Financial Reporting Frequency, Information Asymmetry, and the Cost of Equity

Renhui Fu
Rotterdam School of Management
Erasmus University
Burgemeester Oudlaan 50, T08-39
Rotterdam, 3062 PA, The Netherlands
RFu@rsm.nl

Arthur Kraft*
Sir John Cass Business School
City University London
106 Bunhill Row, room 5063
London, EC1Y 8TZ, United Kingdom
Arthur.Kraft.1@city.ac.uk

Huai Zhang
Nanyang Business School
Nanyang Technological University
Nanyang Avenue, S3-01C-90
Singapore 639798
HuaiZhang@ntu.edu.sg

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1. Introduction

A firm's SEC-filed financial reports are accorded, on the whole, a substantially greater degree of prominence and attention than other firm disclosures. For example, firm executives, under threat of legal sanction, are obligated to certify the accuracy of these reports; independent accountants attest to the consistency of the financial statements' presentation and preparation with GAAP; and a wide variety of firm stakeholders intensely scrutinize and analyze these financial reports as they seek to assess the timing, magnitude, and risk of a firm's future cash flows. While the important role played by financial statements in the economy is well recognized, relatively little academic attention has been paid to how the frequency with which firms issue reports influences the decision making and actions of firm stakeholders. In this study, we examine directly how the frequency of interim reporting affects two linchpins of fair and efficient resource allocation in the economy—information asymmetry and the cost of equity.

Theoretically, the effect of reporting frequency on information asymmetry is unclear. A series of analytical papers show that public disclosures reduce information asymmetry by providing investors equal access to information (Diamond, 1985; Bushman, 1991; Lundholm, 1991). More frequent financial reporting thus could lead to lower information asymmetry if it increases the amount of information available to the public. However, as Verrecchia (2001) points out, one common assumption of these papers is that investors' private information is exogenously endowed. Relaxing this assumption, several studies model private information acquisition as an endogenous decision and show that sophisticated investors have incentives to acquire private information in anticipation of forthcoming disclosures. These incentives increase with reporting frequency because higher reporting frequency offers sophisticated investors more opportunities to profit from private information. It is thus possible that more frequent financial reporting leads to higher information asymmetry due to increased private information acquisition activities by sophisticated investors. Furthermore, whether more frequent reporting increases the amount of information is unclear since it can affect the information from other sources. Gigler and Hemmer (1998) show that more frequent mandatory financial disclosures may reduce firm's voluntary disclosures. In addition, high reporting frequency may encourage or discourage information production of financial intermediaries such as financial analysts and the business press (Bhushan, 1989a and 1989b; Healy and Palepu, 2001; Lang and Lundholm, 1993). Therefore, the question of how financial reporting frequency affects information asymmetry becomes more complicated when we consider its impact on other information sources.

Similarly, it is not clear conceptually how financial reporting frequency affects the cost of equity. While early works, such as Diamond and Verrecchia (1991), have generally suggested that more disclosures lower the cost of equity by reducing adverse selection and estimation risks, later studies have offered different views (Kim and Verrecchia, 1994; Zhang, 2001). In addition, Hughes et al. (2007) and Lambert et al. (2007) consider the force of diversification and suggest that disclosures have no impact on the cost of equity if they convey only information on diversifiable risks.

Consistent with the different views in theoretical works, empirical evidence is mixed on the relation between disclosures and information asymmetry. Leuz and Verrecchia (2000) show that a commitment to more disclosures significantly reduces information asymmetry for two out of the three information asymmetry measures they consider, but not for the third measure, stock return volatility. Van Buskirk (2011) finds no evidence that the information asymmetry is lower for firms that provide more frequent disclosure of sales figures. Similarly, prior empirical studies show mixed evidence on the relation between disclosures and the cost of equity. Botosan (1997) documents a negative relation between her self-constructed disclosure index and the firm's cost of equity for firms with low analyst following, but not for firms with high analyst following. Botosan and Plumlee (2002) find that the cost of equity decreases in the annual report disclosure level but increases in the quarterly report disclosure level. Francis et al. (2008) also show that the relation with the cost of equity takes on different signs for different types of disclosure measures. In sum, the mixed theoretical and empirical evidence suggests that the impact of financial reporting frequency on information asymmetry/cost of equity remains an important empirical issue.

To examine this empirical issue, we collect interim reporting frequency data for the time period 1951-1973. The SEC required annual financial reporting in 1934 and raised the required reporting frequency to semi-annual reporting in 1955. Not until 1970 did the SEC mandate quarterly reporting. However, prior to 1970 a substantial proportion of firms reported more frequently than required by the SEC.[1] For example, Butler et al. (2007) document that at least 70% of firms reported quarterly in the period we examine. By offering substantial cross-sectional and time-series variation in reporting frequency, our sample period provides an ideal setting to investigate our research question. It is difficult to study our research question using more recent data in the U.S. because, since 1970, almost all firms have followed the SEC's requirement of quarterly reporting, resulting in a lack of

[1] More detailed descriptions of the reporting environment from 1951 to 1973 are provided by Leftwich et al. (1981) and Butler et al. (2007).

variation in reporting frequency.

During our sample period, a significant proportion of firms voluntarily choose to report more frequently than the SEC requires. Their decision on reporting frequency is not likely to be random. Thus, it's possible that some unobservable firm characteristics, such as the firm's riskiness, affect both observed reporting frequency and information asymmetry/cost of equity, giving rise to an endogeneity concern. To alleviate this concern, in addition to a simple OLS regression, we use the following three approaches – a firm fixed effects model, a two-stage least squares estimation procedure (2SLS hereafter), and a matched control sample.

Our empirical results can be summarized as follows. In a simple OLS regression of information asymmetry/cost of equity on reporting frequency and other control variables, the coefficient on reporting frequency is negative and significant, suggesting that firms with higher reporting frequency have lower information asymmetry/cost of equity. We obtain similar inferences from both the firm fixed effects model and two-stage procedure. Specifically, results from the two-stage procedure suggest that an increase of one in the reporting frequency on average reduces our information asymmetry measure, the price impact, by 0.216% and the cost of equity measure based on the CAPM model by 0.628%.

Results from the matched control sample approach show that information asymmetry and the cost of equity decrease significantly for firms that increase their reporting frequency relative to control firms, regardless of whether the increase in reporting frequency is voluntary or mandatory. Specifically, the price impact decreases by 0.431% and 0.455% on average and the cost of equity based on the CAPM model drops by an average of 1.217% and 1.279% for firms with a voluntary increase and for firms with a mandatory increase in reporting frequency, respectively. Most increases amount to doubling the reporting frequency (i.e., from semi-annual reporting to quarterly reporting). Our results related to decreases in reporting frequency are much weaker, possibly because decreases in reporting frequency are typically temporary and do not reflect a commitment to reduced disclosures. Statistically, more than 90% of firms with a reduction in the reporting frequency revert back to the original level or higher level of reporting frequency over the three years after the reduction.

Our contribution to the literature is as follows. First, despite calls for more frequent financial reporting in the U.S. and abroad (FASB, 2000; Litan and Wallison, 2000; SEC, 2000; Jopson, 2006; Yiu, 2009), theoretically, it is unclear whether such a requirement actually improves the information environment (Bhushan, 1989a and 1989b; Gigler and Hemmer, 1998; Healy and Palepu, 2001; Lang and Lundholm, 1993). We document empirically that a mandatory increase in financial reporting frequency leads to lower information asymmetry and the cost of equity. By highlighting the benefit of higher reporting frequency, this finding may be relevant to practitioners and security regulators who are interested in exploring the consequences of higher financial reporting frequency.

We also contribute to the broad disclosure literature by adding evidence to the debate on whether more disclosures lead to lower information asymmetry and/or a lower cost of equity. Theoretically, some studies argue that more disclosures reduce the cost of equity by mitigating information asymmetry (e.g., Verrecchia, 1983; Diamond and Verrecchia, 1991), while other studies argue that more disclosures may actually increase information asymmetry and the cost of equity, due to private information acquisition by informed traders (Kim and Verrecchia, 1994; McNichols and Trueman, 1994). Consistent with the mixed theoretical evidence, empirical studies find that the relation between disclosure and information asymmetry/cost of equity varies across different types of firms, different types of disclosures and different measures of information asymmetry (Botosan, 1997; Leuz and Verrecchia, 2000; Botosan and Plumlee, 2002; Francis et al., 2008). To the extent that more frequent financial reporting constitutes more disclosures, we extend prior literature by providing more evidence on this important topic.

Furthermore, we argue that our proxy for disclosure – financial reporting frequency – is not affected by subjective biases. The majority of disclosure-related studies use either a self-constructed index or financial analysts' ratings of firms' disclosures to measure the effectiveness of firms' disclosure practices. A self-constructed index requires subjective assessment on the importance of items disclosed by the firm[2], while analysts' ratings reflect analysts' perceptions of firms' disclosure practices, which may be inaccurate[3]. Our measure of disclosure, financial reporting frequency, is a less noisy measure because it is not based on any subjective assessment. The reduced measurement error helps to detect the effect of more disclosures. Additionally, the information contained in financial reporting, especially earnings, is more relevant to investors and more informative about firm value than other types of voluntary disclosures examined by prior literature, such as conference calls, which are issued mainly to provide supplemental information for results reported in financial statements. To the extent that more informative disclosures are likely to have a greater impact on information asymmetry/cost of equity, financial reporting frequency provides a powerful setting to detect the effect of disclosures.

[2] In addition, Heitzman et al. (2009) point out that self-constructed disclosure measures ignore the materiality of the disclosed items.
[3] For example Lang and Lundholm (1993) find a strong correlation between analysts' ratings on firm disclosure and the firm's past performance, suggesting the possibility that analysts simply assign higher ratings on disclosure to firms with better prospects and financial performance.

Lastly, our study contributes to the specific line of literature that examines disclosure frequency. Leftwich et al. (1981) examine the reporting frequency of firms over the period 1937-1948 and document that firms' choices on interim reporting frequency vary with the stock exchange listing. McNichols and Manegold (1983) show that return volatility around the annual earnings release is reduced for the 34 AMEX firms that change their reporting frequency from annual to semiannual or quarterly in the 1960's. Their evidence is consistent with the notion that interim reports diminish the marginal information provided by the annual earnings report. Using a sample prior to the SEC's requirement on quarterly reporting, Butler et al. (2007) examine the impact of reporting frequency on earnings timeliness. They show that a voluntary increase in reporting frequency improves earnings timeliness, but a mandatory increase has no impact. Van Buskirk (2011) focuses on monthly sales reporting in the U.S. retail trade section. He finds no evidence that information asymmetry is lower for firms that provide more frequent disclosure of sales figures. Our study adds to this line of literature by showing that an increase in financial reporting frequency, regardless of whether it is voluntary or mandatory, reduces information asymmetry and the cost of equity.

The rest of the paper is organized as follows. Section 2 reviews prior literature and develops our main hypothesis. Section 3 discusses measurement for information asymmetry and the cost of equity. Section 4 covers sample formation and descriptive statistics. Section 5 provides details of the estimation procedures. Section 6 discusses the empirical results. Section 7 concludes.

2. Literature review and hypothesis development

It is not clear how financial reporting frequency affects information asymmetry.[4] A series of theoretical papers show that public disclosures reduce information asymmetry by providing investors equal access to information (Diamond, 1985; Bushman, 1991; Lundholm, 1991). More frequent financial reporting thus could lead to lower information asymmetry assuming it increases the amount of information available to the public.

However, as Verrecchia (2001) points out, one common assumption of the theoretical research on disclosure reducing information asymmetry is that investors' private information is exogenously endowed. Relaxing this assumption, several studies model private information acquisition as an endogenous decision and show that it can widen the information asymmetry between informed and uninformed traders prior to scheduled disclosure (Kim and Verrecchia, 1991; Demski and Feltham, 1994; McNichols and Trueman, 1994). Because each financial reporting represents an opportunity to profit from private information, higher reporting frequency encourages informed traders to acquire private information, thereby increasing information asymmetry.[5]

Furthermore, whether more frequent reporting increases the amount of information is unclear since it can affect the information from other sources. Gigler and Hemmer (1998) theoretically show that managers might have lower incentives to disclose voluntarily if they are forced to increase the reporting frequency, suggesting that more frequent reporting

[4] More frequent financial reporting does not necessarily mean an increase in quality. Quarterly reporting differs from annual reporting in two ways. First, quarterly financial statements are not audited, while annual financial statements are. Second, compared to annual earnings, quarterly earnings are based on more subjective managerial estimations. Mendenhall and Nichols (1988) find that managers have an income increasing bias in the interim quarters and defer bad news to the fourth quarter. Das and Shroff (2002) present evidence that fourth-quarter reversal is more likely to be a result of earnings management than mean reversion of earnings or fourth quarter settling up. These academic evidence casts doubt on the claim that more frequent financial reporting is necessarily of higher quality.

[5] Kim and Verrecchia (1994) predict that in the short term, information asymmetry increases with disclosure frequency. We are unable to test this prediction because of a lack of data on earnings announcement dates for our sample period (1951-1973).

could lead to less voluntary disclosures. In addition, more frequent reporting could alter the information production of financial intermediaries such as financial analysts and the business press. On one hand, the effect could be negative because more frequent financial reporting enriches the information set publicly available to investors, thereby reducing the value of information provided by information intermediaries (Healy and Palepu, 2001). On the other hand, the effect could be positive because more frequent financial reporting reduces the information acquisition costs of those intermediaries, enabling them to provide information of higher quality (Bhushan, 1989a and 1989b; Lang and Lundholm, 1993). The consideration of the impact on other information sources further complicates the question of how financial reporting frequency affects information asymmetry.

Consistent with the different views present in theoretical work, empirical work offers mixed evidence on the relation between disclosures and information asymmetry. Leuz and Verrecchia (2000) examine the impact of adopting more onerous financial standards on information asymmetry using a sample of German firms that voluntarily switch from German GAAP to an international reporting regime (i.e., IAS or U.S. GAAP). They find that switching firms have smaller bid-ask spreads and higher trading volume following the switch, relative to German GAAP firms. However, they fail to find significant results for stock return volatility, their third measure of information asymmetry. Van Buskirk (2011) examines monthly sales reporting in the U.S. retail trade section. He finds no evidence that information asymmetry is lower for firms that provide more frequent disclosure of sales figures.

The mixed theoretical and empirical evidence leads to the following two hypotheses.

H1a: Higher reporting frequency reduces information asymmetry.

H1b: Higher reporting frequency increases information asymmetry.

Turning to the relation between disclosures and the cost of equity, theoretical work suggests that this relation is context-specific.[6] While early works, such as Barry and Brown (1985), Amihud and Mendelson (1986), Diamond and Verrecchia (1991) and Handa and Linn (1993), have generally suggested that more disclosures lower the cost of equity by reducing adverse selection and estimation risks, later studies have offered different views. Kim and Verrecchia (1994) model a setting where more voluntary disclosures lead to greater private information acquisition, resulting in higher cost of equity. Zhang (2001) shows that the relation between the cost of equity and voluntary disclosures may be positive or negative, depending on what causes variation in disclosure levels. Hughes et al. (2007) and Lambert et al. (2007) consider the force of diversification, and their results suggest that disclosures have an impact on the cost of equity only if they convey information on non-diversifiable risks. Given the mixed prior findings, it is not clear how financial reporting frequency affects the cost of equity.

Consistent with the different theoretical views, the empirical evidence is mixed on the relation between more frequent disclosures and the cost of equity. Botosan (1997) finds a negative relation between her self-constructed disclosure index and the firm's cost of equity for firms with low analyst following but not for firms with high analyst following. Botosan and Plumlee (2002) and Francis et al. (2008) document that the relation between the cost of equity and voluntary disclosures varies across different disclosure measures.

[6] Armstrong et al. (2011) show that the impact of information asymmetry on the cost of equity depends on the competitiveness of the market. In a perfectly competitive market, information asymmetry does not affect the cost of equity. Without a clear understanding of the competitiveness of the market, we cannot infer the impact of reporting frequency on the cost of equity from the results based on information asymmetry.

Specifically, Botosan and Plumlee (2002) find that the cost of equity decreases in the annual report disclosure level but increases in the quarterly report disclosure level. Their interpretation is that more detailed quarterly reports attract the attention of transient investors, whose trading activities elevate the cost of equity by increasing the return volatility. Francis et al. (2008) show that the cost of equity is negatively related to the disclosure measure based on annual reports and 10-K filings, positively related to disclosure measures based on management forecasts and conference calls, and unrelated to press-release based disclosure measures. In sum, the mixed theoretical and empirical evidence suggests that whether financial reporting frequency affects the cost of equity remains an intriguing empirical issue.

Our discussions yield the following two hypotheses.

H2a: Higher reporting frequency reduces the cost of equity.

H2b: Higher reporting frequency increases the cost of equity.

3. Measurement of information asymmetry and the cost of equity

3.1. Measures of information asymmetry

We use two measures – bid-ask spread and price impact – to proxy for information symmetry.[7] The first measure, bid-ask spread, is a common measure of information asymmetry. The more severe the information asymmetry, the wider the spread necessary to cover higher expected market-maker losses from trading with informed investors. We follow Mohd (2005) and Silber (2005) and calculate daily bid-ask spread as $(Ask - Bid)/((Ask + Bid)/2)$. Since in our sample period the bid-ask spread may capture the

[7] Our inferences are robust to using either share turnover or stock return volatility as measures of information asymmetry. Results are available upon request.

difference between the daily low and daily high prices, we regress our raw spread measure on the daily absolute return for each firm year and use the estimated intercept term in our empirical analyses.[8] Our second measure of information asymmetry, price impact, is a measure of illiquidity suggested by Amihud (2002). This proxy is intended to capture the ability of an investor to trade in a stock without moving its price. Following Daske et al. (2008) we calculate illiquidity as the yearly median of the Amihud (2002) illiquidity measure (i.e., daily absolute return divided by the $ trading volume).[9]

3.2. Measures of cost of equity

Generally, there are two categories of methods to estimate the cost of equity. One is based on analysts' forecasts and the other is based on stock returns. Given the lack of analysts' forecast data during the sample period, we use return-based measures and the earnings-to-price ratio used in Francis et al. (2005) as proxies for the cost of equity. Specifically, our measures of cost of equity include ex-post realized returns, expected returns based on the CAPM model, expected returns based on the Fama-French three-factor model, and earnings-to-price ratio. We discuss each next.

3.2.1. Cost of equity based on realized returns

Gebhardt et al. (2001) suggest that ex-post realized returns should be an unbiased estimator of the unobservable cost of equity in an efficient market. We thus include the

[8] Our results are similar if we use the raw value of the spread or if we adjust the spread by deducting the absolute value of daily return. To alleviate the concern that our bid-ask spread results are subject to measurement error, we rely more on the price impact results to draw inferences related to information asymmetry.

[9] Our results are robust to measuring the bid-ask spread using the yearly median value and measuring the price impact using the yearly mean value. Results are available upon request.

realized annual returns as a proxy for the cost of equity due to its theoretical appeal. It is computed by compounding the twelve monthly returns in the calendar year.[10]

3.2.2. Cost of equity based on the CAPM model

Realized returns are a potentially noisy measure of cost of equity because they are affected by unexpected cash flow news and discount rate news, according to Vuolteenaho (2002). To reduce such noise, we use expected returns based on the CAPM model and the Fama-French three-factor model as measures of cost of equity.

We compute the cost of equity based on the CAPM model by running the following regression: $r_t = \alpha + \beta r_{M,t} + \varepsilon_t$, where $r_{M,t}$ indicates the market return and r_t indicates the stock return. For each firm-year observation, the parameters of the model α and β are estimated using daily data in the past year.[11] We use value-weighted CRSP returns as proxies for the market return. After the parameters are estimated, we plug in the market return in year t to obtain the estimated expected return r_t, which is our cost of equity estimate.

3.2.3. Cost of equity based on the Fama-French three-factor model

The Fama-French three-factor model argues that the expected returns are decided by three risk factors: market return, size and book-to-market ratio (e.g., Fama and French, 1996). Our regression model is specified as $r_t = \alpha + \beta_1 r_{M,t} + \beta_2 SMB_t + \beta_3 HML_t + \varepsilon_t$, where r_t indicates the stock return, $r_{M,t}$ indicates the market return, and SMB_t and HML_t

[10] We obtain quantitatively similar evidence when we exclude observations with negative cost of equity values.

[11] We obtain similar results when we use the monthly data in the past five years to estimate these parameters in the CAPM model and Fama-French model.

respectively indicate the risk premium related to firm size and that related to book-to-market ratio. We use the above model to estimate the factor loadings, α, β_1, β_2 and β_3, using daily data in the past year. After estimating parameters, we plug in the market return and the compounded annual risk factors for year *t* in the regression above to estimate the cost of equity.[12]

3.2.4. Earnings-to-price ratio

Dechow and Dichev (2002), Francis et al. (2005) and Liu et al. (2002) suggest that a higher price-earnings multiple implies a lower cost of equity as investors are willing to pay more for a given dollar of earnings, if the cost of equity is lower. We follow Francis et al. (2005) and use the earnings-to-price ratio (EP ratio hereafter) as one measure of cost of equity. Because a negative EP ratio is difficult to interpret, consistent with Francis et al. (2005), we require earnings to be positive and thus this ratio is missing for non-profitable firms.

4. Sample formation and descriptive statistics

The starting point for the analysis is the firm-level reporting frequency data described in Butler et al. (2007). Their primary source for determining how frequently firms published financial statements during the 1951–73 period was the index to the annual edition of *Moody's Industrial News Reports*. The following firms are excluded from their sample: firms not listed on either the NYSE or AMEX; firms lacking CRSP or Compustat data; and firms in industries with distinctive disclosure requirements (e.g., utilities;

[12] The market return is proxied by the value-weighted CRSP return for the calendar year. We compound daily factors to obtain annual size and book-to-market factor.

financial service, insurance, and real estate firms; and railroad and other transportation companies). For our regression analysis, we require further that all variables included in both the first- and second-stage regressions of the 2SLS procedure to be non-missing, except the EP ratio and the cost of equity (COE hereafter) measure based on the Fama-French three-factor model.[13] Our sample consists of 7,654 firm-year observations for the period 1951 to 1973.[14]

[Insert Table 1 here]

Table 1 reports the frequency distribution by reporting regimes. In the period 1951-1954 when only annual reporting is mandated, 24.09% of our sample firms report semiannually and 68.98% of them report quarterly. During the period 1955-1969, although only semiannual reporting is mandated, 86.97% of our sample firms report quarterly. After 1970, all firms report on a quarterly basis. The mean annual reporting frequency increases from 3.36 in the annual reporting regime to 3.77 in the semiannual reporting regime, and finally to 4 in the quarterly reporting regime. Our sample period offers significant variation in reporting frequency and provides an ideal setting for our investigation.

[13] Requiring the cost of equity measure based on the Fama-French three-factor model reduces our sample size to 6,083 observations. This is because Fama-French daily factors are available only after July 1, 1963. Requiring the EP ratio to be non-missing reduces our sample size to 6,290 observations, because this ratio is missing for non-profitable firms. Imposing these additional requirements not only reduces the sample size significantly but also limits the generalizability of our findings.

[14] Our data exclude observations that do not seem to follow the SEC's requirements. For example, we observe firms reporting annually after 1955 and semi-annually after 1970. These observations are largely related to firms that either became public or were delisted in the middle of the calendar year and firms in industries that were allowed by the SEC to defer adoption of the higher reporting frequency standards. Including these firms does not alter any of our inferences. Firms that report three times a year are mainly firms that decide to switch to quarterly reporting in the middle of the calendar year. Miscoding firms' reporting frequency biases against any significant findings, especially for our matched control sample approach results.

[Insert Table 2 here]

Table 2 provides descriptive statistics for the sample. In Panel A, the descriptive statistics for the full sample are provided. The mean values of the bid-ask spread and price impact are, respectively, 1.393 and 3.266. The mean value of the COE measure based on ex-post return is 12.960%, while the COE measure based on the CAPM model and Fama-French three-factor model have a mean value of 8.663% and 8.024%, respectively.[15] The EP ratio has a mean value of 7.342% and a median value of 6.720%. The mean and median values of Size are 17.660 and 17.539, respectively. The mean values of log(Turnover) and log(Volatility) are -6.786 and -3.748, respectively. The mean and median values of beta are 1.108 and 1.019, respectively. The log value of book-to-market ratio (log(BM)) has a mean value of -0.464 and a median value of -0.398. The mean value of log(Growth) is 0.102. Leverage has a mean value of 37.2%.

Panel B of Table 2 provides descriptive information related to the observations included in our matched sample analysis. Our treatment sample consists of 1,525 observations for the three years before and after a frequency change, while our control sample consists of observations that are matched with treatment observations by industry, year and size (please refer to Section 5.4 for more details). Panel B shows that the distribution of the variables is similar between the treatment sample and the control sample, suggesting that the match is successful.

[Insert Table 3 here]

[15] Although our results show that the mean value of the realized returns is higher than that of expected returns, it is likely driven by extreme values, since the median value of the realized returns (4.854%) is actually lower than that of expected returns based on the CAPM model (8.372%).

Our results reported so far are based on the raw values of information asymmetry/cost of equity measures. These raw values may exhibit time-series pattern, as a result of changing microeconomic conditions or technological improvement in the stock trading. In addition, the reporting frequency is correlated with passing of time since the SEC raised mandatory reporting frequency over the years. If we regress the raw values on the reporting frequency, the coefficient on reporting frequency may simply capture the time-series trend of cost of equity/information asymmetry. To alleviate this concern, we de-trend information asymmetry/cost of equity measures and use the de-trended values in our analyses hereafter.

Table 3 reports Pearson and Spearman correlation coefficients for the variables. Coefficients significant at the 5% level are in bold. We find that reporting frequency exhibits a significant negative correlation with the two proxies of information asymmetry and four proxies of cost of equity, suggesting that higher reporting frequency is associated with lower information asymmetry/cost of equity. In addition, reporting frequency is positively correlated with firm size, turnover, volatility, beta, growth and leverage. Our three return-based measures of cost of equity are all positively correlated with one another.

5. Research Design

Observed reporting frequency may reflect firms' voluntary choices or the SEC's requirement. For example, reporting quarterly in 1953 likely reflects the firm's own discretions while reporting quarterly in 1972 likely reflects the SEC's quarterly reporting requirement in 1970. Whether the reporting frequency is a firm's choice is a matter of concern because it determines the need to control for endogeneity. In the pooled sample, the likelihood of observing voluntary choices is high because the majority of firms report on the quarterly basis during the bulk of our sample period when the SEC did not require quarterly reporting. We identify three approaches to address this endogeneity concern: a firm fixed-effects model, a 2SLS procedure, and a matched sample analysis.[16] Because these approaches aim to control for endogeneity, it is not clear whether they are appropriate for cases in which reporting frequencies reflect the SEC's requirement. For these cases, the OLS regression approach seems adequate.[17] The remainder of this section discusses each of the four approaches, including the OLS regression approach.[18]

5.1 OLS Regression

We estimate a simple OLS regression of our information asymmetry or cost of equity proxy on reporting frequency and control variables.

Prior literature suggests that information asymmetry is associated with the market value of equity, share turnover and return volatility (Stoll, 1978; Chiang and Venkatesh, 1988; Glosten and Harris, 1988). Following Leuz and Verrecchia (2000) and Daske et al.

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[16] Leftwich, et al. (1981) and Butler et al. (2007) show that reporting frequency is related to several firm characteristics, including operating cycle length, firm size, firm performance, NYSE listing status, leverage and industry concentration. One might be concerned that our results are due to those firm characteristics, but not reporting frequency per se. In response, we follow a two-step procedure. We first regress reporting frequency on operating cycle, firm size, NYSE listing status, lagged firm performance, future firm performance, external financing, leverage, industry concentration and industry reporting practice. By construction, the residual is related to reporting frequency but orthogonal to all independent variables in the regression. We then replace reporting frequency with the residual and obtain similar results from the OLS regression, the firm fixed effects regression and the 2SLS procedure. Our finding is inconsistent with the concern that reporting frequency is simply a proxy for underlying firm characteristics. Results are available upon request.

[17] We run the pooled regression separately for the voluntary period (1955 through 1967) and the mandatory period (1968-1970) and find that our inferences hold in both periods.

[18] To examine whether our results are robust to alternative measurements of changes in reporting frequency, we re-define reporting frequency as a fraction of the number of reports in a year (annual reporting would be a 1, semi-annual ½, and quarterly ¼) and find results with similar implication.

(2008), we control for these three variables when the dependent variable is an information asymmetry measure (the bid-ask spread/the price impact). The market value of equity (*Size*) is the log of the average market value of equity at the beginning and end of the prior calendar year. Share turnover – *log(Turnover)* – is the log of the median daily share turnover ratio during the year (i.e. the dollar value of all shares traded during the day divided by the market capitalization on that day). Finally, return volatility – *log(Volatility)* – is computed as the log of the standard deviation of daily returns during the year.

Consistent with Botosan (1997), we control for firm size, beta (estimated by using daily return data in the prior year)[19] and logged value of the book-to-market ratio (the ratio of book value of equity to the market value four months after prior fiscal year end) when the dependent variable is one of the three return-based cost of equity measures. Additionally, when the dependent variable is the COE measure based on ex-post realized returns, we also include the logged value of growth (computed as one plus the percentage change in book value of equity) to control for variation in realized returns due to changes in expected future cash flows (as opposed to changes in risk).

For the EP ratio, we follow Francis et al. (2005) and control for firm size, beta, the logged value of growth, and financial leverage. Financial leverage is calculated as total liabilities divided by the sum of total liabilities and market value of equity at the beginning of the year.

5.2 Firm Fixed Effects

A firm fixed effects model generates unbiased estimates under the assumption that

[19] As a robustness check we also calculate full period betas (see Easley et al. 2002). Results are qualitatively similar and all inferences remain unchanged. Results are available upon request.

those unobservable firm characteristics, which affect both reporting frequency and information asymmetry/cost of equity, are constant over time. A firm fixed effects model is recommended by econometrics textbooks (e.g., Wooldridge, 2000) and is widely used in applied research (e.g., Campa and Kedia, 2002) as a practical approach to address the endogeneity concern.

5.3 2SLS Model

2SLS model is standard approach to control for endogeneity. Its success in addressing the endogeneity concern critically hinges on the quality of the instrumental variable.

The ideal instrumental variables (IV) should be highly related to the endogenous independent variable (in our case, reporting frequency) and unrelated to the dependent variable (the information asymmetry/cost of equity measures) except through the endogenous independent variable. As Larcker and Rusticus (2010) show, IV estimates are more biased and more likely to provide the wrong statistical inference than simple OLS estimates that make no correction for endogeneity, when instrumental variables are of low quality.

Our choice of IV, *Year Index*, is computed as the calendar year of an observation (for example, 1970) minus 1951, the year when our sample period starts. The SEC required annual financial reporting in 1934, semi-annual reporting in 1955 and quarterly reporting in 1970. The SEC's exogenous actions led to an increasing time trend in reporting frequency, implying a positive association between reporting frequency and *Year index*. In addition, by construction, the de-trended measures of information asymmetry/cost of equity have no time trend. Thus, *Year index* is related to reporting frequency but unrelated to the de-trended information asymmetry/cost of equity.

In the first-stage regression, the dependent variable is the observed reporting frequency and the independent variables include the instrumental variable (*Year index*) and control variables in the original OLS model discussed in Section 5.1. [20] We obtain the predicted reporting frequency from the first-stage regression results. The predicted reporting frequency replaces the observed reporting frequency in the second-stage regression. [21]

5.4 Matched Control Sample

The matched control sample approach compares changes in information asymmetry/cost of equity between firms that change their reporting frequency and firms that do not. Essentially, firms that maintain the same reporting frequency serve as controls for intertemporal changes in industry- and market-wide factors.

The details of this approach are provided as following. For each observation under investigation with a change in the reporting frequency (i.e., a treatment observation), we select one matched observation that does not change its reporting frequency, is in the same year-and-industry combination as the change observation, and is closest to the treatment observation in size. We then follow the difference-in-difference analysis approach used in

[20] Although the dependent variable (reporting frequency) in the first-stage regression is categorical, we follow Angrist and Pischke (2008) and estimate an OLS regression instead of a probit regression. Angrist and Pischke (2008) recommend estimating an OLS model as it does not assume a specific non-linear function in the first-stage. In untabulated results, we repeated all of our analysis using a probit regression in the first-stage regression and our inferences remain unchanged.

[21] We detrend only the information asymmetry/cost of equity measures. Since the instrumented reporting frequency is a linear combination of Year Index and other control variables (e.g. firm size, the book-to-market ratio), its coefficient in the second-stage model can be non-zero. Our approach is similar to Campa and Kedia (2002) in which the dependent variables are adjusted to be orthogonal to the instrumental variables.

Daske et al. (2008). Specifically, we obtain observations for six years (three years before and after the change) and regress measures of information asymmetry/cost of equity on indicator variables for the treatment versus control firms (*Treatment*) and the pre- and post-frequency change period (*After*), along with the relevant control variables from Tables 5 and 6 and an interaction term (*Treatment*After*).

Our regression model is as follows:

$$DV_{i,t} = \alpha_0 + \alpha_1 Treatment_{i,t} * After_{i,t} + \alpha_2 Treatment_{i,t} + \alpha_3 After_{i,t} + \alpha_4 CV_{i,t-1} + \varepsilon_{i,t} \qquad (1)$$

where

DV are the information-asymmetry (cost-of-equity) measures;

Treatment is a dummy variable coded as 1 for firms that change their reporting frequency and 0 for control firms;

After is a dummy variable coded as 1 for the three years after the change in the reporting frequency, i.e., (1,+3), and 0 for the three years before the change in the reporting frequency, i.e., (-3,-1);

CV are the control variables included in the information asymmetry regressions (see Table 5) or the cost of equity regressions (see Table 6).

We focus on the interaction between the two dummies (*Treatment* * *After*) because it represents differences in the change in information asymmetry/cost of equity across treatment firms and control firms. If higher reporting frequency reduces information asymmetry and the cost of equity, we expect the interaction term to be negative and significant.

Our matched sample approach offers the following benefits. First, to the extent that the control firm and the treatment firm are in the same industry and year, and they differ little in size, a significant coefficient on the interaction term (*Treatment* * *After*) cannot be attributed to size, industry-factor and year-factor. Also, this approach controls for non-time-varying firm characteristics. For example, firms listed on the NYSE typically provide more frequent interim reports than those listed on AMEX. To the extent that this firm characteristic is stable over time, our research design controls for it.

6. Empirical results

This section discusses the empirical results. Following Petersen (2009), we base inferences from all regressions on standard errors corrected for clustering at the firm and year levels. Furthermore, because some of the dependent variables appear to contain outliers (see Table 1 and Section 4), we winorize the top and bottom 1% of all of the variables other than reporting frequency.[22]

6.1 The first-stage regression results

To assess the quality of our IV, Table 4 shows the empirical results from the first-stage regression of the 2SLS procedure. We report results from five models. The first model only includes the IV (*Year index*), while the remaining four models also include the relevant control variables for the information asymmetry measures (model 2), the COE measure based on ex-post realized return (model 3), the COE measures based on CAPM model and Fama-French three factor model (model 4), and the EP ratio (model 5).

[Insert Table 4 here]

[22] As a robustness check we also estimate our results without winsorizing any of the variables. Results are qualitatively similar and all inferences remain unchanged. Results are available upon request.

We follow Larcker and Rusticus (2010) in evaluating the first-stage results. In all five regressions, the coefficient on our instrument, *Year index*, is positive and highly significant (significant at the 1% level), confirming our conjecture that the SEC's requirements on reporting frequency introduced an upward time-series trend in reporting frequency. In addition, Model 1 reports that the partial F-statistic is 50.59, which is much higher than the critical value of 8.96 suggested by Stock et al. (2002). Overall, our results suggest that we do not have a weak instrument problem and hence there is no need to perform a Wald test.

6.2. Regression results

Panel A of Table 5 reports results from the regression in which the dependent variable is the bid-ask spread. Three sets of results are reported. The "OLS" column reports simple OLS regression results, the "Fixed effects" column reports results from the firm fixed effects model, and the "2SLS" column reports results from our second-stage regression. In all three sets of results we find that the coefficient on each of our control variables (size, turnover, and volatility) is significant at the 1% level, implying that they help explain a firm's bid-ask spread. More importantly, in each of our regressions the coefficient on *Frequency* is significantly negative, indicating that as reporting frequency increases, a firm's bid-ask spread decreases. Its value ranges from -0.085 to -0.146, suggesting that bid-ask spread decreases by between 0.085% to 0.146% when reporting frequency increases by one.[23]

[23] For expositional purposes, the information asymmetry/cost of equity measures are multiplied by 100 in regressions.

[Insert Table 5 here]

Panel B shows results when the dependent variable is the price impact. Again all three sets of results reveal a significant association between price impact and each of the control variables. Furthermore, *Frequency* is significantly negative in each of the specifications. Our results suggest that price impact decreases by between 0.216% and 0.382% when reporting frequency increases by one.

We next compare the OLS results with the 2SLS results, following the advice of Larcker and Rusticus (2010). This comparison shows that the 2SLS results generally are smaller in magnitude (but still statistically significant), suggesting the existence of endogeneity. We also perform Hausman tests to evaluate the significance between the OLS and 2SLS results and find that in all specifications, the 2SLS results are significantly different from the OLS results.

In addition, we assess how severe the endogeneity problem must be to overturn our OLS results. It is well known that the bias induced by the omitted variable is determined by the omitted variable's correlation with the independent variable of interest and its correlation with the dependent variable. The stronger the two correlations, the more biased the coefficient estimate. A product of the two correlations therefore reflects the extent of the bias. This insight leads to the computation of the Impact Threshold for a Confounding Variable (hereafter ITCV). Specifically, the ITCV is the lowest product of the two correlations (the partial correlation between the dependent variable and the confounding variable, and the partial correlation between the independent variable of interest and the confounding variable) that could render the coefficient statistically insignificant (Frank

2000). The larger (smaller) the ITCV, the more (less) robust the OLS results are to omitted variables concerns.

The ITCV for reporting frequency is reported in Table 5, Panels A (the bid-ask spread model) and B (the price impact model). The ITCV of -0.034 in the bid-ask spread regression implies that the correlations between reporting frequency and bid-ask spreads with the unobserved confounding variable each need to be around 0.184 (=$0.034^{0.5}$) in order to overturn the OLS results. For the price impact regression, the ITCV of -0.182 implies that similar correlations need to be around 0.427.

In both cases the ITCV appears large enough to suggest that the OLS results are robust to omitted variable concerns. However, in order to be sure we use our control variables to calculate a benchmark for the magnitude of likely correlations involving the unobserved confounding variable. To do this we calculate Impact for each of our control variables. Impact is defined as the product of the partial correlation between the x-variable and the control variable and the correlation between the y-variable and the control variable (partialling out the effect of the other control variables). We also calculate $\text{Impact}_{\text{raw}}$ for each of the control variables, which is based on the raw correlations instead of the partial correlations.

In both Panels A and B, none of the control variables have an Impact or $\text{Impact}_{\text{raw}}$ with a larger magnitude than the relevant ITCV. This suggests that any unobserved confounding variable must be more highly correlated with the dependent variable and reporting frequency than any of our existing control variables in order to overturn our OLS results. Under the assumption that we have a good set of control variables, it is unlikely that such an unobserved variable exists, suggesting that our OLS results are robust to unobserved confounding variables.

Overall, the results in Table 5 indicate that higher reporting frequency is associated with lower bid-ask spread and price impact, both before and after we consider potential endogeneity issues. Our finding is consistent with the notion that higher reporting frequency increases the amount of the information provided to investors, and this better information environment more than offsets the impact of more intense private information acquisition activities by informed traders, leading to a lower level of information asymmetry.

[Insert Table 6 here]

Table 6 reports the regression results when cost of equity measures are dependent variables. Panel A reports the results when the dependent variable is the cost of equity estimate based on ex-post realized stock returns. We find that it is negatively correlated with size and positively correlated with the beta, book-to-market and growth rate. In each regression the coefficient on the reporting frequency variable is negative and significant (at the 5% level), which indicates that firms reporting more frequently have lower cost of equity.

Panel B and C report, respectively, the regression results when the measure of cost of equity is expected returns based on the CAPM model and expected returns based on the Fama-French three-factor model. Our results are similar to those reported in Panel A. In particular the coefficients on beta and book-to-market are positive and the coefficient on size is negative. More importantly, the coefficient on reporting frequency is always significantly negative in both panels, indicating a reduction in cost of equity when the reporting frequency increases.

To evaluate the economic significance of our finding, we focus on the results from the 2SLS procedure. The coefficient on reporting frequency ranges between -0.628 for the COE measure based on the CAPM model to -0.885 for the measure based on ex post realized returns, suggesting that when reporting frequency increases by one, the cost of equity measure is reduced by between 0.628% and 0.885%.

Panel D reports results when the dependent variable is the EP ratio. Consistent with Francis et al. (2005), the association with the EP ratio is negative for size and growth and positive for financial leverage. Beta shows mixed sign across model specifications. The coefficient on reporting frequency is negative and significant at the 5% level in all model specifications. The results from the 2SLS procedure suggest that the EP ratio is reduced by 0.502% when reporting frequency increases by one.

As in Table 5, we evaluate the robustness of our OLS results to unobserved confounding variables through the calculation of ITCV, Impact, and $\text{Impact}_{\text{raw}}$ scores. In each OLS regression the calculated ITCV has a larger magnitude than the calculated Impact and $\text{Impact}_{\text{raw}}$ scores for all of the control variables. Therefore, any unobserved omitted variable must have higher correlations with both cost of equity and reporting frequency than any of our control variables in order to invalidate the OLS results in Table 6. To the extent that we have a good set of control variables, this result suggests that our OLS results are unlikely overturned by unobserved confounding variables.

Overall, the results in Table 6 are consistent with the notion that the higher the reporting frequency, the lower the cost of equity. If we further consider the force of diversification, our results are consistent with the notion that more frequent financial reporting affects non-diversifiable information risk. A plausible alternative explanation is that financial reporting frequency affects diversifiable idiosyncratic information risk, which is priced in the cost of equity due to investors' failure to diversify their portfolios.

6.3 Results of the matched control sample approach

Following Butler et al. (2007), we examine the effects of mandatory and voluntary changes in reporting frequency. As in Butler et al. (2007) we define mandatory changes as increases to semiannual reporting occurring after 1954 and increases to quarterly reporting occurring after 1967.[24] There are two reasons to examine mandatory and voluntary changes separately. First, to the extent that mandatory changes in reporting frequency are not affected by the firm, results based on mandatory changes are less subject to the endogeneity concern and provide further assurance for our conclusions. Second, this analysis has implications for security regulators, as any potential changes in reporting frequency require mandatory acceptance at the firm level. Thus, it is important to investigate whether the documented benefits also apply to firms that change their frequency involuntarily. Panels A and B of Table 7 report the matched control sample approach results separately for firms that increase their reporting frequency voluntarily and firms that increase their reporting frequency involuntarily.

[Insert Table 7 here]

[24] Our results are similar when we classify increases to semiannual reporting after 1955 and increases to quarterly reporting after 1970 as mandatory increases and other increases as voluntary increases. Results are available upon request.

Panel A shows that, relative to control firms, firms that voluntarily increase reporting frequency experience significantly lower information asymmetry (lower bid-ask spreads and price impact) and lower cost of equity after the increase. The coefficients of the interaction between treatment dummy and period dummy (*Treatment*After*) are all significant at the 5% level. For example, the coefficient of the interaction for IA_{PI} is -0.431, significant at the 1% level, implying that, relative to control firms, firms that voluntarily increase their reporting frequency experience a drop of 0.431% in the price impact. The coefficient estimates in the COE_{CAPM} regression suggest that, on average, the cost of equity based on the expected returns from the CAPM model is lower by 1.217% for firms with a voluntary increase in reporting frequency than for control firms. These results suggest that voluntary increases in reporting frequency reduce information asymmetry and the cost of equity.

Results in Panel B are similar to those in Panel A. Relative to control firms, firms that increase their reporting frequency involuntarily have lower information asymmetry and lower cost of equity. The results are significant at the 5% level for both information asymmetry and cost of equity proxies. Take IA_{PI} for example, the coefficient of the interaction is -0.455, significant at the 1% level, indicating that price impact is reduced by 0.455% on average for firms that increase their reporting frequency involuntarily, relative to control firms. The coefficient estimates in the COE_{CAPM} regression suggest that, on average, the cost of equity based on the expected returns from the CAPM model is lower by 1.279% for firms with a mandatory increase in reporting frequency than for control firms. To summarize, we find that both information asymmetry and the cost of equity are lower for firms that increase their reporting frequency, regardless of whether the increase is voluntary or mandatory.

Panel C in Table 7 provides results for firms opting to report less frequently.[25] We generally find insignificant differences between treatment and control firms in information asymmetry and the cost of equity. This finding is puzzling and warrants more investigation. To that end, we examine the 78 firms that decreased their reporting frequency. Among them, only 5 remained at the reduced reporting frequency over the following 3 years, while the rest reverted back to the original level or an even higher level of reporting frequency. This finding suggests that the drop in reporting frequency is temporary.[26] Prior literature (e.g., Leuz and Verrechia, 2000; Brown et al., 2005) argue that information asymmetry/cost of equity effect is more significant for the commitment to disclose than for a one-time change in firms' disclosure practices. Thus, one possible explanation for this result is that these decreases are rationally viewed as temporary by investors and have a much weaker impact on either information asymmetry or the cost of equity.

7. Conclusion

How financial reporting frequency affects information asymmetry and the cost of equity is an important empirical question not answered by prior literature. Much of the difficulty in tackling this issue lies in the lack of cross-sectional variation in the reporting frequency of U.S. firms after the SEC's requirement of quarterly reporting in 1970.

[25] There are 17 firms that report less than quarterly before 1968 and delist during our mandatory switching period (from 1968 to 1970). To examine this issue, we include an analysis of firms that delisted during our sample period. We find that firms with smaller size, worse market performance and higher leverage ratio are more likely to delist, consistent with Leuz et al. (2008). Furthermore, we find that firms reporting less than quarterly are more likely to delist in the mandatory switch period.

[26] This drop likely reflects changes in a firm's fiscal year-end and missing information in the Moody's Industrial News Reports Index.

We address this difficulty by analyzing firm-level financial reporting frequency data for the period 1951 to 1973, a period during which meaningful cross-sectional variation in reporting frequency still existed. We are aware that firms' decisions on reporting frequency are not random, and we use three approaches to control for the self-selection. Our first approach is to use a firm fixed effects model, which yields unbiased and consistent estimates under the assumption that the unobservable firm characteristics affecting both reporting frequency and information asymmetry/cost of equity do not change over time. The second approach is to use 2SLS, which is widely used in academic studies to control for the self-selection bias. The third approach is to compare changes in information asymmetry and the cost of equity across firms that change their reporting frequency and firms that do not (matched sample approach). The results from all approaches are consistent with the notion that higher reporting frequency reduces information asymmetry and the cost of equity.

By showing that higher financial reporting frequency reduces information asymmetry and the cost of equity, our study documents the benefits of requiring more frequent financial reporting. In particular, our results related to mandatory changes in reporting frequency suggest that these benefits remain, even when firms are forced to deviate from their chosen reporting frequency. We however cannot conclude that firms should be forced to report more frequently, because our analysis does not address the potential costs of increasing reporting frequency (e.g., out-of-pocket costs, proprietary costs). A more detailed analysis of these costs is a fruitful area for future research.

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Table 1
Reporting frequency distribution by reporting regime

Period	N	Freq=1(%)	Freq=2(%)	Freq=3(%)	Freq=4(%)	Mean Freq	Mand. Freq
1951-1954	303	4.62	24.09	2.31	68.98	3.36	1
1955-1969	4,465		10.15	2.89	86.97	3.77	2
1970-1973	2,886				100.00	4.00	4
Total	7,654	0.18	6.87	1.78	91.17	3.84	

The sample includes 7,654 firm-year observations for the period 1951–73. Reporting frequency data were hand-collected from Moody's Industrial News Reports. The following industries are excluded from the sample: utilities (SIC 49); finance, insurance, and real estate (SIC 60–67); railroads and other transportation (SIC 40–41); and firms whose primary SIC code begins with 9. *N* is number of firms. *Freq = 1(%)* indicates the percentage of firms reporting annually. *Freq = 2(%)* indicates the percentage of firms reporting semiannually. *Freq = 3(%)* indicates the percentage of firms reporting three times per year. *Freq = 4(%):* indicates the percentage of firms reporting quarterly. *Mean Freq* is the mean value of reporting frequency. *Mand. Freq* is the reporting frequency mandated by SEC.

Table 2
Descriptive statistics

Panel A: Full sample

Variable	N	Mean	Std Dev	25%	50%	75%
Spread	7,654	1.393	0.917	0.739	1.168	1.821
PriceImpact	7,654	3.266	10.603	0.151	0.567	1.996
COE_{RET}	7,654	12.960	56.971	-22.246	4.854	34.333
COE_{CAPM}	7,651	8.663	20.357	-4.907	8.372	20.740
COE_{FF3}	6,083	8.024	39.539	-12.005	7.970	26.785
EP	6,925	7.342	4.396	4.359	6.720	9.314
Size	7,654	17.660	1.673	16.446	17.539	18.794
log(Turnover)	7,654	-6.786	0.900	-7.397	-6.813	-6.192
log(Volatility)	7,654	-3.748	0.458	-4.093	-3.739	-3.403
Beta	7,654	1.108	0.648	0.637	1.019	1.505
log(BM)	7,654	-0.464	0.739	-0.925	-0.398	0.057
log(Growth)	7,654	0.102	0.186	0.033	0.078	0.145
Lev	7,654	0.372	0.194	0.219	0.357	0.513

Panel B: Matched control sample

Variable	Treatment sample			Control sample		
	N	Mean	Median	N	Mean	Median
Spread	1,525	1.616	1.293	1,525	1.469	1.242
PriceImpact	1,525	3.226	0.535	1,525	3.216	0.556
COE_{RET}	1,525	25.461	11.088	1,525	24.997	10.514
COE_{CAPM}	1,525	9.920	8.786	1,525	10.689	9.310
COE_{FF3}	1,190	9.183	8.364	1,190	9.598	8.952
EP	1,360	6.496	5.948	1,360	6.863	6.551
Size	1,525	17.134	16.962	1,525	17.434	17.331
log(Turnover)	1,525	-6.742	-6.817	1,525	-6.660	-6.655
log(Volatility)	1,525	-3.636	-3.623	1,525	-3.750	-3.687
Beta	1,525	1.043	0.921	1,525	1.105	1.032
log(BM)	1,525	-0.686	-0.601	1,525	-0.449	-0.369
log(Growth)	1,525	0.106	0.073	1,525	0.094	0.074
Lev	1,525	0.334	0.312	1,525	0.373	0.379

The sample includes 7,654 firm-year observations for the period 1951–73. Panel A reports the descriptive statistics for the full sample used in Table 3 to Table 6 and Panel B provides the descriptive statistics for the matched control sample used in Table 7. All variables are based on calendar year. IA_{Spread} is the intercept of the firm-year regressions of daily spread based on CRSP file (i.e., absolute spread divided by average of bid and ask) on absolute daily return. IA_{PI} is the yearly median of the Amihud (2002) illiquidity measure (i.e., daily absolute stock return divided by the $ trading volume (measured in 10,000s). COE_{RET} is the cost of equity measure based on the realized return in a year. COE_{CAPM} is the cost of equity measure based on a modified CAPM model with risk loadings estimated from the daily data in the prior year. COE_{FF3} is the cost of equity measure based on Fama-French three-factor model with risk loadings estimated from the daily data in the prior year. The Fama-French daily factors start from 1 July 1963, which reduces the sample size to 6,083. COE_{EP} is calculated as earnings (#20) divided by month-end price four months after fiscal-year end. The earnings are required to be positive, which reduces the sample size to 6,925. *Size* is the log of average market equity value at beginning and end of 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. *Beta* is calculated by regressing each firm's daily return on the market daily return in the prior year. *log(BM)* is the log of the book-to-market ratio calculated as the ratio of book value of equity (#216) on its market value of equity at the end of four months after fiscal-year end. *log(Growth)* is the log of one plus the percentage change in book value of equity (#216). *Lev* is total liabilities [#181] divided by sum of total liabilities [#181] and beginning-of-year market value of equity. The values of IA_{Spread}, IA_{PI}, COE_{RET}, COE_{CAPM}, COE_{FF3} and COE_{EP} are multiplied by 100 for expositional purpose.

Table 3

Pearson (upper diagonal) and spearman (lower diagonal) correlation coefficients for select variables

	Freq	IA_{Spread}	IA_{PI}	COE_{RET}	COE_{CAPM}	COE_{FF3}	COE_{EP}	*Size*	*Log (Turnover)*	*Log (Volatility)*	*Beta*	*Log (BM)*	*Log (Growth)*	*Lev*
Freq		**-0.061**	**-0.165**	**-0.060**	**-0.048**	**-0.052**	**-0.058**	**0.155**	**0.077**	**0.023**	**0.131**	**-0.047**	**0.043**	**0.026**
IA_{Spread}	**-0.071**		**0.399**	**0.116**	**0.052**	**0.023**	**-0.115**	**-0.241**	**-0.385**	**0.614**	**0.410**	**0.043**	**0.026**	**0.135**
IA_{PI}	**-0.237**	**0.145**		**0.051**	**0.068**	**0.083**	0.007	**-0.421**	**-0.109**	**0.345**	**0.134**	**0.137**	**0.129**	**0.121**
COE_{RET}	**-0.063**	**0.118**	**0.175**		**0.403**	**0.461**	**-0.156**	**-0.028**	**0.265**	**0.053**	**0.133**	**0.194**	**0.241**	**0.074**
COE_{CAPM}	**-0.049**	**0.065**	**0.146**	**0.477**		**0.812**	**-0.162**	**-0.019**	**0.249**	0.005	**0.151**	**0.126**	**0.033**	**0.073**
COE_{FF3}	**-0.053**	**0.028**	**0.169**	**0.475**	**0.821**		**-0.139**	**-0.025**	**0.237**	0.007	**0.093**	**0.112**	**0.045**	**0.068**
COE_{EP}	**-0.057**	**-0.151**	**0.151**	**-0.143**	**-0.152**	**-0.093**		**-0.139**	**-0.085**	**-0.059**	**-0.148**	**0.554**	**-0.073**	**0.350**
Size	**0.159**	**-0.200**	**-0.779**	**-0.077**	-0.013	**-0.034**	**-0.134**		**-0.301**	**-0.551**	**-0.063**	**-0.275**	**0.124**	**-0.284**
log(Turnover)	**0.068**	**-0.448**	**-0.144**	**0.147**	**0.231**	**0.213**	**-0.120**	**-0.291**		**0.563**	**0.576**	**-0.135**	**0.199**	**0.183**
log(Volatility)	0.011	**0.600**	**0.297**	**0.110**	0.004	0.008	**-0.116**	**-0.555**	**0.577**		**0.467**	**-0.050**	0.012	**0.210**
Beta	**0.144**	**0.500**	**0.233**	**0.040**	**0.137**	**0.091**	**-0.192**	**-0.068**	**0.573**	**0.490**		**0.247**	**0.201**	**0.048**
log(BM)	**-0.050**	**0.020**	**0.315**	**0.176**	**0.121**	**0.082**	**0.571**	**-0.267**	**-0.126**	**-0.035**	**0.222**		**-0.387**	**0.549**
log(Growth)	**0.081**	**0.026**	**-0.229**	**0.268**	**0.036**	**0.019**	**-0.092**	**0.148**	**0.184**	**0.032**	**0.228**	**-0.474**		**-0.117**
Lev	**0.032**	**0.164**	**0.161**	**0.037**	**0.077**	**0.094**	**0.357**	**-0.267**	**0.184**	**0.195**	**0.067**	**0.555**	**-0.132**	

Sample includes 7,654 firm-year observations for the period 1951–73. All variables are based on a calendar year. *Freq* is number of times firm issues financial reports annually (based on information hand-collected from *Moody's Industrial New Reports*). IA_{Spread} is the intercept of the firm-year regressions of daily spread based on CRSP file (i.e., absolute spread divided by average of bid and ask) on absolute daily return. IA_{PI} is the yearly median of the Amihud (2002) illiquidity measure (i.e., daily absolute stock return divided by the $ trading volume (measured in 10,000s). COE_{RET} is the cost of equity measure based on the realized return in a year. COE_{CAPM} is the cost of equity measure based on a modified CAPM model with risk loadings estimated from the daily data in the prior year. COE_{FF3} is the cost of equity measure based on Fama-French three-factor model with risk loadings estimated from the daily data in the prior year. The Fama-French daily factors start from 1 July 1963, which reduces the sample size to 6,083. COE_{EP} is calculated as earnings (#20) divided by month-end price four months after fiscal-year end. The earnings are required to be positive, which reduces the sample size to 6,925. *Size* is the log of average market equity value at beginning and end of 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. *Beta* is calculated by regressing each firm's daily return on the market daily return in the prior year. *log(BM)* is the log of the book-to-market ratio calculated as the ratio of book value of equity (#216) on its market value of equity at the end of four months after fiscal-year end. *log(Growth)* is the log of one plus the percentage change in book value of equity (#216). *Lev* is total liabilities [#181] divided by sum of total liabilities [#181] and beginning-of-year market value of equity. The values of IA_{Spread}, IA_{PI}, COE_{RET}, COE_{CAPM}, COE_{FF3} and COE_{EP} are de-trended and are multiplied by 100 for expositional purpose. Correlation coefficients significant at 5%-level or lower are in boldface.

Table 4

The OLS regression of reporting frequency on its determinants

$Freq_{i,t-1}=\alpha+ \beta_1 YearIndex_{i,t-1}+ \beta_2 Size_{i,t-1}+ \beta_3\ log(Turnover_{i,t-1})+\beta_4 log(Volatility_{i,t-1})+ \beta_5 Beta_{i,t-1} +\beta_6 log(BM)_{i,t-1}+ \beta_7 Log(Growth)_{i,t-1}+ \beta_8 Lev_{i,t-1}+ \varepsilon_{i,t}$

Variable	Model 1 (Simplified)	Model 2 (IA_{Spread} + IA_{PI})	Model 3 (COE_{RET})	Model 4 (COE_{CAPM} + COE_{FF3})	Model 5 (COE_{EP})
Year Index	0.032*** (7.11)	0.033*** (8.55)	0.029*** (7.32)	0.028*** (7.32)	0.029*** (7.38)
Size		0.068*** (7.18)	0.061*** (6.94)	0.060*** (6.90)	0.067*** (7.64)
Log(Turnover)		0.068*** (3.95)			
Log(Volatility)		-0.104** (-2.24)			
Beta			0.069*** (4.72)	0.072*** (4.90)	0.053*** (3.42)
Log(BM)			0.026 (1.10)	0.018 (0.82)	
Log(Growth)			0.113*** (2.86)		0.074* (1.83)
Lev					0.187** (2.21)
Fixed effect	None	Industry	Industry	Industry	Industry
Adj. R^2	6.58%	16.54%	17.19%	17.09%	15.92%
F-statistic	50.59***	102.01***	30.42***	18.62***	35.36***

Sample includes 7,654 firm-year observations for the period 1951–73. The dependent variable *Freq* is the number of times firm issues financial reports annually (based on information hand-collected from *Moody's Industrial New Reports*). Model One is used to get partial R^2 and partial F statistics. Model Two is used as 1st-stage regression model for 2SLS with IA_{Spread} and IA_{PI} as the dependent variable. Model Three is used for 2SLS with COE_{RET} as the dependent variable. Model Four is used for 2SLS with COE_{CAPM} and COE_{FF3} as the dependent variable and Model Five is used for 2SLS with COE_{EP} as the dependent variable. *Size* is the log of average market equity value at beginning and end of 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. *Beta* is calculated by regressing each firm's daily return on the market daily return in the prior year. *log(BM)* is the log of the book-to-market ratio calculated as the ratio of book value of equity (#216) on its market value of equity at the end of four months after fiscal-year end. *log(Growth)* is the log of one plus the percentage change in book value of equity (#216). *Lev* equals total liabilities [#181] divided by sum of total liabilities [#181] and beginning-of-year market value of equity. The table reports OLS coefficient estimates and (in parentheses) t-statistics based on robust standard errors that are clustered by firm and year. ***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels (two-tailed), respectively.

Table 5

Regression Results with Information Asymmetry Measures as Dependent Variables

Panel A: Bid-ask spread model

$IA_{Spread\ i,t}=\alpha+\beta_1 Freq_{i,t-1}+\beta_2 Size_{i,t-1}+\beta_3\ log(Turnover_{i,t-1})+\beta_4 log(Volatility_{i,t-1})+\varepsilon_{i,t}$

Variable	OLS	Fixed Effects	2SLS	Impact	$Impact_{raw}$
Frequency	-0.146*** (-3.11)	-0.093*** (-4.66)	-0.085*** (-4.82)		
Size	-0.072*** (-6.85)	-0.131*** (-5.48)	-0.073*** (-6.45)	-0.031	-0.032
Log(Turnover)	-0.089*** (-4.20)	-0.098*** (-5.04)	-0.087*** (-4.99)	-0.008	-0.031
Log(Volatility)	1.245*** (17.81)	0.854*** (13.43)	1.273*** (18.49)	0.031	0.007
Fixed effects	Industry	Firm	Industry		
Adj. R^2	42.97%	74.89%	42.96%		
F-statistic	81.56***	21.85***	97.02***		
Hausman test: F statistics (P value)			8.49*** (<0.001)		
ITCV				-0.034	

Panel B: Price impact model

$IA_{PI\ i,t}=\alpha+\beta_1 Freq_{i,t-1}+\beta_2 Size_{i,t-1}+\beta_3\ log(Turnover_{i,t-1})+\beta_4 log(Volatility_{i,t-1})+\varepsilon_{i,t}$

Variable	OLS	Fixed Effects	2SLS	Impact	$Impact_{raw}$
Frequency	-0.382*** (-4.07)	-0.225*** (-4.57)	-0.216*** (-4.62)		
Size	-2.036*** (-13.75)	-3.172*** (-15.08)	-2.078*** (-18.14)	-0.158	-0.124
Log(Turnover)	-5.421*** (-8.68)	-5.788*** (-14.39)	-5.476*** (-20.24)	-0.029	-0.010
Log(Volatility)	9.929*** (7.03)	9.644*** (12.22)	10.099*** (17.76)	0.006	0.003
Fixed effects	Industry	Firm	Industry		
Adj. R^2	35.55%	56.59%	35.38%		
F-statistic	20.29***	9.54***	70.83***		
Hausman test: F statistics (P value)			5.07** (0.024)		
ITCV				-0.182	

Sample includes 7,654 firm-year observations for the period 1951–73. Information asymmetry is measured by IA_{Spread} or IA_{PI}, respectively. The "OLS" column reports the results based on OLS regressions. The "Fixed effects" column reports the results are based on OLS regressions with firm fixed effects. The "2SLS" column reports the 2^{nd} stage regression results. In the 1^{st} stage, we regress observed reporting frequency on our instrumental variable, Year Index (the difference between the current year and 1951), and control variables in the 2^{nd} stage regression. The predicted value from the 1^{st} stage regression is used in the 2^{nd} stage regression. The "Impact" column is the product of the partial correlation between the x-variable and the control variable and the correlation between the y-variable and the control variable (partialling out the effect of the other control variables). The "$Impact_{raw}$" column is the product of the simple correlation between the x-variable and the control variable and the simple correlation between the y-variable and the controlling variable. IA_{Spread} is the intercept of the firm-year regressions of daily spread based on CRSP file (i.e., absolute spread divided by average of bid and ask) on absolute daily return. IA_{PI} is the yearly median of the Amihud (2002) illiquidity measure (i.e., daily absolute stock return divided by the $ trading volume (measured in 10,000s)). *Freq* is the number of times firm issues financial reports annually (based on information hand-collected from *Moody's Industrial New Reports*). *Size* is the log of average market equity value at beginning and end of 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. The table reports OLS coefficient estimates and (in parentheses) t-statistics based on robust standard errors that are clustered by firm and year. The values of IA_{Spread} and IA_{PI} are de-trended and are multiplied by 100 for expositional purpose. *ITCV*, indicating the minimum impact of a confounding variable that would be needed to render the coefficient statistically insignificant, is defined as the product of the correlation between the x-variable and the confounding variable and the correlation between the y-variable and the confounding variable. ***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels (two-tailed), respectively.

Table 6

Regression Results with Cost of Equity Measures as Dependent Variables

Panel A: Ex-post realized returns as dependent variable

$COE_{RET\ i,t}=\alpha+\beta_1 Freq_{i,t-1}+ \beta_2 Size_{i,t-1}+ \beta_3\ Beta_{i,t-1} +\beta_4 log(BM)_{i,t-1}+ \beta_5 Log(Growth)_{i,t-1} +\varepsilon_{i,t}$

Variable	OLS	Fixed Effects	2SLS	Impact	$Impact_{raw}$
Frequency	-1.482***	-0.854**	-0.885**		
	(-3.03)	(-2.22)	(-2.39)		
Size	-4.675***	-17.299***	-4.734***	-0.004	-0.012
	(-4.45)	(-5.69)	(-4.72)		
Beta	5.961***	6.730***	6.027***	0.011	0.006
	(5.25)	(4.39)	(5.39)		
Log(BM)	8.768***	14.901***	8.744***	-0.000	-0.009
	(2.83)	(2.92)	(3.16)		
Log(Growth)	87.951***	80.820***	87.842***	0.016	0.022
	(6.02)	(7.90)	(9.31)		
Fixed effects	Industry	Firm	Industry		
Adj. R^2	10.42%	23.01%	10.43%		
F-statistic	180.03***	1.98***	13.53***		
Hausman test: F statistics (P value)			8.55*** (<0.001)		
ITCV				-0.027	

Panel B: Expected returns from modified CAPM model as dependent variable

$COE_{CAPM\ i,t}=\alpha+\beta_1 Freq_{i,t-1}+ \beta_2 Size_{i,t-1}+ \beta_3\ Beta_{i,t-1} +\beta_4 log(BM)_{i,t-1} +\varepsilon_{i,t}$

Variable	OLS	Fixed Effects	2SLS	Impact	$Impact_{raw}$
Frequency	-1.085**	-0.655**	-0.628**		
	(-2.04)	(-2.20)	(-2.31)		
Size	-0.286***	-3.127***	-0.286*	-0.003	-0.002
	(-3.17)	(-4.66)	(-1.77)		
Beta	9.202***	11.212***	9.328***	0.023	0.020
	(6.71)	(5.40)	(6.75)		
Log(BM)	2.868***	4.268***	2.766***	-0.001	-0.006
	(4.38)	(3.32)	(3.35)		
Fixed effects	Industry	Firm	Industry		
Adj. R^2	8.00%	14.28%	7.93%		
F-statistic	45.18***	2.10***	10.53***		
Hausman test: F statistics (P value)			9.34*** (<0.001)		
ITCV				-0.025	

Panel C: Expected returns from Fama-French three-factor model as dependent variable

$COE_{FF3\ i,t}=\alpha+\beta_1 Freq_{i,t-1}+\beta_2 Size_{i,t-1}+\beta_3\ Beta_{i,t-1}+\beta_4 log(BM)_{i,t-1}+\varepsilon_{i,t}$

Variable	OLS	Fixed Effects	2SLS	Impact	$Impact_{raw}$
Frequency	-1.014** (-2.21)	-0.662** (-2.25)	-0.658** (-2.40)		
Size	-0.575*** (-2.59)	-6.785*** (-3.23)	-0.489** (-2.52)	-0.001	-0.005
Beta	9.169*** (4.64)	13.974*** (5.19)	9.275*** (4.80)	0.016	0.013
Log(BM)	4.846*** (3.62)	8.143*** (3.06)	4.908*** (3.62)	-0.002	-0.004
Fixed effects	Industry	Firm	Industry		
Adj. R^2	2.68%	12.29%	2.66%		
F-statistic	57.19***	2.74***	3.50***		
Hausman test: F statistics (P value)			10.06 (<0.001)		
ITCV				-0.024	

Panel D: Earnings-price ratio as dependent variable

$COE_{EP\ i,t}=\alpha+\beta_1 Freq_{i,t-1}+\beta_2 Size_{i,t-1}+\beta_3 Beta_{i,t-1}+\beta_4 Log(Growth)_{i,t-1}+\beta_5 Lev_{i,t-1}+\varepsilon_{i,t}$

Variable	OLS	Fixed Effects	2SLS	Impact	$Impact_{raw}$
Frequency	-0.906** (-2.07)	-0.591*** (-3.87)	-0.502*** (-3.42)		
Size	-0.072 (-1.08)	-0.128* (-1.69)	-0.066* (-1.89)	-0.004	-0.021
Beta	-0.530*** (-4.79)	0.188* (1.87)	-0.601*** (-3.27)	-0.010	-0.028
log(Growth)	-1.379*** (-3.58)	-1.725*** (-5.60)	-1.374*** (-4.42)	-0.002	-0.007
Lev	7.210*** (3.70)	5.542*** (3.02)	7.072*** (3.66)	0.022	0.011
Fixed effects	Industry	Firm	Industry		
Adj. R^2	14.99%	43.85%	14.93%		
F-statistic	11.78***	4.88***	19.76***		
Hausman test: F statistics (P value)			9.93*** (<0.001)		
ITCV				-0.031	

Sample includes 7,654 firm-year observations for the period 1951–73. Cost of equity is measured by COE_{RET}, COE_{CAPM}, COE_{FF3} or COE_{EP}, respectively. The "OLS" column reports the results based on OLS regressions. The "Fixed effects" column reports the results are based on OLS regressions with firm fixed effects. The "2SLS" column reports the 2nd stage regression results. In the 1st stage, we regress observed reporting frequency on our instrumental variable, YearIndex (the difference between the current year and 1951), and control variables in the 2nd stage regression. The predicted value from the 1st stage regression is used in the 2nd stage regression. The "Impact" column is the product of the partial correlation between the x-variable and the control variable and the correlation between the y-variable and the control variable (partialling out the effect of the other control variables). The "$Impact_{raw}$" column is the product of the simple correlation between the x-variable and the control variable and the simple correlation between the y-variable and the controlling variable. COE_{RET} is the cost of equity measure based on the realized return in a year. COE_{CAPM} is the cost of equity measure based on a modified CAPM model with risk loadings estimated from the daily data in the prior year. COE_{FF3} is the cost of equity measure based on Fama-French three-factor model with risk loadings estimated from the daily data in the prior year. The Fama-French daily factors start from 1 July 1963, which reduces the sample size to 6,083. COE_{EP} is calculated as earnings (#20) divided by month-end price four months after fiscal-year end. The earnings are required to be positive, which reduces the sample size to 6,925. *Freq* is the number of times firm issues financial reports annually (based on information hand-collected from *Moody's Industrial New Reports*). *Size* is the log of average market equity value at beginning and end of prior calendar year. *Beta* is calculated by regressing each firm's daily return on the market daily return in the prior year. *log(BM)* is the log of the book-to-market ratio calculated as the ratio of book value of equity (#216) on its market value of equity at the end of four months after fiscal-year end. *log(Growth)* is the log of one plus the percentage change in book value of equity (#216). *Lev* equals total liabilities [#181] divided by sum of total liabilities [#181] and beginning-of-year market value of equity. The table reports OLS coefficient estimates and (in parentheses) t-statistics based on robust standard errors that are clustered by firm and year. The values of COE_{RET}, COE_{CAPM}, COE_{FF3} and COE_{EP} are de-trended and are multiplied by 100 for expositional purpose. *ITCV*, indicating the minimum impact of a confounding variable that would be needed to render the coefficient statistically insignificant, is defined as the product of the correlation between the x-variable and the confounding variable and the correlation between the y-variable and the confounding variable. ***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels (two-tailed), respectively.

Table 7

Regression results for treatment and control sample (-3,+3)

Variables	Information asymmetry measures		Cost of equity measures			
	IA_{Spread}	IA_{PI}	COE_{RET}	COE_{CAPM}	COE_{FF3}	COE_{EP}
Panel A: Voluntary increases in reporting frequency (N=1,090)						
Treatment*After	-0.162**	-0.431***	-1.613**	-1.217***	-1.308*	-0.963**
	(-2.07)	(-2.84)	(-2.22)	(-3.76)	(-1.96)	(-2.18)
Treatment	0.325***	0.411***	1.522*	1.221***	1.286*	0.994*
	(3.88)	(2.93)	(1.73)	(4.80)	(1.83)	(1.70)
After	-0.071***	-0.437**	-1.474***	-1.309**	-1.231***	-0.867**
	(-3.69)	(-2.40)	(-4.74)	(-2.05)	(-6.95)	(-2.35)
Panel B: Mandatory increases in reporting frequency (N=1,258)						
Treatment*After	-0.171**	-0.455***	-1.644***	-1.279***	-1.329**	-1.049**
	(-2.26)	(-3.92)	(-2.69)	(-5.94)	(-2.44)	(-2.47)
Treatment	0.189**	0.404***	1.420***	1.257***	1.285**	1.164**
	(2.54)	(4.53)	(2.63)	(6.49)	(2.48)	(2.37)
After	-0.033***	-0.461*	-1.286***	-1.239***	-1.115**	-0.850***
	(-3.24)	(-1.79)	(-6.21)	(-2.95)	(-2.03)	(-2.79)
Panel C: Decrease in reporting frequency (N=702)						
Treatment*After	0.102	0.381	1.053	1.224	1.255	0.874
	(0.64)	(1.59)	(1.34)	(1.03)	(1.38)	(1.58)
Treatment	0.159*	0.407	1.746	1.445	1.875	0.935
	(1.77)	(1.27)	(1.45)	(1.32)	(1.53)	(0.78)
After	-0.134	-0.415	-1.152	-1.500	-1.317	-0.817
	(-1.35)	(-0.88)	(-0.58)	(-1.55)	(-1.02)	(-1.23)

Sample includes 3,050 observations for treatment firms and control firms 3 years before and after the event year during the period 1951–73. Control firms are matched on industry (SIC2), size, and year. The results are based on the model $DV_{i,t}=\alpha + \beta_1 Treatment_{i,t}*After_{i,t} + \beta_2 Treatment_{i,t} + \beta_3 After_{i,t} + \beta_4 CV_{i,t-1} + \varepsilon_{i,t}$. *DV* are the information-asymmetry (cost of equity) measures; *Treatment* is the dummy variable coded as 1, for firms that change their reporting frequency and 0 for control firms; *After* is the dummy variable coded as 1 for three years after the change in the reporting frequency, i.e., (1,+3), and 0 for three years before the change in the reporting frequency, i.e., (-3,-1). *CV* are the control variables included in the information asymmetry regressions (see Table 5) and cost-of-equity regressions (see Table 6). The sample sizes for COE_{FF3} in Panel A, B, and C are 550, 1078 and 596, respectively. The sample sizes for COE_{EP} in Panel A, B, and C are 817, 916 and 605, respectively. For expositional purpose the table reports only OLS coefficient estimates (multiplied by 100) and firm-year-clustered t-statistics (in parentheses) for *Treatment*After*, *Treatment* and *After*. ***, **, and * indicate statistical significance at the 1%, 5%, and 10% levels (two-tailed), respectively.