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ScienceDirect
Journal of Accounting and Economics 43 (2007) 153–180
JOURNAL OF Accounting & Economics
www.elsevier.com/locate/jae

Earnings announcement premia and the limits to arbitrage ☆

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Received 23 March 2003; received in revised form 30 December 2006; accepted 15 January 2007
Available online 28 March 2007

Abstract

We examine the factors underlying the presence of earnings announcement premia. We find that the premia persist beyond the sample period examined in prior studies (ending in 1988), although they decline in magnitude after 1988. Further, premia are lower on the expected than the actual earnings announcement dates. We document that increases in voluntary disclosures result in lower premia, despite the increase in return volatility over time. Finally, our evidence suggests that the premia are not completely eliminated because of the costs of arbitrage.

JEL classification: G12; G14; M41; M45

Keywords: Earnings announcements; Announcement premium; Preannouncements; Disclosure; Limits to arbitrage

☆A previous version of this paper was titled: "Blinded by the light: Are earnings announcements worth the risk?" We would like to thank Yonca Ertimur, Thomas Dyckman, Tom Fields, Emre Karaoglu, Rick Mendenhall, Margaret Neale, Craig Nichols, Doug Skinner (the editor), Linda Vincent, an anonymous referee and seminar participants at the 2004 meetings of the American Accounting Association, Columbia University, Cornell University, the Massachusetts Institute of Technology, the University of Illinois at Chicago, the University of Southern California, and the Zell Brown Bag Seminar Series at the Kellogg School for helpful comments on previous drafts. Financial support from the Zell Center for Risk Research at the Kellogg School is gratefully acknowledged. All remaining errors are our own responsibility.
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0165-4101/$ - see front matter
doi:10.1016/j.jacceco.2007.01.008

1. Introduction

Prior literature documents significant positive abnormal returns around periodic news announcements (Penman, 1984; Kalay and Loewenstein, 1985; Chari et al., 1988; Ball and Kothari, 1991).[1] The explanation offered for these abnormal returns is that it is compensation for ''disclosure'' risk incurred when holding securities during a period when valuation relevant information is expected to be released.[2] Assuming mean–variance pricing, investors require higher announcement returns when a newsworthy announcement is expected and the associated risk is non-diversifiable.

The first of these conditions is satisfied because firms make periodic earnings announcements and voluntary disclosures that resolve uncertainty. As for the second condition, on the surface it appears that the risk should be diversifiable and not reflected in higher returns. However, Ball and Kothari (1991, hereafter BK) document a significant excess return of 0.24% on earnings announcement dates. We explore the factors underlying the presence of these earnings announcement premia.

Several factors motivate our re-examination of earnings announcement-period premia. First, much has changed since BK. The disclosure environment is richer with more frequent and detailed voluntary disclosures, such as earnings guidance, preannouncements, and conference calls (Soffer et al., 2000; Anilowski et al., 2007). Consequently, earnings announcement premia should decrease since there is more frequent resolution of uncertainty preceding the earnings announcement (although these voluntary disclosures may in fact be in response to an increase in newsworthy events). There has also been an increase over time in announcement-period return variances (Campbell et al., 2001; Rajgopal and Venkatachalam, 2005). Several factors are likely to contribute to the increased announcement-period variance, such as increased noise trading, greater news flow, decreased quality of earnings, and increased dispersion in analysts' forecasts (Rajgopal and Venkatachalam, 2005). All of these factors are likely to increase the uncertainty associated with earnings announcements, and hence are likely to result in higher announcement-period premia.

The overall effect of these trends on the announcement-period volatility is unclear. If firms are unable to completely offset the increased uncertainty with supplemental voluntary disclosures, then there is likely to be a net increase in announcement-period return volatility. Consistent with this, we find that while on average, the variance of the announcement-period abnormal returns increases significantly from 0.031 in the 1980–1988 period to 0.047 in the 1989–2001 period, there is a significantly smaller increase in announcement-period return variances for firms that preannounce earnings.

[1]A few empirical studies fail to find the evidence of earnings announcement-day premia (Peterson, 1990; Brown and Kim, 1993). However, the majority of the evidence supports the presence of higher returns on predictable disclosure events indicating that investors require an announcement-day premium.

[2]Robicheck and Myers (1966) illustrate this phenomenon with the following story: A ship sets out on a 2-year voyage in search of gold. At its departure, the prices of all financial claims on the payoffs of this journey reflect all the available information. Suppose no information reaches the market while the ship is away. Until the point the ship reaches the port, expected return on all financial claims related to the payoffs from the journey will be free of additional risk, because there is no information that would lead investors to change their valuations. Hence, during the voyage, the investment should earn the risk-free rate. However, the uncertainty is resolved once the ship returns with cargo. Therefore, if the risk were not diversifiable, the expected return would be higher on the day the market receives information about the claims' likely payoff.

Our work is also motivated by the fact that prior research does not control for announcement timing. For example, BK use the actual rather than the expected announcement dates to measure the premia. As we discuss in detail in Section 3, this approach is likely to result in an upward bias of the measured announcement-period premia. Thus, controlling for timing is important in order to obtain accurate measures of the magnitude of the premia.[3]

Our results indicate that the premia continue to persist beyond the period studied by BK but decline by a factor of two (0.07% versus 0.04%) from the BK to the post-BK period. Part of this decline can be attributed to the increased frequency of earnings preannouncements. We find that preannouncements reduce the earnings announcement premia by a factor of five, from 0.05% to 0.01%. However, the net decline in premia is not consistent with the concurrent increase in announcement-period volatility because a net increase in idiosyncratic risk over time would imply an increase in disclosure risk and hence result in higher earnings announcement premia.

This seemingly inconsistent result is the motivation for our second inquiry, namely whether the decrease in announcement premia over time (despite the increase in announcement-period volatility) is partly due to changes in the amount and cost of arbitrage.

We first document the benefits of arbitraging those premia by showing that portfolios comprising announcer firms earn a significant excess return (Jensen's alpha) of between 0.008% and 0.039% per day (corresponding to annualized abnormal returns between 2.12% and 10.63%). Further, excluding announcing firms from an otherwise fully diversified portfolio, while reducing the portfolio total risk (standard deviation) by 3.1%, lowers the excess return by enough to result in a lower Sharpe ratio. Thus, excluding announcing firms results in a less favorable return-risk tradeoff suggesting that arbitrage of these returns is attractive.

The decline in the premia over time could have resulted from either an increase in arbitrage capital and/or a decrease in the costs of arbitrage. As to the former, total assets under management at hedge funds increased from $26 billion in 1988 (the end of the BK period) to about $1,000 billion in 2004 (Avellaneda and Besson, 2005). As to the latter, the cost of arbitrage is likely to have gone down as the number of stocks with exchange-traded options has increased over time.

Consistent with the cost of arbitrage hypothesis, we find that announcement premia are positively associated with the costs of arbitrage, including idiosyncratic risk and bid-ask spreads. There is a negative association between the premia and the float on announcement days, which is consistent with arbitrage being more likely when a sufficient number of shares are available for trading. We also find lower premia for firms with exchange-traded options, consistent with the explanation in Skinner (1989) that traded options impart liquidity for the underlying stocks. In addition, we document that premia are higher on days when greater concentrations of firms announce earnings, consistent with the fact that

[3]Consistent with Chambers and Penman (1984), we find that the news is overwhelmingly "good" for early announcers: firms that announce earlier than the expected announcement date have positive earnings surprises 75% of the time. As a result, the timing of earnings announcements (early versus late) is associated with the content of the news. Therefore, the returns on actual earnings announcement dates are likely to reflect both the content of the news as well as the premium due to the resolution of uncertainty. Indeed, this effect results in a roughly three times larger estimate of the premia. We avoid this problem by measuring announcement premia on the expected earnings announcement date. See Section 3 for a discussion.

arbitrage capital is limited and cannot be deployed to exploit all arbitrage activities. Finally, we find a negative association between the announcement premia and trading volume on announcement days, indicating that large trading volumes, possibly a result of arbitrage activity, result in lower premia.[4]

Taken together, our results are important for several reasons. First, they establish the continued existence of an earnings announcement premia, albeit of a reduced magnitude. Second, we show that voluntary disclosures are associated with reduced premia and that such disclosures have partially mitigated the effect of the over time increase in idiosyncratic risk documented in extant studies. Finally, the magnitude of the premia is related to limits to arbitrage, suggesting that while arbitrage may have eliminated a significant part of the premia, the phenomenon is likely to persist.

The remainder of this paper is organized as follows: Section 2 describes the data used in the analysis. Section 3 contains an analysis of the premia after controlling for the timing of the announcements, the trends in the premia over time, and the effect of the changing disclosure environment on the premia. Section 4 presents an analysis of the relation between the premia and arbitrage activity and Section 5 concludes.

2. Data

Quarterly earnings announcement dates are collected from the COMPUSTAT quarterly file for the period 1980–2001 to yield a maximum of 88 quarters for each of the sample firms. We collect data for all firms with a December fiscal year end and available quarterly earnings announcement dates on COMPUSTAT, resulting in 12,377 firms and 297,426 firm-quarter observations. We only retain firms with at least 10 firm-quarter observations, resulting in 8,493 firms and 275,820 firm-quarter observations. Finally, after merging the COMPUSTAT sample with CRSP daily files, we are left with 7,260 firms and 227,281 firm-quarter observations.

Per COMPUSTAT, the earnings announcement date corresponds to "the date in which quarterly earnings and earnings per share figures are first publicly reported in the various news media (such as the Wall Street Journal or newswire services)".[5] As a result, the information is likely to have been impounded into security prices on days -1 or 0 for events where the announcement date is from the news media and on days 0 or $+1$ when the earliest date is from newswire services. Therefore, we define the earnings announcement period as the 3 days centered on the COMPUSTAT earnings announcement date (days -1, 0, and $+1$).

[4]However, this result is sensitive to the specification of trading volume. We discuss this issue in more detail in Section 4.

[5]Ball and Bartov (1995) state that COMPUSTAT relies on data sources such as, the Wall Street Journal, for earnings dates and they suggest that the errors are likely to be "small in number". Nevertheless, to investigate the issue of errors in COMPUSTAT earnings announcement dates, we obtain quarterly earnings announcement dates from IBES for the period 1984–2001 (we begin with 1984 since this is the first year for which report dates are available in IBES). We compare the dates from the two sources and find that prior to 1989, the COMPUSTAT announcement date is identical to the IBES announcement date 63% of the time. Further, whenever there is an inconsistency, the dates are off by only one trading day on average. For the period 1989–1994, the accuracy rate is even higher (82%), and in 1995–2001 the accuracy rate increases to 97%. This is consistent with the findings in DellaVigna and Pollet, (2005). Given the above evidence, and since we conduct our analyses using the 3-day window centered on the announcement date, our results are unlikely to be affected by any errors in the COMPUSTAT announcement dates.

For our analyses of the effect of the disclosure environment on announcement premia, we collect data on preannouncements of quarterly earnings from the Company Issued Guidance (CIG) database maintained by First Call. First Call collects data about earnings preannouncements from press releases and interviews by company officials.[6] We restrict our preannouncements sample to observations over the period 1998–2001 because the data for prior periods are less complete (Anilowski et al., 2007). This sub-sample consists of 70,073 firm-quarter observations for 5,178 firms. Of these, 46,184 firm quarters representing 5,121 firms had preannouncements.

3. Analysis of the announcement risk premia

We begin by investigating the effect of the timing of announcements on the earnings announcement premia by measuring the premia on the actual versus the expected announcement dates in Section 3.1. In Section 3.2 we study the trend in the premia over time by computing the premia over the entire sample period, the sub-period examined by BK and the sub-period subsequent to BK. Finally, in Section 3.3 we discuss the effect of changes in the disclosure environment on the earnings announcement premia.

3.1. Premia relative to the timing of announcements

Prior studies measure the premia at the actual announcement dates rather than the expected announcement dates. This approach is likely to lead to an upwardly biased estimate of the premia because the timing of the announcement conveys information (Chambers and Penman, 1984): firms with good news are more likely to announce early while late announcers tend to have bad news (McNichols, 1988; Begley and Fischer, 1998). As a result, when an earnings announcement is made (unexpectedly) prior to its expected date, the return on the actual announcement date reflects the good news (Chambers and Penman, 1984) plus an announcement-period premium.[7] Conversely, firms with unfavorable news are more likely to announce late. However, the unfavorable news is likely to be, at least partially, anticipated when firms fail to announce on the expected earnings announcement date. As a result, the effect of the bad news is likely to be excluded from the short-window announcement-period return. Therefore, a combined portfolio of early, on-time, and late announcers on the actual announcement date is likely to overstate the announcement-period premia.

We estimate the expected earnings announcement date for each firm quarter using the procedure described in Appendix A. In brief, we use the median announcement date for each firm quarter as the proxy for the expected announcement date. Note that the use of a model to estimate the expected announcement date introduces measurement error in the analysis by inclusion of some "expected dates" when investors actually did not expect an

[6]Other studies that have used the database are Soffer et al. (2000) and Cotter et al. (2002). Anilowski et al. (2007) explain in detail the origin and data collection approaches of First Call in compiling these data.

[7]To illustrate, assume that the expected value of the firm is \$100 per share. To allow for an announcement-period premium, the firm will trade at a discount form its expected value, say \$98 per share. Assume now that a firm with good news announces earnings early and, upon announcement, the stock price is \$104 per share. The increase in value of \$6 represents two components, a disclosure risk premium of \$2 and an increase with respect to the good news of \$4. Not controlling for the timing of the news, announcement would result in overstating the announcement premium to \$6 rather than \$2 in this case.

announcement. Inclusion of these dates biases the measured announcement premia towards zero.

To provide greater confidence in the performance of our expectation model, we hand collect expected earnings announcement dates from the ''Earnings Calendar'' published in the Wall Street Journal for July 2005.[8] We find a total of 2,047 expected earnings announcement dates compared to 2,735 actual announcements in COMPUSTAT in the same period in 2004. Thus, the Earnings Calendar provides expectations for roughly 75% of announcing firms. We find that 59% of the 2,047 firms announce within 1 day of the expected date as published in the Earnings Calendar, as compared to 62% on-time announcements using our expected announcement date model (see Appendix A). Moreover, this statistic is actually biased against our model, because the 62% includes all four quarters while the 59% is for the first quarter only, and fourth quarter announcement dates are the most difficult to predict.[9] Overall, data from the Earnings Calendar reinforce the reliability of the expected earnings announcement date prediction methodology used by us.

Bagnoli et al. (2002) show that the Chambers-Penman effect is stronger for bad-news announcing firms than documented in prior literature, and that good news firms do not tend to announce early. As a result, the upward bias in measuring the premia on the actual date is likely to be more important than the downward bias of measuring premia on the expected date (we provide evidence on these effects below). As a result, the research design hinges on the reliability of the expectation model, and the best estimate of the premia lies between the (lower) premia measured on the expected announcement date and the (higher) premia measured on the actual announcement date. To provide a complete analysis of the upper and lower bounds of the premia, we present all results using both the expected and the actual announcement dates.

We compute the premia for the 3-day expected earnings announcement period (days -1, 0 and $+1$) where day 0 represents the expected earnings announcement day.[10] For each firm j and quarter q, we subtract the mean return in non-announcement periods using the firm as its own control. We also use the returns on all non-announcing firms on the announcer's date as a second control: for each announcement in period t and quarter q, we compute the mean daily return for firms that did not announce.[11]

We define an earnings announcement to be on-time when it occurs within 1 day of the expected date predicted by our model. In Table 1, we report the announcement-period abnormal returns on the actual and expected announcement dates based on whether firms announce early, on-time, or late. For each of these categories, we also report the returns for positive surprises (which include zero surprises) and negative surprises. We measure

[8] ⟨http://online.wsj.com/public/Markets_Calendar.htm⟩.

[9] However, using proprietary data from First Call for a sub-set of 4,434 firms in the January 1995–July 1998 period, Bagnoli et al. (2002) report that 85% of firms report within 1 day of the expected announcement date collected by First Call as of 2 weeks of the expected date.

[10] Our results remain qualitatively unchanged when we define on-time announcements as those occurring on the expected announcements date (as opposed to the 3-day window centered on the expected announcement date).

[11] Our results are materially the same using either method, and thus for brevity we only discuss the results using the firm as its own control and for the 3-day announcement period. While our major results are invariant to the specific benchmark, the following trade-off exists: Using the firm as its own control (the first benchmark) has the advantage that it is likely to provide a better control for risk to the extent that firms within an industry announce in close proximity (and hence the non-announcer portfolio is likely to consist of firms from other industries). However, the second benchmark (using non-announcing firms as a benchmark) controls for market movements.

Table 1
Announcement-return premia relative to the timing of the earnings announcement 1980–2001

Portfolio	N	Expected announcement date (−1, 0, +1)	Actual announcement date (−1, 0, +1)
Early announcers	24,028	0.01 (2.21)	0.66 (22.84)
Positive surprise	19,040	0.02 (2.76)	0.86 (12.87)
Negative surprise	4,988	−0.03 (−3.87)	−0.08 (−7.67)
On-time announcers	141,414	0.12 (4.96)	0.12 (4.96)
Positive surprise	101,139	0.22 (5.27)	0.22 (5.27)
Negative surprise	40,275	−0.13 (−4.05)	−0.13 (−4.05)
Late announcers	61,839	−0.09 (−4.26)	−0.02 (−2.99)
Positive surprise	17,648	0.03 (2.67)	0.07 (3.42)
Negative surprise	44,191	−0.14 (−5.17)	−0.06 (−3.54)
On-time and late announcers	203,253	0.05 (4.14)	0.07 (2.84)
Positive surprise	118,787	0.19 (5.19)	0.20 (4.08)
Negative surprise	84,466	−0.14 (−4.61)	−0.09 (−3.97)

This table reports the announcement-day premia for the expected announcement window and the actual announcement window. The premia are the percentage daily return cumulated over the announcement period. The premia are computed for all firm years (Entire Portfolio), for those firms that announce ahead of the expected announcement date (Early Announcers), for those firms that announce either on the expected announcement date or after the expected dates (On-time and Late Announcers), for those firms that announce on the expected announcement date (On-time Announcers), and those firms that are neither early nor on-time announcers (Late Announcers). Earnings surprise is measured as the seasonal difference in earnings, defined as the earnings at quarter q less the earnings at quarter ($q-4$). We use the return on the firm on non-announcement days as a benchmark. Virtually identical results are obtained when we use the return on non-announcing firms on the same day as a control.

Two sample t-tests for differences between abnormal returns at the expected announcement dates for the different sub-samples, results in the following t-statistics: Early announcers versus on-time announcers (0.01 versus 0.12): $t = -7.22$. Early announcers versus late announcers (0.01 versus −0.09): $t = 6.54$. Early announcers versus on-time and late announcers (0.01 versus 0.05): $t = -3.52$. On-time announcers versus late announcers (0.12 versus −0.09): $t = 10.56$.

Two sample t-tests for differences between abnormal returns at the actual announcement dates for the different sub-samples, result in the following t-statistics: Early announcers versus on-time announcers (0.66 versus 0.12): $t = 14.69$. Early announcers versus late announcers (0.66 versus −0.02): $t = 13.74$. Early announcers versus on-time and late announcers (0.66 versus 0.07): $t = 11.67$. On-time announcers versus late announcers (0.12 versus −0.02): $t = 9.63$.

earnings surprise as the seasonal difference in earnings, defined as earnings at quarter q less earnings at quarter $q-4$.[12]

The first observation is the proportion of positive versus negative surprises for the early versus the late announcers. For the early announcers, approximately 79.2% of the earnings announcements comprise positive surprises, whereas for the late announcers the earnings announcements comprise positive surprises about 28.5% of the time. The on-time announcers have positive surprises approximately 58.4% of the time.

Using a χ^2 test, we reject the hypothesis (at the 0.001 level) that the proportions of positive and negative surprises are independently distributed across the early, on-time, and

[12]We also repeat the analysis with earnings surprise defined as actual earnings less the consensus analyst forecast. The results are similar to those reported in Table 1.

late announcers are randomly distributed. This supports the Chambers and Penman (1984) argument that firms tend to accelerate the announcement of good news and defer the announcement of bad news. Further, this analysis confirms the hypothesis that the magnitudes of the premia measured on the actual (versus the expected) announcement dates are likely to be overstated, and strengthens that argument for constructing an estimate of the expected announcement date and carrying out the analyses using expected announcement dates (see Frazzini and Lamont (2006), for more discussion on this issue).

Next, we find that on-time announcers earn an abnormal return of 0.12% in the expected (and, by definition, the actual) announcement period. However, this return is also a biased estimate of the announcement premium, as some firms that were expected to announce ended up announcing late and those firms are more likely to have had ''bad news''. Consistent with this observation, their return is significantly lower than that of the on-time announcers in the expected announcement period (difference significant at the 0.01 level). Thus, the return of on-time announcers is comprised of two effects: first, a positive announcement risk premium and, on average, a positive effect because these on-time announcers did, in hindsight, not announce late.

For late announcers, the measured return in the expected announcement period is the sum of the announcement-date premium and the informational effect of announcing late. Because the evidence indicates that the premium is positive, the measured return of −0.09 in the expected announcement period suggests that the late announcement effect is negative, consistent with the Chambers and Penman hypothesis. The on-time announcers' estimate of the announcement premium (i.e., 0.12%) implies that announcing late is associated with an abnormal return of roughly −0.21% (−0.09% minus 0.12%).[13] Finally, from an ex-ante perspective, the best approach to measuring the premia is to focus on on-time and late announcers. Our analysis indicates that premia for these announcers is 0.05% ($t = 4.14$).

These results show that the magnitude of the announcement premia documented in prior research was most likely overstated because these studies use actual rather than expected announcement dates. Indeed, the abnormal return for the entire portfolio is higher when computed for the actual announcement periods as compared to the expected announcement periods (0.14% compared to 0.05%): a large component of the premia on the actual announcement period is due to early announcers (0.66%). Because the premia in the expected announcement period is likely to be biased towards zero, while the premia measured in the actual announcement period are likely to be biased upwards, the best estimate of the premium is likely to lie between these two estimates and the regression coefficients in the subsequent analysis provides bounds on the likely magnitude of the underlying economic constructs.

To allow comparison with the results reported in BK, in the subsequent analyses we report results using returns for on-time and late announcers measured on the expected announcement window and returns for the entire sample on the actual announcement dates.

Finally, we find a significant positive abnormal return for early announcers in the expected announcement period. Because this result is inconsistent with market efficiency,

[13]The reason for stating ''roughly'' is because the 0.12% includes the positive effect of not having announced late.

we perform the following two diagnostic checks. One possible reason for the positive reaction is that for early announcers that announce on day -2 relative to the expected date, the abnormal return on the expected announcement window includes day $+1$ of the actual announcement window. To check whether this causes the positive reaction, we recompute the abnormal return on the expected window excluding early announcers that announced on day -2 and obtain a premia of 0.01 ($t = 2.14$). These results are similar to those obtained earlier, and we conclude that including announcers on days -2 relative to the expected date is not causing the positive reaction.

Second, we investigate whether the positive reaction for early announcers on the expected date results from a drift to (predominantly) good news for early announcers, similar to that observed in case of the post-earnings announcement drift documented in Ball and Brown (1968) and others. In unreported analysis, we divide the early announcers into very early announcers (those that announced 15 or more trading days early), somewhat early announcers (those that announced 7–14 trading days early), and slightly early announcers (those that announced 1–6 days early). We find that the magnitude of returns (and their significance) of the announcers that had positive surprises are greater when the expected announcement dates are closer to the actual dates (0.07 for the slightly early versus 0.01 for the very early announcers; although the magnitude of the return for the somewhat early announcers is slightly higher by 0.001 than the very early announcers). The magnitude of returns of the announcers that had negative surprises is more negative for the slightly early announcers than the very early announcers (-0.02 versus -0.01). These findings provide some support that drift in the reaction to the earnings news is responsible for the significantly positive market reaction for early announcers on the expected announcement date.

For the sake of completeness, we also conduct the above two diagnostic tests for the late announcers. First, for late announcers that announce on day $+2$ relative to the expected date, the abnormal return on the expected announcement window includes day -1 of the actual announcement window. To verify whether this overlap affects our results, we compute the abnormal returns on the actual announcement date excluding late announcers that announced on day $+2$ relative to the expected date. We obtain a premia of -0.10 ($t = -3.91$) on the expected announcement window, which is similar to the result obtained earlier. Thus, as before, we conclude that including announcers on days $+2$ relative to the expected date does not affect our results.

We then investigate the reaction to (predominantly bad) news on the actual announcement dates for late announcers based on the lateness of the announcements. As before, we divide the late announcers into very late announcers (those that announced 15 or more trading days late), somewhat late announcers (those that announced 7–14 trading days late), and slightly late announcers (those that announced 1–6 days late). We find that there is a monotonic decrease in the magnitude of returns of the announcers that had negative surprises when the actual announcement dates are further away from the expected dates (-0.01 for the slightly late, -0.02 for the somewhat late and -0.07 for the very late announcers). This provides further support for the theory that the worse is the news the more firms are likely to delay its announcement. This theory is also supported in the case of announcers that had positive surprises: the later is the announcement, the less positive is the return (0.01 for the very late, 0.02 for the somewhat late and 0.03 for the slightly late announcers). These results also provide further support to the theory of a drift in the reaction to earnings news.

Table 2
Earnings announcement return premia in the 1980–2001 period on expected and actual earnings announcement dates

Sample period	N	Expected announcement date (−1, 0, +1)	Actual announcement date (−1, 0, +1)
1980–2001	227,281	0.05 (2.99)	0.14 (11.24)
1980–1988 (Ball and Kothari)	57,974	0.07 (3.81)	0.24 (12.74)
1989–2001 (time period subsequent to Ball and Kothari)	169,307	0.04 (2.91)	0.11 (13.21)

This table reports the announcement-day premia and the associated t-statistics in parenthesis for the entire time period and two sub-periods: the time period studied by Ball and Kothari (i.e.,1980–1988), and the time period subsequent to Ball and Kothari (i.e.,1989–2001). The premia are the percentage daily return cumulated over the announcement period. We use the return on the firm on non-announcement days as a benchmark. Virtually identical results are obtained when we use the return on non-announcing firms on the same day as a control.
Performing two sample t-tests for differences across columns: Testing for equality between the abnormal return at the expected and the actual announcement dates (0.05 and 0.14) results in a t-statistic of 10.18. Testing for equality between the abnormal return at the expected and the actual announcement dates in the Ball and Kothari sub-period (0.07 and 0.24) results in a t-statistic of 8.75. Testing for equality between the abnormal return at the expected and the actual announcement dates in the post Ball and Kothari sub-period (0.04 and 0.11) results in a t-statistic of 6.41.
Performing two sample t-tests for differences across rows: Testing for equality of the abnormal return at the expected announcement dates between the Ball and Kothari sub-period and the post Ball and Kothari sub-period (0.07 and 0.04) results in a t-statistic of 3.65. Testing for equality of the abnormal return at the actual announcement dates between the Ball and Kothari sub-period and the post Ball and Kothari sub-period (0.24 and 0.11) results in a t-statistic of 7.93.

In the next section, we investigate the trend in the announcement risk premia over time and focus on the effect of the changing disclosure environment on the premia.

3.2. Premia over the sample period

3.2.1. Trends in premia in the BK and post-BK time periods

We begin our analysis by computing the earnings announcement premia in the entire 1980–2001 period and two sub-periods: 1980–1988 (which corresponds to the period investigated by BK) and the 1989–2001 period.

Row 1 of Table 2 reports the premia for the entire sample period, for the expected and the actual earnings announcement periods. For both windows, we find a positive and statistically significant (at conventional levels) average announcement premium of 0.05% ($t = 2.99$) and 0.14% ($t = 11.24\%$). While the premia are small they represent the arbitraged premia, i.e., the returns that remain after arbitrage activity has been undertaken.

The premia are significantly larger (at the 0.01 level) in the earlier sub-period for both expected and actual announcement windows by roughly a factor of two, indicating that the premia decline over time;[14] see rows 2 and 3 of Table 2.

[14]In earlier versions of the paper we limited our focus only to firms that we had observations for the entire sample period (88 quarters). All our results remain unchanged when we used a constant sample across different

Finally, consistent with our expectations, we find that the return on the actual announcement date is significantly larger than the return on the expected announcement date (by a factor of roughly three) in both sub-periods. Because the premia on the expected announcement dates are likely to be biased downwards (due to estimation error of the expected announcement dates) and the premia on the actual announcement dates are likely to be biased upwards (due to the fact that announcing per se conveys information), the unbiased estimate of the announcement-period premia is likely to be between those two measures.

3.2.2. *The effect of changing disclosure environment on premia*

The disclosure environment has changed considerably since the period studied by BK. One significant change is an increase in voluntary disclosures, especially earnings preannouncements (e.g., Skinner, 1994, 1997; Soffer et al., 2000; Anilowski et al., 2007). Preannouncements reduce announcement risk on the final earnings announcement date and therefore are likely to reduce the earnings announcement premia. However, while preannouncements and earnings guidance reduce the magnitude of earnings surprises, there is also evidence of an overall increase in idiosyncratic volatility over time (Rajgopal and Venkatachalam, 2005). Campbell et al. (2001) show that returns of individual firms became more volatile since 1960. Thus, there are two countervailing developments: increases in the frequency of preannouncements which reduce the announcement-period risk, and increases in stock return volatility which increase the announcement-period risk.

We use two tests to investigate the impact of the changes in the disclosure environment on announcement-period premia. In the first series of analyses, we investigate the effect of earnings preannouncements on the announcement premia. Next, we analyze the effects of increasing returns volatility on the announcement premia.

We report the results of tests of the effects of earnings preannouncements in Table 3. We first analyze announcement-period returns for on-time and late announcing firms in the 1998–2001 period on the expected announcement dates (Column 3 of Table 3). We find a total of 40,224 preannouncements for the 203,253 on-time and late announcement firm-quarter observations. For firms that issue preannouncements, the announcement-period premium is 0.01% ($t = 2.41$). In contrast, the returns are 0.05% ($t = 3.75$) for firms that did not issue preannouncements in the same period. Row 3 reports that differences between rows 1 and 2 are statistically significant at conventional levels ($t = -4.02$). We obtain virtually identical results for the entire sample on the actual announcement date (51,486 preannouncement firm-quarter observations out of 227,281 firm-quarter observations; see Column 4 of Table 3).

As a robustness check, we investigate whether the differences between preannouncing and non-preannouncing firms are due to firm-specific differences. We compare the announcement-period returns for the sample of firms that did and did not preannounce earnings in the 1998–2001 period with returns in the period 1980–1997 when the preannouncing firms are unlikely to have preannounced. As reported in Row 4 of Table 3, the announcement-period return in the 1980–1997 period for those firms that did preannounce in 1998–2001 are 0.06% ($t = 4.12$). The announcement-period returns in

(*footnote continued*)
sub-periods. This is also true for the findings documented in Table 3, where we compare the Ball and Kothari period (1980–1988) and the subsequent period (1989–2001).

Table 3
The relation between announcement return premia and preannouncements

Row	Portfolio	Expected announcement date (−1, 0, +1)	Actual announcement date (−1, 0, +1)
1	Abnormal returns for firms that preannounced 1998–2001	0.01 (2.41)	0.02 (2.89)
2	Abnormal returns for firms that did not preannounce 1998–2001	0.05 (3.75)	0.07 (3.84)
3	Difference between rows 1 and 2	−0.04 (−4.02)	−0.05 (−4.92)
4	Abnormal returns in 1980–1997 for firms that preannounced in 1998–2001	0.06 (4.12)	0.06 (4.12)
5	Abnormal returns in 1980–1997 for firms that did not preannounce in 1998–2001	0.07 (5.03)	0.07 (5.87)
6	Difference between rows 4 and 5	−0.01 (−1.22)	−0.01 (−1.27)

This table analyses the relation between the earnings announcement premia, and preannouncements for the sample of on-time and late announcers on the expected announcement date and for the entire sample on the actual announcement date. The premia are the percentage daily return cumulated over the announcement period. We use the return on the firm on non-announcement days as a benchmark. Virtually identical results are obtained when we use the return on non-announcing firms on the same day as a control. Row 1 (Row 2) reports the returns for preannouncing (non-preannouncing) firms on the expected and the actual announcement dates in the 1998–2001 period. Row 4 reports the returns on expected announcement dates in the 1980–1997 for firms that preannounced earnings in the 1998–2001 period. Row 5 reports the returns on expected announcement dates in the 1980–1997 period for firms that did not preannounce earnings in the 1998–2001 period.

1980–1997 for firms that did not preannounce earnings in the 1998–2001 are 0.07 ($t = 5.03$, Row 5). However, as reported in Row 6 of Table 3, the difference across these two portfolios is not statistically significant at conventional levels. Thus, we find no evidence that the lower announcement-period return of preannouncing firms is due to firm-specific characteristics.

We report the results with respect to the effect of returns volatility on announcement premia in Table 4. We begin by analyzing the differences in abnormal return variances across periods for expected and actual announcement event windows, using both parametric F-tests and non-parametric χ^2 tests.[15] For each sub-period and each condition (expected and actual announcement windows), we compare the average daily abnormal return variances to the average daily return variances in the days not falling either in the expected or actual announcement windows of that quarter (referred to as the non-event period). Panel A of Table 4 presents these results.

Consistent with prior research (Beaver, 1968; Patell and Wolfson, 1979), both the parametric F-test and the non-parametric χ^2 tests indicate that relative to the non-event periods, the variance in both the expected and the actual announcement periods are significantly larger than the variances in non-event periods at the 1% level or better. Thus, for the expected (actual) announcement period the mean and median daily abnormal return variance is 0.031 and 0.029 (0.039 and 0.035) in the BK period compared to 0.047 and 0.042 (0.051 and 0.048) in the post-BK period. Moreover, for 81.3% (84.2%) of the firm quarters in the BK period and 86.1% (89.6%) of the post-BK firm quarters the

[15]We use both parametric (F) and non-parametic (χ^2) tests because parametric F-tests may over-reject the hypothesis of no differences in variances.

Table 4a
Announcement abnormal return variances on the actual and expected earnings announcement dates (Panel A)

Time period	N	Expected announcement date (−1, 0, +1)	Actual announcement date (−1, 0, +1)
1980–1988 (Ball and Kothari)	57,974	0.031	0.039
		0.029	0.035
		81.3%	84.2%
1989–2001 (subsequent to Ball and Kothari)	169,307	0.047	0.051
		0.042	0.048
		86.1%	89.6%

This table reports the abnormal returns variances for the expected and the actual announcement dates (the event windows) for two sub-periods: the Ball and Kothari period (1980–1988) and the post Ball and Kothari period (1989–2001). For each cell, the first two numbers are the mean, median abnormal return variances. The third number represents the frequency that the respective abnormal return variances in the event windows exceed the abnormal return variance in the remaining days of the quarter excluding the expected and the actual return windows (non-event windows).

Two tests are performed to investigate whether the average daily abnormal returns variances in expected and actual earnings announcement windows exceed the return variances in the remaining days of the quarter: a parametric F-test (comparing the variances in the event periods and the non-event periods) and a non-parametric χ^2-test (comparing the frequencies that the return variances in the respective event periods exceed the return variances in the non-event periods).

Expected earnings announcement dates: Ball and Kothari period: $F = 8.92$ and $\chi^2 = 76.56$; Post Ball and Kothari Period: $F = 7.25$ and $\chi^2 = 87.65$.

Actual earnings announcement dates: Ball and Kothari period: $F = 9.65$ and $\chi^2 = 89.24$; Post Ball and Kothari Period: $F = 11.23$ and $\chi^2 = 92.77$.

All F-statistics and the χ^2-test are significant at the 1% level or lower, indicating that the variances in both the expected and actual announcement periods exceed the variances when earnings were neither expected to be announced nor were announced.

We test whether the average daily abnormal returns variances in the expected and actual earnings announcement periods increased from the Ball and Kothari period to the post-Ball and Kothari period by comparing the variances (0.031 versus 0.047, and 0.039 versus 0.051) using an (parametric) F-test and by comparing the frequencies that the respective variances exceed the respective variances in the non-event periods (i.e., 81.3% versus 86.1% and 84.2% versus 89.6%) using a (non-parametric) χ^2-test.

Expected earnings announcement dates (BK versus post-BK): $F = 8.91$ and $\chi^2 = 35.22$.

Actual earnings announcement dates (BK versus *post-BK)*: $F = 6.99$ and $\chi^2 = 31.09$.

All F and χ^2-statistics are significant at the 1% level or lower, indicating that the abnormal stock return volatility has increased from the period studied by Ball and Kothari to the post Ball and Kothari period.

variances in the expected (actual) event periods exceed the average abnormal returns variance in the non-event periods.

A comparison across the rows of Panel A indicates that the return variance for both the expected and the actual announcement periods have increased by approximately 50% from the BK to the post-BK period. Moreover, both the parametric and non-parametric tests indicate that those increases are statistically significant at the 1% level. This result is consistent with Rajgopal and Venkatachalam (2005), who document an over-time increase in return variances in earnings announcement periods.

We next examine how the over-time increase in the announcement return variances affects announcement premia. We divide our sample into three portfolios of firms based on their level of relative return volatility at earnings announcements (RV_{jq}).

Table 4b
Announcement return variance on the actual versus expected earnings announcement dates, preannouncement sample vs. non-preannouncement sample, 1989–2001 (Panel B)

Time period	N	Expected announcement date (−1, 0, +1)	Actual announcement date (−1, 0, +1)
Non-preannouncement sample	139,800	0.047 0.034 82.6%	0.051 0.039 84.1%
Preannouncement sample	29,507	0.048 0.021 76.2%	0.052 0.027 78.1%

This table reports the abnormal returns variances for the expected and the actual announcement dates (event windows) in the 1989–2001 periods for two sub-samples: firms that did preannounce earnings and firms that did not preannounce earnings. For each cell, the first two numbers are the mean, median abnormal return variances. The third number represents the frequency that the respective abnormal return variance exceeds the abnormal return variance in the remaining days of the quarter, excluding the expected and the actual return windows (non-event windows).

Testing whether the average daily abnormal returns variances in expected and actual earnings announcement windows exceed the return variances in the remaining days of the quarter.

Expected earnings announcement dates: Non-preannouncement sample: $F = 8.02$ and $\chi^2 = 69.03$; Preannouncement sample: $F = 6.91$ and $\chi^2 = 65.01$; *Actual earnings announcement dates*: Non-preannouncement sample: $F = 9.93$ and $\chi^2 = 81.64$; Preannouncement sample: $F = 10.03$ and $\chi^2 = 73.09$.

The F-statistics compare the average variances between the event windows and the non-event windows while the χ^2-statistics investigate whether the frequencies of the event window variances exceeding the return variances in the non-event windows are higher than 50%. All F and χ^2-statistics are significant at the 1% level or lower.

Testing whether the average daily abnormal returns variances in expected and actual earnings announcement windows are higher for firms that preannounced relative to firms that did not preannounce earnings.

Expected earnings announcement dates: $F = 1.68$ and $\chi^2 = 47.23$; *Actual earnings announcement dates*: $F = 1.54$ and $\chi^2 = 51.37$.

The F-statistics compare the average variances between the event windows for firms that did and did not preannounce earnings (e.g., 0.047 versus 0.048). The χ^2-statistics investigate whether the frequencies of the event window variances exceeding the return variances in the non-event windows are higher across the two disclosure regimes (e.g., 82.6% versus 76.2%). The F-statistics are not significant at conventional levels, while all the χ^2-statistics are significant at the 1% level or lower.

Relying on Beaver (1968), we compute a measure of relative volatility of stock returns, i.e. volatility in announcement periods relative to the volatility in non-announcement periods. For each firm quarter, we define the relative return volatility at an announcement (RV_{jq}) as the ratio of the sum of squares of returns in the announcement period relative to the sum of squares of returns in the entire quarter, or

$$RV_{jq} = \sum_{t \in A_j} [R_{j,t \in A_j} - \bar{R}_{jq}]^2 \Big/ \sum_{t \in q_j} [R_{j,t \in q_j} - \bar{R}_{jq}]^2, \qquad (1)$$

where for firm j and quarter q, A_j represents the 3-day announcement period and q_j represents the entire quarter, commencing on day $+2$ relative to the $q-1$ earnings announcement and ending on day $+1$ of quarter q's earnings announcement.[16]

[16]Because the number of days between quarterly announcements varies, we normalize (1) by dividing the denominator by the actual number of days and multiplying by 63 (the average number of trading days in a

We compare the difference in premia between the high and low relative volatility portfolios. The results (untabulated) indicate that the premia are significantly higher for the high return volatility portfolio. Second, we compute the correlation between return volatility and the announcement-period premia, and obtain positive and significant (at the 1% level) correlations for both benchmarks. These results support the claim that the higher is the return volatility at the time of earnings announcements (i.e., the higher the risk associated with the announcements), the higher are announcement premia.

The above results jointly illustrate the opposing effects of changes in the disclosure environment on the earnings announcement risk. While preannouncements reduce announcement premia, increasing announcement return volatility increases the announcement premia. Recall from Table 2 that the overall magnitude of the announcement-period premia has declined from the 1980–1988 to the 1989–2001 period. This suggests that for preannouncing firms the announcement return volatility is likely to be lower, resulting in a net decrease in the premia. We test this conjecture below by comparing the return variances for firms that preannounce versus those that did not preannounce.

We find that the abnormal return variances are significantly higher (at the 0.01 level using either parametric or non-parametric tests) in the event windows when compared to non-event days for both the non-preannouncement and the preannouncement sub-samples (Table 4, Panel B). However, only the non-parametric tests indicate a significantly lower (at the 0.01 level) event-period variance for the preannouncement sub-sample than the non-preannouncement sub-sample. To investigate these results further, we first note that the median variance is lower for the preannouncement sub-sample. This suggests that there may be some outliers that affect the results. When we eliminate the top and bottom 1% of the observations in each cell, the respective means for the non-preannouncement sub-sample are 0.044 and 0.048 for the expected and the actual announcement dates, and 0.042 and 0.046 for the preannouncement sub-samples. Consistent with the non-parametric results, preannouncements result in a significantly lower variance (at the 0.01 level). This supports that claim that preannouncements lessen the effect of increasing volatility of returns on announcement days, which is likely to explain the overall decline in the announcement premia in the post-BK period.

We conclude this section with a more formal analysis of the impact of the disclosure environment on the announcement premia. We estimate the following regression to measure the effect of preannouncements and the return volatility on earnings announcements on the announcement premia:

$$AP_{jq} = a + b \times RM_{jq} + c \times PreAnn_{jq} + d \times RV_{jq} + \varepsilon_{jq}, \tag{2}$$

where for security j and quarter q, AP_{jq} represents the announcement premium, RM_{jq} is the value-weighted CRSP return on the market on the announcement period, $PreAnn_{jq}$ is a dummy variable that equals 1 if a company made a preannouncement in that quarter, and 0 otherwise, and RV_{jq} is relative return volatility at earnings announcements defined in (1) above.

Recall that our measure of dependent variable is the firm's event period return minus its non-event period returns. Therefore, since BK document that betas increase in the

(*footnote continued*)
quarter). We also repeat all our analyses using simply the variance in the 3-day announcement period (instead of the relative measure). The results are similar to those obtained on using RV_{jq}.

announcement window, we include the market return to control for the beta in the event window. We expect a negative association between $PreAnn_{jq}$ and AP_{jq}, and a positive association between RV_{jq} and AP_{jq}.

Table 5 summarizes the results of estimating regression (2) using both the returns on the expected (on-time and late announcers) and the actual (total sample) announcement dates as the dependent variables. The coefficient of RM (a control variable) represents the systematic (beta) risk on earnings announcement days. Our estimated coefficient of 0.88 ($t = 18.42$) is generally consistent with this interpretation, although it is significantly less than 1.00 ($t = -2.51$). For the main variables of interest, the results of the multivariate analysis are consistent with the univariate results. We find that the coefficient estimate corresponding to $PreAnn_{jq}$ is negative and significant (-2.16, $t = -5.51$), suggesting that the announcement risk premium is lower for firms that preannounce their earnings. Consistent with the univariate results in Table 3, we find that the sum of the intercept and the $PreAnn_{jq}$ dummy variable does not differ from zero at conventional levels (F-value $= 0.96$ for the expected announcement period and F-value $= 0.14$ for the actual

Table 5
Relation between announcement return premia, preannouncements and volatility of returns on earnings announcements for on-time and late announcers, 1998–2001

$$AP_{jq} = a + b \times RM_{jq} + c \times PreAnn_{jq} + d \times RV_{jq} + \varepsilon_{jq}$$

Independent variables		Dependent variable: announcement premia on	
		Expected ($N = 40{,}224$)	Actual ($N = 51{,}486$)
	Constant ($\times 10^{-3}$)	2.41 (4.81)	2.84 (4.92)
Value-weighted CRSP return on the market in the announcement period	RM_{jq}	0.88 (18.42)	0.94 (29.94)
Dummy variable which takes the value of 1 if the firm made a preannouncement in that quarter, and 0 otherwise	$PreAnn_{jq}$ ($\times 10^{-3}$)	−2.16 (−5.51)	−2.98 (−7.67)
Sum of squared returns in the 3-day announcement period divided by the sum of squared returns in the entire quarter	RV_{jq} ($\times 10^{-2}$)	7.77 (23.92)	8.95 (25.74)
F-statistic (p-value)		473.28 (<0.0001)	498.69 (<0.0001)
Regression R^2		0.035	0.048

This table analyses the relation between the earnings announcement premia, preannouncements and the volatility of returns on earnings announcements for the sample of on-time and late announcers. The dependent variable, AP_{jq}, is the return on announcement dates minus the return of the firm on non-announcing days. Virtually identical results are obtained when we use the return on non-announcing firms on the same day as a control. The table reports the regression coefficients and the associated t-statistics in parenthesis. The expected (actual) announcement period corresponds to days -1, 0, and $+1$ relative to the date predicted by the expected date model (actual date as per COMPUSTAT).

Testing for the sum of the intercept and $PreAnn$ dummy to equal zero results in F-values of 0.96 in column 3 and 0.14 in column 4, neither of which is significant at conventional levels.

announcement period, and neither of these are significant at conventional levels). This indicates that preannouncements eliminate the premia during the announcement period for our sample firms. The coefficient estimate corresponding to RV_{jq} is positive and significant (7.77, $t = 23.92$), supporting the argument that the greater the relative volatility of returns on earnings announcements, the higher the announcement risk premia.[17,18]

We obtain similar results when we measure announcement premia on the actual announcement date, although the magnitudes of the coefficients on the independent variable are larger than those reported using the expected announcement date (including the slope coefficient of RM which is no longer significantly less than 1.00 at conventional levels). This suggests that apart from the scale effects, measuring premia using the actual or the expected announcement dates does not affect the conclusions drawn regarding the relation between announcement premia and over time changes in the disclosure environment.

Finally, we re-estimate regression (2) using the Fama and MacBeth (1973) approach. Because our sample has minimal time-series clustering (the mean, median, and maximum number of firms announcing on a given day being 0.75%, 0.45%, and 6.47% respectively), none of our conclusions change.

In summary, we document that the earnings announcement premia are significantly lower for firms that preannounce earnings and are significantly higher for firms with greater volatility of returns on earnings announcements. Further, we find that preannouncements mitigate the effect of increasing idiosyncratic volatility, and are likely to have contributed to the over-time decline in the magnitude of the earnings announcement premia.

4. Earnings announcement premia and arbitrage activity

The continued existence of the earnings announcement premia raises the question of the impact of arbitrage activity on the announcement-period premia. Specifically, does the compensation for the increased risk on announcement dates provide sufficient incentives to arbitrage these premia? Also, if arbitrageurs are trading to eliminate the premia (and are likely to have contributed to the decline in the premia over time), then do they find it costly to take positions in the announcement premium stocks to arbitrage away the excess returns? Section 4.1 examines the risk-return tradeoff of earnings announcements and Section 4.2 examines whether the limits to arbitrage are related to the premia.

4.1. Risk-return tradeoff: are earnings announcements premia worth the risk?

To examine the risk-return tradeoff on earnings announcements, we compare 3-day announcement-period returns to their standard deviations. There are likely to be some days when very few firms are expected to announce, resulting in an announcer portfolio of

[17] In order to ensure that the above results are not being driven by unobservable characteristics of firms that pre-announce in this period but are a function of the pre-announcement, we repeat the analysis for the period prior to 1998, and include a dummy variable for the firms that pre-announced in the 1998–2001 period. In unreported results we find that the dummy variable is not significant, which provides greater confidence in the above results. Finally, in untabulated results we find, consistent with BK, a small but significant increase in beta-risk on expected earnings announcements dates. However, the overall tenor of our results remains unchanged.

[18] Our results were not materially affected when we use $\log RV_{jg}$ as the independent variable.

Table 6
Analysis of the on-time and late announcer portfolio on the expected announcement dates

		(1) entire portfolio	Trading days with the market capitalization of expected announcers as a percentage of the entire portfolio of at least						
			(2) 0.1%	(3) 0.3%	(4) 0.5%	(5) 0.7%	(6) 0.9%	(7) 1.1%	(8) 1.3%
Trading days		5,427	3,234	2,565	2,192	1,908	1,678	1,500	884
Average Market Cap. (%)		1.457	2.087	2.547	2.914	3.012	3.241	3.657	3.841
Average market cap. ($\$10^3$)		938,509	1,449,021	1,707,814	1,937,677	2,115,691	2,186,927	2,276,663	2,353,049
Return (%)	Portfolio on $t-10$	0.0293	0.0301	0.0324	0.0328	0.0348	0.0354	0.0361	0.0381
	Portfolio on $t=0$	0.0511	0.0651	0.0692	0.0711	0.0748	0.0811	0.0839	0.0891
St. dev. (%)	Portfolio on $t-10$	0.0452	0.0472	0.0481	0.0483	0.0482	0.0483	0.0482	0.0483
	Portfolio on $t=0$	0.0491	0.0528	0.0515	0.0503	0.0493	0.0485	0.0479	0.0472
Sharpe ratio	Portfolio on $t-10$	0.0481	0.0574	0.0681	0.0742	0.0796	0.0847	0.0892	0.0928
	Portfolio on $t=0$	0.0512	0.0614	0.0768	0.0795	0.0869	0.0912	0.0981	0.1029
Jensen's Alpha	Daily (%)	0.008	0.012	0.019	0.023	0.028	0.031	0.033	0.039
($t=0$)	Annualized (%)	2.12	3.16	5.04	6.14	7.52	8.36	8.92	10.63
	t-statistic	0.45	1.48	1.64	1.83	1.97	2.09	2.18	2.32

This table reports results for value weighted portfolios on days when at least 0.1, 0.3, 0.5, 07, 0.9, 1.1, and 1.3% of firms by market capitalization are expected to announce earnings. Column (1) reports returns by including the portfolio of expected announcers independent of the market capitalization of the announcing firms. Columns (2)–(8) report results of excluding expected announcers when the announcers constitute at least 0.1–1.3% by 0.2% of the market capitalization of the sample. The last three rows report Jensen's Alpha, the corresponding annualized returns (assuming 259 trading days per annum) and the associated t-statistics for the announcer portfolio. The expected announcement period corresponds to days −1, 0, and +1 relative to the date predicted by the expected date model.

only one or two securities. Of the 5,638 trading days in the sample period, 211 trading days include none of the expected 3-day announcement periods for any of the sample firms. While announcement periods overlap for an additional 1,014 trading days, the market capitalization of the announcing firms comprise less than 0.1% of the market capitalization of the sample. Therefore, we analyze announcer portfolios consisting of trading days which include at least one firm's expected announcement period—we form seven portfolios consisting of trading days on which the market capitalization of the expected announcers range from 0.1% to 1.3% (by 0.2 percentage point increments) of the market capitalization of the entire sample. To allow for a meaningful comparison, we value-weight all securities in each portfolio, with 100% of the funds invested in the respective securities.[19]

We report the results of the analysis using the expected announcement dates in Table 6. Column (1) reports the results for all 5,427 trading days that include an expected 3-day announcement period of at least one firm. Columns (2)–(8) summarize the results for the seven portfolios. For comparison, we also report all statistics on day −10 relative to the

[19]The resulting analysis does not correspond to an implementable trading strategy because we exclude days on which either no firms announced, or days when less that 0.1–1.3% of the sample firms (by market capitalization) announced.

expected announcement date. We chose a date that precedes the expected announcement date to avoid using the actual announcement date for late announcers.

The average combined market capitalization of the announcing firms is 1.46% of the entire sample, corresponding to $938 million for the entire sample period. This figure ranges from 2.09% corresponding to $1,449 million for portfolio (2) to 3.84% and $2,353 million for portfolio (8).

The first three rows report the value-weighted returns for the entire sample and each of seven announcer portfolios. The premia increase monotonically with the concentration of announcers. This is consistent with the theory that arbitrage capital is limited: the larger the number of announcers, the greater is the portion of the unarbitraged premia because the limited capital is spread across more securities. We formally test this conjecture in Section 4.2 when we analyze whether limits to arbitrage are related to announcement-period premia. Another possible explanation is that it is harder to diversify on days when a greater number of firms announce. However, note that the number of announcing firms on any given day is small compared to the total market; even in the most extreme portfolio the announcing firms comprise only 1.3% of the market capitalization of the total sample. Therefore, we feel that the limits to arbitrage theory is a more likely explanation for the increase in the premia with the concentration of announcers.

Next we analyze portfolio risk. The results in the next two rows in Table 6 indicate that for columns (1) through (5) the standard deviations of the returns are significantly higher (at the 0.05% level or better) on expected announcement dates than on non-announcement dates (Beaver, 1968). However, the differences in risk between announcement and non-announcement dates decline monotonically and, beginning with column (6), is no longer significantly higher on expected announcement dates. In fact, for the last column, we find a standard deviation that is significantly lower (at the 0.10 level, but this may be simply sampling variation) on the expected announcement date. This indicates that while announcement risk is diversifiable, it takes relatively large announcer portfolios to diversify that risk.[20]

Is the increase in return worth the risk? To answer this question, we compute Sharpe Ratios for companies in announcement and non-announcement periods.[21] First, we find that Sharpe ratios increase as the number of securities in the portfolio increases (diversification effect). Second, we show that Sharpe ratios in the announcement periods (days -1, 0, $+1$) always exceed Sharpe ratios in the non-announcement comparison periods (days -9, -10, -11). This indicates that excluding announcers reduces the return-risk tradeoff.

Finally, we compute Jensen's alpha for the announcer portfolios.[22] The results, reported in the last three rows of Table 6, indicate that the expected announcer portfolios earn a significant excess return of from 0.008% to 0.039% per day (significant at the 0.05 level of better). Ignoring transactions costs, these daily excess returns correspond to annualized excess returns from 2.12% to 10.63%. Once again, it is interesting to note the monotonic

[20]Note that this analysis is from the perspective of an arbitrageur taking positions in announcing firms alone. In other words, the computations do not allow for diversification through positions in non-announcing firms.

[21]The Sharpe ratio is defined as the ratio of the portfolio excess return to the portfolio standard deviation, i.e., $SharpeRatio = (R_P - R_E)/\sigma(R_P)$, where R_P, R_F and $\sigma(R_P)$ are, respectively, the portfolio return, the risk-free rate and the standard deviation of the portfolio return.

[22]Jensen's Alpha is defined as the intercept in a regression of a securities excess return (return minus risk-free rate) on the market's excess return.

increase in the Jensen's alpha with the increase in earnings announcement concentrations. We offer the same two possible explanations as in the earlier case (limited arbitrage capital and difficulty of diversification), but again feel that the theory on limits to arbitrage is more plausible. These results imply that it is worthwhile to arbitrage the announcement risk premia (Frazzini and Lamont, 2006 also report positive abnormal returns in earnings announcement periods. It is important to point out that our estimates of announcement premia and returns to announcer's portfolios are those that remain after arbitrage. Thus the estimates of announcement premium and the Jensen's alpha are smaller in magnitude than likely earnings announcement premia prior to arbitrage activity.[23]

4.2. Limits to arbitrage and the persistence of announcement risk premia

Shleifer and Vishny (1997) point out that frictions and risks in execution of the arbitrage in practice results in "risk arbitrage" that requires substantial commitment of capital on part of arbitrageurs. In turn, arbitrageurs are likely to face capital constraints or have to raise capital from investors who may not have the required appetite to sustain risk positions. These factors deter arbitrage activities and can lead to persistent anomalous returns in the market. Mashruwala et al. (2006) argue that for a riskless hedge to exist, the arbitrageur needs to find close substitute stocks whose returns are highly correlated with the returns of the firms subject to anomalous mispricing. Identifying such substitutes is a difficult task in practice. In their study, they find that the well documented "accruals anomaly" is persistent because it is difficult to arbitrage. We investigate whether the costs of arbitrage are a potential explanation for the persistence of the announcement-period premia.

We use five proxies to capture the costs of executing arbitrage transactions for each firm j and quarter q: idiosyncratic risk (Div_Risk_{jq}), bid-ask spread ($Spread_{jq}$), float ($Float_{jq}$), the concentration of announcing firms ($Weight_{jq}$) for a given announcement-period window, and the presence of exchange-traded options ($Option_{jq}$) on the firm's stock.[24]

Following Mashruwala et al. (2006), Wurgler and Zhuravskaya (2002), and Pontiff (1996), we use the idiosyncratic part of a stock's volatility to proxy for the absence of close substitutes while assuming arbitrage positions. Idiosyncratic risk is relevant to arbitrageurs because arbitrageurs can only hold relatively few positions at a time due to limited capital. Several papers (e.g., Pontiff, 1996; Shleifer and Vishny, 1997; Wurgler and Zhuravskaya, 2002; Ali et al., 2003) that explore explanations related to barriers to arbitrage make similar assumptions. Thus, we predict that the higher the idiosyncratic risk associated with a firm, the harder it will be for arbitrageurs to diversify the risk. We measure idiosyncratic risk as the residual variance from a regression of firm-specific stock returns on the value-weighted CRSP stock index for 12 months preceding the quarterly earnings announcement.

[23]We also repeat the analysis using the actual announcement day, similar to the methodology followed in BK. The results (untabulated) are materially similar to results when the announcement date is the expected date. Consistent with announcement premia being large when using the actual announcement date, we find that the announcement-period return for the announcers (0.0921%), the Sharpe Ratio (0.0784) and the Jensen's Alpha (4.4) are larger when using the actual announcement date as compared to using the expected announcement date. In the last section, we analyze whether the continued existence of the premia is due to limits to arbitrage.

[24]The bid-ask spread, volume, and float have been commonly used in the literature as proxies for limits to arbitrage, e.g., Ali et al. (2003).

The bid-ask spread ($Spread_{jq}$) captures a large portion of the round-trip transactions costs. We measure $Spread_{jq}$ in the announcement window by computing the average daily bid-ask spread scaled by the mid-point of the spread reported by CRSP for each firm and quarter.

The variable $Float_{jq}$ measures the number of shares that are available for trading in a particular stock. Depending on the nature of investors and the characteristics of the firm, certain stocks may have few shares that are actively traded in the market. Arbitrageurs would then be limited in their ability to take positions if there were not enough floating stock available for a particular firm. We measure $Float_{jq}$ as the average trading volume of a firm in each quarter scaled by the average shares outstanding for the firm.

The next variable, $Weight_{jq}$, represents the market capitalization of all firms announcing earnings concurrently with firm j. We include $Weight_{jq}$ as a measure of the number of arbitrage positions available on a given day. We assume that arbitrage capital is limited. Therefore, higher values of $Weight_{jq}$ imply that a given amount of arbitrage capital is spread across more deals, leaving a larger portion of the announcement period return un-arbitraged.[25]

Our final variable, $Option_{jq}$, takes the value of 1 if the firm had exchange traded options at the date of the quarterly earnings announcement and zero otherwise. Options are likely to reduce the costs of arbitrage, either directly by offering a convenient vehicle for arbitrage or due to the associated increase in liquidity (Skinner, 1989). Since the development of options markets is a relatively new trend, data for all firms are not available.[26]

We also include an additional variable, $Volume_{jq}$, measured as the average of the daily trading volume in the earnings announcement window reported by CRSP for each firm and quarter. This measure could represent three phenomena. First, it could serve as a proxy for the impact of arbitrageurs' taking advantage of the premia. By purchasing the stock, arbitrageurs would drive down the premia, and we would expect a negative association between $Volume_{jq}$ and the observed announcement-period premia. Second, volume is also a proxy for the arrival of information (Beaver, 1968). Finally, according to the attention-grabbing hypothesis, individual investors trade heavily on earnings announcements and are net buyers, unconditional on the news (Frazzini and Lamont, 2006). In the latter two cases, volume will be positively correlated with the announcement-period premia. Thus, we have no ex-ante predictions for the relation between volume and the announcement-period premia. All other variables are as defined before.

We test whether limits to arbitrage are associated with the announcement-period premia by estimating the following pooled time-series, cross-sectional regression for the sample of on-time and late announcers:

$$\begin{aligned} AP_{jq} = {} & a + b \times RM_{jq} + c \times Div\_Risk_{jq} + d \times Spread_{jq} \\ & + e \times Float_{jq} + f \times Weight_{jq} + g \times Option_{jq} \\ & + h \times Volume_{jq} + k \times PreAnn_{jq} + l \times RV_{jq} + \varepsilon_{jq}. \end{aligned} \tag{3}$$

[25] An alternative metric for concentration of announcers could be constructed by counting the total number of firms that announced as a percentage of all the sample firms traded on that day (equal-weighted measure of clustering). This metric is frequently used in the literature (Chambers and Penman, 1984; Brown et al., 1992). This equal-weighted metric, however, overemphasizes firms with small capitalizations and hence is not representative for investors holding diversified portfolios.

[26] We obtain data for whether the firm has traded options from the OptionMetrics database. When an observation for traded options is available in the database for a firm, the dummy variable takes on a value of 1. Data are available from 1996. For 1996 there are 2278 firms, in 1997 there are 2767 firms, in 1998 there are 3092 firms, 1999 there are 3213 firms, in 2000 there are 3413 firms, and in 2001 there are 2798 firms.

If limits to arbitrage contribute to the existence of the announcement premia, then after controlling for other factors that affect announcement premia, firms with higher values for Div_Risk_{jq}, higher values for $Spread_{jq}$, lower values for $Float_{jq}$, higher values for $Weight_{jq}$ and with no exchange-traded options will have higher premia. Pearson correlations among the independent variables (untabulated) in (3) indicate that the correlations among them while statistically significant are fairly low.

The regression results reported in Table 7 are consistent with the hypothesis that the arbitraged announcement-period premia are higher when limits to arbitrage are higher. We report results for announcement premia computed using the expected as well the actual announcement date but discuss the results for the expected announcement date (we point out differences in results, if any). First we discuss the results for the entire sample. The coefficient for Div_Risk_{jq} is positive and significant (6.84, $t = 5.04$) indicating that announcement premia are higher for firms with higher idiosyncratic risks. Similarly, the coefficient estimate for $Spread_{jq}$ is positive and significant (8.69, $t = 5.35$), consistent with our expectation that announcement premia are higher for firms with higher trading costs (and greater information asymmetries). The coefficient for $Float_{jq}$ is significant (-6.83, $t = -5.94$), indicating that announcement premia are higher when there are not enough floating stocks available for arbitrage. Finally, the coefficient for $Volume_{jq}$ is significant (-3.66, $t = -2.99$), indicating that announcement premia are lower when a large volume of shares has been traded, possibly as a result of arbitrage.[27]

Consistent with the limits to arbitrage hypothesis, the coefficient on $Weight_{jq}$ is positive and significant (7.94, $t = 4.19$), indicating arbitrage capital is not sufficient to exploit the premia when there are numerous available arbitrage positions. With respect to the presence of exchange-traded options, we find that the variable $Option_{jq}$ is negative and statistically significant (-2.18, $t = -7.36$), suggesting that the earnings announcement premia are lower if the firm had exchange-traded options. This finding is consistent with the argument behind the results in Skinner (1989), and further supports the hypothesis on the relation between limits to arbitrage and the earnings announcement premia.

Finally, consistent with the earlier results, the coefficients on RV_{jq} and RM_{jq} are positive and significant (9.07, $t = 56.97$ and 7.89, $t = 65.48$, respectively), indicating that premia are higher when the volatility of returns on earnings announcements is high and when earnings announcements have a higher β-risk relative to non-announcement days (BK).

Next, we repeat the analysis for the 1998–2001 sub-sample where we have information on preannouncements. Again, the correlation among the six independent variables is statistically significant but small in magnitude (untabulated).

The results for the preannouncement sample reveal that the dummy variable $PreAnn_{jq}$ has a negative and significant coefficient (-1.44, $t = -3.88$), indicating that preannouncers have lower announcement premia. The variables Div_Risk_{jq} (1.34, $t = 4.79$), $Volume_{jq}$ (-5.84, $t = -4.83$), $Spread_{jq}$ (3.14, $t = 6.81$), $Float_{jq}$ (-4.67, $t = -7.55$), $Weight_{jq}$ (3.09,

[27] Frazzini and Lamont (2006) find that announcement premium is positively associated with abnormal trading volume in the stock. In their study, abnormal trading is defined as the trading volume in the announcement month scaled by non-announcement volume adjusted for market-wide scaled trading volume in the announcement month. In order to reconcile our results with Frazzini and Lamont (2006), we repeat our tests using their measure of scaled volume and we also obtain a positive coefficient. However, our remaining results are unaffected by this change in the volume measure. Thus, the result on the relation between volume and the premia is sensitive to the specification of the volume measure. Given that there is no authoritative theory on how to measure volume, and this measure could proxy for two opposing phenomenon, we leave this as an unresolved issue.

Table 7
Earnings announcement premia and the limits to arbitrage

$$AP_{jq} = a + b \times RM_{jq} + c \times Div\_Risk_{jq} + d \times Spread_{jq} + e \times Float_{jq} \\ + f \times Weight_{jq} + g \times Option_{jq} + h \times Volume_{jq} + k \times P \times r \times eAnn_{jq} + l \times RV_{jq} + \varepsilon_{jq}$$

Independent variable		Announcement premia			
		Expected date		Actual date	
		Total sample (1996–2001) ($N = 95{,}219$)	Pre-announcement sample (1998–2001) ($N = 40{,}224$)	Total sample (1996–2001) ($N = 95{,}219$)	Pre-announcement sample (1998–2001) ($N = 51{,}486$)
	Constant ($\times 10^{-3}$)	2.67 (3.21)	3.12 (2.72)	2.14 (3.87)	3.68 (3.01)
Value-weighted CRSP return on the market in the announcement period	RM_{jq} ($\times 10^{-1}$)	7.89 (65.48)	7.01 (29.83)	9.77 (73.24)	8.45 (36.04)
Diversifiable risk of security j measured in the 12 month period prior to quarter q	Div_Risk_{jq} ($\times 10^{-4}$)	6.84 (5.04)	1.34 (4.79)	6.45 (5.94)	1.95 (5.49)
Ratio of the high ask price minus the low bid price divided by the average of the high ask and the low bid price	$Spread_{jq}$ ($\times 10^{-3}$)	8.69 (5.35)	3.14 (6.81)	9.67 (7.74)	4.67 (6.86)
Average trade volume during quarter (excluding announcement window) divided by total number of shares outstanding	$Float_{jq}$ ($\times 10^{-5}$)	−6.83 (−5.94)	−4.67 (−7.55)	−4.96 (−3.88)	−4.08 (−5.96)
Market capitalization of the firms announcing concurrently with firm j	$Weight_{jq}$ ($\times 10^{-2}$)	7.94 (4.19)	3.09 (5.67)	8.47 (7.14)	5.68 (6.83)
Dummy variable which equals 1 if the company had exchange traded options	$Option_{jq}$ ($\times 10^{-2}$)	−2.18 (−7.36)	−3.54 (−8.67)	−1.26 (−5.64)	−2.84 (−6.71)
Average trading volume during the announcement period	$Volume_{jq}$ ($\times 10^{-10}$)	−3.66 (−2.99)	−5.84 (−4.83)	−4.17 (−3.54)	−5.77 (−4.95)
Dummy variable which equals 1 if the firm made a preannouncement in that quarter, and 0 otherwise	$PreAnn_{jq}$ ($\times 10^{-3}$)	NA	−1.44 (−3.88)	NA	−2.98 (−4.06)
Relative volatility of returns on earnings announcement dates measured as the sum of squared returns in the 3-day announcement period divided by the sum of squared returns in the entire quarter	RV_{jq} ($\times 10^{-2}$)	9.07 (56.97)	7.57 (27.22)	9.28 (57.54)	8.22 (32.12)
F-statistic (p-value)		1014.05 ($<.0001$)	367.14 ($<.0001$)	1687.72 ($<.0001$)	457.36 ($<.0001$)
Regression $R^2 \times 10^{-2}$		5.57	4.28	8.01	6.14

This table analyses whether limits to arbitrage arguments are a potential explanation for the existence of the announcement premia for the sample of on-time announcers as well as on-time and late announcers. The dependent variable, AP_{jq}, is the return on the 3-day period centered on the expected (columns 1 and 2) and the actual (columns 3 and 4) announcement dates minus the return on non-announcement dates for the same firm. Virtually identical results are obtained when we use the return on non-announcing firms on the same day as a control. The expected announcement date is as computed by our model (see Appendix A) and actual announcement date is as per COMPUSTAT.

$t = 5.67$) and $Option_{jq}$ (-3.54, $t = -8.67$) continue to have the expected signs and are significant at conventional levels in explaining announcement premia. The coefficient estimates for RV_{jq} and RM_{jq} also continue to be positive and significant (7.57, $t = 27.22$ and 7.01, $t = 29.83$).

In summary, the above evidence indicates that it is attractive to arbitrage the earnings announcement returns, and arbitrageurs are likely to have contributed to the decline in the premia over time. However, the difficulty that arbitrageurs face in trying to exploit the anomaly prevents the premia from being eliminated completely, which also suggests that this phenomenon is likely to persist. [28,29]

5. Summary and conclusions

We examine the factors underlying the presence of earnings announcement premia. As a first step, we introduce a methodological refinement of measuring the premia on the expected announcement date and not the actual announcement date as in BK. We find that the magnitude of the premia are smaller when measured on the expected rather than the actual announcement date, but remain statistically significant. We document that the premia persist beyond the time period studied in BK, although the magnitude declines by nearly half.

We investigate whether changes in the disclosure environment contribute to the decline in the premia. Theory suggests that the increase in disclosure activity of firms over time leads to reductions in disclosure risk, and hence in announcement premia. Our evidence is consistent with this claim. However, extant empirical research also documents increases in idiosyncratic risk over time. We confirm this finding, and also show that greater idiosyncratic volatility is related to higher announcement premia. However, we find that preannouncements mitigate the increase in idiosyncratic volatility and are likely to have contributed to the over-time decline in the premia.

Next, we investigate whether the arbitrage of these returns is attractive, and whether limits to arbitrage explain why the premia have not been completely eliminated by arbitrageurs. We document that a portfolio strategy of excluding announcers from a daily rebalanced market portfolio reduces the standard deviation of the portfolio. However, this strategy also reduces the Sharpe ratio, which indicates that excluding announcers worsens the return-risk tradeoff. Moreover, returns on expected earnings announcement dates "earn" significantly positive Jensen's Alphas, corresponding to an annualized return between 2.1% and 10.6%. Thus announcer firms are potentially attractive targets for arbitrageurs which could also explain reductions in announcement premia over time. Since the measured premia already incorporate the effects of arbitrage activity, we would expect that stocks which are harder for arbitrageurs to take positions in would have higher announcement premia. Consistent with this we find that stocks with greater limits to

[28]As an additional sensitivity test, we re-estimate the regressions by including size as a control variable (defined as the market value of equity). BK find that smaller firms have larger announcement premia. While the coefficient of size is negative and significant at the 5% level, inclusion of this variable does not materially affect any of the coefficients of our variables of interest.

[29]When we compute announcement premia using the actual announcement date the results are similar to those discussed above. However, the coefficients for our proxies indicating difficulty in taking arbitrage positions (Diversifiable Risk, Spread, Volume, Float, Weight and Option) are larger than those reported when announcement premia are computed using the expected announcement date.

arbitrage have higher announcement premia. This also suggests that the premia are likely to continue to exist.

Overall, our results provide important insights into earnings announcement premia. Our study broadly speaks to the process of price formation in response to earnings announcements, especially variation in firm-level returns driven by firm-specific news, in a similar spirit to Vuolteenaho (2002). Our evidence implies that factors affecting announcement date excess returns play a role in explaining the variation in firm-level returns.

Appendix A. Analysis and prediction of quarterly earnings announcement dates

This Appendix provides a detailed discussion of a model of expected quarterly earnings announcement dates, which we develop and use in our analyses. In practice, firms can announce earnings on a particular day of the week or choose the day of the week based on the content of news to be disclosed (Watts, 1978; Patell and Wolfson, 1982; Penman, 1987; Ball and Bartov, 1995; Brown et al., 1992). Additionally, firms often follow complicated algorithms such as "the first Tuesday, 3 weeks following the end of the fiscal quarter," and it may not be easy to detect those rules because actual announcement dates often deviate from expected announcement dates.[30]

Various models for prediction of earnings announcement dates have been used in the literature. A widely used model, introduced by Givoly and Palmon (1982), uses a firm's prior-period announcement date as a proxy for the current year's announcement date.[31,32] However, this approach results in measurement errors for any announcement that deviates from the 'normal' disclosure strategy. For example, if Q_t were to be late, then Q_{t+4} is likely to be classified as early, even when in reality it were announced 'on time.' Similar problems are present with the other models used in the literature. Essentially, the use of any expectation model that relies on actual announcement dates introduces measurement error to the analysis.

To develop the expected announcement date model, we analyze the distribution of actual earnings announcement dates.[33] First, the distribution of the fourth quarter announcement days differs from those of the preceding three fiscal quarters. While using separate models for Q_4 and for the other three quarters (Q_1, Q_2, and Q_3) could be efficient, the expected announcement dates for the first three fiscal quarters would be more precise (by virtue of a larger sample size) and possibly affect some of the subsequent analyses. Therefore, we decided to estimate a separate model (by firm) for each of the four fiscal quarters. Second, there is an upward trend in the announcement dates for Q_1 through Q_3 and an increase in the standard deviation of the Q_4 announcement dates. To account for these over-time changes in the distribution of the announcement dates, we divide the 22 year period into 6 sub-periods, 5 sub-periods of 4 years and 1 sub-period of 2 years. We

[30]We thank the referee for pointing this out.

[31]Analysis in Givoly and Palmon (1982) is based on the reporting lag relative to the fiscal year end, rather than the reporting date. However, because the fiscal year end does not vary in their analysis, there is a one-to-one relation between the expected reporting lag and the expected announcement date.

[32]Chambers and Penman (1984) and Begley and Fischer (1998) use the same model to analyze quarterly as well as annual earnings announcements. Chambers and Penman (1984) use two additional models to estimate the expected earnings announcement date. They report that their results were not sensitive to which model they used.

[33]Detailed results of the analysis are available from the authors and are not reported here in the interest of brevity.

also repeat our analysis by dividing the period into 4 sub-periods, 3 sub-periods of 6 years, and 1 sub-period of 4 years.

For each sample firm and each fiscal quarter, we use the median announcement date as the proxy for the expected announcement date. We select the median because this statistic is least likely to be affected by individual deviations for the normal disclosure schedule. For each firm and each quarterly earnings announcement, we compute the median announcement date for each 4-year sub-period. Each quarter is divided into 63 trading days. Using the quarterly earnings announcement data from COMPUSTAT, we identify each firm-quarter earnings announcement date with the day of the quarter (i.e., day 1–day 63). For each sample firm and each fiscal quarter, we compute the median day of announcement (Med_{jq}).

Using the median as a proxy for the expected announcement date, we compute the deviation from the expected announcement day (Dev_{jq}) as the absolute difference between the actual announcement day and the median announcement day:

$$Dev_{jq} = \left| D_{jq} - Med_{jq} \right|,$$

where for firm j and quarter q, D_{jq} represents the actual announcement day. Thus, $Dev_{jq} = 0$ corresponds to firm j having announced quarter q earnings on the expected day, $Dev_{jq} = 1$ represents instances where corporations have announced on either a day earlier or a day later than expected, etc.

Overall, in our sample 37.61% of the firm-quarter announcements are on the expected announcement day, and 62.22% are within 1 day of the expected date. In fact, 86.64% of the firm-quarter announcements fall into the 11-day window centered on the expected announcement day. The announcement dates for the first three fiscal quarters are more predictable than the announcement dates for the fourth fiscal quarter. For the first three fiscal quarters, the percentages of announcements that are released on the expected date range from, 39.90% (Q_1) to 39.71% (Q_3), while approximately 64% are released in the 3-day window centered on the expected announcement date. In contrast, only 37.54% of Q_4 announcements are on the expected release date, and 57.14% are within 1 day of the expected release date.

To investigate the sensitivity of our proxy to the length of the estimation period, we also use 4 sub-periods (3 sub-periods of 6 years and 1 sub-period of 4 years) and one sub-period (of 22 years) to compute the expected announcement date and Dev_{jq}. The results indicate a drop of overall on-time announcements to 35.03% for four sub-periods and to 23.98% for one sub-period. The decrease in on-time announcements is most pronounced for Q_4. Thus, the percentage of on-time announcements is approximately 36% for Q_1 through Q_3 for 4 sub-periods and 25.01% for one sub-period. In contrast, for Q_4, the frequencies are 32.12% and 20.39%, respectively. We get similar results (although slightly higher for on-time announcements) when we use the integer value of the mean (rather than the median) as the proxy for the expected announcement date.

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