The Impact of Options Listing Upon the Optimistic Bias in Analysts' Quarterly Earnings Forecasts

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Abstract

We provide evidence that firms' options listing increases the divergence of analysts' opinions and, at the same time, leads to a reduction in the systematic optimistic bias in an important element of the market expectation of earnings, analysts' consensus earnings forecasts. Our contribution is added insight into how the increased divergence of analysts' opinions following options listing drives a reduction of systematic optimistic bias in consensus analysts' forecasts.

Introduction

The purpose of this study is to investigate whether firms' options listing reduces the systematic optimistic bias in analysts' earnings forecasts by *decreasing* the consensus among forecasting analysts. This study integrates the research results reported by DeBondt and Thaler (1990) and Kang, O'Brien, and Sivaramakrishnan (1992) documenting analysts' systematic optimistic earnings forecast bias with the research results reported by Skinner (1990) and Ho (1993) documenting the decrease in the security price response to earnings news following firms' options listing to provide a basis from which to examine the impact of options listing upon analysts' optimistic earnings forecast bias. The results reported herein provide empirical evidence that firms' options listing reduces the systematic optimistic bias in analysts' earnings forecasts, resulting in forecasts better approximating rational expectations (Muth (1960)) earnings forecasts.

Skinner and Ho conjecture that firms' options listing increases the incentives for investors, and, therefore, analysts, to acquire and trade upon additional pre-earnings release information. However, differential information search and processing costs are likely to cause these incentives to have a differential impact upon the behavior of securities analysts, inducing a divergence in their opinions. Empirical results reported by Imhoff and Lobo (1992) indicate that the security price response to firms' earnings news decreases in relation to the divergence of analysts' earnings forecasts, concluding that the divergence of analysts' opinions is a reflection of noisy predisclosure information described in theoretical results provided by

Holtheausen and Verrecchia (1988), Lev (1989), and Choi and Salamon (1990). While abundant theoretical research supports the contention that the security price response to earnings news decreases in relation to the precision of investors' prior expectations (e.g., Verrecchia (1980), Marshall (1980), Robichek and Myers (1966), and Epstein and Turnbull (1980)), the result derived from the Holtheausen and Verrecchia (1988) (hereafter HV), Lev (1989), and Choi and Salamon (1990) (hereafter CS) analyses are especially appealing because, among other reasons, it provides separate security price adjustment implications for sources of predisclosure uncertainty attributable to (1) uncertainty associated with investors' prior expectations, and (2) noise in earnings as a proxy for price-relevant cash flows. These two attributes of earnings forecast uncertainty have different implications for the security price response to earnings news. Results derived in the HV, Lev, and CS analyses suggest that (1) the security price response to earnings news increases in relation to uncertainty associated with investors' prior expectations, and (2) the security price response to firms earnings news decreases in relation to the noise in earnings as a proxy for price-relevant cash flows. Cho and Jung (1991, p.89) reconcile the HV, Lev, and CS results and demonstrate that the implications for the earnings/returns relation are virtually identical.

This research proposes that divergence of analysts' opinions following upon firms' options listing attenuates a "portfolio effect" documented by Conroy and Harris (1987) and manifesting as less biased consensus earnings

forecasts. We employ a qualitative variable variant of the DeBondt and Thaler (1990) and Kang, O'Brien, and Sivaramakrishnan (1992) rational expectations method to examine whether the systematic optimistic bias which they observe in analysts' earnings forecasts decreases subsequent to firms' options listing. Their results provide convincing evidence that analysts' systematic optimistic earnings forecast bias decreases significantly subsequent to firms' options listing. Extant theory suggests that the observed decrease in analysts' systematic optimistic bias subsequent to firms' options listing is consistent with a reduction in the uncertainty associated with investors' prior earnings expectation. Consequently, these results provide empirical evidence that firms' options listing decreases the uncertainty associated with an important element of the market's pre-earnings release expectations, investors prior earnings expectations, and one theoretically consistent explanation for the observed reduction in the security price response to firms' earnings news following upon options listing as reported by Skinner (1990) and Ho (1993).

The Impact of Options Listing upon the Security Price Response Earnings News

The observed association between accounting earnings and security price changes adds credence to the proposition that reported earnings provide investors with relatively good proxies for long run average cash flows to firms. Easton and Zmijewski (1989, p. 118) provide a discussion of the intuition underlying the association between unexpected earnings and unexpected security price changes in terms of unexpected earnings providing investors with a signal regarding prospective changes in investment-related cash flows. Following their rationale, the discounted implications of investment-related cash flow changes signaled by unexpected earnings explain, in part, the observed security price adjustment associated with earnings news. Beaver, Clarke, and Wright (1979), for example, report correlations between earnings forecast errors and risk-adjusted security returns as high as 90% at the portfolio level, although the relation is weaker and exhibits considerable variation at the firm level. Comparing Table 7 and Table 8 from Beaver, Clarke, and Wright (1979, pp. 331-332), the mean Spearman Rank Correlations reported for the ten years examined based upon 25 portfolios per year are 0.7381, 0.6607 (percentage forecast error), and 0.6949 and 0.6902 (standardized forecast error) for the two earnings forecast models employed. The mean Spearman Rank Correlations reported for the ten years examined based upon firm-specific observations are 0.3783, 0.3161 (percentage forecast error), and 0.3284, and 0.3303 (standardized forecast error) for the two earnings forecast models used. Because of the degree of variation in the earnings/returns relation at the firm level, an extensive body of literature

has developed explaining the cross-sectional variation in the earnings/returns relation. In addition to firms' options listing (discussed subsequently), some of the factors which have been identified as important in describing the cross-sectional variation in the earnings/returns relation are: (1) earnings persistence -- positive association (Kormendi and Lipe (1987), Easton and Zmijewski (1989), Collins and Kothari (1989), and Lipe (1990)); systematic risk -- negative association (Collins and Kothari (1989), and Easton and Zmijewski (1989)); earnings growth -- positive association (Collins and Kothari (1989)); and firm size -- negative association (Grant (1980), Atiase (1985, 1987), Freeman (1987), and Collins, Kothari, and Rayburn (1987)).

Extant research indicates that options trading is one factor explaining the cross-sectional variation in the security price response to firms' earnings news. That is, whether a firm has options traded is one important factor determining the magnitude of the security price response to its earnings releases, and, therefore, the information content of its earnings releases to investors. Black (1975), Cox and Rubinstein (1985), Diamond and Verrecchia (1987), and Ross (1976) suggest that options listing increases the amount of pre-earnings release price-relevant information by providing incentives for private information acquisition which are not available from trading equity securities alone. This reduces the amount of new information and, therefore, reduces the security price response accompanying the earnings news release. For example, options can be used to obtain a leveraged position in a security by taking a long position in call options which involves less cash outlay and is less restricted (in dollar amount) than borrowing to purchase the underlying security. Options trading can also be used to circumvent the "Uptick Rule" in securities trading, because no similar restriction exists for options trading. A sell order is marked a short sale if the seller does not own the shares offered for sale. The tick test states that a short sale may be executed only following a plus tick or a zero plus tick. A price increase is called a plus tick. A zero plus tick is plus tick followed by two consecutive trades at the same price. See Schwartz (1988, p.50) for a thorough discussion of trading constraints.

Empirical research such as Skinner (1990) and Ho (1993) establishes that firms' options listing decreases the security price response to earnings news, suggesting that the incentives to acquire incremental pre-earnings release information provided by options listing reduces the uncertainty associated with market's prior expectation of earnings. Using longitudinal earnings response regression analysis, Skinner documents a significant reduction in the security price response to firms' earnings news subsequent to their options listing. Ho employs cross-sectional

analyses to obtain evidence that (1) options listed firms have significantly smaller security price responses to earnings releases, and (2) security prices anticipate the relevant information conveyed by firms' earnings news at significantly earlier points in time (relative to the earnings release) for options listed firms than non-options listed firms. The intuition underlying the observed decrease in the security price response (at the time of the earnings release) to options listed firms' earnings releases is that the precision of the market expectation of earnings (for which analysts are one important source) is greater for options listed firms. However, existing literature provides little conjecture or evidence as to how the improved precision of pre-earnings release earnings expectation occurs. While the Skinner (1990, p.204) and Ho (1993, p.373) studies provide evidence of an increased level of analyst following pursuant to firms' options listing, the impact of this increased interest on the properties of analysts' earnings forecasts is not clear. Skinner (1990, p.199) reports finding no significant increase in analysts' earnings forecast accuracy (i.e., reduction in absolute percentage forecast errors) subsequent to firms' options listing.

Table 1 shows the mean number of analysts contributing forecasts and mean standard deviation of analysts earnings forecasts before and after firms' options listing, providing an indication of whether an increase in analyst following subsequent to firms options listing is accompanied by an increase in the standard deviation of their earnings

forecasts. We also examined the impact of options listing upon analysts' forecast accuracy employing several measures of absolute and squared forecast errors. Like Skinner (1990), we found no significant difference at conventional confidence levels in mean forecast errors before and after firms' options listing. Our forecast accuracy results are not shown, as they are established elsewhere in the literature. Two-tailed t-tests of the null hypotheses of equality of means are rejected at the $\alpha=0.05$ confidence level for both mean number of analysts contributing forecasts and mean standard deviation of analysts earnings forecasts. In addition, we regressed the standard deviation of analysts' earnings forecasts on the number of analysts, using a dummy variable D_i to distinguish pre-options listing and post-options listing quarters. Using ordinary least-squares techniques, the following estimated regression equation resulted:

$$STD_{it} = 0.5311 + 0.0067 \cdot NUM_{it} + 0.0320 \cdot D_{it} \cdot NUM_{it} + u_{it} \quad (1)$$

where:

- STD_{it} = standard deviation of analysts' firm i earnings per share forecasts
 NUM_{it} = number of analysts contributing firm i earnings per share forecasts
 D_{it} = 0 if pre-options listing quarter, 1 otherwise
 u_{it} = normal distributed error term with zero mean and constant variance

Table 1
Summary Statistics for Number of Analysts Contributing Opinions and
Standard Deviation of Analysts' Earnings Forecasts Before and After Firms Options Listing ^a

Variable / Time Periods	Mean Value	Std. Deviation	Max Value	Min Value
Number of Analysts Contributing Earnings Forecasts (NUM_{it}):				
Pre-Options Listing Periods	1.510158	0.903367	6.000000	1.000000
Post-Options Listing Periods	1.733480	1.149672	8.000000	1.000000
t-test Null Hypothesis of Equality of Means (Unequal Variances ^b): $t = -3.2391$ ($p = 0.0012$) †				
Standard Deviation of Analysts' Earnings Forecasts (σ_{Ait}):				
Pre-Options Listing Periods	0.476618	0.367017	2.190000	0.010000
Post-Options Listing Periods	0.592739	0.428107	2.450000	0.010000
t-test Null Hypothesis of Equality of Means (Unequal Variances ^b): $t = -6.4949$ ($p = 0.0001$) †				

- †: The null hypothesis of no difference in means rejected at the $\alpha=0.05$ confidence level employing two-tailed tests.
a: The statistics shown are after deletion to get equal numbers of pre-options listing and post-options listing observations.
b: The null hypothesis of no difference in variances is rejected at the $\alpha=0.05$ confidence level employing two-tailed F-tests. The tests are performed for the purpose of calculating pooled variances for use in t-tests.

The pre-options listing coefficient (-0.0067) is not significantly different from zero and the post-options listing coefficient (0.0320-0.0067=0.0253) is significantly greater than zero ($t=2.143$) at the $\alpha=0.05$ confidence level. These results suggest that the earnings forecasts of the enlarged analyst group contributing opinions following upon firms' options listing are extreme observations, lying further from the mean than before options listing and, as such, are likely forecasts incorporating incremental information made available due to additional and more wide spread analyst interest after firms' options listing. This research contends that the marginal contributions of this enlarged analyst group attenuates the "portfolio effect" in consensus forecasts observed by Conroy and Harris, in that the combined forecast (i.e., the mean forecast in this case) is less biased due to the disparity of the individual forecasts. We investigate the impact of firms' options listing upon the extent to which forecasted earnings are realized as actual earnings, believing that the observed reduction in systematic forecast bias accompanying firms' options listing corresponds to a sufficient reduction in earnings forecast uncertainty (discussed further in the following section) to provide one theoretically consistent explanation for the observed reduction in the security price response to earnings news in periods following upon firms' options listing.

Bias in Analysts' Earnings Forecasts

Fried and Givoly (1982) and O'Brien (1987) were among the first researchers to observe that security analysts exhibit a tendency, on average, to be optimistically biased in their earnings forecasts. That is, analysts' forecasted earnings consistently overestimate actual earnings. Consequently, less than 100% of analysts' forecasted earnings per share are realized as actual earnings per share reported by firms. More recently, DeBondt and Thaler (1990) document security analysts' systematic optimistic bias in their earnings forecasts. Other researchers observe a positive first-order autocorrelation in analysts' earnings forecast errors, indicative of analysts' systematic under reaction to the most recent past forecast error in their current forecasts. Ali, Klein, and Rosenfeld (1992), Lys and Sohn (1990), Klein (1990), and Abarbanell (1991) observe a systematic under reaction of analysts' current earnings forecasts to the implications of past equity security returns for future earnings. Abarbanell and Bernard (1992) examine the properties of analysts' earnings forecast errors as one explanation for the post earnings announcement equity security price drift phenomenon. Schipper (1991) provides a review of the literature regarding the properties of analysts' earnings forecasts. Regressing actual earnings changes on forecasted earnings changes, DeBondt and Thaler (1990) observe estimated slope coefficients roughly one-half the

value hypothesized under rational expectations. Table 1 from DeBondt and Thaler (1990, p.54) shows that the ordinary least-squares regression of actual earnings changes (deflated by the standard deviation of analysts' forecasts) onto forecasted earnings changes (deflated by the standard deviation of analysts' forecasts) produces the following estimated regression equation:

$$\frac{AC_{it}}{\sigma_{A_{it}}} = -0.094 + 0.648 \cdot \frac{FC_{it}}{\sigma_{A_{it}}} + v_i \quad (2)$$

Where:

AC_{it} = $EPS_{it} - EPS_{it-1}$

FC_{it} = $\bar{F}_{it} - EPS_{it-1}$

$\bar{F}_{it}$ = mean analysts' forecasts

σ_A = standard deviation of analysts' forecasts

v_{it} = regression residual

A one-tailed t-test of the slope coefficient (0.648) indicates that it is significantly less than unity ($t=-21.7$) at the $\alpha=0.05$ confidence level, and a one-tailed t-test of the intercept (-0.094) indicates that it is significantly less than zero ($t=-3.7$) at the $\alpha=0.05$ confidence level. Because the estimated slope coefficient is significantly less than unity (as hypothesized under the null) and the estimated intercept is not positive, analysts appear to overestimate forecasted earnings changes. The extent to which forecasted earnings changes are realized in actual earnings changes is less than one would expect under rational expectations. Kang, O'Brien, and Sivaramakrishnan (1992) expand upon the DeBondt and Thaler (1990) results by (1) providing additional evidence regarding the impact of the horizon effect observed by DeBondt and Thaler (1990) upon analysts' optimistic forecast bias, and (2) identifying a price effect in analysts' systematic optimistic forecast bias.

The observed analysts' optimistic bias introduces noise into one important element of the market's prior expectations of earnings which may be measurably reduced by the incremental incentives afforded investors and analysts to acquire and process additional firm-specific information accompanying firms' options listing. Intuitively, rational expectations regressions of actual changes on forecasted changes (i.e., employing the Muth (1960) (0, 1) null hypothesis) for which the intercept is constrained to equal zero and the slope coefficient is constrained to unity will produce larger sums of squared errors employing firm-specific earnings forecasts which exhibit analysts' optimistic bias than employing firm-specific earnings forecasts which better approximate rational expectations earnings forecasts. In this sense, market expectations of earnings, taking optimistically

biased analysts' forecasts as rational expectations forecasts, have more uncertainty associated with them than expectations based upon less biased analysts' forecasts. As a result, one consequence of firms' options listing may be a marked reduction in the systematic optimistic bias in analysts' earnings forecasts, an increase in the precision of an important component of the markets' expectation of firms' earnings which more than compensates for the opposite effect of increased disparity (standard deviation) of analysts' forecasts, and the observed decrease in the security price response to firm' earnings news.

Sample Firms and Firm Specific Data

The firms used in this research meet the following data availability criteria:

They had options listed between the years 1984 and 1987 (inclusive) on one of the following five U.S. options exchanges: (1) The American Stock Exchange; (2) The Chicago Board of Options Exchange; (3) The New York Stock Exchange; (4) The Pacific Stock Exchange; or (5) The Philadelphia Stock Exchange; and

They have non-missing quarterly primary earnings-per-share (before discontinued operations and extraordinary items), quarterly mean and median analysts' earnings forecasts, and standard deviation of mean analysts' earnings forecasts in the 1991 Institutional Brokers Estimate Service (IBES) data files from 1984 and extending through 1990.

A list of 793 firms identified as having options listed on one of the five options exchanges listed above was obtained courtesy of the Chicago Board of Options Exchange. A total of 532 (63%) of these 793 companies are carried in the IBES data files during the study period. The firm-quarter observations are truncated, if necessary, to result in equal numbers of pre- and post-options listing observations, although the number of observations is not the same for each firm. These requirements resulted in a total of 1900 firm-quarter observations for analysis.

Options Listing and the Rational Expectations Regression

This research estimates the rational expectations regression employed by DeBondt and Thaler (1990) and Kang, O'Brien, and Sivaramakrishnan (1992) using qualitative variables to distinguish pre- and post-options listing observations. We regress measures of actual earnings and actual earnings changes onto analysts' forecasted earnings and forecasted earnings changes, employing these qualitative variables, to investigate

whether analysts' optimistic bias decreases subsequent to firms' options listing. The DeBondt and Thaler (1990) and Kang, O'Brien, and Sivaramakrishnan (1992) results indicated that the slope coefficient obtained by regressing actual earnings changes on forecasted earnings changes is less than unity at conventional significance levels, indicating that analysts are (on average) optimistically biased in forecasting earnings changes (i.e., only a portion of forecasted earnings changes are realized as actual earnings changes). Ordinary least-squares estimation techniques are applied to pooled quarterly earnings data. Using Equation (3) through Equation (8), we examine whether, controlling for intercept changes as well, the slope of the rational expectations regression increases toward unity subsequent to firms' options listing by employing the qualitative variable D_i . Each firm-quarter observation is assigned to one of two strata employing a qualitative variable (D_i). The first stratum contains the pre-options listing firm-quarter observations ($D_i=0$); the second stratum contains the post-options listing firm-quarter observations ($D_i=1$). A percentage rank transformation is employed for the firm-quarter observation values. The rank transformation substitutes the value of the variable with the value of its sample rank. This technique provides additional confidence in the statistical results because: (1) the results are independent of assumptions regarding the distribution of the data (i.e., it is a distribution free technique); (2) the transformation generalizes the functional form of the regression equation, since it provides the same results as all ordinal transformations; and (3) it mitigates the impact of measurement error, outliers, and residual heteroscedasticity on the regression results. Rank ties are replaced with the mean rank value, rather than the high or low value. Percentile ranks are used rather than the raw numerical ranks because they are independent of the maximum rank, and, therefore, more general. Percentile ranks express the variable rank as a fraction of the maximum raw rank. Consequently, the data range from zero to one. See Iman and Conover (1979) for added details regarding the uses of rank transformations in regression analysis. The earnings coefficient for pre-options listing firm-quarter observations is δ_{2j} ($\forall j \in (3, \dots, 8)$), and is expected to be significantly greater than zero but significantly less than one at the $\alpha=0.05$ confidence level, as observed by DeBondt and Thaler (1990) and Kang, O'Brien, and Sivaramakrishnan (1992). The earnings coefficient for the post-options listing firm-quarter observations is $\delta_{2j} + \delta_{3j}$ ($\forall j \in (3, \dots, 8)$), and is expected to be significantly greater than δ_{2j} , and perhaps not significantly different from unity, at the $\alpha=0.05$ confidence level. The regression technique used is suggested by Johnston (1984, p.227). With this design, the regression coefficient for the qualitative variables represent only the incremental impact of the

change in the regression regimes upon the regression coefficients. An increase in the coefficient toward unity for the post-options listing firm-quarter observations provides an indication that firms' options listing improves investors' prior earnings expectations by reducing the degree of observed systematic optimistic bias in forecasted earnings and earnings changes relative to actual earnings and earnings changes (i.e., a greater portion of forecasted earnings and earnings changes are realized as actual earnings and earnings changes subsequent to firms' options listing). The result provides direct evidence that firms' options listing alters firms' information environment by improving analysts' and other investors' prior expectations, and, as a result, reducing the observed security price reaction to firms' earnings releases.

In Equation (3) through Equation (5) we regress actual earnings levels on analysts' forecasted mean (or average) earnings levels. In Equation (3) we regress (non-deflated) actual earnings onto (non-deflated) forecasted earnings. In Equation (4) we regress actual earnings deflated by the standard deviation of the earnings forecast onto forecasted earnings deflated by the standard deviation of forecasted earnings. In addition to the standard deviation of the earnings forecast deflator, we also estimated the qualitative variable rational expectations regression employing a security price deflator. DeBondt and Thaler (1990) also tested both deflators. Like DeBondt and Thaler (1990, p.54 (fn.1)), we obtained virtually identical results using the security price deflator as for the standard deviation of the earnings forecast deflator. Because the similarity of the results using these two deflators is documented by DeBondt and Thaler (1990), we do not report the security price deflator results. In Equation (5) we regress actual earnings deflated by the absolute value of previous actual earnings on forecast earnings deflated by the absolute value of the previous quarters' actual earnings. We also estimated Equation (5) and Equation (6), Equation (7), and Equation (8) (discussed subsequently) using seasonal quarter earnings (i.e., EPS_{it-4}) and earnings differences rather than adjacent quarter earnings (i.e., EPS_{it-1}), and found that seasonal earnings and earnings differences produced virtually identical results. This result is also established in prior research (e.g., Kang, O'Brien, and Sivaramakrishnan (1992) and Abarbanell and Bernard (1992)). We use last actual earnings as a deflator for quarter t actual and forecasted earnings and

earnings changes following Elton, Grueber, and Gultekin (1984). The μ_{it} , v_{it} , and ξ_{it} are error terms with zero mean and constant variance. The severity of the heteroscedasticity of the residuals of the pooled qualitative variable rational expectations regressions is evaluated using the White (1980, pp. 824-825) χ^2 test. The White test regresses the earnings response regression squared residuals onto the independent variables and their cross-products (eliminating duplicates). The product of the R^2 of this regression and the number of observations used is a test statistic which is approximately $\chi^2_{\frac{1}{2}(k+k^2)}$ (where k is the number of independent variables) distributed under the null hypothesis of homoscedasticity (consequently, one wants to not reject the null hypothesis using this test).

$$EPS_{it} = \delta_{03} + \delta_{13} \cdot D_{it} + \delta_{23} \cdot \bar{F}_{it} + \delta_{33} \cdot D_{it} \cdot \bar{F}_{it} + \mu_{it} \quad (3)$$

$$\frac{EPS_{it}}{\sigma_{F_{it}}} = \delta_{04} + \delta_{14} \cdot D_{it} + \delta_{24} \cdot \frac{\bar{F}_{it}}{\sigma_{F_{it}}} + \delta_{34} \cdot D_{it} \cdot \frac{\bar{F}_{it}}{\sigma_{F_{it}}} + v_{it} \quad (4)$$

$$\frac{EPS_{it}}{|EPS_{it-1}|} = \delta_{05} + \delta_{15} \cdot D_{it} + \delta_{25} \cdot \frac{\bar{F}_{it}}{|EPS_{it-1}|} + \delta_{35} \cdot D_{it} \cdot \frac{\bar{F}_{it}}{|EPS_{it-1}|} + \quad (5)$$

When the null hypothesis of homoscedasticity is rejected at the $\alpha=0.05$ confidence level, the reported t-statistics pertaining to each of the qualitative variable rational expectations regressions are calculated with consistent variance estimates.

In Equation (6) through Equation (8) we regress actual earnings changes on forecasted earnings changes. In Equation (6) we regress (non-deflated) actual earnings

$$AC_{it} = \delta_{06} + \delta_{16} \cdot D_{it} + \delta_{26} \cdot FC_{it} + \delta_{36} \cdot D_{it} \cdot FC_{it} + o_{it} \quad (6)$$

$$\frac{AC_{it}}{\sigma_{F_{it}}} = \delta_{07} + \delta_{17} \cdot D_{it} + \delta_{27} \cdot \frac{FC_{it}}{\sigma_{F_{it}}} + \delta_{37} \cdot D_{it} \cdot \frac{FC_{it}}{\sigma_{F_{it}}} + \pi_{it} \quad (7)$$

$$\frac{t}{-1} = \delta_{08} + \delta_{18} \cdot D_{it} + \delta_{28} \cdot \frac{FC_{it}}{|EPS_{it-1}|} + \delta_{38} \cdot D_{it} \cdot \frac{FC_{it}}{|EPS_{it-1}|} \quad (8)$$

changes on (non-deflated) forecasted earnings changes. In Equation (7) we regress actual earnings changes deflated by the standard deviation of the earnings forecast on forecasted earnings changes deflated by the standard deviation of forecasted earnings. In Equation (8) we regress actual earnings changes deflated by the absolute value of previous actual earnings on forecast earnings changes deflated by the absolute value of previous actual earnings. The ϵ_i , π_i , and ϖ_i are error terms with zero mean and constant variance.

Empirical Results

Table 2 shows summary statistics for the actual earnings and forecasted earnings data used to estimate regression Equation (3) through Equation (8) for combined (i.e., both pre- and post-options listing periods) observations, pre-options listing observations, and post-options listing observations. Because one important reason for using the rank transformation on the data is to mitigate the impact of outliers, extreme observations are not deleted. Two-tailed t-tests (not shown) of the null hypothesis of the equality of means (unequal variances) for pre-options listing observations and post-options listing observations are not rejected at any conventional confidence level for any of the variables shown.

Equation (3) regresses (non-deflated) actual earnings levels on (non-deflated) forecasted earnings levels. While the intercept (δ_{03}) is significantly different from zero at the $\alpha=0.05$ confidence level, the coefficient for the qualitative variable indicating shifts in the intercept subsequent to firms' options listing (δ_{13}) is not significantly different from zero at the $\alpha=0.05$ confidence level. Indeed, δ_{23} is significantly greater than zero and significantly less than one at the $\alpha=0.05$ confidence level, as reported by DeBondt and Thaler (1990) and Kang, O'Brien, and Sivaramakrishnan (1992). However, for Equation (3), the coefficient for the post-options listing firm-quarter observations ($\delta_{23} + \delta_{33}$) is not significantly greater than the coefficient for the pre-options listing stratum (δ_{23}) at the $\alpha=0.05$ confidence level (although it is significantly greater than the coefficient for the pre-options listing firm-quarter observations at the $\alpha=0.10$ confidence level), and is significantly less than unity at the $\alpha=0.05$ confidence level. Consequently, the results for Equation (3) do not support the contention that firms' options listing increases the extent to which forecasted earnings changes are realized in actual earnings changes.

Table 3 shows the results of the qualitative variable rational expectations regression shown in Equation (4), regressing actual earnings levels deflated by the standard deviation of forecasted earnings on forecasted earnings

levels deflated by the standard deviation of forecasted earnings. As in Equation (3), δ_{24} is significantly greater than zero at the $\alpha=0.05$ confidence level and significantly less than one at the $\alpha=0.05$ confidence level, as reported by DeBondt and Thaler (1990) and Kang, O'Brien, and Sivaramakrishnan (1992). The coefficient for the post-options listing stratum ($\delta_{24} + \delta_{34}$) is significantly greater than the coefficient for the pre-options listing firm-quarter observations (δ_{24}) at the $\alpha=0.05$ confidence level but is still significantly less than unity at the $\alpha=0.05$ confidence level. As a result, the results for Equation (4) support the contention that firms' options listing decreases the systematic optimistic bias in analysts' earnings forecasts observed by DeBondt and Thaler (1990) and Kang, O'Brien, and Sivaramakrishnan (1992).

Table 3 shows the results of the qualitative variable rational expectations regression shown in Equation (5), which regresses actual earnings levels deflated by the previous quarters' actual earnings on forecasted earnings levels deflated by the previous quarters' actual earnings. The coefficient for forecasted earnings levels, δ_{25} , is significantly greater than zero at the $\alpha=0.05$ confidence level and significantly less than one at the $\alpha=0.05$ confidence level, as reported by DeBondt and Thaler (1990) and Kang, O'Brien, and Sivaramakrishnan (1992). Again, the coefficient for the post-options listing firm-quarter observations ($\delta_{25} + \delta_{35}$) is significantly greater than the coefficient for the pre-options listing stratum (δ_{25}) at the $\alpha=0.05$ confidence level but is still significantly less than unity at the $\alpha=0.05$ confidence level. Consequently, the results for Equation (5) support the contention that firms' options listing increases the extent to which forecasted earnings changes are realized in actual earnings changes.

Table 3 shows the results of the qualitative variable rational expectations regression shown in Equation (6), which regresses (non-deflated) raw actual earnings changes on (non-deflated) raw forecasted earnings changes. As expected from prior research, the coefficient for forecasted earnings changes, δ_{26} , is significantly greater than zero at the $\alpha=0.05$ confidence level and significantly less than one at the $\alpha=0.05$ confidence level. As with the previous regression equations, the coefficient for the post-options listing stratum ($\delta_{26} + \delta_{36}$) is significantly greater than the coefficient for the pre-options listing stratum (δ_{26}) at the $\alpha=0.05$ confidence level but is still significantly less than unity at the $\alpha=0.05$ confidence level. Consequently, the results for Equation (6) support the contention that firms' options listing decreases the systematic optimistic bias in analysts' earnings forecasts observed by DeBondt and Thaler (1990) and Kang, O'Brien, and Sivaramakrishnan (1992).

Table 2
Summary Statistics for Earnings Data Used to Estimate Rational Expectations Regressions

Dependent Variable	Time Period	Actual Forecast	Arithmetic Mean	Std Deviation	Max Value	Min Value
EPS_{it}	Overall	Actual: Forecast:	0.4928 0.5644	0.8887 0.5025	5.7400 2.6200	-20.6700 -4.8300
	Pre-Options Period	Actual: Forecast:	0.4318 0.5095	0.9217 0.4441	2.8500 2.3300	-20.6700 -4.7600
	Post-Options Period	Actual : Forecast:	0.5596 0.6229	0.8466 0.5521	5.7400 2.6200	-10.0100 -4.8300
$\frac{EPS_{it}}{\sigma_{F_{it}}}$	Overall	Actual: Forecast:	1.1292 1.2169	1.4082 0.5144	21.7500 10.0000	-34.8571 1.0000
	Pre-Options Period	Actual: Forecast:	1.1228 1.2162	1.3301 0.4797	21.7500 7.0000	-6.2800 1.0000
	Post-Options Period	Actual: Forecast:	1.1362 1.2177	1.4901 0.5489	14.0384 10.0000	-34.8571 1.0000
$\frac{EPS_{it}}{ EPS_{it-1} }$	Overall	Actual: Forecast:	1.1604 0.9397	11.4564 5.2308	379.0000 52.0000	-203.0000 -176.0000
	Pre-Options Period	Actual: Forecast:	0.7946 0.8376	7.4804 6.6220	64.3333 40.0000	-203.0000 -176.0000
	Post-Options Period	Actual: Forecast:	1.5594 1.0502	14.6013 3.0839	379.0000 52.0000	-46.0000 -45.6666
AC_{it}	Overall	Actual: Forecast:	0.0046 0.0678	1.0996 0.9084	21.1400 21.1800	-18.3600 -8.5800
	Pre-Options Period	Actual: Forecast:	-0.0194 0.0490	0.9795 0.6899	10.7200 9.4400	-18.3600 -8.5800
	Post-Options Period	Actual: Forecast:	0.0310 0.0881	1.2170 1.0970	21.1400 21.1800	-11.7700 -6.5900
$\frac{AC_{it}}{\sigma_{F_{it}}}$	Overall	Actual: Forecast:	-0.0961 -0.0113	4.7979 4.3472	105.7000 105.9000	-105.5000 -100.5000
	Pre-Options Period	Actual: Forecast:	-0.1363 -0.0446	3.5084 3.0262	41.7500 21.0000	-76.4000 -71.0000
	Post-Options Period	Actual: Forecast:	0.0521 0.0244	5.8976 5.4243	105.7000 105.9000	-105.5000 -100.5000
$\frac{AC_{it}}{ EPS_{it-1} }$	Overall	Actual: Forecast:	0.3177 0.5387	11.4531 5.2033	232.5000 177.0000	-378.0000 -81.0000
	Pre-Options Period	Actual: Forecast:	0.5194 0.6030	7.4651 6.5964	204.0000 177.0000	-63.3333 -81.0000
	Post-Options Period	Actual: Forecast:	0.0977 0.4690	14.6117 3.0484	232.5000 51.0000	-378.0000 -9.1063

Table 3
Results of Qualitative Variable Rational Expectations Regression Analyses Assessing the Impact of Firms' Options Listing Upon the Systematic Optimistic Bias in Analysts' Earnings Forecasts

Rational Expectations Regression ¹	Adjusted R ² Observations	Whites's $\chi^2_{1/(k+k^2)}$ (p-value)	δ_{0j} Coefficient t-stat.H ₀ : $\delta_{0j}=0$	δ_{1j} Coefficient t-stat.H ₀ : $\delta_{1j}=0$	δ_{2j} Coefficient ^a t-stat.H ₀ : $\delta_{2j}\leq 0$	δ_{3j} Coefficient ^b t-stat.H ₀ : $\delta_{3j}\leq 0$
Equation (j=3)	0.7947 (1900)	9.2660 (0.0989)	5.6171 (6.897)†	-0.9191 (-0.724)	0.8831 (60.392)□	0.0317 (1.491)
Equation (j=4)	0.0917 (1900)	148.4941 (0.0000)‡	35.0798 (20.893)†	-2.1950 (-0.889)	0.2698 (8.1347)□	0.0897 (1.831)■
Equation (j=5)	0.6520 (1900)	14.8400 (0.0111)‡	10.5678 (9.896)†	-2.7706 (-1.899)†	0.7724 (38.620)□	0.0801 (3.027)■
Equation (j=6)	0.6424 (1900)	40.3373 (0.0000)‡	12.3683 (11.372)†	-4.1084 (-2.935)†	0.7509 (37.545)□	0.1037 (3.919)■
Equation (j=7)	0.6424 (1900)	31.5405 (0.0000)‡	11.3332 (10.456)†	-2.8300 (-1.999)†	0.7580 (37.900)□	0.0901 (3.185)■
Equation (j=8)	0.6517 (1900)	51.7692 (0.0000)‡	10.9074 (11.647)†	-2.1072 (-1.629)	0.7764 (44.825)□	0.0710 (2.683)■

‡: Whites's $\chi^2_{1/(k+k^2)}$ test rejected at the $\alpha=0.05$ confidence level. t-statistics shown are calculated using covariance consistent estimates.

†: The null hypothesis that coefficient is equal to zero is rejected at the $\alpha=0.05$ level using two-tailed t-test.

□: The null hypothesis that $\delta_{2j} \leq 0$ ($\forall j \in (3, \dots, 8)$) is rejected at the $\alpha=0.05$ confidence level using one-tailed t-test.

■: The null hypothesis that $\delta_{2j} \leq \delta_{2j} + \delta_{3j}$ ($\forall j \in (3, \dots, 8)$) is rejected at the $\alpha=0.05$ level using one-tailed t-test.

a: The null hypothesis that $\delta_{2j} \geq 1$ is rejected at the $\alpha=0.05$ confidence level using one-tailed covariance consistent adjusted (where appropriate) t-tests (not shown in Table 3) $\forall j \in (3, \dots, 8)$. The six p-values for these tests are entirely zeros to four decimal places.

b: The null hypothesis that $\delta_{2j} + \delta_{3j} \geq 1$ is rejected at the $\alpha=0.05$ confidence level using one-tailed covariance consistent adjusted (where appropriate) t-tests (not shown in Table 3) $\forall j \in (3, \dots, 8)$. The six p-values for these tests are entirely zeros to four decimal places.

1: The six regression equations used to obtain the results shown in Table 3 are presented below:

$$EPS_{it} = \delta_{03} + \delta_{13} \cdot D_{it} + \delta_{23} \cdot \bar{F}_{it} + \delta_{33} \cdot D_{it} \cdot \bar{F}_{it} + \mu_{it}$$

$$\frac{EPS_{it}}{\sigma_{F_{it}}} = \delta_{04} + \delta_{14} \cdot D_{it} + \delta_{24} \cdot \frac{\bar{F}_{it}}{\sigma_{F_{it}}} + \delta_{34} \cdot D_{it} \cdot \frac{\bar{F}_{it}}{\sigma_{F_{it}}} + v_{it}$$

$$\frac{EPS_{it}}{[EPS_{it-1}]} = \delta_{05} + \delta_{15} \cdot D_{it} + \delta_{25} \cdot \frac{\bar{F}_{it}}{[EPS_{it-1}]} + \delta_{35} \cdot D_{it} \cdot \frac{\bar{F}_{it}}{[EPS_{it-1}]} + \xi_{it}$$

$$AC_{it} = \delta_{06} + \delta_{16} \cdot D_{it} + \delta_{26} \cdot FC_{it} + \delta_{36} \cdot D_{it} \cdot FC_{it} + o_{it}$$

$$\frac{AC_{it}}{\sigma_{F_{it}}} = \delta_{07} + \delta_{17} \cdot D_{it} + \delta_{27} \cdot \frac{FC_{it}}{\sigma_{F_{it}}} + \delta_{37} \cdot D_{it} \cdot \frac{FC_{it}}{\sigma_{F_{it}}} + \pi_{it}$$

$$\frac{AC_{it}}{[EPS_{it-1}]} = \delta_{08} + \delta_{18} \cdot D_{it} + \delta_{28} \cdot \frac{FC_{it}}{[EPS_{it-1}]} + \delta_{38} \cdot D_{it} \cdot \frac{FC_{it}}{[EPS_{it-1}]} + \varpi_{it}$$

Table 3 shows the results of the qualitative variable rational expectations regression shown in Equation (7), which regresses actual earnings changes deflated by the standard deviation of analysts' forecasted earnings on forecasted earnings changes deflated by the standard deviation of analysts' forecasted earnings. As observed by DeBondt and Thaler (1990) and Kang, O'Brien, and Sivaramakrishnan (1992), the coefficient for forecasted earnings changes, δ_{27} , is significantly greater than zero at the $\alpha=0.05$ confidence level and significantly less than one at the $\alpha=0.05$ confidence level. As expected, the coefficient for the post-options listing stratum ($\delta_{27} + \delta_{37}$) is significantly greater than the coefficient for the pre-options listing stratum (δ_{27}) at the $\alpha=0.05$ confidence level but is still significantly less than unity at the $\alpha=0.05$ confidence level. Consequently, the results for Equation (7) support the contention that analysts' forecasted earnings changes are realized in actual earnings changes to a greater extent following firms' options listing.

Table 3 shows the results of the qualitative variable rational expectations regression shown in Equation (8), which regresses actual earnings changes deflated by the previous quarters' actual earnings on forecasted earnings changes deflated by the previous quarters' actual earnings. As expected, the coefficient for forecasted earnings changes, δ_{28} , is significantly greater than zero and significantly less than one at the $\alpha=0.05$ confidence level. Again, the coefficient for the post-options listing stratum ($\delta_{28} + \delta_{38}$) is significantly greater than the coefficient for the pre-options listing stratum (δ_{28}) but is still significantly less than unity at the $\alpha=0.05$ confidence level. Consequently, the results for Equation (8) are consistent with the proposition that firms' options listing decreases analysts' systematic optimistic bias documented by DeBondt and Thaler (1990) and Kang, O'Brien, and Sivaramakrishnan (1992).

To summarize the results shown in Table 3 for the qualitative variable rational expectations regression shown in Equation (3) through Equation (8), each of the regressions (1) replicates the results reported by DeBondt and Thaler (1990) and Kang, O'Brien, and Sivaramakrishnan (1992) demonstrating that only a portion of analysts' forecasted earnings changes are realized as actual earnings changes, and (2) provides evidence supporting the contention that firms' options listing increases the portion of analysts' forecasted earnings which are realized as actual earnings (i.e., the systematic optimistic bias in analysts' mean earnings forecasts is reduced). The later result is not too surprising considering the increase in analyst following subsequent to firms' options listing. It is, however, surprising that, following upon an increase in analysts' following, we find

that (1) no significant increase in analysts' forecast accuracy is observed, and (2) that the standard error of analysts' earnings forecasts increases subsequent to firms' options listing. However, insight may be gained from considering the results reported by Conroy and Harris (1987), suggesting that combinations of disparate forecasts may result in improved consensus forecasts. For options listed firms, the increase in divergence of analysts' earnings forecasts subsequent to their options listing may be exactly the force reducing the optimistic bias in analysts' earnings forecasts, reducing the uncertainty associated with an important element of the market's expectation of earnings, and reducing the security price changes associated with earnings.

Conclusion

Extant research documents an observed reduction in the security price response to firms' earnings news subsequent to their options listing, conjecturing that, because options trading provides added incentives for investors to acquire firm-specific information, options listing improves the precision of investors expectations prior to firms' earnings releases. However, existing research provides little insight as to how this improvement may occur. The purpose of this study is to provide more direct evidence that options listing improves investors' prior expectations. Employing a qualitative variable variant of the rational expectations regression employed by DeBondt and Thaler (1990) and Kang, O'Brien, and Sivaramakrishnan (1992) to document an observed optimistic bias in analysts' earnings forecasts, we document that analysts' forecasted earnings are realized in actual earnings to a significantly greater extent subsequent to firms' options listing. We consider these results to be convincing evidence that firms' options listing significantly reduces the uncertainty associated with an important element of the market's expectation of earnings, analysts' consensus earnings forecasts, an interpretation which is theoretically consistent with the observed reduction in the security price response to firms' earnings news subsequent to their options listing reported in the literature.

Suggestions for Future Research

Future research regarding composite forecasts will provide additional insight into the process by which consensus forecasts are formed, and particularly the phenomenon whereby increasingly divergent forecasts, when combined into composite forecasts, produce earnings forecasts with smaller forecast errors. Two aspects of this "portfolio effect" which may be investigated in future research are (1) the marginal contribution of outlying forecasts to forecast accuracy, and (2) the impact of combining functional form, e.g., linear vs. nonlinear

combinations of forecasts, upon forecast accuracy. Inquiries along either of these avenues will provide accounting and financial economics researchers and practitioners with added insight into the institutional processes by which accounting earnings are impounded in equity security prices.

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