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The Impact of Analyst Coverage Initiation on the Market's Ability to Anticipate Future Earnings

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Abstract

In this study, we examine whether the *first-time* initiation of analyst coverage of firms is associated with an increase in the stock market's ability to predict firms' future earnings. We find that stock prices are related to future earnings after the initiations, but not before. When we test for alternative explanations of the observed effect, we find that the increase in the market's ability to predict future earnings is not associated with selection biases such as reduction in risk, firm performance, and increase in liquidity, change in institutional interest, or a natural maturity effect. We conclude that analysts' activities make market prices reflect more future earnings information.

Keywords: Analyst coverage, future earnings response coefficient, market efficiency, liquidity, institutional investors
JEL Classification: G14, G20

I. Introduction

In this paper, we examine whether the initiation of analyst coverage of firms is associated with an increase in the stock market's ability to predict future earnings. Specifically, we investigate the association between current year stock returns and following year corporate earnings, for periods both before and after the *first-time* initiation of coverage. Our investigation is important because the effects of analyst coverage on the amount of information reflected in stock prices are largely unknown, despite the fact that analysts comprise a multi-billion dollar industry whose existence is based on its information benefits. Only two papers, Ayers and Freeman (2001) and Piotroski and Roulstone (2004) examine the impact of variation in coverage (defined as the number of analysts covering a firm) on the amount of future earnings information impounded into prices, and no papers study the effects of coverage initiations. It is possible that analysts will actually make prices less informative about future earnings by crowding out investors who would potentially search for private information. Whether this crowding out effect dominates the analyst-induced increase in price informativeness is an empirical question, which our tests address.

Initiation of coverage is a particularly interesting event, because analysts are perhaps the market's most important information intermediaries. Indeed, their reason of existence is to uncover and analyze information that better enables the market to forecast a firm's prospects and achieve a more accurate firm valuation. After the initiation, analysts continue to provide recommendations for the firms they cover. These recommendations change from time to time reflecting the change of analyst opinion regarding the future prospects of the firm vis-à-vis its price. Thus, it is important to know whether coverage initiation is associated with more future earnings information reflected in returns. Testing this expectation is the focus of our investigation.

There are two reasons why *first-time* coverage initiations provide a good sample for our tests. First, Dhiensiri et al. (2004) find that the importance of additional analysts decreases as the number of analysts covering a firm increases. Furthermore, Irvine (2003) finds that analyst coverage initiations are more important events compared to recommendation changes. Hence, *first-time* initiations will be more likely to provide a powerful test of the effects of analyst coverage. Second, the *first-time* initiation represents a structural change in the information structure of firms; i.e., firms transition from having no coverage to having coverage of analysts. Therefore, the effects observed in prices are likely to be attributed to the analysts and no other concurrent events.

Our proxy for the stock market's ability to predict firms' future earnings is the future earnings response coefficient (*FERC* hereafter), which is estimated by regressing current year stock returns against following year corporate annual earnings plus control variables.[1] *Ceteris paribus*, a change that better enables investors to predict future earnings will result in stock prices that better anticipate future earnings, and thus will result in higher *FERC*s. Our tests are similar to those in Gelb and Zarowin (2002) and Lundholm and Myers (2002), who relate measures of voluntary corporate disclosure to the *FERC*. In their case, the disclosure metric is based on AIMR analysts' rankings. In our case, the "disclosure" metric is based on the initiation of analyst coverage. A positive shift in *FERC* coincident with the initiation of coverage implies the coverage results in additional information that better enables investors to predict future earnings.

We regress current returns against future earnings for a sample of firms before (the pre-period) and after (the post- period) the initiation of analyst coverage, and we compare the change in the coefficients on the future earnings between the pre- and post- periods. We find that in the pre-period, information about future earnings is not reflected in stock prices, i.e. *FERC* is not significant. However, in the post-period *FERC* is positive and significant, indicating that information about future earnings is reflected in stock prices. Our findings are robust to various alternative explanations of the observed effect such as selection biases arising from reduction in risk, firm performance, reduction in the cost of capital associated with an increase in liquidity, increased institutional ownership around the time of initiation, and firms' natural maturity. All of these alternative explanations can cause an increase in *FERC*. Thus, our study provides the first empirical evidence that the initiation of analyst coverage is associated with an increase in the ability of stock prices to anticipate future earnings, attesting to the important role of analysts in capital markets.

The rest of the paper is organized as follows Section II reviews extant literature. Section III discusses the measures of stock price informativeness and provides the primary empirical model of the relations among variables and specifies hypotheses. Section IV discusses the sample and data. Section V reports the results of the primary empirical tests. Section VI reports the results of robustness checks. Section VII concludes.

[1] Several studies, such as Berger and Hann (2003), test for changes in earnings predictability by examining analyst forecast accuracy before vs. after a particular event. Since we deal with coverage initiations, such a test is infeasible.

II. Prior Literature

A number of papers examine the effects of either analyst coverage initiation (usually defined as coverage by an *additional* analyst of an already covered firm), or the effects of variation in coverage (defined as the number of analysts covering a firm). Research on coverage initiation has focused on the stock price response to the additional coverage or the effect of the additional coverage on liquidity. For example, Kim, Lin, and Slovin (1997), Jurgens (2000), Barber et al. (2001), and Irvine (2003) all find positive abnormal returns in response to coverage initiations, and Irvine (2003) finds an increase in liquidity associated with initiations.

Dhiensiri and Sayrak (2004) is the first study to examine first-time coverage of a firm by *any* analyst; i.e., the transition from no coverage to coverage. Like the papers on additional coverage, Dhiensiri and Sayrak also find a positive stock price response and increased liquidity associated with initiation.

However, no paper on coverage initiation, however defined, examines the information benefits of coverage initiation, and in particular how initiation affects the amount of information about future fundamentals impounded in prices. Irvine (2003), Dhiensiri and Sayrak (2004), and Demiroglu and Ryngaert (2005) all focus on the stock price reaction and change in liquidity around the announcement of coverage initiations.

The closest papers to ours are Piotroski and Roulstone (2004) and Ayers and Freeman (2001). These authors analyze the effect on *FERC* of cross-sectional variation in analyst coverage, defined as the number of analysts covering a firm, and they find that greater coverage is associated with increased informativeness. While their studies are similar to ours, we believe that initiation of coverage provides a superior test of analysts' information benefits than cross-sectional variation in coverage for a number of reasons. First, while it is obvious *prima facie* that initiation implies greater information search and analysis, it is not necessary that an increase from n to $n + k$ have a similar implication, since the additional analysts might well be redundant. Even if additional analysts are beneficial in this sense, we argue that their marginal benefit will be decreasing.

From a research design perspective, our results are less likely to be driven by problems in model specification. Since, in our tests, each firm acts as its own control (pre vs. post), the chance of correlated omitted variable bias is reduced. When analyzing firms cross-sectionally, there is always the possibility that correlated omitted variables are driving the results.

In summary, we contribute to literature on the effects of analyst coverage by being the first paper to examine the information benefits of coverage initiation.

III. Measure of Stock Price Informativeness

Our stock price informativeness measure (how much information about future earnings is capitalized into price) is based on Collins, Kothari, Shanken, and Sloan (CKSS, 1994). CKSS assume revisions in expected dividends to be correlated with revisions in expected earnings, which allows them to express current stock returns as a function of the current period's unexpected earnings and (discounted) changes in expected future earnings. Of course, the expectations imbedded in the returns are unobservable. The goal of this paper is to see whether the market's future earnings expectations, as implied in stock returns, are closer to future earnings realizations for firms when they are covered by analysts, i.e., whether analyst coverage results in current returns that are more highly associated with future earnings. CKSS proxy for current unexpected earnings using observed current change in earnings, and for changes in expected future earnings using changes in reported future earnings. These results in a regression of current annual stock returns, R_t on current and future annual earnings changes (firm subscripts omitted):[2]

$$R_t = a + b_0 D\,E_t + b_1\,DE_{t+1} + u_t \qquad (1)$$

where the earnings variables are in per share form and are scaled by the share price at the beginning of the current year (to avoid having to delete firms with negative or zero beginning-of-period earnings), and the stock returns are total annual stock returns, defined as capital gain plus dividend yield (measured over the period from nine months prior to fiscal year end to three months after fiscal year end).

Using earnings changes as explanatory variables assumes that earnings follow a random walk. Rather than impose this condition, we follow Lundholm and Myers (2002) and estimate the levels form of the regression:

$$R_t = a + b_0E_{t-1} + b_1E_t + b_2E_{t+1} + u_t \qquad (2)$$

As Lundholm and Myers note, (2) allows the random walk as a special case, if $b_0 = -b_1$, that is the more mean reverting earnings are, the smaller (in absolute value) is b_0/b_1. In (2), b_2 is the future earnings response coefficient (*FERC*), and is hypothesized to be positive; b_1, often referred to as the contemporaneous earnings response coefficient, is also hypothesized to be positive, and b_0 is hypothesized to be negative.[3] CKSS argue that using the actual future earnings introduces an error-in-variables bias in estimates of the *FERC*s, since the theoretically correct regressor is the unobservable expected future earnings. To help mitigate the errors in variables bias, we follow CKSS and include the future return as a control variable and estimate the following model:

$$R_t = a + b_0E_{t-1} + b_1E_t + b_2E_{t+1} + b_3R_{t+1} + u_t \qquad (3)$$

The hypothesized coefficient on R_{t+1} is negative.

Because their goal is to maximize the R^2 of the returns-earnings model, CKSS include three future years of earnings and returns in their regression. Our goal is to test whether the future response coefficient changed for firms after the initiation of coverage. Since our goal is not to maximize R^2, we include only one future year in our regression. Using a one-year horizon is also consistent with analysts, who focus on forecasting next year's earnings. Furthermore, since longer horizon earnings are harder to forecast, if there is a change in the future response coefficient related to coverage initiation, it is most likely to be detectable for the nearest year.

[2] We show only one future year for ease of exposition. Liu and Thomas (2000) estimate a model similar to (1) using analysts forecasts as proxies for market expectations. We use actual future earnings as the regressors and not analysts' forecasts, because we want to know how much information about future earnings is reflected in current returns (i.e., how close future earnings realizations are to the unobservable expectations implicit in stock prices).

[3] Contemporaneous ERC is often estimated as the sum of the coefficients on the current and lagged earnings, $b_0 + b_1$.

The goal of this paper is to see whether current returns became more highly associated with future earnings for firms after coverage initiation. Therefore, we estimate equation (3), using pooled data before versus after the initiation of coverage. Our null hypothesis is: b_2 (*POST*) = b_2 (*PRE*), where *PRE* and *POST* designate the pre-coverage and post coverage periods. Our alternative hypothesis is: b_2 (*POST*) > b_2 (*PRE*). Our tests are based on those in Gelb and Zarowin (2002) and Lundholm and Myers (2002), who show that increased voluntary disclosure results in higher *FERC*, implying that the disclosure reveals information that results in returns impounding more information about future earnings. In their case, the disclosure metric was based on analysts' rankings. In our case, the disclosure metric is based on the existence of analyst coverage.[4]

To examine the change in *FERC*, we estimate model (4) for our sample firms using data pooled over the pre- and post- periods. A dichotomous variable, *POST*, is defined as 'one' in the post-initiation period, and as 'zero' in the pre-initiation period:

$$R_t = a_0 + a_1 POST + b_0 E_{t-1} + b_1 E_t + b_2 E_{t+1} + b_3 R_{t+1} + b_4 POST' E_{t-1}$$

$$+ b_5 POST' E_t + b_6 POST' E_{t+1} + b_7 POST' R_{t+1} + u_t \qquad (4)$$

If the initiation of coverage enables investors to predict future earnings more accurately, we expect the shift in *FERC*, b_6, to be positive for our firms. Although we are primarily interested in the coefficient of $POST' E_{t+1}$, we interact variable *POST* with all of the regressors, thereby allowing all of the coefficients to vary by period. If the other coefficients do vary, but are constrained to be equal, any shift in *FERC* that we estimate might be due to the inappropriate restriction. On the other hand, if the coefficients on the other variables do not vary by period, estimating additional parameters decreases the power of the test, by reducing degrees of freedom. This would bias against our ability to find a relation between coverage initiation and informativeness. Given our large sample sizes this should not be a problem. Appendix 1 provides definitions of the variables in equation (4) together with definitions of additional variables used in supplemental analyses.

IV. Data and Sample

Information on analysts' coverage initiations is obtained from the *First Call Recommendation Database* (*FCRD* hereafter). *FCRD* provides access to a history of recommendations by leading Wall Street and regional brokerage firms. *FCRD* converts various brokerage recommendations into a consistent scale from one to five, where (1) is strong buy, (2) is buy, (3) is hold, (4) is under perform, and (5) is sell. *FCRD* also provides a field for previous recommendation so that a change in recommendations can easily be recognized. A record of (0) for the previous recommendation field indicates that there is no previous record for the firm by the analyst. Hence, the set of records that have a (0) for the previous recommendation provides us with the coverage initiation database. Our data set covers the time period from 1995 to 2000.

[4] Other recent papers using the informativeness measure are Durnev, Morck, Yeung, and Zarowin (2003), Ayers and Freeman (2001), Jiambalvo, Rajgopal, and Venkatachalam (2002), and Piotroski and Roulstone (2004).

One issue with *FCRD* is that a (0) for the previous recommendation does not necessarily guarantee that the firm has not been covered by some other analyst before. To address this issue, we perform the following data selection criteria to ensure that our sample of data is comprised of first-time coverage initiations: First, the recommendation record has to be the first one available for a given firm to be a part of our sample. Second, previous recommendation field should be equal to zero. Third, we filter out the recommendation records *FCRD* classifies as a deleted record to ensure data integrity. In addition, to identify the date and time of a record accurately, we impose the constraint that the origin of a recommendation be classified as real-time by *FCRD.*

Yet, note that the procedure described above may still result in some data points that are not first-time coverage initiations. For example, a firm may have been covered at some point in time, but had lost coverage prior to 1995. What we can claim is that our sample is comprised of coverage initiations that took place from 1995 to 2000, and no other analyst firm in *FCRD* had covered them before. Any data points that are not true initiations that infiltrate our sample would bias our results such that the relationship between current period returns and future earnings is weakened. Thus, any observed improvement in *FERC* based on our tests would be strong evidence for our argument.

Our analysis requires that the initiation take place at least one year after the IPO because we need E_{t-1} for our analysis (see equations 3 and 4). In order to fulfill this requirement, we match the original initiation sample of firms with the Center for Research in Securities Prices (CRSP) NYSE/AMEX/NASDAQ daily data tape by CUSIP. Data is retained in the sample if the number of cumulative trading days post IPO is at least 250 days. These restrictions provide us with a set of 967 firms.

To obtain financial information, we further require that the sample firms exist in COMPUSTAT. We match 930 of the 967 firms with records in COMPUSTAT. However, we can obtain both pre and post-initiation financial data for only 905 firms. One issue with the firms in our sample is that they have relatively small market capitalization (median value of 92.68 million), and several firms have relevant financial data classified as missing by COMPUSTAT. Because of data limitations in COMPUSTAT, we cannot compute E_t, E_{t-1}, or E_{t+1} for 305 of the 905 firms for either the pre or post initiation period. In addition, we eliminate 24 firms that have fiscal year end price of less than 1 dollar either in the pre or post initiation period. Our final sample is comprised of 576 firms, for which we have complete return and accounting data.

V. Primary Empirical Results

Figure I Panel A describes the distribution of analysts' coverage initiations that are matched with the stock return data on CRSP for the sample of 576 firms from 1995 to 2000. As observed in the histogram in Panel A, we do not observe time clustering in these coverage initiations. This is important, because it indicates that cross-sectional correlation (Bernard, 1987) is not a problem in our tests.

Figure I Panel B describes the frequency distribution of coverage initiation by the length of time between the coverage initiation and the security's first day of trading. We observe that the sample firms are not clustered with the recent IPO firms. This finding is consistent with Dhiensiri and Sayrak (2004)'s observation that most coverage initiations occur within 3 months after the IPO date. After the 3 months window has passed, the coverage initiations are not driven by the IPO process.

Figure I Panel C depicts the distribution of the initiations by industry. More than 12 percent of the initiations are for firms that operate in the Depository Institutions (SIC: 60). Other major industries represented in the sample comprise of Business Services Industry (SIC: 73), Electrical, Other Equipment (SIC: 36), Measurement Instruments, Photo Goods, Watch (SIC: 38), Holding, Other Investment Offices (SIC: 67), Chemical and Allied Products (SIC: 28), Computer Equipment (SIC: 35), Health Services (SIC: 80.), Electric, Gas, and Sanitary Services (SIC: 49), and Oil and Gas Extraction (SIC: 13). In sum, our sample represents 61 distinct two-digit SIC codes.

In Table 1 Panel A, we report the descriptive statistics of the 576 firms in our sample based on the end of the year information as of one-year before the coverage initiation takes place. Not surprisingly, the sample consists of relatively small firms. Median value for assets is 106.39 million in the year of coverage initiation. Median sales and market capitalization are 52.90 million and 92.68 million respectively.

In Table 1 Panel B we compare the average quarterly and annual (raw) returns of the sample. Consistent with McNichols and O'Brien (1997), it appears that analysts choose to initiate coverage for firms that are recent winners. Average quarterly returns prior to the coverage initiation are high at over 0.10 and are significantly higher than the average quarterly returns after the coverage initiations. For example, average quarterly return in Q–1 is 11.24 percent, whereas average quarterly return in Q+1 is 6.69 percent. The difference of 4.56 percent is significant at 5 percent level. Similarly, the average annual return in the year before coverage initiation is over 0.50, and is significantly greater than the average annual return in the following year.[5, 6]

In Table 2, we report the results obtained by regressing current returns against the earnings (past, current and future) and the future return in the pre- and post-initiation periods (regression model 3) as well as the pooled sample (regression model 4.) In Panel A, the coefficient of E_{t+1} is not significantly different from zero in the pre-initiation period, implying that stock returns in our sample did not reflect future earnings before the coverage initiations. However, in the post-initiation period, the coefficient of E_{t+1} is positive and significantly different from zero at one percent level, implying that after coverage initiations information about future earnings is impounded into stock prices.

[5] Note that our selection criteria differs from McNichols and O'Brien's in that we only select firms when they are first covered by any analyst, whereas they select firms that begin coverage by a particular analyst, even if the firm was already covered by other analysts. We address the issue of selection bias in Section 6.

[6] It is important to point out that despite the fact that coverage initiation is endogenous (i.e., analysts select which firms to cover), selection bias related to initiation of coverage for recent winners cannot be an explanation for our finding that FERC increases after initiation, because as discussed in section 7, high contemporaneous returns should produce a higher FERC in the pre-initiation period. Since returns are lower after initiation, FERC should be lower, not higher.

To test whether the coefficient of E_{t+1} after the coverage initiation is greater than the coefficient of E_{t+1} before the coverage initiation, we ran the pooled regression (4), where we combine the pre-initiation sample with the post- initiation sample. In this model, *POST* is a dummy variable, which is set equal to one if end of firm fiscal year is at least 6 months after initiation takes place. Note that in the pooled regression, the coefficient of $Post´E_{t+1}$ is positive and significant at one percent level, indicating that *FERC* increases significantly after the coverage initiations.[7] Thus, our primary results find that when analysts initiate coverage, the firms' stock returns come to reflect more future earnings information. This attests to the informational benefits of analysts.

The results in Table 2 indicate not only that stock prices reflect the information about the future earnings after the coverage initiations, but also suggest that this relationship (*FERC*) increase is concurrent with analysts' coverage initiations.[8]

VI. Alternative Explanations and Additional Analyses

The previous section showed that coverage initiation is associated with more future earnings information reflected in returns. While we interpret this change as evidence of analysts having beneficial information effects on the market, our results might be due to some other effect associated with the coverage initiation.

In particular, since analysts choose to cover recent winners, our results might be due to selection bias, or to firm success interacting with accounting conservatism. Likewise, the increase in *FERC* might be the result of a decrease in the rate used to discount future earnings, since coverage initiations are associated with increases in liquidity (decreases in risk) and earnings response coefficients are negatively associated with risk (Collins and Kothari, 1989). The increase in *FERC* might also be due to increased institutional interest that may accompany the coverage initiation. Finally, the increase in *FERC* might be due to the market's increased knowledge about the firm and its prospects independent of the coverage initiation as the firm matures. We now test these alternative explanations.

a. Reduction in Volatility

A crucial issue for any test like the one ours based on choice is self-selection (endogeneity of the choice). In particular, analysts might be skilled in identifying young firms whose earnings have become less volatile (more forecastable), which would strengthen the relation between current returns and future earnings. It is extremely difficult to test for such a selection bias, because the signals that analysts use to detect such firms are not observable to the researcher. Furthermore, detection of a shift in earnings variability is infeasible for our sample since there are not enough years of data to measure earnings volatility in both the pre and post-coverage periods.

[7] In a separate analysis (not reported here), we control for the effects of calendar years and industries. The results are virtually identically to what we report in Table 2.

[8] To ascertain that the results in Table 2 are not driven by outliers, we eliminate firms that are in the top and bottom 0.5 percent in term of asset, sales, incomes and market capitalization. The results in pooled and winsorized regressions are still qualitatively unchanged. The results (reported in the last column) are qualitatively unchanged, when we omit the future returns used to remedy the measurement error problem.

Thus, to test whether a decline in earnings variability is responsible for the observable increase in *FERC*, we examine whether the *FERC* shift is associated with a decline in return variability. We use return variability as a proxy for earnings variability, because both we have enough return observations in the pre-coverage period, and because Rosenberg and McKibben (1973) find that both variability measures are positively correlated.

The reduction in volatility explanation implies that the increase in *FERC* is driven by the initiation firms that experienced a reduction in return volatility. To examine this, we create a dummy variable *STDV* that is equal to 1 for firms that had high reduction in daily return volatility from *PRE* to *POST* (above the median reduction in total return volatility), and 0 otherwise, and we expand (4) as follows:

$$R_t = a_0 + a_1POST + b_0E_{t-1} + b_1E_t + b_2E_{t+1} + b_3R_{t+1} + b_4POST\acute{}E_{t-1} + b_5POST\acute{}E_t + b_6POST\acute{}E_{t+1} + b_7POST\acute{}R_{t+1} + a_2POST\acute{}STDV + b_8POST\acute{}STDV\acute{}E_{t-1} + b_9POST\acute{}STDV\acute{}E_t + b_{10}POST\acute{}STDV\acute{}E_{t+1} + b_{11}POST\acute{}STDV\acute{}R_{t+1} + u_t \quad (5)$$

Note that in (5), *STDV* is only interacted with *POST*. This is because the coefficients of the two groups cannot differ in the pre period based on a change in the post period. If a reduction in return volatility is driving our results, then b_6 will be zero and b_{10} will be positive; i.e., *FERC* will only increase for firms with the biggest reductions in volatility. If analyst coverage initiation is driving our results, then b_6 will be positive and b_{10} will be zero; i.e., *FERC* will increase for all of our sample firms.

The results of regression (5) are shown in Table 3. Firms that experienced a reduction in return volatility in the pre period did not experience a differential increase in *FERC*, as shown by b_{10}, the coefficient on $POST\acute{}STDV\acute{}E_{t+1}$. The coefficient on $POST\acute{}E_{t+1}$, b_6, remain significantly positive, consistent with our original results in Table 2. These results imply that reduction in return volatility is not driving the increase in *FERC* for our coverage initiation firms, which is inconsistent with the selection bias explanation.[9]

b. Firm Success and Accounting Conservatism

As pointed out in Table 1, coverage initiation firms tend to be recently successful. This is important for our tests, because in addition to informativeness, the factors that affect the relation between current returns and future earnings are related to success. In particular, Basu (1997) shows that due to accounting conservatism earnings timeliness is related to good vs. bad news. Accounting conservatism implies that firms tend to delay recognition of good news in their financial statements while they immediately recognize bad news. Since, all news is already reflected in stock prices (returns), good news (more successful) firms have less timely earnings (a weaker relation between current returns and current earnings) but have a stronger relation between current returns and future earnings (higher *FERC*).

Since we are comparing *FERC* in the pre-coverage vs. post-coverage period for the same set of firms, the crucial issue for us is not whether coverage initiation firms are successful *per se*, but whether they are *more* successful in the post period than in the pre period; i.e., did analysts select firms that experienced improved performance in the post- initiation period. We have reason to worry about this. For example, Sun (2003) finds that there is a positive relationship between the initial analyst rating of a firm and one- month post-rating market performance. As Dhiensiri and Sayrak (2004) note, majority of initial ratings are either Buy or Strong Buy. Thus, we must compare the firms' stock return performance before vs. after coverage initiation.

As shown in Table 1, initiation firms have significantly lower annual returns after initiation than before. Based on the evidence in Table 1, it is unlikely that firm success is driving our results. Nevertheless, we perform an additional test to increase our confidence. Since the successful firm explanation implies that the increase in *FERC* is driven by the initiation firms that experienced the most improved performance, we create a dummy variable *INC* that is equal to 1 for firms that had higher annual returns in POST than in PRE, and 0 otherwise, and we expand (4) as follows:[10]

$$R_t = a_0 + a_1POST + b_0E_{t-1} + b_1E_t + b_2E_{t+1} + b_3R_{t+1} + b_4POST \times E_{t-1} + b_5POST \times E_t + b_6POST \times E_{t+1} + b_7POST \times R_{t+1} + a_2POST \times INC + b_8POST \times INC \times E_{t-1} + b_9POST \times INC \times E_t + b_{10}POST \times INC \times E_{t+1} + b_{11}POST \times INC \times R_{t+1} + u_t \quad (6)$$

Note that in (6) like in (5), the dummy variable is only interacted with *POST*. This is because the coefficients of the two groups cannot differ in the pre period based on performance in the post period. Like in (5), if improved performance is driving our results, then b_6 will be zero and b_{10} will be positive; i.e., *FERC* will only increase for the best performers. If analyst coverage initiation is driving our results, then b_6 will be positive and b_{10} will be zero; i.e., *FERC* will increase for all of our sample firms.

The results of regression (6) are shown in Table 4. Firms that performed better in the pre period did not experience a differential increase in *FERC*, as shown by b_{10}, the coefficient on $POST \times INC \times E_{t+1}$. The coefficient on $POST \times E_{t+1}$, b_6, remains significantly positive, consistent with our original results in Table 2. These results imply that successful performance is not driving the increase in *FERC* for our coverage initiation firms.

c. Change in Liquidity

Another way that coverage initiations can increase *FERC* is by reducing the discount rate for future earnings, by increasing liquidity. Prior literature relates liquidity to discount rates, that is, firms with higher liquidity benefit from lower discount rates as indicated by relatively higher prices.[11] As Irvine (2003) and Dhiensiri and Sayrak (2004) find, analyst coverage is associated with significant increase in liquidity. Table 1 Panel A showed that for our sample, liquidity increases for newly covered firms.[12] Since this is the case, the increase in *FERC* in the post period may simply be due to the decrease in the discount rate, which is due to the increase in liquidity of the firms that had coverage initiations.

[9] We also estimated (5) using BETA instead of STDV, with similar results.

[10] We also estimated (6) with INC equal to 1 for firms that had above median change in annual returns from PRE to POST, with similar results.

To test for the effects of change in liquidity, we form a dummy variable *LIQ* that is set to 1 if the firm experienced above median percentage increase in liquidity, as measured by average daily transaction volume, and 0 otherwise. Expanding Equation 4 with this dummy variable, we obtain the following regression model:

$$R_t = a_0 + a_1POST + b_0E_{t-1} + b_1E_t + b_2E_{t+1} + b_3R_{t+1} + b_4POST \times E_{t-1} + b_5POST \times E_t$$

$$+ b_6POST \times E_{t+1} + b_7POST \times R_{t+1} + a_2POST \times LIQ + b_8POSTxLIQ \times E_{t-1} + b_9POST \times LIQ \times E_t + b_{10}POST \times LIQ \times E_{t+1} + b_{11}POST \times LIQ \times R_{t+1} + u_t \quad (7)$$

If change in liquidity is driving our results, we expect b_6 to be zero and b_{10} to be positive, which would indicate that only firms with relatively large increases in liquidity experienced an increase in *FERC*. Results in Table 5 indicate that this is not the case; b_6 remains significantly positive (at the 0.05 level, for a two-tailed test). Consistent with the liquidity-risk interpretation b_{10} is significantly positive (at about the .05 level for a two-tailed test). Hence, we conclude that while the increase in liquidity increases *FERC*, it is not the increase in liquidity that makes prices related to future earnings following coverage initiations.

d. Increase in Institutional Interest

Over the last decade, the importance of institutional investors in the equity markets has constantly increased. Institutional investors play multiple roles and potentially effect the information structure of firms in which they invest. Therefore, it is possible that our results are driven by institutional interest that tends to increase following analyst initiations.[13]

Among others, one role for large institutional investors is to provide a credible mechanism for transmitting information to the financial markets, that is, to other investors. According to Chidambaran and John (2000), large institutional investors can convey private information that they obtain from management to other shareholders. Another role suggested by Hartzell and Starks (2003) is that institutional investors serve a more effective monitoring role. The key point is that institutional investors are sophisticated investors whose analyses may cause more information to be reflected in prices.

To test for the effects of increase in institutional interest, we use a dummy variable *INST*, which is set to 1 if the firm experienced above median percentage increase in institutional ownership (percentage) and 0 otherwise obtaining the following regression model:

$$R_t = a_0 + a_1POST + b_0E_{t-1} + b_1E_t + b_2E_{t+1} + b_3R_{t+1} + b_4POST \times E_{t-1} + b_5POST \times E_t$$

$$+ b_6POST \times E_{t+1} + b_7POST \times R_{t+1} + a_2POST \times INST + b_8POSTxINST \times E_{t-1} + b_9POST \times INST \times E_t + b_{10}POST \times INST \times E_{t+1} + b_{11}POST \times INST \times R_{t+1} + u_t \quad (8)$$

[11] See, for example, Datar et al. (1998). Their evidence suggests that liquidity, as measured by the turnover rate (number of shares traded as a fraction of the number of shares outstanding,) plays a significant role in explaining the cross-sectional variation in stock returns.

[12] Note that this liquidity effect is not a selection bias, since it is caused by the analyst coverage.

If an increase in institution interest is driving our results, we expect b_6 to be zero and b_{10} to be positive, which would indicate that only firms with an increase in institutional interest experienced an increase in *FERC*. Results in Table 6 indicate that this is not the case; b_6 remains significantly positive (at the 0.05 level, for a two-tailed test). Hence, we conclude that, it is not the increase in institution interest that makes prices related to future earnings following coverage initiations.[14]

e. Natural Maturity Effect

Finally, our results might be due to a "natural maturity effect". After the IPO takes place, firms experience increased trading activity and transparency. This in turn makes the prices more efficient, independent of whether coverage was initiated or not. If prices reflect the expectations about future earnings more efficiently, we would expect to see higher *FERC*. In this view, increased *FERC* is not necessarily related to analyst coverage, rather to this "natural maturity" effect.[15]

To address the natural maturity issue, we examine whether the firms in our sample experienced significant increase in *FERC* due to the passage of time before or after the coverage initiation. Specifically, we estimate model (3) in in years *PRE*-1 versus year *PRE* (pre initiation period), and year *POST* versus year *POST*+1 (post initiation period). If prices become more efficient simply due to passage of time, we should see improvements in *FERC* in periods other than the initiation period. The reason we want to repeat the test for both before and after the initiation is that the natural maturity effect may be nonlinear. One reasonable form of this non-linearity is a concave monotone increasing function of time, where the improvements in price efficiency tend to be greater in the earlier periods after the IPO.

Table 7 reports our findings from the analysis of the natural maturity effect. In the before-initiation period, *FERC* is not significant at any conventional levels.[16] In the after-initiation period, however, *FERC* is significantly positive in year *POST* and *POST*+1. Thus, once coverage begins, the improvement in *FERC* appears to be permanent at its new, higher level.

The findings in Table 7 together with those in Table 2 indicate not only that stock prices reflect the information about the future earnings after the coverage initiations, but also that the increase in *FERC* is concurrent with analysts' coverage initiations. The results not only show that there is no significant link between prices and future earnings in the before-initiation period, but also that the improvement in *FERC* is permanent after the coverage initiation takes place, supporting the importance of analysts in the improvements of *FERC* during coverage initiations. Based on the results in Table 7, we conclude that the "natural maturity effect" does not explain the improvement in *FERC* for coverage initiation firms.

[13] In this sample, institutional ownership does not change significantly on average, although the median value for it increases (see Table 1). For a more detailed analysis, see Dhiensiri and Sayrak (2004).

[14] We also estimated (8) with INST set to 1 if the firm experienced an increase in institutional ownership (percentage) and 0 otherwise; the results are qualitatively unchanged.

[15] For example, Lang (1991) finds that young firms' earnings have a low association with stock returns, and that this association increases as the firms mature.

[16] In the pooled regression (not reported), there is no significant increase in FERC from year t-2 to year t-1.

In summary, we find no other potential alternative explanations can explain the increase in *FERC* after the initiation of analyst coverage. This increases our confidence that it is the activities of the analysts themselves that are responsible for this change.[17, 18]

VII. Conclusion

In this paper, we examine whether the initiation of analyst coverage of firms affects the stock market's ability to predict the firms' future earnings. Our proxy for the stock market's ability to predict firms' future earnings is the future earnings response coefficient, *FERC*, which is estimated by regressing current year stock returns against following year corporate annual earnings plus control variables.

We regress current returns against future earnings for a sample of firms before (the pre-period) and after (the post-period) the initiation of analyst coverage, and we compare the change in the coefficients on the future earnings between the pre- and post-periods. We find that in the pre-period, information about future earnings are not reflected in stock prices, i.e., *FERC* is not significant. However, in the post-period, *FERC* is positive and significant indicating that information about future earnings is reflected in stock prices. This result is robust to various alternative procedures, such as controlling for the effects of selection bias, firm performance, changes in risk (liquidity), institutional interest, and firms' natural maturity. While the possibility always remains that there exists some unidentifiable selection bias that might be driving our results, our numerous tests rejecting alternative explanations increase our confidence that it is the activities of the analysts themselves that are responsible for the observed increase in *FERC*.

Thus, our study provides the first empirical evidence that the initiation of analyst coverage is associated with an increase in the ability of stock prices to anticipate future earnings. This provides support for the important role that analysts play in the capital markets.

[17] Using First Call's Company Issue Guidelines database of management forecasts, we also examined whether increased FERC might be due to increased voluntary disclosures by management, which result in more forward looking information being impounded into price. We found no support for this explanation. This is not surprising, since less than 20% of our sample firms have a management forecast in the Post initiation period, and more importantly, management rarely gives a forecast for the future period. They are for the current period. Therefore, management forecasts are not a guideline for future earnings.

[18] Since negative earnings are more transitory and thus harder to forecast, we also included a dummy variable to control for negative (current or future) earnings. None of our conclusions were affected.

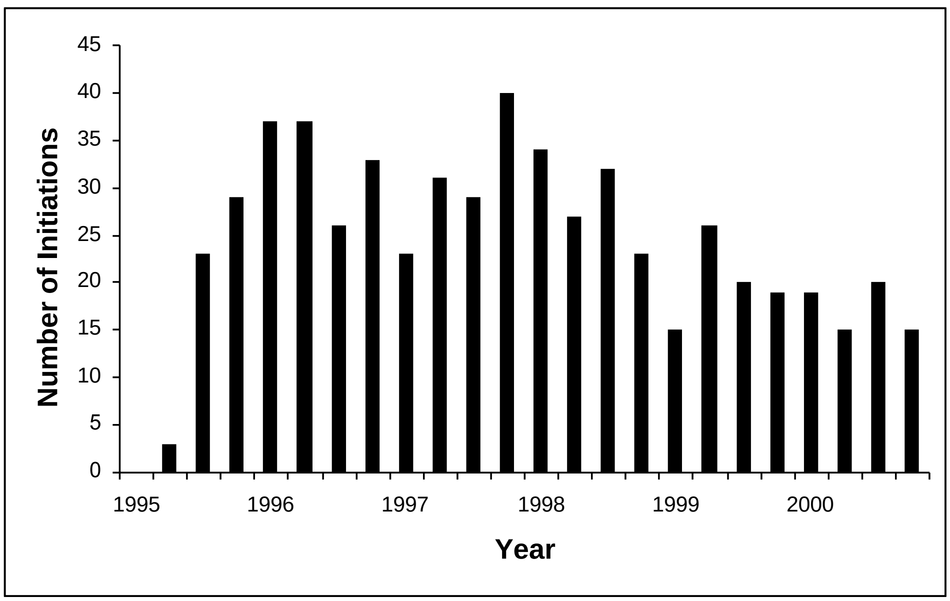
Appendix: Variable Definitions

E	Net income before extraordinary items (Annual COMPUSTAT Item# 18) divided this by #of shares*adjustment factor (Annual COMPUSTAT Item #25 multiplied by #27) and deflated by end of previous fiscal year Price.
Price	End of fiscal year share price (Annual COMPUSTAT Item #199 divided by #27)
Dividend	Dividend per share (Annual COMPUSTAT Item #26 divided by #27)
R	Annual market return for the fiscal year, defined as [(*Price* + *Dividend*) / $Price_{t-1}$]– 1
Liquidity	Liquidity as measured by average daily transactions volume.
INST	Institutional ownership as measured by percentage of shares held by institutional investors.
STDV	Total return volatility, based on daily returns for a one-year time frame in the PRE and POST periods.
POST	Dummy variable set equal to one, if end of firm fiscal year is at least 6 months after initiation takes place.

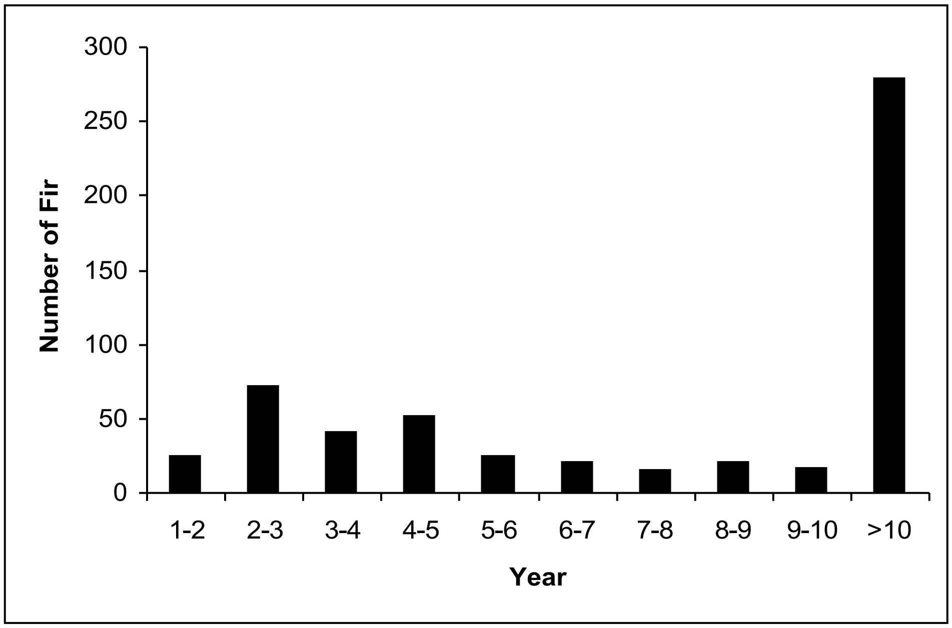
Figure I. Distribution of Coverage Initiations

Panel A. Initiations by time period

We describe the frequency distribution of coverage initiations by time period (quarters).



Panel B. The length of time between the coverage initiation and IPO date
We describe the frequency distribution of coverage initiation by the length of time between the coverage initiation and the security's first day of trading.



Panel C. Initiations by industry. We describe the frequency distribution of coverage initiations by industries. The distribution is based on two-digit SIC codes. We list only the first ten two-digits SIC codes separately: Depository Institutions (SIC: 60), Business Services Industry (SIC: 73), Electrical, Other Equipment (SIC: 36), Measurement Instruments, Photo Goods, Watch (SIC: 38), Holding, Other Investment Offices (SIC: 67), Chemical and Allied Products (SIC: 28), Computer Equipment (SIC: 35), Health Services (SIC: 80.), Electric, Gas, and Sanitary Services (SIC: 49), and Oil and Gas Extraction (SIC: 13). The rest are summed up into the "Other" category. There are 51 distinct two-digit SIC codes under "Other" category.

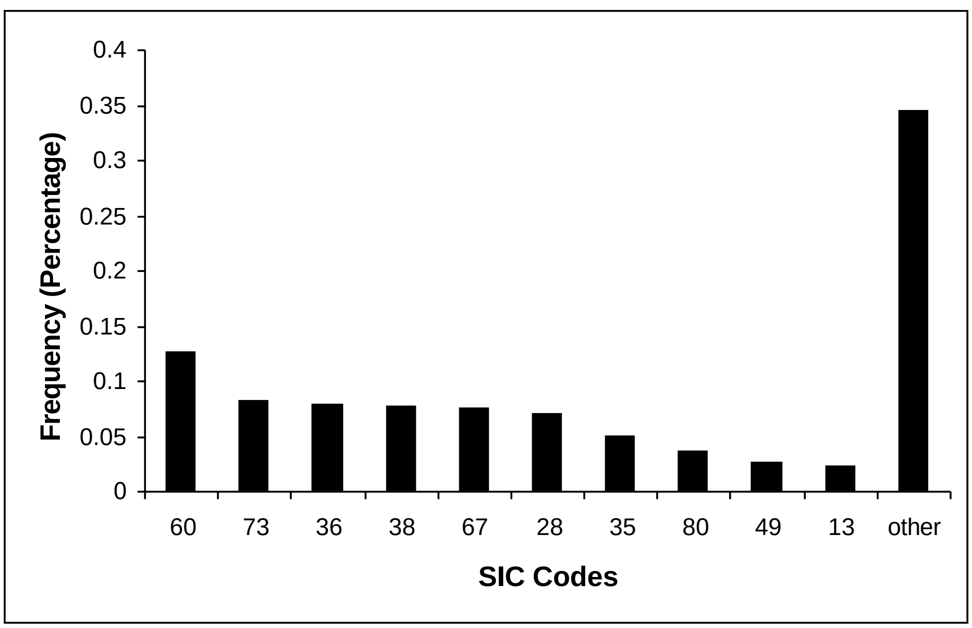


Table 1 Descriptive Statistics

Panel A. Sample Description
We report the descriptive statistics of the sample firms based on the end of the fiscal year information one year before the initiation took place (pre) and one year after the initiation took place (post). The figures are in millions except average daily volume. Sample size is 576.

	Pre Initiation		Post Initiation	
Variables	Mean	Median	Mean	Median
Total Assets	1,375.32	106.39	1,511.39 ^{aaa}	142.86 ^{xxx}
Sales	640.91	52.90	762.65 ^{aaa}	69.38 ^{xxx}
Income Before Extraordinary Items	41.50	3.59	40.67	4.35 ^{xxx}
Market Capitalization	1,177.39	92.68	1,285.53	102.69 ^{xxx}
Average Daily Volume (# of Shares)	74,080.44	20,465.48	94,798.40 ^{aaa}	28,416.99 ^{xxx}
Institutional Ownership	0.29	0.19	0.29	0.22 ^{x}

^{aaa}, ^{aa}, and ^{a} indicates that the difference between pre and post mean values is significant at 1, 5, and 10% respectively.
^{xxx}, ^{xx}, and ^{x} indicates that the difference between pre and post median values is significant at 1, 5, and 10% respectively.

Panel B. Market Returns
We compare average quarterly and annual market returns in the *PRE* and *POST* period. Q0 is the quarter that coverage is initiated.

Quarter	Return	Quarter	Return	Difference	t-stat	# of Pairs
Q-4	0.1054	Q+4	0.0313	0.0741	3.70	562
Q-3	0.1169	Q+3	0.0139	0.1029	4.79	566
Q-2	0.1811	Q+2	0.0252	0.1559	4.06	568
Q-1	0.1124	Q+1	0.0669	0.0456	2.39	571
Q0	0.1106					
Year	Return	Year	Return	Difference	t-stat	# of Pairs
PRE	0.5466	*POST*	0.2106	0.3360	4.59***	576

*** significant at 1% level in two-tailed test
** significant at 5% level in two-tailed test
* significant at 10% level in two-tailed test

Table 2 Measures of Stock Price Informativeness around the Coverage Initiations

We report the results of regression analysis where the dependent variable is the current stock return and the independent variable are normalized past, current and future earnings, future return and interaction variables. *Post* is a dummy variable set equal to one, which is set equal to one if end of firm fiscal year is at least 6 months after initiation takes place. The t-statistics of the coefficients are in the parentheses. For the pooled regression, we implement the White heteroscedasticity consistent estimator.

Model: $R_t = a_0 + a_1 POST + b_0 E_{t-1} + b_1 E_t + b_2 E_{t+1} + b_3 R_{t+1} + b_4 POST \acute{} E_{t-1}$
$+ b_5 POST \acute{} E_t + b_6 POST \acute{} E_{t+1} + b_7 POST \acute{} R_{t+1} + u_t$

Variables	Pre	Post	Pooled	Pooled Winsorized	and Pooled without R_{t+1}
Intercept	0.5407***	0.2294***	0.5407***	0.5502***	0.5113***
	(8.70)	(5.88)	(6.75)	(6.72)	(7.30)
E_t	1.8500***	0.4642***	1.8500*	1.8300*	1.8832*
	(4.28)	(2.69)	(1.67)	(1.64)	(1.70)
E_{t-1}	-1.1565***	-0.4000	-1.1565*	-1.1552*	-1.1388*
	(-3.93)	(-1.51)	(-1.84)	(-1.84)	(-1.81)
E_{t+1}	-0.0922	0.7343***	-0.0922	-0.0876	-0.1740
	(-0.34)	(3.92)	(-0.36)	(-0.34)	(-0.67)
R_{t+1}	-0.1370**	-0.0376	-0.1370	-0.1395	
	(-2.12)	(-0.88)	(-1.15)	(-1.14)	
POST			-0.3113***	-0.3231***	-0.2853***
			(-3.29)	(-3.34)	(-3.31)
$POST^* E_t$			-1.3857	-1.3625	-1.4130
			(-1.23)	(-1.20)	(-1.25)
$POST^* E_{t-1}$			0.7565	0.7805	0.7345
			(0.94)	(0.97)	(0.91)
$POST^* E_{t+1}$			0.8265***	0.8459***	0.8804***
			(2.91)	(2.96)	(3.08)
$POST^* R_{t+1}$			0.0994	0.0990	
			(0.80)	(0.78)	
# Observations	576	576	1152	1126	1152
F-Stat	7.64	6.85	9.22	9.26	10.84
Adjusted R^2	0.0442	0.0391	0.0604	0.0620	0.0564

***, **, * indicates that the coefficient is significantly different from zero at 1%, 5%, and 10% levels in two- tailed test respectively

Table 3 Alternative Explanation: Reduction in Return Volatility

Since the reduction in return volatility explanation implies that the increase in *FERC* is driven by the initiation firms that experienced a reduction in return volatility, we create a dummy variable *STDV* that is equal to 1 for firms that had high reduction in return volatility from *PRE* to *POST* (above the median reduction in total return volatility), and 0 otherwise. *STDV* is based on daily returns for a one-year time frame in the *PRE* and *POST* periods. For the pooled regression, we implement the White heteroscedasticity consistent estimator.

Model: $R_t = a_0 + a_1POST + b_0E_{t-1} + b_1E_t + b_2E_{t+1} + b_3R_{t+1} + b_4POST'E_{t-1} + b_5POST'E_t + b_6POST'E_{t+1} + b_7POST'R_{t+1} + a_2POST'STDV + b_8POST'STDV'E_{t-1} + b_9POST'STDV'E_t + b_{10}POST'STDV'E_{t+1} + b_{11}POST'STDV'R_{t+1} + u_t$

Variables	Pooled with *STDV* Dummy
Intercept	0.5428***
	(6.75)
E_t	1.8527*
	(1.67)
E_{t-1}	-1.1648***
	(-1.85)
E_{t+1}	-0.1092
	(-0.42)
R_{t+1}	-0.1376
	(-1.15)
POST	-0.2960**
	(-2.38)
$POST^* E_t$	-1.0124
	(-0.87)
$POST^* E_{t-1}$	0.1950
	(0.14)
$POST^* E_{t+1}$	0.8865***
	(2.76)
$POST^* R_{t+1}$	0.1119
	(0.90)
*POST*STDV*	-0.0434
	(-0.42)
$POST^*STDV^* E_t$	-0.5942
	(-1.50)
$POST^*STDV^* E_{t-1}$	0.9611
	(0.76)
$POST^*STDV^* E_{t+1}$	-0.0639
	(-0.20)
$POST^*STDV^* R_{t+1}$	-0.0923
	(-0.83)
# Observations	1148
F-Stat	6.12
Adjusted R^2	0.0588

***, **, * indicates that the coefficient is significantly different from zero at 1%, 5%, and 10% levels in two- tailed test respectively; the number of observations is less than in Table 2 due to missing daily returns.

Table 4 Alternative Explanation: Firm Success and Accounting Conservatism
Since the successful firm-accounting conservatism explanation implies that the increase in *FERC* is driven by the initiation firms that experienced the most improved performance, we create a dummy variable *INC* that is equal to 1 for firms that had higher annual returns in POST than in PRE, and 0 otherwise For the pooled regression, we implement the White heteroscedasticity consistent estimator.

Model: $R_t = a_0 + a_1 POST + b_0 E_{t-1} + b_1 E_t + b_2 E_{t+1} + b_3 R_{t+1} + b_4 POST\acute{} E_{t-1} + b_5 POST\acute{} E_t + b_6 POST\acute{} E_{t+1} + b_7 POST\acute{} R_{t+1} + a_2 POST\acute{} INC + b_8 POST\acute{} INC\acute{} E_{t-1} + b_9 POSTxINC\acute{} E_t + b_{10} POSTxINC\acute{} E_{t+1} + b_{11} POSTxINC\acute{} R_{t+1} + u_t$

Variables	Pooled with *INC* Dummy
Intercept	0.5407***
	(6.75)
E_t	1.8500*
	(1.67)
E_{t-1}	-1.1565*
	(-1.84)
E_{t+1}	-0.0922
	(-0.36)
R_{t+1}	-0.1370
	(-1.15)
POST	-0.7106***
	(-8.55)
$POST^* E_t$	-1.3034
	(-1.14)
$POST^* E_{t-1}$	1.4767**
	(2.26)
$POST^* E_{t+1}$	0.5283*
	(1.95)
$POST^* R_{t+1}$	0.0751
	(0.60)
*POST*INC*	0.7890***
	(8.35)
$POST^*INC^* E_t$	-0.0348
	(-0.10)
$POST^*INC^* E_{t-1}$	-1.5420
	(-1.16)
$POST^*INC^* E_{t+1}$	0.0042
	(0.02)
$POST^*INC^* R_{t+1}$	-0.0243
	(-0.43)
# Observations	1152
F-Stat	10.58
Adjusted R^2	0.1044

***, **, * indicates that the coefficient is significantly different from zero at 1%, 5%, and 10% levels in two- tailed test respectively

Table 5 Alternative Explanation: Change in Liquidity

Since the liquidity-risk explanation implies that the increase in *FERC* is driven by the initiation firms that experienced a reduction in the discount rate due to increased liquidity, we create a dummy variable *LIQ* that is equal to 1 for firms that had high liquidity increases from *PRE* to *POST*, and 0 otherwise (i.e., above vs. below median liquidity change). For the pooled regression, we implement the White heteroscedasticity consistent estimator.

Model: $R_t = a_0 + a_1POST + b_0E_{t-1} + b_1E_t + b_2E_{t+1} + b_3R_{t+1} + b_4POST \times E_{t-1} + b_5POST \times E_t + b_6POST \times E_{t+1} + b_7POST \times R_{t+1} + a_2POST \times LIQ + b_8POST \times LIQ \times E_{t-1} + b_9POST \times LIQ \times E_t + b_{10}POST \times LIQ \times E_{t+1} + b_{11}POST \times LIQ \times R_{t+1} + u_t$

Variables	Pooled with *LIQ* Dummy
Intercept	0.5458***
	(6.77)
E_t	1.8654*
	(1.68)
E_{t-1}	-1.1903*
	(-1.83)
E_{t+1}	-0.1139
	(-0.43)
R_{t+1}	-0.1388
	(-1.16)
POST	-0.5140***
	(-5.94)
$POST^* E_t$	-1.6415***
	(-4.61)
$POST^* E_{t-1}$	1.1682*
	(1.67)
$POST^* E_{t+1}$	0.6661**
	(2.32)
$POST^* R_{t+1}$	0.1068
	(0.87)
*POST*LIQ*	0.3819***
	(4.34)
$POST^*LIQ^* E_t$	0.5345
	(1.39)
$POST^*LIQ^* E_{t-1}$	-0.6165
	(-0.76)
$POST^*LIQ^* E_{t+1}$	0.4579*
	(1.95)
$POST^*LIQ^* R_{t+1}$	0.0357
	(0.46)
# Observations	1148
F-Stat	7.20
Adjusted R^2	0.0704

***, **, * indicates that the coefficient is significantly different from zero at 1%, 5%, and 10% levels in two- tailed test respectively; the number of observations is less than in Table 2 due to missing daily data.

Table 6 Alternative Explanation: Increase in Institutional Interest

Since the institutional interest explanation implies that the increase in *FERC* is driven by the initiation firms that experienced an increase in institutional interest, we create a dummy variable *INST* that is equal to 1 for firms that had higher increase in percentage of shares owned by institutions from *PRE to POST*, and 0 otherwise (i.e., above versus below median increase in percentage of shares owned by institutions). For the pooled regression, we implement the White heteroscedasticity consistent estimator.

Model: $R_t = a_0 + a_1POST + b_0E_{t-1} + b_1E_t + b_2E_{t+1} + b_3R_{t+1} + b_4POST\acute{}E_{t-1} + b_5POST\acute{}E_t + b_6POST\acute{}E_{t+1} + b_7POST\acute{}R_{t+1} + a_2POST\acute{}INST + b_8POST\acute{}INST\acute{}E_{t-1} + b_9POSTxINST\acute{}E_t + b_{10}POSTxINST\acute{}E_{t+1} + b_{11}POSTxINST\acute{}R_{t+1} + u_t$

Variables	Pooled with *INST* Dummy
Intercept	0.5121***
	(5.93)
E_t	2.7612**
	(2.16)
E_{t-1}	-1.2452*
	(-1.77)
E_{t+1}	-0.3055
	(-0.80)
R_{t+1}	-0.0589
	(-0.53)
POST	-0.4726***
	(-4.95)
$POST^* E_t$	-1.9690
	(-1.47)
$POST^* E_{t-1}$	1.0379
	(1.32)
$POST^* E_{t+1}$	0.8772**
	(2.15)
$POST^* R_{t+1}$	0.0695
	(0.61)
*POST*INST*	0.4315***
	(3.52)
$POST^*INST^* E_t$	-0.3823
	(-0.78)
$POST^*INST^* E_{t-1}$	-0.7857
	(-0.72)
$POST^*INST^* E_{t+1}$	0.4798
	(1.29)
$POST^*INST^* R_{t+1}$	-0.1269
	(-1.36)
# Observations	880
F-Stat	7.16
Adjusted R^2	0.0894

***, **, * indicates that the coefficient is significantly different from zero at 1%, 5%, and 10% levels in two- tailed test respectively; the number of observations is less than in Table 2, because only 440 pairs of firms have information on institutional ownership for the pre and post period.

Table 7 Alternative Explanation: Natural Maturity

In this table we test whether *FERC* improves simply due to the passage of time, i.e. the natural maturity effect. We denote the year of coverage initiation as *t.* Our test is designed to detect the change in *FERC* before and after coverage initiation takes place, namely from *PRE*-1 to *PRE*, and from *POST* to *POST*+1. The t-statistics of the coefficients are in the parentheses. For the pooled regression, we implement the White heteroscedasticity consistent estimator.

		Before Initiation		After Initiation
Variables	Year *PRE* –1	Year *PRE*	Year *POST*	Year *POST*+1
Intercept	0.3956***	0.5686***	0.2227***	0.1231***
	(8.55)	(7.81)	(5.38)	(3.08)
E_t	0.5923***	1.9597***	0.7644***	0.4336**
	(3.00)	(4.06)	(3.21)	(1.96)
E_{t-1}	-0.1248	-1.2123***	-0.5045***	-0.3288
	(-0.94)	(-3.82)	(-1.80)	(-1.54)
E_{t+1}	-0.3471	-0.1053	0.7883***	0.2674***
	(-1.18)	(-0.35)	(3.87)	(2.89)
R_{t+1}	-0.0270	-0.1269*	-0.0184	-0.0816**
	(-0.97)	(-1.77)	(-0.42)	(-2.03)
# Observations	492	492	526	526
F-Stat	3.01	6.94	7.18	6.40
Adjusted R^2	0.0161	0.0462	0.0450	0.0396

***, **, * indicates that the coefficient is significantly different from zero at 1%, 5%, and 10% levels in two- tailed test respectively; the number of observations is less than in Table 2, because of missing information due to firms being too new or being delisted.

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