−3.13***)	(−3.24***)
REGULATE	−0.4850	−0.4056	−0.4559	−0.4922
	(−1.94*)	(−1.64)	(−1.84*)	(−1.98**)
	(−1.87*)	(−1.57)	(−1.71*)	(−1.83*)
HIGHTECH	0.2006	0.0303	−0.0686	−0.0515
	(1.77*)	(0.26)	(−0.56)	(−0.42)
	(1.26)	(−0.15)	(−0.79)	(−0.69)
lnEarnVar		0.2266	0.1936	0.2030
		(5.24***)	(4.30***)	(4.07***)
		(4.92***)	(4.15***)	(4.00***)
Beta			0.2530	0.2308
			(2.60***)	(2.35**)
			(2.02**)	(1.77*)

Table 3 continued

Instrument	Coefficient estimate (t-statistic) (TOBIT t-statistic)	Coefficient estimate (t-statistic) (TOBIT t-statistic)	Coefficient estimate (t-statistic) (TOBIT t-statistic)	Coefficient estimate (t-statistic) (TOBIT t-statistic)
lnSize				−0.0422
				(−0.97)
				(−0.45)
lnBM				0.1544
				(2.00**)
				(1.63)
N	1,355	1,355	1,355	1,355
Adjusted R^2	0.1317	0.1484	0.1521	0.1558
Partial F-statistic (unique instruments)	23.82	26.58	23.08	21.32

*, **, and *** indicate statistical significance at $p < 0.1$, $p < 0.05$ and $p < 0.01$, respectively using a two-tailed test

Variable definitions, distributions, and intercorrelations appear in Tables 1 and 2

The regression results in Table 3 indicate that several *incremental* associations exist between the unique instruments and management earnings forecast policy (i.e., the dependent variable *lnMFDiscPol*). The quality of the disclosure policy is greater for younger, profitable firms, with fewer shareholders, more analyst coverage, smaller percentage change in equity shares from new issues, and in more concentrated industries, which tend to be high-tech and not regulated.[17] The adjusted R^2 in the first-stage regression ranges from 13.17 to 15.58%, and the partial F-statistic relating to the instruments unique to the first stage ranges from 21.32 to 26.58. The strength of the first-stage instruments is of primary concern in a two-stage least squares regression. Larcker and Rusticus (2010) demonstrate the potential problems with two-stage least squares in the presence of weak instruments, and they suggest a simple way to detect the presence of weak instruments, an examination of the partial F-statistic in the first stage which, if low, would indicate weak instruments. They reference Stock et al. (2002), who develop benchmarks for the necessary size of the F-statistic. When the number of instruments is nine, the critical F-statistic is 19.71. F-statistic values below these indicate that the instruments are weak and inference problems are potentially serious. Our F-statistic

[17] Again, these results are incremental associations in a multiple regression, and we are not concerned with the signs of the relations, only the joint ability of the instruments to capture management forecast policy choice (Maddala 1977). For example, although the negative sign on number of shareholders is counterintuitive, number of shareholders is highly positively associated with number of analysts (see Table 2), and number of analysts loads heavily in the expected direction in the regression. Thus, the result on number of shareholders is after control for the number of analysts. The results on all variables are consistent when using TOBIT regressions with the exception of HIGHTECH, which loses significance.

Table 4 Second stage regression results: association between and cost of equity capital (dependent variable) and the fitted value for the management forecast policy measure from the Table 3 regression, *fitMFDiscPol* (independent variable of interest); $N = 1{,}355$ firms; *t*-statistics for this second stage regression are presented for both a first stage OLS and a first stage TOBIT estimation

Variable	Expected sign	Univariate association with cost of capital Coefficient estimate (*t*-statistic) (TOBIT *t*-statistic)	Association with control for earnings quality measure Coefficient estimate (*t*-statistic) (TOBIT *t*-statistic)	Association incremental to CAPM control Coefficient estimate (*t*-statistic) (TOBIT *t*-statistic)	Association incremental to Fama–French risk factors Coefficient estimate (*t*-statistic) (TOBIT *t*-statistic)
Intercept	±	0.1287	0.1629	0.1398	0.1675
		(25.81***)	(26.41***)	(21.13***)	(23.68***)
		(60.25***)	(37.27***)	(22.41***)	(23.92***)
Hypothesized effect					
fitMFDiscPol	−	−.0107	−0.0094	−0.0132	−0.0056
		(−4.49***)	(−4.36***)	(−5.61***)	(−2.49***)
		(−4.89***)	(−4.46***)	(−5.77***)	(−2.45***)
Control variables					
lnEarnVar	+		0.0121	0.0084	0.0046
			(10.46***)	(6.69***)	(3.15***)
			(10.54***)	(6.82***)	(3.15***)
Beta	+			0.0212	0.0220
				(7.13***)	(8.00***)
				(7.17***)	(7.99***)
lnSize	−				−0.0066
					(−6.42***)
					(−6.26***)
lnBM	+				0.0070
					(2.97***)
					(2.92***)

Table 4 continued

Variable	Expected sign	Univariate association with cost of capital Coefficient estimate (*t*-statistic)(TOBIT *t*-statistic)	Association with control for earnings quality measure Coefficient estimate (*t*-statistic)(TOBIT *t*-statistic)	Association incremental to CAPM control Coefficient estimate (*t*-statistic)(TOBIT *t*-statistic)	Association incremental to Fama–French risk factors Coefficient estimate (*t*-statistic)(TOBIT *t*-statistic)
Adjusted R^2		0.0139	0.0779	0.1007	0.1547

*, **, and *** indicate statistical significance at $p < 0.1$, $p < 0.05$ and $p < 0.01$, respectively using one-tailed tests or two-tailed tests if no clear sign prediction

Variable definitions, distributions, and intercorrelations appear in Tables 1 and 2

fitMFDiscPol is the fitted-value from the first-stage regression in Table 3. This table's estimates and statistics are based on a first stage OLS regression (see Table 3). The second stage regression standard errors are corrected for the use of a first stage fitted value in the second stage regression for nine instruments rejects the weak instrument null in each of the Table 3 regressions.[18]

The fitted value from the Table 3 first-stage regression (*fitMFDiscPol*) is used as the primary independent variable of interest in the Table 4 second-stage regression.[19] In the first column of Table 4, we estimate the management earnings forecast disclosure policy effect without control variables to gauge the effect in isolation. The coefficient on *fitMFDiscPol* is significantly negative ($p < 0.01$), indicating lower cost of capital for firms with a higher quality management earnings forecast policy during the 4-year period. The second column of Table 4 adds the earnings quality control variable in the spirit of Francis et al. (2008). *EarnVar* is positively associated with cost of equity capital, as expected ($p < 0.01$). The coefficient on *fitMFDiscPol* remains significantly negative ($p < 0.01$) and changes little.

The third column of Table 4 adds CAPM beta. Beta is positively related to cost of equity capital as expected ($p < 0.01$). After control for CAPM factors, *fitMFDiscPol* remains significantly negatively related to the cost of equity capital ($p < 0.01$). The final column adds Fama–French factors, firm size and book-to-market. Firm size is negatively related to cost of equity capital ($p < 0.01$), as expected, and book-to-market is positively associated, as expected ($p < 0.01$). Again, *fitMFDiscPol* is negatively related to cost of equity capital ($p < 0.01$) after control for CAPM and additional Fama–French risk factors. Our conclusions are not altered if we use TOBIT in the first stage estimation.[20]

In summary, after taking into account the endogeneity of management forecasting policy, the historical earnings quality from mandated reports, and known correlates with cost of equity capital established by the CAPM and Fama–French analyses, higher quality management forecast disclosure policy is associated with a lower cost of equity capital.

To provide evidence on the intertemporal stability of the negative association between the quality of management forecast disclosure policy and cost of equity

[18] The adjusted R^2s of 13.17–15.58% are also relatively high when compared with recent published work. As examples, Brown and Hilligeist (2007) report a pseudo R^2 of 8.2% for their disclosure quality first-stage regression, and Barton and Waymire (2004) report an R^2 of 11% and F-statistic of 4.05 with 14 instruments.

[19] All results are reported after truncating values of independent and dependent variables at the 1st and 99th percentiles. Results also are not affected by the inclusion or exclusion of financial firms. The Table 4 estimation is also replicated using TOBIT in the first stage.

[20] Easton and Monahan (2005) show that higher long-term growth forecasts by analysts are associated with their forecast errors, and thus are a good proxy for financial analyst forecast quality. They show that ex ante cost of equity capital estimates are more reliable, though still fraught with measurement error, in a sub-sample of firms with lower long-term growth forecasts. In a supplemental test (not tabulated), we replicate our PEG-based tests with a sample constructed to match the low long-term growth forecast (and hence more reliable cost of equity capital) sub-sample in Easton and Monahan's paper. We discard firms with high long-term growth estimates (>15%) to obtain a sample of 884 firms with mean and median long-term growth of 8.8 and 9.2%, respectively. In this sub-sample, the second-stage t-statistic on the management forecast disclosure policy variable is significantly negative, as expected ($t = -1.88$). We also extended the PEG-based tests by including an interaction of *MFDiscPol* with the analysts' long-term growth forecast to examine whether our results are driven by analyst forecast quality. The coefficient on *MFDiscPol* remains significantly negative, and the coefficient on the interaction term is insignificant.

capital, we recompute all research variables on a 1-year basis and estimate the results using two-stage least squares for individual years 2001 through 2006 separately (results not tabulated). We find the expected significant negative coefficient on *fitMFDiscPol* in 4 of the 6 years, 2003 through 2006. We do not expect single year tests to be the most powerful tests of the hypothesis, especially for years immediately following a regulatory change. Strongest results in 2003 through 2006 are also consistent with those years being more representative of a forecasting policy. Given the passage of Reg FD in late 2000, it is not surprising that firms would have to reestablish their disclosure policies in light of the new regulation (Wang 2007). Given that disclosure policy is not typically announced, the market would need to observe forecasting behavior after the new law to infer any new policy.

In untabulated results, we regress *MFDiscPol* for each year on the same variable for the other years to assess the ability to predict a given year's disclosure with the other years' disclosures. Using each regression's coefficient of determination to measure ease of prediction, management forecast disclosure in 2003, 2004, and 2005 are the easiest to predict given management forecast disclosure in other years. The fact that these are the 3 years with the strongest negative relations between management forecast disclosure policy and cost of capital, combined with the fact that these years have management forecast disclosure that is easiest to predict given other years' disclosures, suggests that the years' forecast activity is most consistent with a management forecast policy. Finally, we examined whether particular patterns of disclosure within the 4-year period would affect our results. Our main test's approach of summing disclosure over a four-year period does not capture disclosure timing within that period. Essentially, each year's disclosure is given equal weighting. We tried several alternative disclosure weightings that would give higher or much higher weight to more recent management forecast disclosures. Our results (not tabulated) are consistent with our main finding that management forecast disclosure policy is negatively correlated with cost of equity capital.

4.3 Effects of disclosure costs and management quarterly forecast relevance (H2 and H3)

In this section, we test whether the negative association between the quality of a firm's management earnings forecasting policy and its cost of equity capital is increasing in its disclosure costs and the relevance of its quarterly management earnings forecasts. Prior research has not tested directly for cross-sectional differences in the disclosure/cost of capital relation based on the expected costs and benefits of disclosure. Because disclosure is costly, the relation should be stronger for firms with higher disclosure costs. Further, the capital market benefits of a policy to disclose a particular piece of accounting information are increasing in the usefulness of the information in security pricing.

Following Cohen (2006), we use four proxies for disclosure costs: current product market competition as captured by the Herfindahl-Hirschman index (*HHI*), capital intensity (*CapIntensity*), the expected litigation costs of operating in a high-tech industry (*HIGHTECH*), and growth opportunities as reflected in the book-to-market ratio (*BM*). We examine whether these costs individually affect the association between management earnings forecast policy and cost of equity capital.

To test our disclosure benefit hypothesis, we identify the firms with higher information content of management quarterly earnings forecasts and examine whether the strength of the relation is stronger for these high disclosure benefit firms. We measured information content of a firm's management forecasts by the average absolute price reaction to its quarterly management forecasts (*Avgabscar*) during the sample period preceding the end-of-2004 cost of capital measurement.

We modify second-stage equation (4) to include intercept and slope-shifts for disclosure costs and information content of quarterly management forecasts. High disclosure costs are indicated by higher values of *HIGHTECH* (i.e., $HIGHTECH = 1$) and lower values of *HHI*, *CapIntensity*, and *BM*. Therefore, we multiply the latter three proxies by a negative one to cause higher values to indicate high disclosure costs. Thus, the expected sign of each slope-shift is negative to indicate a stronger negative relation for higher disclosure costs. A stronger negative relation for more relevant quarterly management forecasts is indicated by a negative coefficient on the *Avgabscar* slope-shift.

Table 5's first four columns present the results on H2 for each individual disclosure cost proxy. The final column presents results on the test the disclosure relevance hypothesis (H3). The results are consistent with our predictions that higher disclosure costs and higher disclosure relevance strengthen the negative relation between management earnings forecast disclosure policy and cost of equity capital. All slope shift coefficients in the table are significantly negative, as expected.[21]

5 Robustness tests

In the sections that follow, we perform a battery of additional tests to address the robustness of our main results on management forecast disclosure and cost of equity

[21] We instrument each interaction term using the set of instruments for *MFDiscPol* described earlier except that we do not, for example, use *lnHHI* as a first stage instrument for its interaction with *MFDiscPol*. Also, because the variable of primary interest, *fitMFDiscPol*, is the predicted value from the first stage, it is correlated with a given first-stage variable to the extent that the first-stage variable explains it. The results we report in Table 5 assume that multicollinearity does not affect the coefficient estimates. To make sure that it does not, we (a) re-estimated the regression using OLS (given that OLS estimation yields the same conclusions in other tests), (b) estimated the first-stage of the two-stage least squares with the given cost proxy omitted so that the fitted first-stage variable is not correlated with the cost proxy, and (c) estimated separate regressions for high and low disclosure cost cases. Our conclusions do not vary across these alternative estimation techniques except for the Herfindahl Index variable (proxy for product market competition). The negative significant relation between management forecast disclosure quality and cost of equity capital is similar in both the high and low cost subsamples. However, we use continuous variables to measure disclosure costs in our primary tests, and partitions at the median are made to form the high and low cost groups for the separate regression tests. Thus, the separate regression tests also discard information and ignore the possibility that sufficiently high disclosure costs occur at a place other than the median. We do not test all disclosure cost proxies jointly because we are not interested in incremental effects, and instrumenting each interaction with the same instruments leads to severe collinearity.

Table 5 Second stage results on regressions of cost of equity capital on the fitted value for the management forecast policy measure, *fitMFDiscPol*, conditioned on firm-specific disclosure costs and the value-relevance of quarterly management earnings forecasts ($N = 1{,}355$ firms)

Variable	Expected sign	Disclosure cost proxies				
		Herfindahl index Coefficient estimate (*t*-statistic)	Capital intensity Coefficient estimate (*t*-statistic)	High-tech industry Coefficient estimate (*t*-statistic)	Book to market Coefficient estimate (*t*-statistic)	Quarterly forecast disclosure relevance proxy Coefficient estimate (*t*-statistic)
Intercept	±	0.1751	0.1718	0.1691	0.1307	0.1492
		(17.69***)	(19.25***)	(22.84***)	(11.22***)	(8.03***)
Hypothesized effects						
fitMFDiscPol	−	0.0103	−0.0297	−0.0064	0.0231	0.0141
		(1.43)	(−4.15***)	(−2.68***)	(3.54)	(2.15)
fitMFDiscPol × *lnHHI*	−	−0.0068				
		(−2.77***)				
fitMFDiscPol × *lnCapIntensity*	−		−0.0848			
			(−3.56***)			
fitMFDiscPol × *HIGHTECH*	−			−0.0111		
				(−2.40***)		
fitMFDiscPol × *lnBM*	−				−0.0297	
					(−4.85***)	
fitMFDiscPol × *Avgabscar*	−					−0.0843
						(−2.33***)
Intercept shifts						
lnHHI	±	0.0019				
		(0.95)				
lnCapIntensity	±		0.0060			
			(0.60)			

Table 5 continued

Variable	Expected sign	Disclosure cost proxies				
		Herfindahl index Coefficient estimate (*t*-statistic)	Capital intensity Coefficient estimate (*t*-statistic)	High-tech industry Coefficient estimate (*t*-statistic)	Book to market Coefficient estimate (*t*-statistic)	Quarterly forecast disclosure relevance proxy Coefficient estimate (*t*-statistic)
HIGHTECH	±			−0.0065		
				(−1.66*)		
Avgabscar	±					−0.0505
						(−1.56)
Control variables						
lnEarnVar	+	0.0064	0.0056	0.0064	0.0178	0.0030
		(3.94***)	(3.13***)	(4.01***)	(5.32***)	(1.53*)
Beta	+	0.0288	0.0400	0.0336	0.0103	0.0209
		(8.50***)	(6.93***)	(7.38***)	(2.41***)	(6.13***)
lnSize	−	−0.0057	−0.0074	−0.0065	0.0052	−0.0093
		(−5.15***)	(−5.93***)	(−6.10***)	(1.87)	(−6.68***)
lnBM	+	0.0071	0.0091	0.0068	0.0027	0.0071
		(2.85***)	(3.14***)	(2.80***)	(0.88)	(2.38***)
Adjusted R^2		0.1503	0.1210	0.1540	0.1112	0.1346

*, **, and *** indicate statistical significance at $p < 0.1$, $p < 0.05$ and $p < 0.01$, respectively using one-tailed tests or two-tailed tests if no clear sign prediction. *fitMFDiscPol* is the fitted-value from a first-stage regression instrumented as in Table 3. All interactions are fitted values from a first-stage regression instrumented as in Table 3, omitting the conditioning variable as an instrument. The second stage regression standard errors are corrected for the use of first stage fitted values in the second stage regression. *Avgabscar* is the average absolute price reaction (3-day abnormal returns) to the firm's quarterly management earnings forecasts issued during the sample period. Remaining variable definitions, distributions, and intercorrelations appear in Tables 1 and 2

capital to various research design issues. We perform (a) ordinary least squares tests that ignore endogeneity, (b) Heckman (1979) tests to control for self-selection, (c) ordinary least squares tests that treat instruments as additional control variables, (d) tests employing ex post realized returns tests as an alternative to ex ante cost of capital estimation for a much larger sample of firms, and (e) the aforementioned Easton (2009) dummy variable approach in a cross-sectional ordinary least squares regression.

5.1 Ordinary least squares and Heckman approaches

In our main analysis, we appeal to cost of equity capital theory to exclude our instrumental variables on a priori grounds. We rely on a priori reasoning given van den Berg's (2006) argument that exclusion restrictions are identifying restrictions that cannot be tested empirically.[22] In this section, we examine our results using three alternative specifications. First, we ignore endogeneity by estimating the relation using OLS regression. Second, we address endogeneity using the alternative Heckman (1979) approach to model the first stage regression with a Probit choice model and then include the Inverse Mills ratio from the first stage as an additional control variable in a second stage.[23] Third, one could view the first-stage variables omitted in the OLS approach as a set of potential correlated omitted variables which, through their likely association with other disclosure practices, might be associated with cost of equity capital incremental to the management earnings forecast policy choice. Accordingly, we estimate the OLS model with the first stage determinants of voluntary disclosure simply appended as additional control variables. In contrast to the OLS approach, which might be biased against the null in the presence of correlated omitted variables, this approach is biased in favor of the null to the extent that the additional control variables explain the management forecast disclosure policy choice.

The conclusions from these additional tests are identical to conclusions from the two-stage least squares approach (results not tabulated). The quality of management earnings forecast policy is negatively associated with cost of equity capital, as expected.[24]

[22] Larcker and Rusticus (2008) do provide an over-identifying restrictions test of the appropriateness of the instruments that can be applied when the number of instruments exceeds the number of endogenous regressors. The over-identifying restrictions test regresses the second stage residuals of a 2SLS estimation on all exogenous instruments. If the instruments are valid, then the R^2 from the model should be close to zero. Larcker and Rusticus (2008) note that nR^2 in this test is distributed χ^2 with K–L degrees of freedom where K is the number of exogenous variables unique to the first stage, and L is the number of endogenous explanatory variables. We ran this test on the Table 4 model with all controls and obtained an R^2 of 4%. The χ^2 statistic is significantly different from zero at the 0.001 level. However, Larcker and Rusticus note that this test nearly always rejects in large samples.

[23] The independent variable of interest in the OLS regression is *lnMFDiscPol* (described previously). In the interest of brevity, we do not provide the details of the Heckman choice model, which may be found in Heckman (1979). Also, a recent application of the model appears in Feng et al. (2009). The Heckman approach uses a Probit model in the first stage. Therefore, we transform *MFDiscPol* to obtain a first stage dummy dependent variable *MFDiscDummy* that equals one if *MFDiscPol* is greater than its median and zero otherwise. We use all exogenous variables to estimate the first stage.

[24] The inverse Mills ratio is significant in three of the four Heckman-type regressions.

5.2 Tests based on ex post realized returns

5.2.1 Description of test

Although Botosan and Plumlee (2005) find that the cost of capital measure based on the PEG method is most highly associated with known risk factors, Easton and Monahan (2005) provide evidence that calls into question the reliability of the PEG method and many other ex ante cost of equity capital measures by showing that the measures are not associated with ex post realized returns unless financial analyst forecasts are of high quality. Accordingly, we sidestep the low analyst forecast quality issue with a robustness test using ex post realized returns, which does not require the availability of financial analysts and hence does not suffer from the low quality financial analyst forecast problem. Also, the relaxation of the financial analyst data requirement permits the use of a much larger sample.[25]

Core et al. (2008) present a two-step procedure using ex post returns to jointly test individual firms' factor loadings and the significance of the factor premia. As noted by Francis et al. (2008), the test does not presume the validity of standard asset pricing models. Following the approach outlined by Francis et al. (2008), in the first step we estimate firm-specific regressions of excess daily returns, measured as each firm's daily return (R_{it}) less the risk free rate (R_{Ft}), on several potential pricing factors:

$$\begin{aligned} R_{it} - R_{Ft} = {} & a_i + d_i HiLoMFDiscPol_t + e_i HiLoEarnVar_t + b_i RMRF_{it} + s_i SMB_{it} \\ & + h_i HML_{it} + \zeta i_{it} \end{aligned} \tag{5}$$

where *HiLoMFDiscPol* is a portfolio return resulting from ranking all firms on the fitted values (*fitMFDiscPol*) from an instrumental variables regression analogous to the one described previously and taking a long (short) position in firms within the top (bottom) 20% of *fitMFDiscPol*. *HiLoEarnVar* is a similar portfolio formed using *EarnVar*. *RMRF*, *SMB*, and *HML* are also portfolio returns supplied by Kenneth French, as described in Fama and French (1993). Because this is a firm-specific model, we run this regression 3,686 times using, on average, 354 observations for each firm and retain the firm-specific loadings for use in the second step.

The dependent variable in our second step regression is the mean daily excess return over the sample period for each of the 3,686 firms. This mean is measured by taking the average of the firm-specific excess daily returns, the step one dependent variable, over the entire sample period. We then regress the mean daily excess return on the parameter estimates (3,686 sets of firm-specific factor loadings) from the first stage regression:

$$\overline{R_i - R_F} = \alpha_i + \delta\hat{d}_i + \varepsilon\hat{e}_i + \beta\hat{b}_i + \sigma\hat{s}_i + \theta\hat{h}_i + \xi_i \tag{6}$$

[25] As noted by Francis et al. (2008), the choice between ex ante and ex post (i.e., realized returns) cost of capital proxies remains controversial. Asset pricing tests in the finance literature are based on broad samples and time periods, and thus a caveat is warranted for any study that employs limited samples and time periods in the analysis.

The parameter estimates from Eq. 6 represent estimates of the factor premia over the sample period. Our hypothesis of a negative association between cost of equity capital and management earnings forecast disclosure quality suggests $\delta < 0$.

A deviation between ex post returns and expected returns exists in the presence of cash flow news (Fama and French 2002; Easton and Monahan 2005), which is defined as the revisions during the period of expected benefits accruing to equity shareholders (Campbell 1991). Ogneva (2009) demonstrates that realized returns tests can yield incorrect inferences if cash flow news is ignored. If firms issue management earnings forecasts because of poor performance (e.g., to manage analysts' expectations downward or to avoid legal liability), realized returns may be lower, not due to the effect of disclosure on the cost of capital component of realized returns but due to the "cash flow news" component of realized returns (Campbell 1991). To guard against the possibility that management earnings forecast policy is associated with firm performance during the period in which returns are realized, we also include cash flow news as an additional control variable in Eq. 6. We calculate the cash flow news variable for each year in our sample following Easton and Monahan (2005):

$$\begin{aligned} \hat{cn}_{it+1} &= (roe_{it} - froe_{it,t}) + (froe_{it+1,t+1} - froe_{it,t+1}) \\ &\quad + \frac{\rho}{1 - \rho \times \omega_t} \times (froe_{it+1,t+2} - froe_{it,t+2}) \end{aligned} \tag{7}$$

$roe_{it} = \ln(1 + ROE_{it})$, where $ROE_{it} = eps_{it}/bps_{it-1}$; eps_{it} is reported earnings per share for year t per I/B/E/S; bps_{it-1} is equity book value at the end of year $t - 1$ divided by common shares outstanding at the end of year $t - 1$. For our study, year t is 2004, so we use $eps_{i,2004}$ and $bps_{i,2003}$. $froe_{it,t} = \ln(1 + ROE_{ij,k})$, where $ROE_{ij,k}$ is the forecasted return on equity for fiscal year k based on consensus analyst forecasts made at the end of year j. For our study, we use forecasted 2004 earnings at the end of 2004 divided by 2003 book value per share: $feps_{i2004,2004}/bps_{i,2003}$. ω_t comes from the following regression:

$$roe_{i\tau+1} = \omega_{ot} + \omega_t \times roe_{i\tau} \tag{8}$$

estimated by Fama and French (1992) industry code, where τ is a number between t and $t - 9$.

To obtain ρ, consistent with Ogneva (2009), we form five price-to-dividend portfolios consisting of one made up of nondividend payers and four quartiles of dividend payers, then match them to the values of ρ in Easton and Monahan (2005, Appendix A). In computing the cash flow news variable, we delete observations consistently with Easton and Monahan (2005, Table 1 notes).

5.2.2 Results of ex post returns tests

Comparing the descriptive statistics on the 3,686 firms used in the ex post realized returns tests (not tabulated) to those presented for the 1,355 firms used in our ex ante cost of capital tests based on the PEG ratio (Table 1) reveals that removing the need for analyst following results in a much larger sample of smaller, younger firms with greater earnings variability, fewer shareholders, far less analyst coverage, more occurrences of losses (e.g., negative ROA's), and a far smaller average management earnings forecast disclosure score. Following Leuz and Verrecchia's (2000) argument that less developed information environments likely provide the greatest chance to document the effects of disclosure on cost of capital, we expect relatively powerful tests of our hypothesis with this diverse set of firms.

Table 6, presents the first stage instrumental variable-based estimation. Higher quality management earnings forecast disclosure policy is (incrementally) associated with younger firms, more profitable firms, with lower capital intensity, fewer shareholders, and higher analyst following, in concentrated, nonregulated, high-tech industries. The adjusted R^2 of 24.51% and F-statistic of 133.95 with nine instruments indicates relatively strong instruments.[26] In this much larger sample, over half of the firms do not issue management forecasts. Therefore, the censoring of our sample is more severe, and using OLS instead of TOBIT in the first-stage might be problematic. Accordingly, we provide t-statistics using TOBIT, and we replicate our results in Table 8, using first-stage TOBIT estimates. Note that use of TOBIT to estimate the first-stage yields nearly identical conclusions about the relation of the instruments to management forecast disclosure policy.

Table 7, shows the returns over the period for which we collect realized returns (July 1, 2005, to December 31, 2006) for each portfolio. Consistent with our hypothesis, the portfolio, which goes long in the highest 20% of our management earnings forecast disclosure policy quality measure (the fitted value from the first-stage instrumental variables estimation from Table 7) and short in the lowest 20%, has a compounded annual return of −2.79%. In order to create *HiLoMFDiscPol* from the TOBIT procedure, we rank the predicted values from the first stage TOBIT regression. For the 3,686 firms, 1,899 had predicted values of zero, which is well over the 20% needed for the portfolio. We use simple random sampling (proc surveyselect in SAS) to generate 737 firms from the 1,899 firms with a predicted value of zero. These 737 firms make up the low 20% of the sample in creating *HiLoMFDiscPol*. Table 6 reports that the annual return for the portfolio using TOBIT first-stage estimates is a relatively large −3.3%. To calibrate the magnitude of these returns, note that the market return portfolio return is 10.91% over the same period.

Table 8, shows the results of the two-step test as described in Eqs. 5 and 6. Results for both OLS and Tobit appear in five columns to correspond to various versions of Eq. 6. The first two columns exclude other risk factors. The factor premium on the *HiLoMFDiscPol* portfolio is significantly negative ($t = -9.72$) when considered in isolation, as expected. Controlling for an historical proxy for information precision (the *HiLoEarnVar* portfolio) does not remove the disclosure

[26] The second stage is a trading strategy based solely on the fitted management forecast policy variable from the first stage, which alternatively could be formulated as a linear regression with a single indicator variable that captures the trading strategy returns. Therefore, no additional exogenous variables appear in the second stage for inclusion in the first stage.

Table 6 Ex post realized returns tests. First stage OLS (TOBIT) regression with *lnMFDiscPol* (*MFDiscPol*) as the dependent variable (3,686 firms)

Instruments	Coefficient estimate	*t*-statistic/TOBIT *t*-statistic
Intercept	1.5443	7.77***/1.93*
lnAge	−0.0487	−4.90***/−4.62***
SignROA	0.4167	7.93***/7.67***
lnCapIntensity	−0.4748	−3.39***/−4.00***
lnShareholders	−0.0232	−1.88**/−1.98**
lnHHI	0.1840	5.38***/4.94***
lnOffer	−0.0165	−0.42/−0.73
lnAnalyst	0.6877	29.18***/25.86***
Regulate	−0.3249	−2.74***/−2.97***
Hightech	0.2984	4.83***/4.53***
Adjusted R^2	0.2451	
F-value	133.95***	

*, **, and *** indicate statistical significance at $p < 0.1$, $p < 0.05$ and $p < 0.01$, respectively using a two-tailed test

Variable definitions appear in Table 1. As in prior tests, natural logs are used in the empirical analysis. For example, *lnMFDiscPol* equals the natural log of *MFDiscPol* + 1

Table 7 Ex post realized returns tests. Returns for factor portfolios (*HiLoMFDiscPol* portfolio formed using fitted values from Table 6)

Portfolio	Mean daily return	Annual return (compounded)
HiLoMFDiscPol	−0.0111	−2.7972
		−3.3012 (using TOBIT-formed portfolios)
HiLoEarnVar	0.0032	0.8064
RMRF	0.0433	10.9116
SMB	0.0039	0.9828
HML	0.0307	7.7364

HiLoMFDiscPol represents the portfolio of the return of the 20% of firms with the largest value of *MFDiscPol* minus the return of the 20% of firms with the smallest values of *MFDiscPol* overall trading days between July 1, 2005, and December 31, 2006. *HiLoEarnVar* represents the portfolio of the return of the 20% of firms with the largest value of *EarnVar* minus the return on the 20% of firms with the smallest values of *EarnVar* overall trading days between July 1, 2005, and December 31, 2006. *RMRF* is the average daily and annualized excess market return over all trading days between July 1, 2005, and December 31, 2006. *SMB* is the Fama and French SMB factor over all trading days between July 1, 2005, and December 31, 2006. *HML* is the Fama and French HML factor over all trading days between July 1, 2005, and December 31, 2006

policy effect. Adding the market portfolio (CAPM controls column) does not affect the results, and the market portfolio return is positive and significant as predicted by the CAPM. Adding size and book-to-market effects (Fama/French controls) weakens the results somewhat, but the coefficient on the disclosure quality portfolio remains significantly negative and has a factor premium magnitude that is similar to

Table 8 Ex post realized returns tests. Factor premia estimated from the second step of two-step regression (Text Eq. 6; estimates and *t*-statistics are presented for the case where OLS was used to instrument the management forecast disclosure policy variable. *T*-statistics using TOBIT to instrument are presented for comparison.)

Coefficients on the following factors estimated in text Eq. 5	Excluding other risk factors		CAPM controls	Fama/French controls	Add Ogneva cash flow news control
Intercept	0.0010	0.0007	−0.0002	−0.0004	0.002
	(14.66***)	(12.32***)	(−2.11**)	(−4.50***)	(2.41***)
	(TOBIT:12.60)	(TOBIT: 10.59)	(TOBIT: 0.95)	(TOBIT: −0.04)	(TOBIT: 2.37)
HiLoMFDiscPol	−0.0005	−0.0004	−0.0003	−0.0001	−0.0002
	(−9.72***)	(−8.67***)	(−9.03***)	(−2.37***)	(−2.81***)
	(TOBIT: −6.26)	(TOBIT: −6.34)	(TOBIT: −7.67)	(TOBIT: −3.20)	(TOBIT: −3.40)
HiLoEarnVar		0.0001	0.0004	0.0001	0.0002
		(8.72***)	(8.90***)	(2.12**)	(3.84***)
		(TOBIT: 6.96)	(TOBIT: 7.59)	(TOBIT: 3.11)	(TOBIT: 3.81)
RMRF			0.0006	0.0011	0.0003
			(7.87***)	(16.10***)	(3.89***)
			(TOBIT: 4.37)	(TOBIT: 9.58)	(TOBIT: 3.92)
SMB				−0.0002	0.0002
				(−3.31)	(4.87***)
				(TOBIT: −3.16)	(4.93)
HML				−0.0001	0.0029
				(−3.45)	(0.08)
				(TOBIT: 3.68)	(TOBIT: −0.21)
Cash flow news					0.0006
					(4.31***)
					(TOBIT: 4.31)

Table 8 continued

Coefficients on the following factors estimated in text Eq. 5	Excluding other risk factors		CAPM controls	Fama/French controls	Add Ogneva cash flow news control
Adjusted R^2	0.0247	0.0199	0.1146	0.1802	0.1079
Adjusted R^2 (TOBIT)	0.0103	0.0127	0.0541	0.0859	0.1112

*, **, and *** indicate statistical significance at $p < 0.1$, $p < 0.05$ and $p < 0.01$, respectively using a one-tailed test (two-tailed on intercept)

In the first step we estimate firm-specific regressions of excess daily returns, measured as each firm's daily return (R_{it}) less the risk free rate (R_{Ft}), on several potential pricing factors:

$R_{it} - R_{Ft} = a_i + d_i HiLoMFDiscPol_t + e_i HiLoEarnVar_t + b_i RMRF_{it} + s_i SMB_{it} + h_i HML_{it} + \zeta i_{it}$ (5)

Using the portfolio returns summarized in Table 7. We run this regression 3,686 times using, on average, 354 observations for each firm and retain the firm-specific loadings for use in the second step. The dependent variable in the second step regression is the mean daily excess return over the sample period for each of the 3,686 firms. This mean is measured by taking the average of the firm-specific excess daily returns, the step one dependent variable, over the entire sample period. We then regress the mean daily excess return on the parameter estimates (3,686 sets of firm-specific factor loadings) from the first stage regression:

$\overline{R_i - R_F} = \alpha_i + \delta\hat{d}_i + \varepsilon\hat{e}_i + \beta\hat{b}_i + \sigma\hat{s}_i + \theta\hat{h}_i + \xi_i$ (6)

The parameter estimates from Eq. 6 represent estimates of the factor premia over the sample period. These are the estimates reported in Table 8

Cash flow news is estimated via methods specified in Easton and Monahan (2005) and Ogneva (2009) the earnings quality portfolio. Adding the cash flow news variable as in Ogneva (2009) slightly improves the significance of the disclosure quality variable. Consistent with Ogneva's findings, the cash flow news variable is significantly positive, and the coefficient on earnings quality improves substantially.[27] The results using TOBIT are very similar to the OLS results.

In summary—and subject to the caveat relating to the interpretation of asset pricing tests using limited subsamples and time periods—the ex post realized returns tests support our main analysis using the ex ante implied cost of equity capital proxy derived from analysts earnings forecasts. Our results remain after controlling for cash flow news in the realized returns tests.

5.3 Easton (2009) approach

Our goal is to examine whether firms with higher quality management earnings forecast policies have lower costs of equity capital. We use firm-specific cost of capital estimates in our primary tests and thus are subject to the criticism that firm-specific estimates contain significant measurement error. While measurement error tends to bias regression coefficients toward the null of no association, measurement error does raise the possibility that an observed relationship is spurious. A source of measurement error in estimating firm-specific cost of capital is poor quality financial analyst forecasts.

Easton (2009) notes that hypotheses such as ours do not require firm-specific estimation of implied cost of equity capital and presents a method to assess how a variable of interest affects cost of equity capital without relying on firm-specific estimates. His approach uses a dummy variable to partition the sample into portfolios of firms on the variable of interest. In our case, we can use a dummy variable that is equal to one for firms with high *MFDiscPol*, based on the median *MFDiscPol*, and zero otherwise. Easton's approach is based on a model used in Easton et al. (2002) that simultaneously estimates cost of equity capital and growth in residual earnings implied by current stock price, current book value of equity, and aggregate forecasted cum-dividend earnings. Following Easton's (2009) suggested design, we estimate the following model for a sample of 1,860 firms:

$$\frac{X_{icT}}{BPS_{i,0}} = \xi_0 + \xi_1 \frac{P_{i,0}}{BPS_{i,0}} + \xi_2 D + \xi_3 D \frac{P_{i,0}}{BPS_{i,0}} + \upsilon_{i,0} \tag{9}$$

The dependent variable is the aggregate of forecasted cum-dividend earnings for 2005 through 2008 (X_{icT}) scaled by the book value per share at the end of 2004

[27] The significance of the book-to-market and size effects fluctuate in this last pair of columns, while the market portfolio return and the earnings quality return remain positive and significant. Francis et al. (2008) find the market portfolio and the book-to-market effect to be insignificant in their tests and note that Core et al. (2008) find that only the book-to-market effect is priced as expected in a longer window monthly returns test.

($BPS_{i,0}$).[28] $P_{i,0}$ is the price at the end of 2004, and D represents a dummy variable equal to one for firms with above median *MFDiscPol* and zero otherwise. Note also that the form of the model places forecasted earnings as the dependent variable, which mitigates measurement error concerns. According to Easton et al. (2002), the model without the intercept and slope-shift dummy variables yields coefficient estimates $\xi_0 = [(1+g)^4 - 1]$ and $\xi_1 = [(1+r)^4 - (1+g)^4]$, where g equals growth and r equals implied expected rate of return. The sum of the two coefficient estimates can be used to derive r. Easton (2009) notes that the sum of the dummy variable coefficients, $\xi_2 + \xi_3$, is the effect of being in the portfolio of interest on expected rate of return, r. Our hypothesis suggests that being in the portfolio with higher management forecast disclosure quality is associated with a lower implied expected return, $\xi_2 + \xi_3 < 0$.

Easton (2009) also provides guidance on incorporating control variables into the model:

$$\frac{X_{icT}}{BPS_{i,0}} = \xi_0 + \xi_1 \frac{P_{i,0}}{BPS_{i,0}} + \xi_2 D + \xi_3 D \frac{P_{i,0}}{BPS_{i,0}} + \xi X_{i,0} + \xi \frac{X_{i,0} \times P_{i,0}}{BPS_{i,0}} + \upsilon_{i,0} \quad (10)$$

where $X_{i,0}$ represents a vector of variables that includes *lnEarnVar*, *Beta*, and *lnSize*. *lnBM* is not a control variable because it is the inverse of a regressor needed to estimate cost of capital and growth in the model.

Table 9 presents the estimation of various forms of Eq. 10. The first column estimates the difference in cost of equity capital between two portfolios (high versus low quality management earnings forecast policy) without consideration of additional control variables. We focus our discussion on this simple estimation because inspection of the other columns in which controls are added yields the conclusion that the simple estimation yields results that are robust to the inclusion of controls. Our hypothesis (H1) predicts $\xi_2 + \xi_3 < 0$. The estimate of the sum of the two coefficients is the highly statistically significant −0.1760. The expected rate of return for the low disclosure quality portfolio is $\xi_0 + \xi_1$, which, using the formulae given earlier, equals $[(1+g)^4 - 1] + [(1+r)^4 - (1+g)^4]$. Using Table 9's estimates of $\xi_0 + \xi_1 = 0.6222$ and solving for r yields a 12.86% expected rate of return for the low disclosure quality portfolio. Using the formula and the fact that $\xi_0 + \xi_1 + \xi_2 + \xi_3 = 0.6222 - 0.1760 = .4462$ (i.e., the sum of coefficients for low quality disclosure and the difference between high and low quality disclosure is the estimate for high quality disclosure) yields a 9.66% expected rate of return for the high quality disclosure portfolio.

In summary, the Easton (2009) dummy variable design supports the conclusion that a statistically and economically significant difference in expected rate of return

[28] We use a rate of 12% to estimate cum-dividend earnings and assume that dividends in the current period are equal to dividends in the next four periods. This is consistent with Easton et al. (2002). We use the forecasted EPS for 2005 and 2006 from I/B/E/S and then use growth rates from I/B/E/S to estimate EPS for 2007 and 2008. For firms with only 2005 forecasted EPS available on I/B/E/S, we use the growth rate from I/B/E/S to calculate EPS for years 2006 through 2008. For firms with 2005 and 2006 forecasted EPS but no growth rates available, we calculate their growth rate from the two forecasted earnings numbers and use this rate to estimate 2007 and 2008 EPS. All forecasts are from the I/B/E/S Summary Statistics file at the end of 2004.

Table 9 Application of Easton (2009) dummy variable reverse regression to estimate cost of equity capital for two portfolios (high versus low management forecast disclosure quality; $N = 1{,}774$ firms)

Variable	Coefficient	Univariate association with cost of capital Coefficient estimate (*t*-statistic)	Association with control for earnings quality measure Coefficient estimate (*t*-statistic)	Association incremental to CAPM control Coefficient estimate (*t*-statistic)	Association incremental to Fama–French risk factors (*lnBM* excluded) Coefficient estimate (*t*-statistic)
Intercept	ξ_0	0.5598	0.5095	0.5013	0.4743
		(30.26***)	(21.78***)	(20.40***)	(15.81***)
$\frac{P_{i,0}}{BPS_{i,0}}$	ξ_1	0.0624	0.0615	0.0615	0.0666
		(30.50***)	(13.38***)	(13.37***)	(10.79***)
D	ξ_2	−0.2685	−0.2844	−0.2919	−0.2931
		(−8.26***)	(−8.75***)	(−8.74***)	(−8.77***)
$D\frac{P_{i,0}}{BPS_{i,0}}$	ξ_3	0.0925	0.0905	0.0905	0.0905
		(15.04***)	(14.77***)	(14.69***)	(14.62***)
Hypothesis of interest	$\xi_2 + \xi_3 < 0$	−0.1760	−0.1939	−0.2014	−0.2026
		(−4.44***)	(−4.93***)	(−4.91***)	(−5.04***)
lnEarnVar	ξ_4		0.1284	0.1234	0.1348
			(4.34***)	(4.02***)	(4.27***)
lnEarnVar x $\frac{P_{i,0}}{BPS_{i,0}}$	ξ_5		−0.0002	−0.0012	−0.0030
			(−0.04)	(−0.21)	(−0.53)
Beta	ξ_6			0.0258	0.0246
				(0.85)	(0.80)
Beta x $\frac{P_{i,0}}{BPS_{i,0}}$	ξ_7			0.0017	0.0007
				(0.42)	(0.17)
lnSize	ξ_8				0.0425
					(1.45)

Table 9 continued

Variable	Coefficient	Univariate association with cost of capital Coefficient estimate (*t*-statistic)	Association with control for earnings quality measure Coefficient estimate (*t*-statistic)	Association incremental to CAPM control Coefficient estimate (*t*-statistic)	Association incremental to Fama–French risk factors (*lnBM* excluded) Coefficient estimate (*t*-statistic)
lnSize $x \frac{P_{i,0}}{BPS_{i,0}}$	ξ_9				−0.0056 (−1.31)
Adjusted R^2		0.4807	0.4874	0.4873	0.4874

*, **, and *** indicate statistical significance at $p < 0.1$, $p < 0.05$ and $p < 0.01$ (two-tailed tests)

Variable definitions, distributions, and intercorrelations appear in Tables 1 and 2

$D = MFDiscDummy$, which is equal to 1 if *MFDiscPol* is greater than its median and 0 otherwise

Model : $\frac{X_{icT}}{BPS_{i,0}} = \xi_0 + \xi_1 \frac{P_{i,0}}{BPS_{i,0}} + \xi_2 D + \xi_3 D \frac{P_{i,0}}{BPS_{i,0}} + \xi X_{i,0} + \xi \frac{X_{i,0} \times P_{i,0}}{BPS_{i,0}} + \upsilon_{i,0}$

exists between high and low management forecast quality disclosure portfolios. The method is less susceptible to the measurement error attributed to firm-specific estimation.[29]

6 Conclusion

We extend the literature by examining a specific type of voluntary disclosure rather than disclosure in the aggregate. We choose a high profile, relatively precise, voluntary disclosure of a direct input into equity valuation models, management's quarterly earnings forecasts. We find robust evidence that the quality of management earnings forecasting policy is negatively associated with cost of equity capital, and we document that the strength of the relation is greater for both higher disclosure costs and for firms with more relevant quarterly management earnings forecasts. We use multiple estimation methods to address endogeneity and measurement error in firm-specific estimates of implied cost of equity capital. Evidence of the negative association of management earnings forecasting policy and cost of equity capital does not depend on whether we estimate the association using firm-specific estimates of implied cost of capital with two-stage least squares, ordinary least squares, or a Heckman-type cross-sectional design, whether we derive ex ante implied cost of capital from accounting fundamentals, whether we test the association with ex post realized returns tests both before and after control for cash flow news, and whether we employ a portfolio-based dummy variable regression approach suggested by Easton (2009) to both estimate growth and implied cost of capital simultaneously and address issues with measurement error in firm-specific cost of capital estimates.

Our results highlight the need to refine the definition of "disclosure," which is most often treated in aggregate terms and is frequently measured at a point in time. Not all types of disclosure are equal, and it is unlikely that the capital market consequences of different types of disclosures are equal. Further, forecast disclosures vary, even within a given type of disclosure. Management earnings forecasts are unconstrained in terms of existence, frequency, form, timing, placement with other disclosures, venue, and reason for release, and forecast policy formulation must consider these and other dimensions. The choices managers make in developing forecast policy are likely to lead to fundamentally different effects on cost of equity capital, some of which we document herein. A limitation of our study is that it is an initial and likely incomplete characterization of the richness

[29] Easton (2006) adapts a method in O'Hanlon and Steele (2000) to examine cost of capital differences across regimes. The method, which is similar in spirit to Easton et al. (2002), relies on actual rather than forecasted earnings per share to obtain growth and cost of capital estimates and thus is independent of analyst forecast quality. As an additional test, we estimated the Easton (2006) adaptation of O'Hanlon and Steele (2000) in our sample and obtained similar conclusions (results not tabulated). The estimate of r for low disclosure quality is 11.12%, and the estimate of r for high disclosure quality is a statistically significantly ($p < 0.0001$) lower 5.2%. While this method does not suffer from potential analyst forecast bias, it does require a choice of "actual" earnings to include in the model (presumably some estimate of permanent actual earnings). Note that, consistent with Easton and Sommers (2007), the estimates of r from the O'Hanlon and Steele (2000) model are lower than estimates from the Easton et al. (2002) model.

of management forecast policy. Further research is needed to enhance management forecast policy characterization and to discover the effects of further dimensions of policy on cost of equity capital. Focus on management earnings forecasts should be particularly fruitful due to the clear relation between what is being disclosed, earnings, and equity valuation.

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