Financial Reporting Transparency and the Cost of Equity: Evidence From Newly Listed Firms

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This paper explores the economic effects of financial reporting transparency of newly listed firms and outcomes associated with enhanced disclosure and financial reporting activities. We find a negative correlation between financial reporting transparency and information asymmetry in newly listed firms during their first five years of public trading. Further, we find a significant positive link between perceived business risk and implied cost of equity for these new public companies. Furthermore, we find a positive association between financial reporting transparency and the cost of equity. Our study helps to extend the research on the consequences of increased disclosures of newly listed firms.

Keywords: financial reporting transparency, business risk, perceived business risk, implied cost of equity

INTRODUCTION

A firm's cost of capital plays a fundamental role in the corporate decision-making process because it affects the operations of the company and the company's subsequent survival in the market (Easley and O'Hara, 2004). Given this importance, exploring how reporting quality (e.g., informativeness of earnings) impacts a firm's cost of capital is of substantial interest to policy makers, financial analysts, and researchers. The former SEC Chairman Arthur Levitt contended that "high-quality accounting standards …improve liquidity [and] reduce capital costs" (Admati and Pfleiderer, 2000). Recent studies investigate the properties of financial information of public firms related to the cost of equity capital (e.g., Easley and O'Hara, 2004; Hughes, Liu, and Liu, 2007; Christensen, De la Rosa, and Feltham, 2010; Lambert, Leuz, & Verrecchia, 2007, 2012; Fu, Kraft, & Zhang, 2012; Barth, Konchitchki, & Landsman, 2013), there has been limited academic attention on how the newly listed firms' reporting quality affects their perceived business risk and the cost of equity.

A careful analysis of the effects of newly listed companies' financial reporting information could improve our understanding of the relation between information characteristics and cost of capital for two reasons. First, theoretically, the effect of reporting quality on the cost of equity capital is unclear (Beyer,

Cohen, Lys, & Walther, 2010). Analytical papers argue that more value-relevant accounting information can reduce information risk and adverse selection and, therefore, mitigate the cost of capital (Brown, 1979; Barry and Brown, 1984, 1985; Easley and O'Hara, 2004). However, a series of recent papers have provided alternative views on the effect of accounting information on the cost of capital (Hughes et al., 2007; Lambert et al., 2007; Francis, Nanda, & Olsson, 2008; Christensen et al., 2008; Heitzman, Wasley, & Zimmerman, 2010). Hence, empirical evidence on the relation between financial reporting quality attributes, information asymmetry, and cost of capital is still inconclusive (Hughes et al., 2007; Beyer et al., 2010). Second, the research settings adopted by the previous studies are subject to limitations on the weakness of the proxies for reporting quality and the cost of equity capital, making it difficult to infer how reporting quality impacts the cost of equity capital (Beyer et al., 2010; Leuz and Wysocki, 2016). Nonetheless, questions remain about whether prior inferences regarding the association between reporting quality and the cost of capital will hold if the association is examined by employing a new financial reporting transparency measure in the post-IPO period. This paper aims to exploit the features of newly listed companies' financial information and investigate the empirical relation between reporting quality, perceived business risk, and the cost of equity capital in an initial public offering sample.

Previous studies on newly listed firms do not provide clear predictions on the relation between earnings transparency and cost of equity capital. The most fundamental purpose of initial public offerings is to raise external capital for future operations. When seeking the lowest possible equity capital cost, newly listed firms' managers are likely to have strong incentives to employ high-quality standard disclosure policies to influence market expectations favorably (Beyer et al., 2010; Hamm, Li, & Ng, 2016). Because all IPO firms must obey and follow strict financial reporting regulations and standards (e.g., Securities Exchange Act of 1933), these companies' accounting information should indicate their underlying economic value and real business activities.1 Additionally, due to the lack of financial analysts' coverage of newly listed firms, the marginal effect of public disclosures on reducing information asymmetry among investors should be greater than the public disclosure of existing firms. Thus, more public information disclosure, such as increased financial reporting transparency that reflects a firm's economic value, will lead to lower information asymmetry if such high-quality public information reduces investors' private information acquisition activities and costs. This view predicts that newly listed firms should have greater financial reporting transparency and that these firms should enjoy a lower cost of equity capital as there is less information asymmetry between the firm and the investors and among investors.

Since competitors will likely obtain newly listed companies' proprietary information from the companies' financial statements, these companies' managers will reluctantly disclose valuable and sensitive information in the financial statements (Verrecchia, 1983; Dye, 2001; Hoberg, Phillips, & Prabhala, 2014; Boone, Floros, & Johnson, 2016). Indeed, disclosing proprietary information led to the high costs of information reduction, increasing information asymmetry, adverse selection concerns, and the cost of equity capital. Moreover, Pittman and Fortin (2004) contend that newly public companies usually need to continually provide several years of high-quality financial reporting to convince the market of the credibility of their financial statements. That means newly listed companies with more transparent financial statements may still have a high cost of equity capital during the year after the companies' initial public offering (IPO). To the extent that a firm cannot produce persuasive information that meets the market's expectations, the financial statements of IPO firms could engender a higher cost of equity capital.

The ambiguousness of previous studies regarding the net effect of reporting quality on IPO firm's cost of equity capital suggests that additional research is warranted. All analysis in our study is based on a sample of 1,754 IPO firms (3,240 firm-year observations) from 2004 to 2012 that were listed on major U.S. stock exchanges. The full sample includes 33,715 firm-year observations, containing 30,475 existing public firm-year observations and 3,240 newly listed firm-year observations. Given that the existing public firms and newly listed firms are generally unequal in sample size and exhibit significant differences in firm characteristics, we use size- industry-year matching to select comparable existing public firms for our primary tests. We employ Barth et al.'s (2013) earnings transparency measure to assess financial information disclosure quality, whereas the measures of the cost of equity capital come from previous

studies on implied cost of equity capital (Gebhardt, Lee, & Swaminathan, 2001; Gordon and Gordon, 1997; Claus and Thomas, 2001; Easton &Monahan, 2005; Hou et al. 2012).

To examine our predictions, we compare newly listed firms' financial reporting transparency with existing public firms. Next, we explore the information effect of accounting reporting disclosure on perceived business risk. Finally, we investigate the transparency of newly listed firms' mandatory financial disclosure by examining the association between financial reporting transparency and the cost of equity capital. The results show that newly listed firms' financial statements have higher transparency than existing firms. Furthermore, the results document that newly listed firms' financial reporting transparency is negatively related to the firm's information asymmetry and positively related to their cost of equity capital.

There is a possibility that greater financial information transparency of newly listed firms could be driven by unobservable firm heterogeneity, such as the firm's risk-taking preference, which also could impact perceived business risk and the cost of equity. To alleviate this endogeneity issue, we execute a firm fixed effect model and a two-stage least squares estimation procedure (2SLS hereafter) as additional tests to OLS regressions. Specifically, we apply an alternative two-stage Probit-based approach and use path analysis to decompose the correlation between the financial reporting transparency and the cost of equity capital into direct and indirect (mediated) paths. Our results based on 2SLS support our previous findings. Path analysis demonstrates that financial reporting transparency is directly (positively) related to the average cost of equity, and the total mediated path is indirectly related to the average cost of equity.

The rest of this paper is organized as follows. Related research and a summary of our predictions are in Section 2. In Section 3, we describe the data and research design. In Section 4, empirical results and additional analysis are presented. Conclusions are provided in Section 5.

PRIOR RESEARCH AND HYPOTHESES DEVELOPMENT

Prior Research on Information Disclosure and the Cost of Equity Capital

Beginning with Easley and O'Hara (2004), researchers have been interested in whether the quality of information disclosure impacts the cost of equity capital. Early analytical models (Easley and O'Hara, 2004; Francis et al., 2005a, 2005b) hypothesized that high-quality information could decrease the cost of equity capital by reducing information asymmetry, a proxy for information risk. In these models, information asymmetry arises when informed investors have private information which uninformed investors do not have. Thereby, informed investors can profit by trading their private information. To compensate for such information asymmetry, uninformed investors demand higher returns on investments, which increases the cost of capital. Easley and O'Hara (2004) suggest a firm can attempt to reduce its cost of equity capital by enhancing the quality of accounting information disclosure that mitigates relative information risk in estimating future cash flows and earnings.

Several analytical studies argue that increased disclosure may not impact the cost of equity capital since information risk represented by private information is not a systematic risk and, therefore, it may be diversified in a large economy (i.e., Hughes et al.,2007). Lambert et al. (2007), recasting the returns-based CAPM model, provide empirical evidence that although accounting reporting quality can affect the cost of equity capital directly and indirectly, the information risk can be diversified away. Christensen et al. (2008) address the connection between earnings quality and the cost of capital from ex-ante and ex-post perspectives. They document that the ex-post information effect on the cost of capital could be offset by the ex-ante information effect on the cost of capital, suggesting that earnings quality does not affect the overall cost of capital covering the full-time span of the firm. Lambert et al. (2012) analyze investors' information acquisition process and show that investors' mutual assessment of firms' future cash flows is a critical determinant factor of the cost of equity capital if all market participants hold similar relevant information to make trading decisions. In their model, when informed investors acquire more public information, any incremental information gets partially communicated through price, thereby decreasing the uncertainty of other investors. Alternatively, providing more information to more investors affects the cost of equity only because the additional information increases the average level of information precision. Lambert et al.'s (2012) analysis implies that there should be an indirect link between earnings quality and the cost of equity

mediated by information asymmetry, given that the capital market is not perfectly competitive. While analytical studies employ different models to investigate the association between the quality of accounting information and the cost of equity capital, the underlying argument of these studies is largely similar in that accounting information reflects a firm's underlying economic activities in terms of future cash flows and earnings (Hughes et al., 2007; Lamber et al., 2007; Christensen et al., 2008; Heitzman et al., 2010; Bertomeu, Beyer, & Dye. 2011; Lambert et al., 2012).

Seeking to investigate the information effect of financial reporting quality, recent empirical studies mainly focus on the association between financial reporting quality and the cost of capital using corporate event-driven or cross-sectional settings (Beyer et al., 2010). Concerning corporate event-driven studies, Schrand and Verrecchia (2005) document a negative association between pre-IPO disclosures and underpricing, and Leone, Rock, and Willenborg (2007) demonstrate that IPO proceeds are associated with first-day underpricing. However, empirical evidence to date on association between disclosures and the cost of capital in the cross-sectional setting is mixed and controversial. Several lines of studies document a negative association between characteristics of accounting information, such as accruals quality (Francis et al., 2005; Ecker et al., 2006), voluntary disclosure levels (Botoson, 1997, Lang and Lundholm, 2000), and market-based measures of accounting quality (Francis et al., 2004; Francis et al. 2005; Ecker et al. 2006; Bhattacharya, Ecker, & Olsson, 2011; Barth et al., 2013) to the cost of equity capital. Whereas empirical results from other studies (Botoson and Plumlee, 2002; Aboody et al., 2005; Liu and Wysocki, 2017; Cohen, Dey, & Lys, 2008) show that the observed negative association between disclosure and the cost of equity capital is sensitive to research design choices. More specifically, based on a sample of manufacturing firms, Botosan (1997) shows that companies with a higher index of disclosure level enjoy a lower cost of equity capital. However, Botoson and Plumlee (2002) document that the disclosure quality of quarterly reports is positively associated with the cost of capital, while the same companies' annual report's disclosure quality is negatively related to the cost of capital.

The review of previous documented evidence in this section suggests some limitations. First, the documented negative association between information disclosure quality and the cost of equity capital is subject to endogeneity and selection bias, which could lead to a biased estimation of the relation between disclosure quality and the cost of capital (Leuz and Wysocki, 2016). Second, while previous studies have been largely confined to disclosure quality within mandatory reporting settings, the measures of the cost of equity capture the effects of both mandatory and voluntary disclosures (Beyer et al., 2010). Third, although previous studies document a negative association between disclosure quality and the cost of capital, they provide no direct evidence why public firms do not select complete disclosure models if full disclosures have significant benefits in terms of reducing information asymmetry and the cost of capital (Cohen, 2008). Bignon and Breton (2006) explain that more precise public information disclosed by newly listed firms could deepen the information asymmetry between investors and, in turn, increase the companies' cost of capital. Therefore, it is not straightforward to conclude that greater financial disclosure quality could decrease information asymmetry and the cost of equity capital.

Collectively, a long line of research examines the association between reporting quality, information asymmetry, and the cost of capital. Nevertheless, questions remain about whether previous findings on the relation between information disclosure quality and the cost of equity capital maintain their robustness when we change the measures, sample sets, and research settings that have not been investigated in prior studies. To address these issues, we extend existing research by investigating the association between financial reporting transparency and cost of capital within a newly listed firm's reporting environment. The objective of this paper is to investigate the empirical relation between financial reporting disclosure, perceived business risk, and the cost of equity capital during the post-IPO period. This period was missed by most of the literature.

Hypotheses Development

Through an initial public offering (IPO) a company can raise external capital, expand the scale of operations, and broaden its investors' base. To be publicly traded in the U.S. market, all IPO firms must obey and follow financial reporting regulations and rules to disclose their financial information to the

public. Because of these strict reporting standards, newly public companies are likely to push financial statements for higher financial reporting quality. However, since the newly listed firms generously do not have a previous trading history, well-developed financial disclosure reputation, and a group of financial analysts, these companies are likely to have higher information asymmetry and higher cost of capital. Seeking survival in the market after IPO, managers of newly public firms have stronger incentives to employ higher financial reporting standards and policies to efficiently influence market expectations towards the companies' economic value. In this paper, the disclosure quality is defined as the extent to which informed investors, before trading, have a relatively higher level of information compared to uninformed traders. Particularly, newly listed firms' managers have incentives to disclose financial information that reflects underlying economic activities, which investors can easily understand. Meanwhile, because available public and private information about newly listed firms in the market are limited, any public disclosure made by those companies should have a greater marginal information effect to decrease information asymmetry. Lang (1991) provides theory and evidence that the magnitude of stock price reactions to financial reporting disclosures diminishes with the age of the firm, which he attributes to the gradual revelation of firm-specific information over time and leads to the first alternate hypothesis:

H1: *Compared with similar-sized public firms within the same industry year, newly listed firms' financial reporting is likely to be more transparent than that of existing firms.*

Consistent with the disclosure hypothesis (Verrecchia, 2001), the litigation cost hypothesis (Rogers 2008), and the information precision hypothesis (Lambert et al. 2012), we posit that managers of newly listed firms are more likely to disclose high-quality information to narrow the gap between informed and uninformed investors' expectations about future earnings. Furthermore, to reduce the threat of high litigation costs, managers who desire to reduce the information asymmetry between informed and uninformed investors are more likely to disclose value-relevant information to mitigate the informed investors' information advantage over uninformed traders.

In line with expectations that high-quality information disclosure reduces information asymmetry, Fu et al. (2012) find that bid-ask spread decreases after each public disclosure, and they provide further evidence that managers issue more frequent disclosures to lower the cost of equity capital. If managers issue truthful information in communicating with potential investors, we expect that disclosed information will reduce information asymmetry and the cost of equity capital. Additionally, besides impacting investors' perceptions, the information provided as a part of a mandatory public disclosure reduces the managers' information advantage and their incentives to engage in insider trading behavior (Rogers, 2008).

However, studies on post-IPO performance indicate that some IPO firms engage in earnings management during the IPO process, potentially leading to manipulated earnings in subsequent periods (Teoh et al., 1988, 1989). Healy and Palepu (2001) argue that managerial motivations for self-serving disclosures complicate understanding their discretionary use of financial reporting. Furthermore, the corporate governance structures of newly listed firms might not be robust enough to steer managers towards optimal disclosure policies, especially when their performance is evaluated based on earnings (Fischer and Verrecchia, 2000; Healy and Palepu, 2001; Hermalin and Weishach, 2012). Considering that managers of newly listed firms are often viewed as informed investors due to their substantial shareholdings and deeper insight into the firm's growth opportunities (Certo et al., 2001), they may have strong incentives to retain private information about the firm's value. Consequently, while newly listed firms are obligated to provide both mandatory and voluntary disclosures, these may not be sufficiently informative to bridge the gap in information collection and processing abilities between informed and uninformed investors (Noe, 1999), potentially exacerbating perceived business risks.

Considering the arguments above, the highly transparent information provided by newly listed firms could alleviate or exacerbate information asymmetry among capital market participants. Therefore, our second hypothesis is framed as a null hypothesis, reflecting:

H2: *Newly listed firms' highly transparent information is not associated with perceived business risk between informed* ***and*** *uninformed investors.*

Recent theories have raised questions about the universally beneficial nature of high-quality disclosure and its consistent effect in reducing the cost of capital. Dye (1985) posits that disclosing too much public information, especially proprietary details, can inadvertently benefit competitors and negatively impact a company's value and cash flows. Almazon, Suarez, and Titman (2003) identify another potential downside: high-quality disclosure could escalate a company's human resources expenses. Bignon and Breton (2006) caution against equating increased public information with decreased private information. They argue that more precise public accounting information might amplify informed investors' information advantage. This increase in information asymmetry can then adversely affect market liquidity. Consequently, firms with greater disclosure quality or higher financial reporting transparency could face a higher cost of capital due to these dynamics. Similarly, Han (2022) found that investors often perceive IPO firms as riskier, attributing a higher implied cost of equity to them compared to their same-size matched firms. Thus, our proprietary cost hypothesis, formulated alternatively, suggests that while high-quality disclosure has its merits, it can also inadvertently elevate a firm's cost of capital, particularly in the context of newly listed companies. This hypothesis acknowledges disclosure practices' complex and sometimes contradictory effects in capital markets.

H3: *The level of transparency of newly listed film's financial disclosure will be positively related to the cost of equity.*

RESEARCH DESIGN

Variables Measurement

Financial Reporting Transparency Measure

We adopt the earnings transparency measure from Barth et al. (2013) as a proxy for financial reporting quality. Barth et al. (2013) constructed the earnings transparency based upon the sum of explanatory powers of two return-earnings relations. Specifically, they utilized adjusted R^2 from cross-sectional and intertemporal regressions based on the relation between contemporary annual stock return and earnings and earnings change amount between year t-1 and year t deflated by the price at the beginning of year. The reporting quality measure, *TRANSi,t*, is the sum of the adjusted R^2s derived from the firm i's industry (*TRANSI*) and industry-neutral (*TRANSIN*) return-earnings regressions in year *t* (Barth et al. 2013).

$$TRANSi,t = TRANSIj,t + TRANSINp,t\ ;\ (TRANSi,t = \ Adj\ R2EQ2\ +\ Adj\ R2EQ3) \quad (1)$$

TRANSIi,t is adjusted R^2 estimated from annual returns-earnings regressions by industry *j*. Since companies within the same industry use similar accounting practices, *TRANSI* will be the same for those in the same industry (Barth et al. 1999). Firms in an industry may apply different accounting practices that influence their return-earnings relations, and firm-specific earnings may indicate a different level of management information and economic value. Therefore, *TRANSIN*, an industry-neutral transparency measure, is about the annual return-earnings regressions by groups of firms. Group candidate is determined by the residuals from the prior industry regressions. *TRANSIN* compensates the weakness of *TRANSI* as it captures cross-sectional differences in the previous returns-earnings relation.

We use the following regressions to calculate *TRANSI and TRANSIN.*

$$RETi,j,t = \alpha 0I + \alpha 1IEi,j,t/Pi,j,t-1 + \alpha 2I\ \Delta Ei,j,t/Pi,j,t-1 + \varepsilon i,j,t\ [TRANSI] \quad (2)$$

$$RETi,j,t = \alpha 0IN + \alpha 1INEi,j,t/Pi,j,t-1 + \alpha 2IN\ \Delta Ei,j,t/Pi,j,t-1 + \varepsilon i,j,t\ [TRANSIN] \quad (3)$$

where *RET* is firm i's annual stock return. *Et/Pt-1* is earnings before extraordinary items and discounted operations deflated by the stock price at the beginning of the year t, and ΔE is change in earnings from year *t-1* to year *t*. We estimate the *TRANS* model over nine years (*t*=2004 to 2012) and for 15 industries, provided there were at least 10 observations for the industry year. This estimation procedure constrained the coefficients in equation (2) to be the same for firms within industry *j* in year *t*.

When estimating equation (3), we followed Barth et al.'s (2013) methodology. Specifically, we assigned the observations into four sub-samples based on the value of associated residuals from each industry-year regression for specific industries. We ranked the residuals from the smallest to the largest and place them in groups across the sample period and all 15 industries. In group 1, we placed the largest first quarter residuals, and then we put the observations with second largest 25% residuals in group 2, and so on, until each observation was placed to a group. As such, group 1 contained the observations with the largest negative residuals. In contrast, group 4 included the samples with the largest positive residuals from each of the 15 industries in each year, and so on.

Perceived Business Risk Measures

Our information asymmetry measure is based on bid-ask spread, a common measure of information asymmetry. Following Mohd (2005) and Silber (2005), we first obtained a daily bid-ask spread (*ASK-Bid)/((Ask+Bid)/*2) and then regressed the raw spread on the daily absolute return for each firm year. The estimated that the intercept term from this regression serves as a proxy for information asymmetry.

Cost of Equity Capital Measures

Previous studies almost exclusively employ two categories of methods to estimate the cost of equity. One is based on ex-post stock returns (Fama and French, 1993, Fama and French, 1997; Elton, 1999; Vuolteenaho, 2002), and the other is based on analysts' forecasts to estimate the implied cost of capital (Gordon and Gordon, 1997; Claus and Thomas, 2001; Gebhardt, Lee, & Swaminathan, 2001; Easton, 2004; Ohlson and Juettner-Nauroth, 2005). However, many papers point out that ex-post realized return is a biased estimator of expected return (Blume and Friend, 1973; Froot and Frankel, 1989; Hou, Van Dijk, & Zhang, 2012). Also, the analysts' forecast approach heavily relies on their coverage and forecasts. Given the lack of analysts' forecast data on newly listed firms during the sample period, we followed Hou et al. (2012) to estimate the expected return of equity capital. We used earnings forecasts generated by a cross-sectional model instead of analysts' forecasts to proxy cash flow expectations. We estimated model-based earnings forecasts for up to five years into the future. Then we used those earnings forecasts to compute the expected return of equity capital for newly listed firms. Based on previous studies (Gebhardt, Lee, & Swaminathan, 2001; Gordon and Gordon, 1997; Ohlson and Juettner- Nauroth, 2005), we computed three individual costs of equity capital estimates and a composite cost of equity capital estimate (the average of three individual costs of equity capital estimates).

Specifically, we estimated the following pooled cross-sectional regression for each year between 2004 and 2012 by using the available previous years' data:

$$Ej,t+\gamma = \beta 0 + \beta 1 EVj,t + \beta 2 TAj,t + \beta 3 DIVj,t + \beta 4 DDj,t + \beta 5 Ej,t + \beta 6 NEGEj,t + \beta 7 ACCj,t + \varepsilon j,t + \gamma \quad (4)$$

where *Ej,t+ɤ* (ɤ=1,2, or 3) denotes the earnings before extraordinary items of each firm in *ɤ* year after *t*; *EVj,t* measures the enterprise value of each firm, calculated as total assets plus the market value of equity minus the book value of equity at the end of year *t*; *TAj,t* denotes the total assets, *DIVj*,t is the dividend payment in year t; *DDj,t* is a dummy variable equals to 1 for non-dividend payers and zero for dividend payers; *NEGEj,t* is an indicator variable equals to 1 if firms' earnings are negative and 0 for positive earnings companies; and *ACCj,t* denotes total accruals, calculated as the change in current assets minus the change in current liabilities plus the difference between the change in debt in current liabilities and the change in cash and short-term investments) scaled by total assets. We also control year and industry-fixed effects.

The estimated annual coefficients from equation (4) are listed in Appendix III. The average estimated parameters are qualitatively similar to those coefficients documented in Hou et al. (2012). Follow previous papers, we estimated the cost of equity capital at the end of June of each year. First, we reconcile the market price with the present value of forecasted future earnings at the end of June to determine the discount rate needed in the later estimation models. Second, we set values of the cost of equity capital above 100% as missing. Third, we assumed that the expected *ROE* mean reverts to the historical industry median value. Following Gebhardt et al. (2001), we excluded loss firms when calculating the industry median *ROE*. We used the dividend payout ratio to calculate the dividend for each year for companies with positive earnings or using current dividends divided by 0.06 times total assets as an estimator of the payout ratio for firms with negative earnings. Please see Appendix IV for the estimated cost of equity capital models and assumptions.

Models

To test our hypothesis H2 on how financial disclosure quality of newly listed firm is related to information asymmetry between informed and uninformed investors, we used Fu et al. (2012) model:

$$IAj,t+1 = \alpha + \beta 1\, TRANSj,t + \beta 2\, Sizej,t + \beta 3\, log(Turnover)j,t + \beta 4\, log(Volatility)j,t + \varepsilon t \quad (5)$$

where *IA* measures information asymmetry for company j in fiscal year *t+1*. *TRANS* refers to financial reporting transparency for the company *j* in fiscal year *t*. We included other control variables in the model. *Size* is the natural log of average market equity value at the beginning and end of the prior calendar year. *Log(turnover)* is the log of the median daily turnover ratio in a year (i.e., value of all shares traded divided by the capitalization). *Log(Volatility*) is the log of the standard deviation of daily return in a year.

To examine our hypothesis H3 about the association of earnings transparency and the cost of equity capital, we adopted the model for regression analyses from Barth et al. (2013):

$$COEj,t+1 = \alpha + \beta 1\, TRANSj,t + \beta 2\, Sizej,t + \beta\, Betaj,t + \beta\, GRj,t + \beta\, Levj,t + \beta\, MBj,t + \beta\, Lossj,t + \varepsilon j,t \quad (6)$$

where *COE* is a composite implied cost of equity capital measured as the average of the following three individual COE estimates: Gebhardt et al. (GLS, 2001), Gordon and Gordon (Gordon, 1997), and Ohlson and Juettner-Nauroth (Oj, 2005). *Size* is the natural log of average market equity value at the beginning and end of the prior calendar year. *Beta* is calculated by regressing each firm's daily return on the market daily return in the current year. *GR*, the growth rate, is calculated as the natural log of one plus the percentage change in book value of equity (BE). *MB* is the ratio of the market value of equity (CRSP: Abs (PRC)/SHROUT) to the book value of equity of fiscal year-end (Compustat CEQ*1000). *Leverage* is the ratio of total debt to total assets. *Loss* is an indicator variable equal to 1 if the company reports negative income for year *t*, 0 if it has positive income.

Empirical Results

Sample Selection

We follow Field and Karpoff (2002) and Ritter (2014) to collect IPO firms' information from Thomson Reuters' Securities Data Corporation (SDC) New Issues database. The sample contains 1,455 IPOs of common stock in the U.S. from 2004-2012. The list of variables contains offer dates, firm identified, and firm founding dates. Following Barry and Mihov (2015), we excluded American Depository Receipts (ADRs), closed-end funds, real estate investment trusts (REITs), unit offerings, and IPOs with an offer price below $ 5.00 per share. Additionally, we eliminated financial institutions (SIC codes 6000-6999) and reverse leveraged buyout (LBO) IPOs from the sample because reverse LBO firms have a previous reputation and trading history before going public the second time. After excluding financial firms and reverse LBOs, the sample consisted of 970 IPOs. We also obtained firm-level financial information from

the Compustat North America Fundamentals Annual File and collected stock return and daily price data from CRSP monthly stock file. Table 1 presents the details of the sample used in this study. Panel A shows the sample distribution of firms by year, Panel B shows firm-year observations for the entire sample, and Panel C gives the firm-year observations for IPO firms.

TABLE 1
SAMPLE CONSTRUCTION

Panel A: Distribution of firms

Year	Number of firms	Number of IPO firms
2004	6479	194
2005	6567	202
2006	7097	217
2007	6867	251
2008	6492	34
2009	6327	60
2010	6284	219
2011	6255	129
2012	5811	149
Total	10864	1455

Panel B: Sample Construction (Newly listed and existing firms)

Firm-year observations available in North American Fundamentals Annual for the year 2004-2012	83153
Less	
Firms not listed on major U.S. Stock Exchanges (Only Exchange code=11, 12, or 14)	21991
Firms in the financial and utility industries (SIC=4900-4999 and 6000-6999)	24712
Firm-years without daily security information due to delisted within one year	450
Firms with missing financial variables.	5525
Firm-years available for regression analysis	30475

Panel C: Sample Construction for newly listed firms

Firms-years observation available in North America Fundamentals Annual for the year 2004-2012	1455
Less	
Firms not listed on major U.S. Stock Exchanges (Only Exchange code=11, 12, or 14)	254
Firms in the financial and utility industries (SIC=4900-4999 and 6000-6999)	41
Firm-years without daily security information due to delist	44
Firms with missing financial variables.	109
Firms' offer price below $ 5.00 per share	36
ADRs REITs, unit offerings, and reverse LBO	80
Newly listed firms available for regression analysis	891
Newly listed firm-years available for regression analysis	3240

Due to the presence of outliers, we winorize *EBITDA* to *Total assets*, *EBITDA* to *S*ales, and *Market-to-book* ratio. Dollar levels throughout the paper are adjusted for inflation using the annual consumer price

index (CPI) obtained from the Federal Reserve Research Database (FRED). They are presented in January 2013 constant dollars.

Univariate Tests Firms

Table 2 presents the summary statistics for the variables used in the samples of all and newly listed firms. Panel A of table 2 presents the distribution statistics of variables in the estimating models. *TRANS_all* is consistent with the empirical evidence demonstrated in Barth et al. (2013), it averages 27.7% and ranges from 0.1% to 112.5%. For newly listed companies, the measure of financial reporting transparency, *TRANS-newly listed firms*, has the mean of 45% and ranges from 4.9% to 120.5%, and the average implied cost of equity capital, *Avg_COE-new listed firms*, has the mean and median of 6.5% and 9.2%.

The yearly statistics in Panel B of Table 2 indicate that there is no monotonous change in *TRANS, TRANSI, and TRANSIN*, but that the average magnitude of TRANS within pre-recession period (2004-2007) is higher than the mean magnitude of *TRANS* during the post-recession period (2009-2012).

Panel C indicates that the industry component of *TRANS, TRANSI* is, on average, substantially greater than the industry-neutral component, *TRANSIN*.

Panel D of Table 2 compares the financial information transparency variables' mean between newly listed and existing public firms. Column (1) shows the descriptive statistics of all public firms in the U.S. market. Column (2) shows the equivalent for the size-year-industry-matched existing firms. Column (3) illustrates the descriptive statistics for newly listed firms. There are 30,475 firm-year observations in the existing public firms' sample and 3,240 in the newly listed firms' sample. First, the mean of *TRANS* is significantly greater in newly listed firms with 0.4496 than in existing public firms with a mean *TRANS* of 0.2381. Similarly, mean *TRANSI* and mean *TRANSIN* tend to be greater in the newly listed firms than matched public firms. The results support H1 that newly listed firms' financial information is more informative than the existing firms with similar sizes in the same industry and same year.

Table 3 presents the Pearson correlations among variables used in the model. The correlation between *TRANS* and *IA_Spread* is negative and significant (r=-0.119, p-value<0.001), indicating that there is a significantly negative association between the transparency of disclosed earnings and the information asymmetry between informed and uninformed investors. This association is consistent with the previous research that the precision of public information could reduce information risk (Luez and Verrcchia, 2004). Also, the correlation between *TRANS* and the average measure of the cost of capital, *Avg_COE*, is positive and significant (r=0.0707, p- value<0.001), suggesting that there is a significant association between the financial reporting quality and the cost of equity capital. Moreover, the three implied costs of equity capital are generally highly correlated with each other and the composite implied cost of equity capital. The correlation ranges from a low of 0.194 (between *GLS* and *GORDON*) to a high of 0.89 (between Gordon and OJ). Consistent with previous literature, *Avg_COE* correlates positively with the log of book-to-market (*MB*), Leverage (*LEVERAGE*), growth rate (*GR*), and Beta (*BETA*). *TRANS* also correlates positively with *LOSS*, indicating that loss firms tend to have lower reporting quality. We do not expect our sample to suffer from multicollinearity as all correlations between main variables in the estimated regressions are lower than 0.5.

TABLE 2
DESCRIPTIVE STATISTICS

Panel A: Based on observations pooled across years and industries

Main variables	**Mean**	**Median**	**Std**	**Min**	**Max**
TRANS-all firms	0.277	0.259	0.210	0.001	1.125
TRANSI-all firms	0.039	0.025	0.046	0.000	0.743
TRANSIN-all firms	0.238	0.226	0.193	0.001	0.662
TRANS-newly listed firms	0.450	0.425	0.243	0.049	1.205
TRANSI-newly listed firms	0.150	0.082	0.164	0.001	0.831
TRANSIN-newly listed firms	0.299	0.305	0.160	0.005	0.664
RET-all firms	0.212	0.055	1.024	-0.990	40.151
Et/Pt-1-all firms	0.919	0.153	5.428	-9.652	16.669
ΔEt/Pt-1-all firms	1.541	0.729	8.164	-14.569	25.105
RET-new listed firms	0.124	-0.028	0.961	-0.884	3.376
Et/Pt-1-new listed firms	-0.792	0.538	6.800	-19.168	11.890
ΔEt/Pt-1-new listed firms	0.201	0.117	5.236	-12.063	14.553
IA_Spread	0.049	0.043	0.028	0.004	0.149
COE_GLS-newly listed firms	0.057	0.053	0.036	0.012	0.166
COE_OJ-newly listed firms	0.016	0.024	0.045	-0.127	0.099
COE_Gordon-newly listed firms	0.007	0.021	0.055	-0.204	0.088
Avg_COE-newly listed firms	0.065	0.092	0.037	-0.012	0.109
Other variables	**Mean**	**Median**	**Std**	**Min**	**Max**
log(Turnover)	1.641	1.758	0.906	-0.322	3.105
log(Volatility)	-3.678	-3.729	0.433	-4.391	-2.788
MVE	6.452	6.506	1.461	3.104	9.581
Size	6.632	6.656	1.359	3.345	9.582
MB	1.192	1.122	0.711	0.041	3.394
Leverage	0.175	0.056	0.209	0.000	0.704
Growth rate	-0.054	-0.050	0.624	-2.561	3.614
Loss	0.295	0.000	0.456	0.000	1.000

Panel B: Across-Industry mean and standard deviation within each fiscal year

	TRANS		**TRANSI**		**TRANSIN**		**IA Spread**		**Avg COE**		
Year	Mean	Std	Mean	Std	Mean	Std	Mean	Std	Mean	Std	No. Obs
2004	0.788	0.367	0.343	0.264	0.432	0.231	0.038	0.012	0.037	0.018	32
2005	0.329	0.135	0.127	0.095	0.202	0.105	0.041	0.011	0.039	0.030	148
2006	0.514	0.195	0.187	0.185	0.327	0.057	0.037	0.009	0.039	0.019	256
2007	0.665	0.204	0.211	0.162	0.456	0.121	0.039	0.009	0.039	0.021	371
2008	0.338	0.264	0.135	0.153	0.203	0.189	0.065	0.016	0.046	0.029	461
2009	0.414	0.163	0.155	0.153	0.257	0.049	0.058	0.016	0.039	0.021	456
2010	0.367	0.171	0.106	0.141	0.261	0.092	0.041	0.011	0.041	0.021	466
2011	0.391	0.264	0.124	0.171	0.268	0.191	0.046	0.013	0.044	0.025	509
2012	0.534	0.255	0.158	0.166	0.378	0.151	0.039	0.151	0.044	0.039	541
Total											3240

Panel C: Across-year mean and standard deviation within each industry

	TRANS		**TRANSI**		**TRANSIN**		**IA Spread**		**Avg_COE**		
Industry	Mean	Std	Mean	Std	Mean	Std	Mean	Std	Mean	Std	Obs
Food	0.643	0.287	0.389	0.211	0.267	0.182	0.048	0.021	0.054	0.039	29
Mining and Minerals	0.914	0.294	0.572	0.272	0.327	0.147	0.043	0.016	0.046	0.019	35
Oil and Petro Products	0.412	0.167	0.114	0.083	0.299	0.154	0.051	0.016	0.041	0.021	280
Apparel and Footwear	0.735	0.372	0.411	0.312	0.312	0.129	0.043	0.015	0.059	0.033	48
Durable manufactures	0.689	0.397	0.358	0.332	0.327	0.118	0.049	0.021	0.059	0.031	25
Chemicals	0.749	0.317	0.443	0.208	0.311	0.166	0.044	0.018	0.038	0.021	68
Drugs, Soap, Tobacco	0.435	0.213	0.145	0.171	0.294	0.161	0.053	0.016	0.037	0.029	127
Construction	0.698	0.283	0.412	0.181	0.285	0.159	0.047	0.017	0.041	0.021	59
Steel	0.741	0.348	0.436	0.319	0.333	0.173	0.055	0.015	0.053	0.030	26
Fabricated Products	0.389	0.111	0.075	0.046	0.314	0.110	0.052	0.013	0.046	0.017	18
Machinery & Business Equipment	0.471	0.191	0.161	0.101	0.311	0.158	0.049	0.015	0.042	0.023	453
Automobiles	0.723	0.325	0.402	0.231	0.328	0.163	0.037	0.013	0.043	0.015	36
Transportation	0.578	0.273	0.275	0.167	0.302	0.164	0.045	0.017	0.045	0.026	301
Retail Stores	0.492	0.021	0.181	0.117	0.311	0.159	0.045	0.016	0.044	0.022	212
Services	0.359	0.181	0.066	0.042	0.291	0.161	0.044	0.016	0.041	0.030	1523
Total											3240

Panel D: Descriptive statistics: new listing firms versus existing firms

	(1) All Public Firms		(2) Matched Public Firms		(3) New Listing Firms		Difference	*p-value*	Difference	*p-value*
Variables	Mean	Std	Mean	Std	Mean	Std	(1)-(3)		(2)-(3)	
TRANS	0.238	0.194	0.267	0.211	0.449	0.243	-0.211	<0.001***	-0.182	<0.001***
TRANSI	0.277	0.209	0.036	0.040	0.151	0.163	0.127	<0.001***	-0.115	<0.001***
TRANSIN	0.039	0.046	0.232	0.196	0.299	0.159	-0.260	<0.001***	-0.067	<0.001***
N	30,475		3,240		3,240					

Sample of newly listed firms and existing public firms: 2004-2012. All variables' definitions are provided in Appendix 1.

TABLE 3
PEARSON CORRELATION

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
(1) TRANS	1	0.761***	0.745***	-0.118***	0.044**	0.056***	0.065***	0.071***
(2) TRANSI		1	0.138***	-0.009	0.173	0.064***	0.072***	0.068***
(3) TRANSIN			1	-0.193**	0.051***	0.019	0.024	0.037**
(4) IA				1	0.067***	-0.211**	-0.034**	-0.241***
(5) COE_CLS					1	0.266***	0.191***	0.544***
(6) COE_OJ						1	0.809***	0.906***
(7) COE_GORDEN							1	0.898***
(8) AVG_COE								1

	(9)	(10)	(11)	(12)	(13)	(14)	(15)	(16)
(1) TRANS	0.092***	0.113***	0.093***	0.032*	0.163	0.076***	0.096***	-0.112***
(2) TRANSI	0.089***	0.152***	0.105***	0.053***	0.039***	-0.094**	0.098***	-0.137
(3) TRANSIN	0.046***	0.018	0.034*	-0.001	-0.017	-0.018	0.043***	-0.015***
(4) IA	-0.402***	0.067***	-0.061**	-0.054**	-0.165***	0.362***	-0.069**	0.910***
(5) COE_CLS	-0.297**	-0.176***	-0.001	0.038**	-0.508***	-0.126**	-0.025	0.089***
(6) COE_OJ	0.114***	0.081***	0.046***	0.131***	-0.033*	-0.655**	0.065***	-0.195***
(7) COE_GORDEN	0.323***	0.153***	0.064***	0.159***	-0.161**	-0.763**	0.184***	-0.341***
(8) AVG_COE	0.116***	0.052***	0.043***	0.146***	-0.081**	-0.693**	0.114***	-0.231***
(9) SIZE	1	0.418***	0.163***	0.102***	0.443***	-0.368**	0.513***	-0.491***
(10) LEVERAGE		1	0.167***	0.091***	0.281***	-0.148**	0.028	0.065***
(11) Beta			1	0.023	0.025	-0.067**	0.062***	-0.071***
(12) Growth rate				1	0.238	-0.461**	0.033	-0.053***
(13) MB					1	0.159***	0.165***	-0.175***
(14) LOSS						1	-0.216***	0.375***
(15) Turnover							1	-0.219***
(16) Volatility								1

Multivariate Analyses

To examine whether newly listed firms' financial reporting provides sufficient information content to reduce information asymmetry, we ran an equation (5) to estimate the association between *TRANS* and *IA_spread*. In Table 4, we present the regression model results using the IPO sample. There are 2,401 firm-year observations in the sample. Year and industry dummies are included in fixed effect models, and the t-statistics are based on Newey-West (1987) standard errors, which adjust for both heteroskedasticity and autocorrelation. The "OLS" column shows OLS regression results with clustered standard errors; the "Fixed effect" columns show results from the year and industry fixed effects models. In all sets of results, we find that the coefficient on each control variable (size, turnover, and volatility) is significant, implying that they help explain a firm's bid-ask spread. Moreover, the coefficient on *TRANS* is negative and significant in both OLS and industry fixed-effect models, indicating that as transparency of financial reporting increases, a firm's bid-ask spread decreases, which does not support our hypothesis H2. Its value ranges from -0.003 to -0.004, suggesting that the bid-ask spread decreases between 0.03% and 0.04% when financial reporting transparency increases by one unit.

Overall, the results in Table 4 indicate that higher financial reporting transparency is associated with lower bid-ask spread. This finding is consistent with the notion that greater informativeness of financial reporting increases the amount of public information provided to investors, and this improved information context results in a lower level of information asymmetry.

TABLE 4
REGRESSION RESULTS ON REPORTING TRANSPARENCY AND PERCEIVED BUSINESS RISK FOR NEWLY LISTED FIRMS

$$IAj,t+1 = \alpha + \beta 1\, TRANSj,t + \beta 2\, Sizej,t + \beta 3\, log(Turnover)j,t + \beta 4\, log(Volatility)j,t + \varepsilon t$$

Variable	OLS	Fixed effect	Fixed effect	Fixed effect
Trans	-0.004***	0.001	-0.003***	-0.001
	(-5.05)	(1.08)	(-5.98)	(-1.49)
Size	-0.001*	-0.001***	-0.001***	`-0.001***
	(-1.69)	(-4.71)	(-3.23)	(-6.44)
TURNOVER	0.002***	0.003***	0.002***	0.002***
	(12.96)	(16.21)	(14.17)	(15.11)
Volatility	0.036***	0.032***	0.034***	0.031***
	(94.12)	(79.07)	(99.46)	(78.53)
Intercept	0.181***	0.162***	0.172***	0.161***
	(136.59)	(114.81)	(154)	(114.26)
Obs	2401	2401	2401	2401
Fixed effects	None	Year	Industry	Year+industry
Adj R^2	83.49%	86.49%	85.49%	87.31%

This table presents the relation between financial reporting transparency and information asymmetry using OLS and Fixed effects models to test hypothesis 2. Definitions of the variables are provided in Appendix 1. Year and Industry dummies are included in the specific regressions, but their coefficients are not tabulated. The standard errors are clustered by firm and by year. Significant levels are based on two-tailed tests, ***. **, and * denote significance at 1%, 5%, and 10% levels, respectively.

To test whether and how the informativeness of newly listed firms' financial reporting affects the firms' cost of equity capital during their early public trading stage, we examine the relation between financial reporting transparency and the cost of equity capital. On the one hand, the litigation risk hypothesis predicts

that newly listed firms' managers would enhance the informativeness of financial reporting to reduce the cost of equity capital to avoid potential litigation costs. On the other hand, the proprietary information hypothesis expects the greater transparent financial reports issued by newly listed firms will benefit their competitors and reduce the stock liquidity and, in turn, increase the cost of equity capital. Specifically, we estimate model (6) using the newly listed firms sample. If the litigation cost hypothesis is supported, we will find a significant negative coefficient on β1. We would observe a significant positive coefficient on β1 if the proprietary information hypothesis is valid.

Table 5 presents the estimation results. In Column (1), *TRANS* (0.004, t-stat=2.24) exhibits a significant positive relation with the cost of equity capital, *Avg_COE*. We find similar results in Column (2) (0.007, t-stat=3.46) where the model is regressed using year fixed effects. Also, Column (3) and (4) show the same results. These results support the proprietary information hypothesis and shed light on the effect of financial reporting informativeness on the cost of equity capital within different information contexts.

TABLE 5
FINANCIAL REPORTING TRANSPARENCY AND THE COST OF EQUITY CAPITAL FOR NEWLY LISTED FIRMS

Dependent variable: Avg_COE

$$COEj,t+1 = \alpha + \beta 1\, TRANSj,t + \beta 2\, Sizej,t + \beta\, Betaj,t + \beta\, GRj,t + \beta\, Levj,t + \beta\, MBj,t + \beta\, Lossj,t + \varepsilon j,t$$

Variable	OLS	Fixed effect	Fixed effect	Fixed effect
Trans	0.004**	0.007***	0.005**	0.006**
	(2.28)	(3.46)	(2.46)	(2.36)
Size	-0.002***	-0.002***	-0.002***	-0.002***
	(-5.45)	(-5.42)	(-5.68)	(-5.07)
Leverage	0.017***	0.015***	0.016***	0.014***
	(7.00)	(6.46)	(6.92)	(5.27)
Beta	0.004***	0.005***	0.003***	0.002***
	(2.61)	(3.87)	(2.98)	(4.06)
MB	-0.01***	-0.009***	-0.012***	-0.009***
	(-13.45)	(-12.6)	(-13.13)	(-12.38)
LOSS	0.188***	0.188***	0.188***	-0.187***
	(57.30)	(57.33)	(57.35)	(56.76)
GR	0.004***	0.003***	0.009***	0.004***
	(5.32)	(5.24)	(5.53)	(5.08)
Intercept	0.047***	0.046***	-0.924**	0.047***
	(18.43)	(16.06)	(-2.17)	(15.57)
Fixed effects	None	Year	Industry	Year+Industry
Adj R2	53.35%	53.59%	41.96%	54.09%
N	3240	3240	3240	3240

This table presents the relation between financial reporting transparency and the cost of equity capital by using OLS and Fixed effects models to test hypothesis 3. Definitions of the variables are provided in Appendix 1. Year and Industry dummies are included in the specific regressions, but their coefficients are not tabulated. The standard errors are clustered by firm and by year. Significant levels are based on two-tailed tests, ***. **, and * denote significance at 1%, 5%, and 10% levels, respectively.

Additional Analysis

Alternative Measures of the Cost of Equity Capital

In this subsection, we investigate whether financial reporting transparency is still positively associated with the individual implied cost of equity capital estimates. Because the measure of the cost of equity capital used in this study is a mathematical average of three individual costs of equity estimates, it may include more measurement errors. Therefore, we performed additional tests to ensure that the main findings were robust to alternative measures of the cost of equity capital.

Table 6 reports results from the regression in which the dependent variable is the implied cost of equity capital from the residual income model (Gebhardt et al. 2001). Column 1 shows the OLS regression results. *TRANS* (0.006, *t-stat*=3.42) illustrates a significant positive relation with the cost of equity capital. The results of the year fixed effect model and industry fixed effect model are listed in columns 2 and 3. The estimated coefficients of *TRANS* in both models are significantly positive, namely 0.008 (t-stat=4.48) and 0.005 (t-stat=2.23). Column 4 shows the results of firm-level fixed effect. The estimated coefficient of *TRANS* in this model is still significantly positive (0.005, *t-stat*=2.80). These results provide consistent evidence of the positive association between financial reporting transparency and the cost of equity.

TABLE 6
ADDITIONAL ANALYSIS: FINANCIAL REPORTING TRANSPARENCY AND INDIVIDUAL COE FOR NEWLY LISTED FIRMS

DV=COE_GLS

Variable	OLS	Fixed effect	Fixed effect	Fixed effect
Trans	0.006***	0.008***	0.005***	0.005***
	(3.42)	(4.48)	(2.23)	(2.80)
Size	-0.004***	-0.006***	-0.003***	-0.002***
	(-15.85)	(-16.19)	(-6.74)	(-2.46)
Leverage	0.003	0.002	0.006***	0.234***
	(1.47)	(0.98)	(2.33)	(4.70)
Beta	0.001	0.001	0.001***	-0.001
	(0.04)	(0.13)	(2.31)	-0.55
MB	-0.009***	-0.001***	0.028***	0.028***
	(35.1)	(-15.45)	(37.13)	(27.1)
LOSS	0.039***	0.04***	0.025***	0.022***
	(39.15)	(39.95)	(18.93)	(16.39)
GR	-0.008***	-0.008***	-0.004***	-0.003***
	(-12.12)	(-12.28)	(-5.14)	(-4.97)
Intercept	0.113***	0.101***	0.105***	0.108***
	(24.96)	(31.3)	(27.92)	(16)
Fixed effects	None	Year	Industry	Firm-level
Adj R2	40.63%	38.7%	39.82%	38.12%
N	3240	3240	3240	3240

This table presents the relation between financial reporting transparency and the cost of equity capital by using OLS and Fixed effects models. Definitions of the variables are provided in Appendix 1. Year and Industry dummies are included in the specific regressions, but their coefficients are not tabulated. The standard errors are clustered by firm and by year. Significant levels are based on two-tailed tests, ***. **, and * denote significance at 1%, 5%, and 10% levels, respectively.

We also tested the documented association by using four additional measures of implied cost of equity capital from previous studies (Easton, 2004; Claus and Thomas, 2001; Ohlson and Juettner-Nauroth, 2005; Gordon and Gordon, 1997). Table 7 shows the empirical results from the year fixed effect regressions. In all regressions (except for *COE_Gordon*), the coefficient on financial reporting transparency, *TRANS*, is positive and significant, confirming the finding that newly listed firms financial reporting transparency is positively related to the firms' cost of equity capital within the first five years of public trading.

TABLE 7
SENSITIVE TEST: FINANCIAL REPORTING TRANSPARENCY AND INDIVIDUAL COE FROM OTHER ICC MODELS

Dependent variables = COE_OJ, COE_CT, COE_MPEG, COE_Gordon

	COE_OJ	COE_CT	COE_MPEG	COE_Gordon
Variable				
Trans	0.006***	0.009*	0.011***	0.004
	(2.36)	(1.75)	(4.33)	(1.39)
Size	-0.004***	-0.009***	-0.009***	0.003***
	(-7.01)	(-8.48)	(-16.95)	(4.77)
Leverage	0.021	0.009	0.201***	0.023***
	(0.11)	(1.63)	(6.42)	(7.10)
Beta	0.02***	0.017**	0.003***	0.001***
	(3.04)	(2.13)	(2.31)	(2.55)
MB	-0.01	-0.019***	-0.003***	-0.002**
	(-1.46)	(-11.96)	(-2.78)	(-2.14)
LOSS	0.204***	0.075***	0.09***	0.283***
	(44.7)	10.06)	(18.93)	(61.48)
GR	0.004***	0.002	0.005***	0.005***
	(4.21)	(1.3)	22.01)	(5.24)
Intercept	0.037***	0.133***	0.143***	0.108***
	(9.13)	(19.90)	(37.78)	(16)
Fixed effects	Year	Year	Year	Year
Adj R2	40.34%	12.97%	35.17%	60.11%
N	3240	3240	3240	3240

This table presents the relation between financial reporting transparency and the cost of equity capital by using Fixed effects models. Definitions of the variables are provided in Appendix 1. Year dummies are included in the specific regressions, but their coefficients are not tabulated. The standard errors are clustered by firm and by year. Significant levels are based on two-tailed tests,
***. **, and * denote significance at 1%, 5%, and 10% levels, respectively. Where COE_OJ is an estimated cost of capital measure derived from Ohlson and Juettner-Nauroth's (2005) model. COE_CT is an estimated cost of capital measure derived from Claus and Thomas's (2001) model. COE_MPEG is an estimated cost of capital measure derived from Easton's (2004) modified model. COE_GORDON is an estimated cost of capital measure derived from Gordon and Gordon's (1997) model.

2SLS Model

Previous theoretical and empirical evidence suggests that disclosure quality is endogenously associated with the degree of information asymmetry (Marquardt and Wideman 1998, Leuz and Verrecchia 2000, Cohen, 2003, Brown and Hillegeist, 2007). Brown and Hillegeist (2007) argue that better disclosure quality

more likely results in less information asymmetry, whereas firms with high information asymmetry will have stronger incentives to choose higher disclosure quality for reducing asymmetry levels. The common way to mitigate the effect of endogeneity on coefficient estimates is to apply three-stage least squares (3SLS). However, it is difficult to find two relevant exogenous variables that are unrelated to dependent variables. Therefore, we followed Brown and Hillegeist (2007) to apply an alternative two-stage Probit-based approach (Wooldridge, 2006).

$$Prob(TRANS > Industry - Year\ Median) = \phi(Size, Return, Surprise, Capital\ , InstOwn, Analysts, Owners, EarnVol) \quad (7)$$

$$IAt = \alpha + \beta 1 PrTranst - 1 + \beta 2 Sizet - 1 + \beta 3 log(turnover)j, t - 1 + \beta 4 log(Volatility)j, t - 1 + \beta 5 Dispersionj, t - 1 + \beta 6 Leveragej, t - 1 + \varepsilon t \quad (8)$$

In the first stage, we used a Probit estimation of disclosure quality where the dependent variable equals 1 if the firm's financial reporting transparency is above the median score for the industry year and equals zero otherwise. The independent variables are exogenous variables that affect either financial reporting transparency or information asymmetry. *Log(turnover)* is the median daily turnover ratio in a year (i.e., the value of all shares traded divided by the capitalization). *Log(Volatility)* is the standard deviation of daily return in a year. In the second stage, the fitted probabilities that the firm's disclosure transparency score is greater than the median industry-year score based on the estimated coefficients from model 7, *PrTrans,* are included as an instrumental variable in the information asymmetry model. In the second stage, estimates of the information asymmetry model, we obtained consistent and asymptotically efficient coefficients, estimated using OLS. Untabulated findings revealed that the negative relation between *TRANS* and information asymmetry were unchanged. This finding indicates that the negative association between disclosure transparency and information asymmetry level remains the same after controlling for the endogenous relation between the two variables.

Path Analysis

Lambert et al. (2012) argue that when financial reports can provide precise public information to more investors, the precise information can directly affect the cost of equity capital. Lambert et al. (2012) also suggest an indirect link from disclosed information to the cost of equity mediated by perceived business risk. Meanwhile, previous studies (e.g., Easley and O'Hara (2004), Hughes et al. (2007), and Francis et al. (2004)) document that financial disclosure could be a source that affects information asymmetry. Therefore, we posit a direct path from financial reporting transparency to the cost of equity capital and predict an indirect effect of financial disclosure on the cost of equity capital, mediated by perceived business risk.

Following Bhattacharya et al. (2012), we used path analysis to decompose the correlation between the financial reporting transparency and the cost of equity capital into direct and indirect (mediated) paths. This decomposition provided evidence on the existence and relative importance of the direct and indirect paths between reporting quality and the cost of equity. Path analysis belongs to a class of structural equation models that provide plausible explanations of correlation structures by decomposing a correlation between two variables into a simple or direct path and a compound or indirect path that indicates a mediating variable (Bhattacharya et al. 2012). The primary path analysis we used was a LISREL-type model.

Table 8 provides evidence of both a direct path and an indirect path, with the direct path having greater importance than the indirect path. Specifically, the results indicate that the total correlation for Avg_COE, for COE_GLS, COE_OJ, and COE_Gordon is significantly associated with the transparency measure (significant at the 0.001 level or better). The direct and mediated paths decompose this correlation into the portion attributed to the direct link between financial reporting transparency and the cost of equity capital and the indirect link, mediated by perceived business risk. The *p* [*TRANS, COE*] is the direct path coefficient; the ratio of this path coefficient to the total coefficient is the portion of the correlation between financial reporting transparency and the cost of equity capital attributable to the direct path. The *p* [*TRANS,*

IA_SPREAD] and *p* [*IA_SPREAD, COE*] are the path coefficients between financial reporting transparency and perceived business risk and between perceived business risk and the cost of equity, respectively.

TABLE 8
DIRECT AND MEDIATED COST OF EQUITY EFFECTS OF FINANCIAL REPORTING TRANSPARENCY

	Avg_COE		***COE_GLS***		***COE_OJ***		***COE_GORDON***	
	Coefficient	t-statistic	Coefficient	t-statistic	Coefficient	t-statistic	Coefficient	t-statistic
r [*TRANS, COE*]	0.037*	1.83	0.068***	3.34	0.072***	3.55	0.079***	3.89
Direct Path *p* [*TRANS, COE*]	0.059***	2.94	0.077***	3.77	0.048**	2.39	0.039**	2.02
Mediated Path *p* [*TRANS, IA*]	-0.118***	-5.89	-0.118***	-5.89	-0.118***	-5.89	-0.118***	-5.89
p [*IA, COE*]	0.187***	9.45	0.075***	3.73	-0.205***	-10.45	-0.336***	-18.51
Total mediated path	-0.022***	-4.97	-0.003***	-5.85	0.024***	5.11	0.039***	5.59

This table reports path analysis of the direct and indirect links between financial reporting transparency and the cost of equity capital within newly public firms. All variables' definitions are listed in Appendix I. *, **, *** indicates statistics significance at the 0.10, 0.05, and 0.01 levels, respectively, in two-tailed tests.

The mediated path is the product of the *p* [*TRANS, IA_SPREAD*] and the *p* [*IA_SPREAD, COE*]. As Table 8 demonstrates, both direct and indirect paths are highly significant. Across all measures of the cost of equity capital, all direct and mediated paths are reliably nonzero, and the direct link is substantially more important than the indirect link. More importantly, the direct path result indicates a positive relationship between financial reporting transparency and the cost of equity capital, while the indirect path result shows a negative association between financial reporting transparency and the cost of equity capital and a positive association between information asymmetry and the cost of equity capital. These results suggest that even though greater financial reporting transparency could reduce newly listed firms' information asymmetry, the enhanced informativeness of financial reporting cannot decrease but increase the cost of equity capital.

CONCLUSIONS

Using a dataset of newly listed companies from 2004 to 2012, this study investigates the relationship between financial reporting transparency, perceived business risk, and the cost of equity capital. The empirical findings reveal that enhanced informativeness in the accounting reports of these newly listed firms can decrease the information risk associated with their earnings and cash flows. Furthermore, the study observes a positive association between the informativeness of these firms' accounting reports and the cost of equity capital. This suggests that in a context where investors possess varying information processing capabilities, more precise public information could paradoxically lead to an increase in the cost of capital.

A key implication of our paper is recognizing the inverse effect of enhanced information disclosure on the cost of capital. Specifically, regulators and investors highly value the improvement of disclosure quality in newly listed firms. The findings suggest that investment decisions concerning these newly public entities rely not only on high-quality financial disclosures, which offer reliable insights and reflect the firms' underlying economic value, but also on the broader financial disclosure environment. This includes factors such as the perceived reporting reputation and the extent of financial analyst coverage.

Moreover, the results indicate that newly listed companies require time to interact with capital market participants to build a solid financial reporting reputation. Additionally, the study highlights that certain unobservable or omitted factors also play a role in influencing the cost of capital. Not fully captured in the current research, these factors open avenues for future studies to further illuminate the complexities surrounding financial disclosure and its implications on the cost of capital.

This study acknowledges certain limitations. A primary constraint is the potential for measurement errors in assessing financial reporting transparency, perceived business risk, and implied cost of capital, despite efforts to control numerous factors. Additionally, the analysis is based exclusively on the survival of newly listed firms, as the financial data of delisted companies is inaccessible. This could potentially skew the findings towards those firms that have successfully remained in the market. It is also important to clarify that our findings do not universally characterize the effects of accounting information from newly listed firms as negative. On the contrary, our results demonstrate that enhancing the informativeness of accounting reports can mitigate information asymmetry, potentially leading to a decrease in the cost of equity capital. However, this does not imply a one-size-fits-all effect; the dynamics may vary across different contexts.

Furthermore, we acknowledge the need for future research to delve into how other firm characteristics, such as ownership structure, the presence of venture capital-backed investments, and the nuances of corporate governance, influence the relationship between the informativeness of accounting reports and the cost of capital. These additional dimensions could provide a more comprehensive understanding of how and why accounting information impacts investor perceptions and the financial standing of newly listed companies. We caution against the interpretation that informativeness of newly listed firms' financial reporting will negatively affect the firms' equity capital cost. Instead, our results should be interpreted subject to the following caveats. First, the implied cost of equity capital estimation models are just proxies of the actual cost of capital, and the use unstable earnings, dividend payout policies, and estimated growth rates may not capture the true business risk. Second, our inference is based on association tests and thus do not depict causal relationship. Finally, our results could be influenced by unknown correlated and omitted variables despite our efforts to utilize the most used variables in our models.

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APPENDIX 1: VARIABLE DEFINITIONS

Financial reporting transparency variables

TRANS	Financial reporting transparency, is the sum of the industry component, TRANSI, and the industry-neutral component, TRANSIN (Barth et al. 2013).
TRANSI	The adjusted R^2 from annual regressions of returns, RET, for year t on earnings before discontinued operations and extraordinary items, deflated by lagged price, Et/Pt-1, and change in earnings, deflated by lagged price, ΔEt/Pt-1, by Fama-French 17 industries (Barth et al. 2013).
TRANSIN	The adjusted R^2 from annual regressions of returns, RET, for year t on earnings before discontinued operations and extraordinary items, deflated by lagged price, Et/Pt-1, and change in earnings, deflated by lagged price, ΔEt/Pt-1, by portfolio based on the quartile of the residual from the industry regression (Barth et al. 2013).

Perceived business risk variable

IA_Spread	The intercept of the firm-year regression of daily spread (CRSP dataset: absolute spread divided by the average of bid and ask prices) on absolute daily return (Fu et al. 2012).

Implied cost of equity capital variables

Avg_COE	A composite cost of equity capital measure is the average of the following Three individual COE estimates: Claus and Thomas (CT, 2001), Gebhardt et al. (GLS, 2001), and Gordon and Gordon (1997).

Control variables

Size	Natural logarithm of the market value of equity at fiscal year-end (log of CRSP Abs (PRC)/ SHROUT).
MB	Ratio of the market value of equity (CRSP: Abs (PRC)/SHROUT) to the book value of equity of fiscal year-end (Compustat CEQ*1000).
Leverage	Ratio of total debt to total assets (Compustat (DLTT+DLC)/AT).
Beta	Calculated by regressing each firm's daily return on the market daily return in a current year.
Loss	Calculated by using earnings before extraordinary items divided by total assets at the beginning of the year.
GR	The log of one plus the percentage change in book value of equity (BE)
Turnover	The log of the median daily turnover ratio in a year (i.e., value of all shares traded divided by the market capitalization)
Volatility	Return volatility is the log of the standard deviation of daily returns in a year.

Other variables

ACC	Ratio of total accruals to total assets. Total accruals are calculated as the change in current asset (ACT) plus the change in debt in current liabilities (DCL) minus the change in cash and short-term investments (CHE) and minus the change in current liabilities (CLI).
Analysts	The average number of analysts covering the firm from 8 months before fiscal year end to 4 months after fiscal year end.
Capital	Indicator variable equal to 1 if the firm issues public debt or equity during the current and following two-year period, and 0 otherwise.
DD	Indicator variable equal to 1 if the company did not pay dividends, 0 otherwise.
Dispersion	The log of standard deviation of forecast earnings per share in the 4^{th} month of the fiscal year scaled by stock price.
DIV	Dividend payment in year t.

E	The earnings before extraordinary items of firm j in year t.
EV	The enterprise value of the firm (defined as total assets plus the market value of equity minus the book value of equity).
InstOwn	The percentage of shares owned by institutional shareholders at the end of the fiscal year.
NEGE	Indicator variable equal to 1 if the company has negative earnings, 0 otherwise.
Owners	The number of registered shareholders at the end of the fiscal year.
TA	Total assets in year t.
Surprise	The difference between the firm's actual earnings per share and the consensus forecast measured eight months prior to the fiscal year end scaled by stock price.
Return	The Market-adjusted stock return of the firm's equity measured over the fiscal year.
EarnVol	The log of the standard deviation of earnings scaled by assets measured over the previous 10 fiscal years.

APPENDIX 2: ADDITIONAL PATH ANALYSIS: FINANCIAL REPORTING TRANSPARENCY AND THE COST OF EQUITY