# Firm Size, Common Stock Offerings, and Announcement Period Returns

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***ABSTRACT***

*Recent research offers evidence that common stock offerings for small OTC firms have a more negative mean announcement period stock return than do large OTC firms. Motivated by this finding, we examine whether the firm size results documented for OTC stock offerings hold for AMEX and NYSE stock offerings. We extend the research on several fronts. First, we show that the firm size effect found for OTC stock offerings holds for AMEX and NYSE stock offerings. Second, we offer evidence that this firm size effect cannot always be explained by issue costs. In addition to the above extensions, we conduct a series of regression tests and establish that firm size, along with signaling variables, are significant factors when accounting for stock returns. Our regression tests demonstrate that firm size is the only explanatory variable that is consistently statistically significant.*

## Introduction

Prior event period research discovers that common stock offerings for small OTC firms have a more negative mean announcement period stock return than large OTC firms. However, most event studies involve samples that include AMEX- and NYSE-listed firms. Thus, there is the need to explore whether the firm size results documented for OTC stock offerings hold for AMEX and NYSE stock offerings.

We extend stock-for-debt studies on several fronts. First, we extend the OTC firm size research of Hull and Pinches (1994/1995) by including AMEX/NYSE observations in our stock-for-debt sample. Prior AMEX/NYSE security offering studies argue for signaling effects but do not link the magnitude of the signaling to firm size.[1] Second, motivated by the work of Hull and Kerchner (1996), we investigate to what extent issue costs can explain differences in stock returns attributable to firm size. If issue costs cannot adequately account for the firm size effect, then there is evidence to support firm size as a proxy for a differential information wealth effect.

**Literature Review**

In this study, we use a sample of "pure" leverage decrease announcements consisting of stock offerings that retire debt. As first noted by Masulis (1980), such a sample is ideal for examining the stock price impact of a firm's change in its security mix. This is because a pure leverage change varies the mix of the firm's securities without directly altering the productive assets. Thus, any wealth impact from altering production is not present. Those who have investigated stock-for-debt transactions (Masulis, 1983; Cornett and Travlos, 1989; Copeland and Lee, 1991; Hull, 1994; Hull and Moellenberndt, 1994) document statistically significant negative announcement period stock returns. They generally conclude that the market response is consistent with asymmetric information theory predicated on insider signaling (Leland and Pyle, 1977; Ross, 1977; Fama, 1985).[2]

Hull and Pinches (1994/1995) and Hull and Kerchner (1996) recently have offered new insight into explaining the negative returns for stock-for-debt transactions. Hull and Pinches (1994/1995) find that firm size does a better job of accounting for OTC stock returns than signaling effects (or other wealth effects cited by the extant literature). To the extent firm size proxies for the amount of information activities, their findings are consistent with differential information theory (Wilson, 1975; Atiase, 1980; Verrechia, 1980; Bhushan, 1989). This theory predicts that announcements by small firms reveal more, and thus induce greater wealth effects (by magnifying effects from insider signaling). Lastly, Hull and Pinches (1994/1995) offer some evidence to suggest that a firm size may be attributed to issue-related costs (in addition to differential information).

Hull and Kerchner (1996) analyze the role of issue costs in accounting for OTC/AMEX/NYSE stock offering announcement returns. In their study issue costs comprise cash flotation costs (consisting of the underwriting spread and fees associated with administration, registration, and legal services) and underpricing. They find that issue costs can aid in explaining the negative announcement period return. This is especially true for samples that contain either OTC or AMEX firms. Because issue costs increase as firm size decreases, their findings suggest that support for a firm size effect may be due to issue costs. Like Hull and Pinches (1994/1995), they note collinearity problems when firm size and issue costs variables are used together in regression analysis. The high degree of correlation between these variables prevents regression tests from giving definitive conclusions concerning the role of differential information as proxied by firm size. Both Hull and Pinches (1994/1995) and Hull and Kerchner (1996) leave it to future research to decide the extent that an issue costs effect, as opposed to a differential information effect, explains announcement period returns for stock offerings.

## Sample and Methodology

We use the *Investment Dealers' Digest* (*IDD*) and *The Wall Street Journal* (*WSJ*) as our primary sources for common stock offering announcements.[3] Major sources for the descriptive statistics are: the latter two sources, Compustat annual files, proxy statements, Moody's Industrial Manuals, and CRSP price files. The years covered by the sources are from 1970 to 1989.

## Sample

For our initial tests (which do not require observations to have issue costs data), we examine 725 observations where a firm announces a common stock offering. Following the general criteria used by Hull and Pinches (1994/1995), we require observations to satisfy these five conditions.

**Table 1 -- Descriptive Statistics for 725 Common Stock Offerings that Reduce Nonconvertible Debt, 1970-1989**

| Descriptive Data | Total Sample (n=725) | Small Firm[a] Size Half (n=362) | Large Firm[b] Size Half (n=363) |
|---|---|---|---|
| **Panel A: Time Profile** | | | |
| Observations for 1970-1973 | 98 (14%[c] | 74 (20%) | 24 (7%) |
| Observations for 1974-1976 | 57 (8%) | 37 (10%) | 20 (6%) |
| Observations for 1977-1979 | 54 (7%) | 45 (12%) | 9 (3%) |
| Observations for 1980-1982 | 270 (37%) | 114 (32%) | 156 (43%) |
| Observations for 1983-1985 | 210 (29%) | 85 (24%) | 125 (34%) |
| Observations for 1986-1989 | 36 (5%) | 7 (2%) | 29 (8%) |
| **Panel B: Selected Characteristics** | | | |
| Market Value of Common Stock[d] | $742M[e] | $71M | $1412M |
| | ($206M) | ($60M) | ($761M) |
| Firm Value[f] | $1497M | $128M | $2864M |
| | ($329M) | ($95M) | ($1566M) |
| Planned Offering Size[g] | $40M | $13M | $68M |
| | ($17M) | ($10M) | ($36M) |
| Par Value of the Planned Debt Reduction Divided by Firm Size | -8.5% (-6.3%) | -13.0% (-11.0%) | -4.0% (-2.9%) |
| Planned Percentage Change in Outstanding Common Stock | 14.5% (11.3%) | 21.7% (18.3%) | 7.3% (5.3%) |

[a] The small firm size half contains those observations in the total sample with the smallest firm sizes.
[b] The large firm size half contains those observations in the total sample with the largest firm sizes.
[c] For the last three columns in Panel A, the parenthesis gives the percentage of the column's total number of observations.
[d] The stock price the day before the announcement times the number of shares outstanding at that time.
[e] For the last three columns in Panel B, the mean is reported in the first row with the median given below in parenthesis. M represents millions.
[f] Firm value includes the market value of common stock, the liquidation value of preferred stock (if applicable), and the book value of long-term debt obligations and current liabilities. Values are taken from sources nearest (yet prior) to the announcement date.
[g] The price the day before the announcement times the planned number of new or primary shares.

1. *The common stock offering must be a pure leverage decrease where the purpose is to reduce nonconvertible debt.*

2. *The firm must not be identified as a utility (because the announcement by a utility is most likely known in advance).*

3. *There must be available data from the sources to calculate values for firm size and for the relative changes in the amounts of stock and debt.*

4. *Each observation's common stock must be listed on the CRSP Return Files and be traded during its announcement and comparison periods.*

5. *The planned percentage change in outstanding common stock must lie between 0.5 percent and 100 percent.*[4]

We find that 78 of the 725 observations have other firm-specific announcements in the *WSJ Index* during event days -3 through +3. In addition, there are 15 observations where small amounts of cash from the current assets account are used to help reduce debt. We retain these latter 93 observations for analysis, as their exclusion does not alter our findings.

The 725 observations contain 179 OTC, 187 AMEX, and 359 NYSE observations.[5] Of the 725 observations, we find 240 observations where the retired debt is specified (by the financial press) as bank debt, and 236 where the debt is unspecified. In total, there are 476 (240 + 236) observations where the debt is not identified as non-bank debt. The remaining 249 observations in our sample are identified by our sources as reducing non-bank debt. The financial press typically depicts these 249 non-bank debt reductions as *bonds*, *debentures*, or *public debt*. For these non-bank debt reductions, we find 26 OTC firms, 23 AMEX firms, and 200 NYSE firms. In contrast, for the 240 bank debt reductions, we have 90 OTC firms, 80 AMEX firms, and 70 NYSE firms.

Additionally, our sample of 725 stock offerings includes 131 offerings that we classify as combination offerings. Consistent with prior research, we categorize an offering as a combination offering if the primary portion of the offering is over 90 percent of the total offering (primary plus registered secondary components). Combination offerings are typically OTC firms (n=60) or AMEX firms (n=57) and rarely NYSE firms (n=14). For the 131 combination offerings, the secondary portion divided by the total offering averages more than 1/3. This proportion suggests that insiders can achieve significant changes in ownership proportions by selling shares through the secondary offering.

Next, of the 725 offerings, we identify 203 non-cash offerings (where cash is not raised for the debt reduction but the new shares are traded for outstanding debt). The financial press refers to non-cash offerings as either *private swap* or *exchange offers*. In our sample, we have 192 private swaps and 11 exchange offers. Large firms typically undertake the 203 non-cash offerings. This explains why non-cash offerings are rarely OTC firms (n=8) or AMEX firms (n=13).

Table 1 reports descriptive statistics for our total sample and its small half and large half.[6] Observations in the small half consist of the 362 observations with the smallest firm sizes. The remaining 363 observations have the largest firm sizes and form the large half. We use the market value of common stock to proxy for firm size.[7]

The time profile in Panel A of Table 1 reveals that 37 percent of the total sample occurs between the years 1980 and 1982. These years contain 32 percent of the small half and 43 percent of the large half. The panel also shows that observations for the large half are more likely to be found in the 1980s where 84 percent of the large half and 58 percent of the small half occur. The greater number of observations for the large half for the 1980s is explained by the fact that 192 private swaps (which are predominantly large companies) occur only during this period.

Panel B reports that the small half's means for the *market value of common stock* and *firm value* are about one-twentieth of the corresponding large half's means. The mean for the *planned offering size* for the small half is approximately one-fifth of the mean for the large half. This panel also reveals that observations in the small half undergo greater changes in levels of either debt or common stock. For example, the small half's means for *par value of the planned debt reduction divided by firm value* and *planned percentage change in outstanding common stock* are about three times greater in amount than the respective means for the large half.

**Methodology**

We follow the ordinary least squares (OLS) market model procedure given by Brown and Warner (1985). We test the hypothesis that a sample's mean daily abnormal stock return (AR) or cumulative abnormal stock return (CAR) is equal to zero.[8] We use the equally weighted CRSP NASDAQ and CRSP AMEX/NYSE indices for respective OTC and AMEX/NYSE observations when calculating alpha and beta parameters. The comparison period consists of the 200 days from +41 to +240 after the announcement date (e.g., after event day 0). We choose a post-announcement comparison period because a stock issue generally occurs after a stock price run-up. Although not reported, similar ARs and CARs are generated when using Scholes and Williams (1977) OLS alphas and betas, a 200-day comparison period before the announcement dates, or the CRSP value-weighted indices.[9]

Our alternative hypothesis is that negative news announcements by a sample of small firms will be more negative than news announcements of a sample of large firms. Thus, we must disprove the null hypothesis that the mean return for a small firm sample is less negative or equal to the mean return for a large firm sample. To test this hypothesis, we calculate a standard parametric $t$ statistic for testing the equality of the means of two non-paired samples. A negative test statistic significant at the conventional 5 percent level for the one-tailed test supports the alternative hypothesis that a small firm sample has a greater negative return (and thus a firm size effect is present). When calculating $t$ statistics, sample stock return variances are assumed unequal if $F$ values reject the hypothesis that variances are equal. To ascertain if outliers are a problem, we also compute non-parametric Wilcoxon rank-sum $z$ statistics.

For the OLS regression tests, we report $t$ statistics, $F$ values, and $R^2$ values (both unadjusted and adjusted). We give one-tailed $t$ statistics for explanatory variables because each variable represents a theory that has a definite prediction concerning the sign of its coefficient. Our results are qualitatively unchanged if the White (1980) heteroskedasticity adjustment procedure is followed when calculating $t$ statistics. We find no evidence that outliers for explanatory variables determine our results.

**Table 2 -- Abnormal Stock Return Results for 725 Common Stock Offerings that Reduce Nonconvertible Debt, 1970-1989**

| Event Day(s) | Total Sample (n=725) | Small Firm Size Half[a] (n=362) | Large Firm Size Half[b] (n=363) | Small Half Versus Large Half |
|---|---|---|---|---|
| **Panel A: Daily Abnormal Return Results** | | | | |
| -3 | 0.11%; 1.29[c]<br>50%;-0.26 | 0.03%; 0.23[c]<br>48%;-0.74 | 0.19%; 1.75[c]<br>51%; 0.37 | -0.93;694[d]<br>-1.56 |
| -2 | -0.18%;-1.90<br>43%;-3.53** | -0.30%;-1.98*<br>39%;-4.42** | -0.05%;-0.46<br>48%;-0.58 | -1.39;630<br>-2.54** |
| -1 | -0.41%;-4.21**<br>43%;-3.97** | -0.46%;-2.79**<br>43%;-2.84** | -0.35%;-3.51**<br>43%;-2.78** | -0.57;594<br>-0.41 |
| 0[e] | -1.61%;-13.94**<br>28%;-11.70** | -1.99%;-10.25**<br>27%; -8.83** | -1.21%;-10.04**<br>30%; -7.72** | -3.33**;605<br>-2.85** |
| 1 | -0.67%;-6.03**<br>40%;-5.39** | -0.91%;-4.81**<br>38%;-4.52** | -0.44%;-3.72**<br>42%;-3.10** | -2.12*;602<br>-1.87* |
| 2 | -0.10%;-1.06<br>45%;-2.56** | -0.12%;-0.80<br>45%;-1.89 | -0.07%;-0.72<br>45%;-1.73 | -0.28;627<br>-0.27 |
| 3 | 0.10%; 1.14<br>50%;-0.11 | 0.17%; 1.19<br>50%; 0.11 | 0.02%; 0.25<br>49%;-0.26 | 0.87;607<br>0.69 |
| **Panel B: Two-Day Cumulative Abnormal Return Results** | | | | |
| 0,+1 | -2.28%;-15.06**<br>26%;-12.89** | -2.90%;-11.64**<br>24%; -9.88** | -1.65%;-9.95**<br>28%;-8.35** | -4.09**;632<br>-3.67** |

Note: Two asterisks (**) and one asterisk (*) denote significance at the 1 and 5 percent levels, respectively.

[a] The small firm size half contains those observations in the total sample with the smallest firm sizes.

[b] The large firm size half contains those observations in the total sample with the largest firm sizes.

[c] The top row for each pair of rows reports the mean abnormal (or cumulative abnormal) stock return followed by the *t* statistic (when testing if the mean return equals zero). The bottom row gives the percent of the sample that has a positive abnormal (or cumulative abnormal) stock return followed by the binomial *z* statistic (when testing if the percent equals 50 percent). Although it is arguably more proper to report one-tailed significant levels (because we expect negative returns), we report traditional two-tailed significant levels in the first three columns. Regardless, the results in these columns are generally robust as to whether one-tailed or two-tailed tests are used.

[d] The top row for each pair of rows reports the one-tailed parametric *t* statistic (when testing the null hypothesis that stock returns for the small group are less negative or equal to returns for the large group) followed by the degrees of freedom. The bottom row reports the *z* statistic for the one-tailed non-parametric Wilcoxon rank-sum test.

[e] Day 0 is the announcement day. Assuming no leakage or late reporting, the announcement is expected to impact the market on day 0 or, if the announcement occurs after the market is closed, on day +1.

## Event Study Results

For brevity's sake, we direct the focus of this section on tests for our total sample, small firm size half, and large firm size half. Nonetheless, our firm size findings hold even when we examine samples consisting solely of AMEX firms or solely of NYSE firms. We offer evidence (at least for samples where firm sizes are more heterogeneous) to show that a firm size effect cannot always be explained by greater issue costs for small firms.

### Abnormal Return Results for Total Sample

Panel A in Table 2 reports daily abnormal stock return (AR) results for a period of seven event days from days -3 through +3. The total sample column reveals that the announcement day (day 0) has the greatest AR at -1.61 percent. We also discover significant negative activity on day +1 where the AR is -0.67 percent (reflecting the fact that some of our firms make announcements after the market has closed on day 0). Additionally, we find evidence of leakage (or late reporting), as negative ARs exist for event days -2 and -1.

The small firm size half and large firm size half columns in Panel A demonstrates that the negative market response is most evident for the small half. These columns indicate that substantial differences in ARs exist between small and large firms for event days 0 and +1. The last column shows statistically significant support for our alternative hypothesis for these two days--differences in ARs between the small and large halves are significant at the 5 percent level or better for both the parametric and non-parametric tests.

Panel B reports two-day CAR results for event days 0 and +1.[10] This panel discloses that there is a -1.25 percent difference when subtracting the mean two-day CAR for the large half (of -1.67 percent) from the mean two-day CAR for the small half (of -2.90 percent). This -1.25 percent difference is significant at the 1 percent level for both the parametric and non-parametric tests ($t$ = -4.09 and $z$ = -3.67). The -1.25 percent difference in two-day CARs between halves decreases to -1.65 percent if we extend the two-day period to the seven days from -3 to +3 (although not reported in Panel B, the seven-day CAR is -3.58 percent for the small half versus -1.91 percent of the large half).

As noted by Hull and Pinches (1994/1995), if a firm size effect is present, then it should be stronger when comparing groups that have greater differences in firm size. We find this to be true. For example, consider the four quartiles ordered in terms of smallest to largest firm sizes: quartile one (n=181), quartile two (n=181), quartile three (n=182), and quartile four (n=181). The four respective mean CARs and firm sizes are: -3.33 percent, $32 million; -2.44 percent, $109 million; -1.92 percent, $418 million; and, -1.40 percent, $2,410 million. In general, the medians for CARs and firm sizes are similar to the means. An exception is for quartile four where the median firm size of $1,387 million is noticeably less than the mean of $2,410 million. We find statistically significant CAR differences (for both the parametric and non-parametric tests) at the 1 percent level when comparing non-bordering quartiles (e.g., quartiles one and three, quartiles one and four, quartiles two and four). We find CAR differences significant at the 5 percent level between quartiles one and two and between quartiles three and four, and at the 10 percent level between quartiles two and three. Consistent with a firm size effect, we find significant differences in CARs at the 1 percent level when comparing non-bordering quartiles where there are greater differences in firm size. Perhaps, the weakest result in support of a firm size effect involves comparing quartiles three and four where a large difference in firm sizes exist. The CAR difference between these quartiles is only significant at the 5 percent level. One interpretation is that a firm size effect becomes increasingly diminished once large firm sizes are reached.

Our examination of halves and quartiles show that the two-day CAR difference of -1.25 percent (when comparing halves) decreases to -1.93 percent (when comparing the smallest and largest quartiles). When we partition the sample into eighths, the two-day CAR difference decreases further to -2.54 percent (-3.65 percent for the smallest eighth versus -1.11 percent for the largest eighth). This -2.54 percent difference between CARs is significant at the 1 percent level ($t$ = -3.97 and $z$ = -3.45).

**Abnormal Return Results for AMEX and NYSE Samples**

The results just given resemble those reported by Hull and Pinches (1994/1995) for OTC firms. Thus, including AMEX/NYSE in the analysis does not alter support for a firm size effect. Because many studies only include AMEX/NYSE firms, we duplicate the above tests for our AMEX/NYSE firms (n=546). For brevity, we do not report these results in detailed table format but simply summarize them as follows. For our AMEX/NYSE sample, the mean two-day CARs for the small half (n=273) and large half (n=273) are -2.48 percent and -1.67 percent, and for the smallest quartile (n=136) and largest quartile (n=136) are -2.93 percent and -1.50 percent. The CAR differences are statistically significant at the 1 percent level or better when comparing either the small and large halves or the smallest and largest quartiles.

We next repeat our tests for both our AMEX sample (n=187) and our NYSE sample (n=359). We find support for our alternative hypothesis when testing each sample. For example, for the AMEX sample, the mean two-day CARs for the small half (n=93) and large half (n=94) are -3.00 percent and -1.96 percent, and for the smallest quartile (n=46) and largest quartile (n=47) are -4.09 percent and -1.66 percent. For the NYSE sample, the mean CARs for the small half (n=179) and large half (n=180) are -2.23 percent and -1.51 percent, and for the smallest quartile (n=89) and largest quartile (n=90) are -2.71 percent and -1.24 percent. For both the AMEX and NYSE tests, CAR differences are statistically significant at the 5 percent level or better when comparing either the small and large halves or the smallest and largest quartiles.

**The Role of Issue Costs**

The above empirical support for a firm size effect appears to be consistent with differential information theory. Stock offering announcements by small firms, however, are more likely to have greater issue costs. If so, a small firm group is more likely to have a greater negative CAR (than a large firm size group) due simply to issue costs. To test if issue costs can explain our firm size findings, we repeat our event study tests after first adjusting for issue costs so that the small and large firm size groups have similar issue costs.

We now describe the process used to equalize issue costs. The first step in our process involves following Hull and Kerchner (1996) in their deletion process. For example, we require each offering to be a completed offering with sufficient information to calculate *ex post* or actual issue costs (namely, cash flotation costs and underpricing). In addition, we only keep observations if the percentage change in outstanding stock value is at least 5 percent. The latter criterion (used by Hull and Kerchner) serves to delete observations where issue costs are not expected to exercise a significant impact on outstanding stock value. After applying these additional screens, our sample size is considerably depleted as only 370 observations remain for analysis.

The descriptive statistics for this sample (n=370) are similar to those given by Hull and Kerchner (1996) for their sample (n=323), and so we do not report them. We have 115 OTC firms, 119 AMEX firms, and 136 NYSE firms for our tests. Because our initial sample has been depleted, we test to see if a firm size effect is still present. After dividing this sample into a small firm size half (n=185) and a large firm size half (n=185), we find that the mean two-day CARs for these two groups are -3.36 percent and -1.92 percent, respectively. When comparing CARs for these two groups, we find statistically significant support for our alternative hypothesis ($t$ = -3.35 and $z$ = -3.13).

After verifying a firm size effect for our current working sample, the next step (in the process to equalize issue costs) involves taking our small firm size half and deleting those one-third observations from our small firm size half (n=62) where the impact of issue costs on outstanding market value of common stock is the greatest. The impact is measured by the procedure used by Hull and Kerchner (1996) as adapted from Hull and Fortin (1993/1994). This procedure multiplies the expected percentage change in outstanding common shares times the expected issue costs per share. The expected issue costs per share is proxied by the *ex post* costs that include both the reported cash flotation costs per share and underpricing per share as calculated by Hull and Kerchner (1996). We then take our large firm size half and delete those one-third observations (n=62) where the impact of issue costs on outstanding market value of common stock is the least. After these deletions, we are left with a small firm size group (n=123) and a large firm size group (n=123) where the negative impact of issue costs is roughly -1.2 percent for each group. In other words, issue costs alone cause the stock value for each group to fall, on average, by about -1.2 percent. When we compare CARs for these two groups, we find that the small firm size half has a CAR of -3.51 percent and the large firm size half has a CAR of -1.99 percent. The over 1.5 percent difference between CARs (even after we equalize issue costs) is significant at the 1 percent level for both the parametric and non-parametric tests ($t$ = 2.91 and $z$ = 2.82). Because these results are similar to the results obtained before equalizing issue costs, it appears that differential information alone (without the aid of issue costs) can explain the CAR differences for our total sample.

We also repeat the above total sample tests for samples based upon listing. Except for the OTC sample or combined samples (e.g., OTC/AMEX, OTC/NYSE, or AMEX/NYSE), our alternative hypothesis could not be accepted at the 5 percent level. This is true with or without issue costs equalized. To understand why we do not find a firm size effect when our AMEX and NYSE samples are tested separately, we examine the difference in firm sizes between the small and large halves and compare these differences to those before the sample is reduced from 725 to 370. We find that the reduction in sample size creates a situation where less variability in firm sizes exists.

For example, consider the three listing samples after issue costs are equalized. The sample sizes are 67 for OTC firms, 80 for AMEX, and 91 for NYSE firms. When comparing these three listing samples with our first three listing samples (where observations were not deleted and for which a firm size effect exists), we find that the average differences in firm size between small and large groups are considerably less. To illustrate, the average difference for OTC firms falls from

$151 million to $86 million. For AMEX firms the decrease is from $628 million to $183 million. For NYSE firms the drop is from $1,889 million to $821 million. Smaller differences in firm sizes between samples tested caused less significant statistics. It seems that the lack of support for a firm size effect (when it occurs) is more likely caused by testing samples with fewer differences in firm sizes than by an issue costs effect. This conclusion (stemming from the comparison of groups with fewer differences in firm sizes) may explain why prior research on AMEX/NYSE equity offerings reports no firm size effect. Simply put, there will be no firm size effect when we examine samples that are relatively homogeneous in terms of firm size.

## Regression Results

We conduct OLS regression tests to further explain the role of firm size. We do this by comparing the wealth effect stemming from firm size with other hypothesized effects (especially those suggested by prior stock offering research). Besides reporting regression results for the total sample, we also give results for AMEX and NYSE samples.

### Regression Model

When conducting our regression analysis, we use the following regression model:

$$\text{CAR} = b_0 + b_1\text{LFS} + b_2\text{LIS} + b_3\text{BAN} + b_4\text{COM} + b_5\text{PCH} + b_6\text{LEV} + E.$$

*CAR is the two-day cumulative abnormal stock return and is expressed in decimal form.*

*LFS is the logarithm of firm size where firm size is proxied by the market value of common stock and is expressed in millions of dollars.*

*LIS is a dummy variable that equals 0 if OTC listed, else 1.*

*BAN is a dummy variable that equals 0 if the stock offering is not identified as reducing non-bank debt, else 1.*

*COM is a dummy variable that equals 0 if the primary offering is over 90 percent of the total offering, else 1.*

*PCH is the planned percentage of total shares expected to be owned by managers after the offering (assuming they will not participate in the offering) divided by their percentage of total shares owned before the offering.*

*LEV is the par value of the planned reduction in debt divided by firm size. (Values for LEV are negative to represent the decrease in leverage.)*

*E is the random error term.*

LFS is selected to capture a differential information effect associated with the amount of available information about a firm. Differential information theory predicts a positive coefficient for LFS. Because market participants know less about small firms, we expect a more negative stock return for a negative news release (such as a stock offering announcement). Hull and Pinches (1994/1995) discover a significant positive coefficient for LFS. LIS controls for listing. However, because AMEX and NYSE firms are larger in firm size than OTC firms, we expect a positive coefficient. Hull and Kerchner (1997) find an insignificant positive coefficient for LIS when examining pure leverage decreases that include junior-for-senior types other than just common stock-for-nonconvertible debt.[11]

The other four explanatory variables (BAN, COM, PCH, and LEV) attempt to test a variety of signaling effects. Testing these four variables is important because prior AMEX/NYSE stock-for-debt research suggests that the signaling effects associated with these variables best explain announcement period CARs.

Signaling models stressing the role of bankers, such as Fama (1985), predict a positive coefficient for BAN. If bank debt offerings impart benign inside information by bankers, then bank debt reductions should convey unfavorable news. For those 476 observations that do not specify the debt reduction as a non-bank debt reduction, investors are more likely to suspect the reductions originate from negative information attained in the bank lending process. Researchers such as Hull and Kerchner (1997) find a significant positive coefficient for BAN or similar variables.

Signaling models premised on changes in inside ownership proportions, such as Leland and Pyle (1977), hypothesize definite coefficient signs for COM and PCH. Such models predict a negative coefficient for COM because the market will be apprehensive that insiders are among those selling secondary shares. The fear that insiders might sell a significant portion of their shares is particularly applicable for the 131 combination offerings in our sample. This is because the secondary sales average about half of the primary sales. COM is similar to the dummy variable used by Masulis and Korwar (1986) and identical to that employed by Hull and Pinches (1994/1995). Each study finds a significant negative coefficient.[12]

Signaling models (that stress inside ownership changes) suggest a positive coefficient for PCH. Greater values are achieved for PCH when smaller suspected changes in inside ownership (and, thus, less negative signaling) exists. Cornett and Travlos (1989) use this variable and find mixed results. The coefficient is insignificant for their pure leverage decrease tests, but significant for their pure leverage increase tests. Hull and Kerchner (1997) find marginal support for PCH.

Signaling models tied to relative changes in debt levels, such as Ross (1977), maintain a positive coefficient for LEV. Because greater negative signaling effects accompany greater negative values for LEV. Greater negative values convey greater investor anxiety about providing adequate cash flows to service a firm's current debt level. The stock offering research (Asquith and Mullins, 1986; Masulis and Korwar, 1986; Cornett and Travlos, 1991; Hull and Moellenberndt, 1994) tests LEV, or similar relative size variables, and finds mixed results.[13]

## Check for Multicollinearity

Before conducting regression tests we perform several tests that check for multicollinearity among our explanatory variables. We calculate correlation coefficients (rhos) to check if pairs of variables present potential collinearity problems. We find that the Pearson and Spearman rhos between LFS and LIS range from 0.22 to 0.49 for the samples that we test. To avoid collinearity we use LFS and LIS separately in our regression tests. We discover that rhos between PCH and LEV are greater than +0.60 for all samples that we test. Thus, we also do not use PCH and LEV together. Furthermore, these latter two variables can be significantly correlated with LFS and LIS for some of our tested samples. Consequently, regression tests using either PCH or LEV are conducted with and without LFS and LIS.

We calculate variance inflation factors (VIFs) for regression tests. For all of our reported tests, VIFs are typically close to one and rarely approach two. As noted by Kennedy (1986), such VIFs are relatively small and thus well below conventional levels for indicating multicollinearity.

## Regression Results for Total Sample

Table 3 reports regression results. Panel A shows results for the total sample when LEV is not used. The first pair of rows in the panel reports results when LFS is tested with COM, BAN, and PCH. The coefficient for LFS has its predicted positive sign and is the only coefficient that is significant at the 1 percent level. Coefficients for BAN and COM have their predicted signs and are significant at the 5 percent level. The coefficient for PCH is insignificant and negative. The negative coefficient (opposite of what is predicted) can be attributed to collinearity between PCH and LFS.

We repeat the test reported in the first pair of rows of Panel A by replacing LFS with LIS. The results are recorded in the second pair of rows. The coefficient for LIS is of its predicted sign but not significant at the conventional 5 percent level.[14] The second pair of rows demonstrate that the coefficients for BAN and COM increase in magnitude and are both significant at the 1 percent level. The coefficient for PCH remains insignificant, but now has its predicted sign. This change in coefficient sign can be explained due to lower correlation, and thus less collinearity, between LIS and PCH than between LFS and PCH (average Pearson and Spearman *rhos* drop to 0.37 from 0.70).

Due to the collinearity that can exist not only between LFS and LIS, but also between either of these two variables with PCH, the last three sets of rows in Panel A report results when each of these three variables is used separately. For these tests, the only variable (among the three) that is significant is LFS. Once again, the coefficient for LFS is positive and significant at the 1 percent level. Coefficients for BAN and COM are significant at the 5 percent level when used with LFS, and significant at the 1 percent level when used with either LIS or PCH.

For all tests reported in Panel A, the $F$ values are significant at the 1 percent level. The largest $F$ and $R^2$ values are found when LFS is used in the tests. The results in Panel A suggest that firm size is the most important explanatory factor and that signaling effects stemming from banker actions and changes in insider ownership proportions are also significant factors. Signaling from changes in insider ownership proportions is not found for PCH, but only for COM (which represents insider sales from secondary offerings).

Our significant finding for COM and insignificant finding for PCH suggest that insiders achieve greater changes in ownership proportion through participating in combination offerings (by selling shares through the secondary offerings) than by not participating in seasoned stock offerings (by not buying new shares). These results are consistent with the make-up of our sample. For example, consider the following illustration that uses share numbers representative of our 725 firms. Assume two firms that are identical with eight millions shares outstanding and each has insiders who own one million of these shares. Further suppose that each firm announces a planned primary offering of one million new shares. If insiders do not buy new shares (which we assume, as the financial press very rarely mentions that they do), then the change in insider ownership proportion for each firm falls from 1/8 to 1/9. The percentage decrease in the proportion is about -11.11 percent. Now assume a registered secondary offering of 0.5 million shares is attached to the primary offering of one million shares for one of the two firms. For this firm suppose that insiders are participating in the secondary offering by selling 0.25 million shares. In this scenario, insider ownership proportion falls from 1/8 to 0.75/9. The percentage decrease in the proportion is now -33.33 percent. This decrease is three times greater than when there is no secondary offering. Thus, we see that combination offerings give insiders a better chance of lowering their proportional holdings. In conclusion, for our sample COM should be a better test of the role of inside ownership proportions than PCH.

Panel B in Table 3 reports results for tests where LEV replace PCH. For these three tests coefficients for LEV (like the previous coefficients for PCH) are insignificant at conventional levels. This finding for LEV does not support the debt level signaling model of Ross (1977). All other regression results reported in Panel B are qualitatively similar to those reported in Panel A.

We also performed the same regression tests (reported in Panels A and B) on the small firm size half (n=362) and large firm size half (n=363). For these tests, there is less variability in firm size. We expect (and find) weaker results for LFS. For example, coefficients for LFS are now only significant at the 5 percent level for all tests. Statistical support for BAN and COM is also weaker. BAN is only significant at the 5 percent level for tests done on the large half, while COM is only significant (at the 5 percent level) for tests conducted on the small half. The insignificant results for COM for the large half should be interpreted cautiously as there are only 16 combination offerings in the large half. Finally, results for LIS, PCH, and LEV remain insignificant when the small and large halves are tested.

**Regression Results for AMEX and NYSE Samples**

We present our results for AMEX and NYSE samples in Panels C and D in Table 3. Results for our OTC sample are not given as these results are reported by Hull and Pinches (1994/1995).[15] Like the latter study's OTC results, we find support for a firm size effect for both AMEX and NYSE samples. Panels C and D report that the coefficients for LFS are negative and significant at the 5 percent level or better for all tests. Our insignificant findings for COM differ from those reported by

**Table 3 -- Ordinary Least Squares Regression Results for Stock Offerings that Reduce Nonconvertible Debt, 1970-1989**

The general model is: CAR = $b_0$ + $b_1$LFS + $b_2$LIS + $b_3$BAN + $b_4$COM + $b_5$PCH + $b_6$LEV + *E*.
CAR is the two-day cumulative abnormal stock return expressed in decimal form.
LFS is the logarithm of firm size (the market value of common stock expressed in millions of dollars.)
LIS is a dummy variable that equals 0 if OTC listed, else 1.
BAN is a dummy variable that equals 0 if the stock offering is not identified as reducing non-bank debt, else 1.
COM is a dummy variable that equals 0 if the primary offering is over 90 percent of the total offering, else 1.
PCH is the planned percentage of total shares expected to be owned by managers after the offering (assuming they will not participate in the offering) divided by their percentage of total shares owned before the offering.
LEV is the par value of the planned reduction in debt divided by firm size. (Values are negative to represent the decrease in leverage.)
*E* is the random error term.

**Panel A: Total Sample (n=725)**

| CONSTANT $b_0$ | LFS $b_1$ | LIS $b_2$ | BAN $b_3$ | COM $b_4$ | PCH $b_5$ | LEV $b_6$ | $R^2$ ($Adj\ R^2$) | $F$ Value |
|---|---|---|---|---|---|---|---|---|
| -0.019 | 0.004 | | 0.007 | -0.007 | -0.031 | | 0.047 | 8.85** |
| -1.07 | 3.25** | | 1.96* | -1.72* | -1.29 | | (0.042) | |
| -0.037 | | 0.002 | 0.008 | -0.011 | 0.013 | | 0.034 | 6.23** |
| -2.17* | | 0.65 | 2.28** | -2.73** | 0.62 | | (0.028) | |
| -0.040 | 0.003 | | 0.006 | -0.008 | | | 0.045 | 11.24** |
| -7.16** | 3.11** | | 1.63* | -2.00* | | | (0.041) | |
| -0.027 | | 0.003 | 0.009 | -0.011 | | | 0.033 | 8.18** |
| -7.86** | | 0.88 | 2.89** | -2.70** | | | (0.029) | |
| -0.038 | | | 0.009 | -0.012 | 0.017 | | 0.033 | 8.17** |
| -2.29* | | | 2.30** | -2.90** | 0.86 | | (0.029) | |

**Panel B: Total Sample (n=725)**

| CONSTANT $b_0$ | LFS $b_1$ | LIS $b_2$ | BAN $b_3$ | COM $b_4$ | PCH $b_5$ | LEV $b_6$ | $R^2$ ($Adj\ R^2$) | $F$ Value |
|---|---|---|---|---|---|---|---|---|
| -0.047 | 0.004 | | 0.006 | -0.008 | | -0.029 | 0.047 | 8.83** |
| -5.97** | 3.31** | | 1.82* | -1.88* | | -1.25 | (0.042) | |
| -0.026 | | 0.003 | 0.009 | -0.011 | | 0.006 | 0.033 | 6.15** |
| -5.67** | | 0.75 | 2.65** | -2.70** | | 0.27 | (0.028) | |
| -0.023 | | | 0.009 | -0.011 | | 0.011 | 0.032 | 8.02** |
| -7.84** | | | 2.71** | -2.87** | | 0.54 | (0.028) | |

Each pair of rows for the explanatory variables' columns report estimated coefficients with OLS *t* statistics below. For columns two through six, we report one-tailed significant levels. Two asterisks (**) and one asterisk (*) denote significance at the 1 and 5 percent levels, respectively.

(Table 3 continued on the next page.)

(Table 3 continued from the previous page.)

| CONSTANT $b_0$ | LFS $b_1$ | LIS $b_2$ | BAN $b_3$ | COM $b_4$ | PCH $b_5$ | LEV $b_6$ | $R^2$ ($Adj\ R^2$) | $F$ Value |
|---|---|---|---|---|---|---|---|---|
| **Panel C: AMEX Sample (n=187)** | | | | | | | | |
| -0.008 | 0.006 | | 0.007 | -0.003 | -0.053 | | 0.040 | 1.88 |
| -0.23 | 2.26* | | 0.74 | -0.33 | -1.14 | | (0.019) | |
| -0.018 | | | 0.008 | -0.009 | -0.005 | | 0.013 | 0.79 |
| -0.52 | | | 0.74 | -1.23 | -0.13 | | (-0.003) | |
| -0.046 | 0.005 | | 0.005 | -0.005 | | | 0.033 | 2.07 |
| -3.72** | 1.96* | | 0.52 | -0.69 | | | (0.017) | |
| -0.024 | | | 0.007 | -0.009 | | | 0.013 | 0.80 |
| -4.87** | | | 0.70 | -1.28 | | | (-0.003) | |
| <u>Panel D: NYSE Sample (n=359)</u> | | | | | | | | |
| -0.064 | 0.003 | | 0.005 | 0.003 | 0.024 | | 0.043 | 3.99** |
| -2.24* | 1.96* | | 1.22 | 0.39 | 0.64 | | (0.032) | |
| -0.080 | | | 0.005 | 0.003 | 0.063 | | 0.033 | 4.01** |
| -2.89** | | | 1.18 | 0.30 | 2.03* | | (0.025) | |
| -0.047 | 0.004 | | 0.006 | 0.004 | | | 0.042 | 5.19** |
| -5.40** | 2.75** | | 1.76* | 0.42 | | | (0.034) | |
| -0.024 | | | 0.009 | 0.003 | | | 0.022 | 2.63* |
| -7.30** | | | 2.58** | 0.34 | | | (0.013) | |

Each pair of rows for the explanatory variables' columns report estimated coefficients with OLS *t* statistics below. For columns two through six, we report one-tailed significant levels. Two asterisks (**) and one asterisk (*) denote significance at the 1 and 5 percent levels, respectively.

Hull and Pinches (1994/1995) for their OTC sample. We can infer that the support for signaling (found in Panels A and B)--from insider sales associated with a secondary offering--is largely attributed to the small OTC listed firms. For OTC observations, the market appears to believe insiders are more likely to unload their outstanding shares.

The results reported in Panels C and D for BAN, like those for COM, are weaker than reported by Hull and Pinches (1994/1995) for OTC firms. The coefficient for BAN is only significant for NYSE tests. Panel D shows that this significance for BAN occurs when PCH is omitted. The lack of significance for BAN (when PCH is jointly used) is caused by collinearity that exists between BAN and PCH for NYSE firms (Pearson and Spearman rhos are both about 0.65). The significant positive coefficients found in Panel D for PCH occur when LFS is not used. This insignificance for PCH when used with LFS stems from collinearity between PCH and LFS (Pearson and Spearman rhos are around 0.4). It is possible that the significance for PCH (when it occurs) is due to an error-in-variables problem where PCH captures wealth effects associated with banker actions and firm size.[16] Collinearity can also explain the reduced significance for LFS when used with PCH (the *t* statistic falls from 2.75 to 1.96). Finally, although not reported in Table 3, LEV is insignificant when used to replace PCH in either the AMEX or the NYSE tests.

In conclusion, our regression results in Panels C and D, together with the findings of Hull and Pinches (1994/1995), indicate several things. First, a firm size effect is not dependent upon the inclusion of OTC firms in the sample. Second, the market's anxiety about changes in insider holdings is attributed largely to OTC firms. Third, the concern due to banker actions is just as likely to be present for NYSE firms as for OTC firms. Finally, signaling, predicated on the relative amount of the debt retired, is not a significant factor regardless of whether we examine OTC firms or AMEX firms or NYSE firms.

**Other Results**

We now summarize the results of other empirical tests. First, like Hull and Pinches (1994/1995), we compare CARs for small and large firm size groups for a variety of samples. These samples include primary offerings, combination offerings, bank debt reductions, non-bank debt reductions, and those designed so that the relative sizes of the leverage change are similar. For all tests, we accept our alternative hypothesis. We find significant parametric and non-parametric support for a firm size effect when comparing CARs.

Next, we try to establish a link between differential information theory and firm size. Because differential information theory predicts a greater negative (positive) CAR if negative (positive) events are announced, we test samples consisting of the 534 observations with negative CARs and the 191 observations with positive CARs. Our tests assume that a negative (positive) CAR corresponds to a negative (positive) event and not just random negative (positive) price behavior. For the negative CAR sample (for which the latter assumption is probably more accurate), the small firm size half (n=267) has a two-day mean CAR of -4.78 percent and the large firm size half (n=267) has a mean CAR of -3.02 percent. The CAR difference is statistically significant at the 1 percent level ($t$ = -6.43 and $z$ = -5.79). For the positive CAR sample, the small half (n=95) has a CAR of 2.52 percent and the large half (n=96) has a CAR of 2.04 percent. The CAR difference is marginally significant approaching the 5 percent level ($t$ = 1.60 and $z$ = 1.47). Consistent with a firm size effect (and independent of the direction of the market response), smaller firms have a greater magnitude than larger firms. When we conduct regression tests on the negative and positive CAR samples, we discover some interesting results. For the negative CAR sample, the explanatory power of our regression model triples. We can attribute this almost totally to LFS as its *t* statistic increases for all tests (becoming as high as 5.54). For the positive sample, our model produces insignificant results. Only for a simple linear regression using LFS do we find a significant statistic. In conclusion, the behavior of the firm size variable is generally consistent with the predictions of differential information theory.

We conduct additional regression tests. For these tests summarized below, we find no evidence that collinearity explains any of the findings. To begin with, we test samples other than those formed according to firm size halves or to listing. For example, we examine samples consisting of combination offerings, primary offerings, bank debt reductions, non-bank debt reductions, cash offerings, and non-cash offerings. For all of these tests, we find statistically significant support for our alternative hypothesis at conventional levels. Like our previously reported regression findings, a firm size effect (as captured by the variable LFS) is the only consistent wealth effect that occurs.

We next analyze explanatory variables other than those reported in Table 3. First, we examine a dummy variable representing the cash versus non-cash nature of the offering. Next, we test leverage ratios (e.g., debt to firm value). We also investigate growth variables (e.g., market value of the firm to book value of assets), risk variables (e.g., standard deviation of stock returns for a 200-day period before the announcements), shifts in risk (e.g., pre-announcement period betas minus post-announcement period betas), and planned offering size. Coefficients for all of the above variables are insignificant at conventional levels.

Next, we conduct regression tests with a variety of variables that capture the change in capital structure (Asquith and Mullins, 1986; Masulis and Korwar, 1986; Cornett and Travlos, 1989; Hull and Moellenberndt, 1994). The results for these variables are similar to those reported for LEV (reflecting their high degree of correlation with LEV). Additionally, we test variables capturing the prior run-up in stock prices for firms in our sample. After controlling for outliers, the coefficients are insignificant at conventional levels. Coefficients are also insignificant for variables that capture the prior run-up in various market indices.

Finally, like Hull and Pinches (1994/1995) and Hull and Kerchner (1996), we conduct regression tests with issue costs as the explanatory variable and find similar results. As documented by Hull and Pinches (1994/1995), the high degree of correlation between issue costs and firm size presents an interpretation problem as either variable can proxy for the other. We recommend that future regression research tackle this problem through a detailed regression analysis that is beyond the scope of this paper.

## Summary

In this study, we test a sample of 725 common stock offerings that retire nonconvertible debt. We find support for our alternative hypothesis, as we discover that small firms experience announcement period stock returns that are significantly more negative than returns for large firms. Support for a firm size effect holds even if we examine just our AMEX firms (n=187) or just our NYSE firms (n=359). Our findings are important because prior research only documents a firm size effect for OTC firms. Consistent with a firm size effect, we find that comparing returns for samples with greater differences in firm size generate higher levels of significance. We also offer evidence to show that issue costs cannot account for a firm size effect, especially when we compare samples where differences in firm sizes are larger.

Using regression tests, we show that a firm size variable does a better job of accounting for stock returns than signaling variables. This regression finding for a firm size effect holds even when AMEX and NYSE firms are examined separately (thus our AMEX and NYSE findings are qualitatively similar to the extant OTC finding). We also offer support for signaling effects stemming from banker actions and changes in insider holdings. The support for these signaling effects is weakened for our AMEX and NYSE tests. For all of our tests, we find no evidence to support signaling tied to the relative amount of the debt retired.

Future research should continue to explore the role of firm size when explaining the market's reaction to security offering announcements. In particular, it needs to examine the relationship between firm size and issue costs. As noted by Hull and Pinches (1994/1995) and Hull and Kerchner (1996), a good starting point when examining this relationship would be senior security offerings (such as debt offerings) where issuance expenses are not as costly. If a firm size effect is found for this sample, then we can offer further evidence for a firm size effect independent of issue costs.

# Notes

[1] A possible exception is Brous and Kini (1994) who examine 379 AMEX/NYSE equity offerings covering the years from 1976 through 1985. Unlike the stock-for-debt sample that we analyze, they appear to focus on offerings that raise cash for purposes other than debt reduction. For their regression tests, they find that a firm size variable has its predicted coefficient sign, but the coefficient is not statistically significant at conventional levels. We do not know if their sample consists of heterogeneous firm sizes. As shown later, the significance of a firm size effect depends upon testing a sample where there is heterogeneity in firm size.

[2] See Hull and Pinches (1994/1995) for a detailed review of the evidence concerning the wealth effects hypothesized for stock offering (especially as pertains to stock-for-debt transactions).

[3] The announcement of a stock offering is typically published by *WSJ* one business day after the date identified and reported by *IDD*. Thus, for those 95 observations for which *WSJ* is the only source for the initial announcement (e.g., either the planning date or the registration date), the business day before the date of publication in *WSJ* is taken as the date for the initial announcement.

[4] Our reported results do not change if we use less extreme parameters such as 2 percent and 50 percent. Later, when we repeat tests used by prior research, we will increase our lower parameter from a 1/2 percent to 5 percent.

[5] The OTC firms in our sample are the same firms analyzed by Hull and Pinches (1994/1995).

[6] For brevity's sake, we do not report the same details for our AMEX sample and our NYSE sample. Regardless, in our empirical sections, we will report results based upon listing.

[7] Freeman (1987) discusses classifying firms based upon size. He notes that prior research (like our study) often uses the market value of common stock to proxy for firm size. The results that we report in this paper are robust for other firm size measurements. Our results are also similar if constant dollars are used.

[8] Because we expect a negative return for stock-for-debt announcements, one-tailed statistics are arguably more appropriate than the traditionally reported two-tailed statistics. Regardless, as will be seen in Table 2, statistics for the two-day announcement period are large generating the same level of significance for either the one-tailed or two-tailed test.

[9] Our findings are even more supportive of our alternative hypothesis if value-weighted returns are used.

[10] For the remainder of the paper, we report two-day CAR results. Our results are qualitatively unchanged if three-day CARs (consisting of days -1, 0, and +1) are used.

[11] We also tested another control variable where LST = 0 if NYSE firms, else LIS = 1. LST is significantly correlated with LIS (rho = -0.57) and performed similarly to LIS.

[12] A negative coefficient for COM is also potentially consistent with agency models, rooted in Jensen and Meckling (1976), as managers (who decrease their residual holdings) are more apt to maximize their wealth at the expense of stockholders. The net agency effect is uncertain because the primary portion of a combination offering reduces debt (and thus can remove negative agency effects caused by onerous debt covenants).

[13] A positive coefficient for LEV is also consistent with tax-based models, such as Modigliani and Miller (1963) that hypothesize greater tax shield losses for greater negative values for LEV.

[14] For a simple linear regression test, the coefficient for LIS is significant at near the 1 percent level. Thus, the insignificant results for LIS may be partially explained by the capacity of other explanatory variables in our model to capture the negative effects (that might otherwise be captured by LIS).

[15] Except for PCH (and LIS which is not applicable), Hull and Pinches (1994/1995) conduct tests similar to ours. We test PCH for OTC firms and find insignificant coefficients.

[16] The significant coefficients for PCH (when collinearity is not present) in Panel D can also suggest that traders of NYSE shares consider the relative amount of the suspected change in insider holdings caused by insiders not purchasing the new shares. However, this interpretation does not explain why OTC and AMEX traders do not react similarly. These traders (as opposed to NYSE traders) are more likely to have larger ownership proportions and thus more likely to be concerned about control issues.

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