Pure Leverage Decreases: A Study of Two Junior-for-Senior Groups

Robert M. Hull
Washburn University

Robert Kerchner
Washburn University

We examine 971 pure leverage decreases consisting of junior-for-senior transactions. This sample—which is unique in size and in the variety of junior and senior forms of borrowing that are included—enables us to extend the junior-for-senior research by testing for a wealth effect linked to adverse selection theory. This theory argues that the market penalizes firms when more junior claims are increased over more senior claims. We find, consistent with this theory, that common-for-nonconvertible debt transactions have significantly more negative announcement period returns than other less extreme junior-for-senior transactions. Regression tests further support an adverse selection effect, indicating that an observation's location within the common-for-nonconvertible debt group is a key explanatory variable in accounting for stock price behavior. We cannot link the significance of this variable to signaling effects premised on changes in fractional insider shareholdings or in relative debt levels.

INTRODUCTION

A pure leverage decrease, such as a junior-for-senior transaction, alters the firm's capital structure mix without directly modifying the asset structure. The new issue is junior—in terms of claims on cash flows—compared to the form of borrowing being retired. The junior claim is more residual and carries less legal power to exact payment. Researchers find significant negative stock returns for firms announcing junior-for-senior transactions.[1] They generally conclude that the negative returns are best accounted for by

[1] Masulis (1983) analyzes 41 noncash offerings (both exchange offers and recapitalizations) that include nine common-for-debt, nine preferred-for-debt, and 23 common-for-preferred transactions. Pinegar and Lease (1986) examine 30 common-for-preferred exchange offers. Cornett and Travlos (1989) investigate 114 noncash offerings (both private swaps and exchange offers) consisting of eight preferred-for-debt and 106 common-for-debt transactions. Copeland and Lee (1991) consider 127 noncash offerings (both private swaps and exchange offers) composed of 88 common-for-debt, 11 preferred-for-debt, and 28 common-for-preferred transactions. Studies consisting entirely of private swaps (e.g., Finnerty, 1985; Owers and Rogers, 1985; and Israel, Ofer, and Siegel, 1989) focus almost exclusively on common-for-debt transactions. Hull

0747-5535/97/1500-0051 $2.50

signaling theory (Fama, 1985; Leland and Pyle, 1977; and Ross, 1977) premised on information asymmetry between insiders and outsiders.

This paper extends the junior-for-senior research by testing for an adverse selection wealth effect. We gather a sample unique in size (n = 971) and in the diversity of junior-for-senior transactions. For example, our transactions consist of both cash and noncash offerings and include five major forms of borrowing. We rank the forms of borrowing according to their claims on cash flows. This ranking system yields two groups of junior-for-senior transactions. The first group consists of common-for-nonconvertible debt transactions. The second group encompasses the other junior-for-senior transactions. For this latter group at least one form of borrowing being exchanged lies (in terms of claims on assets) between common stock and nonconvertible debt. These two groups enable us to test for a wealth effect predicted by adverse selection theory based in Myers (1984) and Myers and Majluf (1984).[2] This theory argues that the market penalizes a firm whenever a more junior form of borrowing is chosen over a more senior form of borrowing. Consequently, common-for-nonconvertible debt transactions are predicted to have more negative stock returns than other less extreme junior-for-senior transactions.

Consistent with adverse selection theory, we find that the common-for-nonconvertible debt group has announcement period returns that are significantly more negative than the other junior-for-senior group. There are potential sample biases, however, that may explain the difference in returns. When we correct for these possible biases, support for our hypothesis (although somewhat weakened) remains statistically significant. Regression tests also support adverse selection theory, as we find that a simple dummy variable associated with an observation's location within the common-for-nonconvertible debt group is a key explanatory variable. We are unable to link the significance of this dummy variable to predictions premised on changes in proportional insider ownership or in relative debt levels. We conclude that the forms of borrowing that are issued and retired convey information beyond that suggested by various mainline signaling theories.

DATA

We identify junior-for-senior announcements between 1970 and 1989 from two sources: *Investment Dealers' Digest* and *The Wall Street Journal*. We obtain descriptive data from Compustat annual files, *Moody's Industrial Manuals*, proxy statements, and the CRSP NASDAQ and CRSP NYSE/AMEX daily price and return files.

We obtain a working sample of 971 observations using the following four screens.[3]

- Each must be a security offering where the stated purpose is to reduce a form of borrowing that is senior to that being issued.

(1994) investigates nine noncash and 147 cash junior-for-senior transactions. (These 156 transactions consist mostly of common-for-debt transactions.) Hull and Moellenberndt (1994) examine 209 noncash and 287 cash common-for-nonconvertible debt transactions. Finally, two studies look only at OTC common-for-nonconvertible transactions. Hull and Fortin (1993/1994) investigate 150 cash transactions, while Hull and Pinches (1994/1995) analyze 12 noncash and 165 cash transactions.

[2] Some extensions of Myers and Majluf (1984) are Narayanan (1988), Noe (1989), Lucas and McDonald (1990), and Choe, Masulis, and Nanda (1994).

[3] We include 108 observations in our sample where *The Wall Street Journal Index* notes other firm announcements (that may or may not impact stock prices) for event days -3 through +3. These observations are included because their deletion does not alter our empirical findings. Like prior research (Masulis, 1983; Hull and Moellenberndt, 1994), we keep observations (n = 23) where sources indicate plans to use cash from the current assets account to partially supplement the reduction of senior securities.

- Each must be listed on the CRSP NASDAQ return file or CRSP NYSE/AMEX return file and have trading data to calculate its cumulative excess return (as described later).
- Each must not be identified as a utility (as the announcement by a utility is typically known in advance).
- Each must have a value between 0.5 and 100 percent when dividing its planned offering value by its equity value. (Equity value is the market value of common stock plus the liquidation value of preferred stock.)[4]

Our pure leverage decrease sample includes five major forms of borrowing that can be ranked—in terms of priority of claims—using the following continuum:

Greatest Junior Security	Common Stock	Convertible Preferred	Nonconvertible Preferred	Convertible Debt	Nonconvertible Debt	Greatest Senior Security

We use this continuum to divide our sample into two junior-for-senior groups based upon the rankings of the junior and senior forms of borrowing for each transaction.

Our first group includes 725 transactions that involve the two forms of borrowing located at the beginning and end of the continuum: common stock and nonconvertible debt. Thus, this first group comprises all of the common-for-nonconvertible debt transactions. Because this group simultaneously involves the greatest junior form and greatest senior form, we refer to it as the *greatest* group. Adverse selection theorists, such as Myers and Majluf (1984), predict that this group will be greeted by a more negative return than other junior-for-senior transactions. Specifically, they assert that it is better to issue safer securities than riskier ones. The Leland and Pyle (1977) and Ross (1977) signaling models make similar predictions, but for different reasons. Leland and Pyle (1977) suggest that greatest transactions signal greater negative news because these 725 transactions (relative to the remaining 246 transactions) are more likely to enable insiders to achieve lower residual ownership proportions. Ross (1977) argues that greatest transactions convey more negative information because managers of these firms are more apt to be lowering their relative debt levels.

We assign the remaining 246 junior-for-senior transactions to a second group. At least one of the forms of financing for this group is ranked between common stock and nonconvertible debt. We refer to this group as the *least* group. This group comprises eight classes of junior-for-senior transactions:[5]

---

[4] We find that our empirical results are unaltered for parameters other than 0.5 percent and 100 percent. For example, a range from 2 percent to 50 percent yields similar findings.

[5] We have zero convertible preferred-for-nonconvertible preferred transactions and 34 junior-for-senior transactions that involve more than two forms of borrowing. (The results that we report do not depend upon the inclusion of these 34 transactions.) We classify each transaction based upon which two forms are dominant. There are 21 of these 34 where one of the two forms is warrants. In each of these 21 cases the warrants are attached to an offering used to retire nonconvertible debt. The 21 junior offerings (with attached warrants) are composed of 11 common, six convertible preferred, and four nonconvertible debt. The latter four are classified as convertible debt. The remaining 13 observations (that include more than two forms of borrowing) consist of eight observations where two different junior forms of borrowing are retiring nonconvertible debt, and five common stock offerings where two different senior forms of borrowing are retired.

27 common-for-convertible debt;
6 common-for-nonconvertible preferred;
6 common-for-convertible preferred;
71 convertible preferred-for-nonconvertible debt;
7 convertible preferred-for-convertible debt;
26 nonconvertible preferred-for-nonconvertible debt;
4 nonconvertible preferred-for-convertible debt; and
99 convertible debt-for-nonconvertible debt.

The least group contains 39 (27 + 6 + 6) offerings where the net amount of common stock is being immediately increased and 135 offerings (27 + 71 + 7 + 26 + 4) where the net amount of debt is being immediately decreased. Thus, given the predictions of the Leland and Pyle (1977) and Ross (1977) signaling models, there is less likelihood that managers of firms in the least group can use transactions to signal negative news predicated on decreases either in fractional residual shareholdings or in debt levels.

Table 1 reports descriptive data for the total sample and the two groups. Panel A reveals that our sample, like the sample of Hull and Moellenberndt (1994), is characterized by the inclusion of combination offerings, bank debt reductions, over-the-counter (OTC) firms, and private swaps.[6] Descriptive statistics for these four characteristics are reported because there is an unequal distribution of these characteristics between the two groups. Unless corrected, this may cause problems for tests that compare announcement period returns between the two groups.[7]

Panel B provides a time profile. Observations for both groups are more prevalent from 1982 to 1985. The larger number of observations for these years is caused by tax legislation that encouraged private swaps between August 1981 and July 1984. The presence of relatively more private swaps for the greatest group explains why this group has 45 percent of its observations during this period compared to 33 percent for the least group.

Panel C reports statistics (mean, median, maximum, minimum, and standard deviation) for seven descriptive variables: firm value, planned offering value, and five relative size variables. The latter five variables are:

PCE = Planned value of primary offering divided by equity value;
SHR = Planned percentage change in outstanding common shares;
SEN = Par value of the planned senior reduction divided by firm value;
LEV = Par value of the planned net reduction in debt divided by firm value; and
PCH = Planned percentage of total common shares expected to be owned by insiders after the announcement divided by their percentage of total common shares owned before the offering.

[6] It is possible for an observation to be a combination offering, a bank debt reduction, and an OTC firm (occurs 32 times). A private swap rarely involves a combination offering (occurs twice) or an OTC firm (occurs 13 times) and is never a bank debt reduction. Like prior research (Hull and Fortin, 1993/1994; Hull and Moellenberndt, 1994), we classify an observation as a combination offering when the primary component of the common stock offering is accompanied by registered secondary component that is at least 10 percent of the combined primary and secondary components.

[7] In our sample we have 33 noncash transactions classified by our sources as exchange offers. This small number makes it more difficult to suspect a sample bias due to exchange offers (e.g., as opposed to private swaps). Nonetheless, we delete these 33 observations and repeat all of our tests. The results are unchanged.

Table 1—Descriptive Data for 971 Junior-for-Senior Transactions, 1970-1989

Descriptive Data	Total Sample (n = 971)	Greatest Group[a] (n = 725)	Least Group[a] (n = 246)
Panel A: Sample Characteristics			
Combination Offerings	138 (14%)[b]	131 (18%)[b]	7 (3%)[b]
Primary Offering	833 (86%)	594 (82%)	239 (97%)
Bank Debt Reductions	300 (31%)	240 (33%)	60 (24%)
Nonbank Debt Reductions	671 (69%)	485 (67%)	186 (76%)
OTC Listing	212 (22%)	179 (25%)	33 (13%)
NYSE/AMEX Listing	759 (78%)	546 (75%)	213 (87%)
Private Swaps	222 (23%)	192 (26%)	30 (12%)
Nonprivate Swaps	749 (77%)	533 (74%)	216 (88%)
Panel B: Time Profile			
1970–1973	145 (15%)[b]	98 (14%)[b]	47 (19%)[b]
1974–1977	97 (10%)	68 (9%)	29 (12%)
1978–1981	274 (28%)	197 (27%)	77 (31%)
1982–1985	406 (42%)	326 (45%)	80 (33%)
1986–1989	49 (5%)	36 (5%)	13 (5%)
Panel C: Descriptive Variables			
Firm Value[c]	1440; 358[d] 31658; 4.4; 3069	1497; 329[d] 31658; 5.7; 3291	1271; 455[d] 15559; 4.4; 2289
Planned Offering Value[e]	40; 19 1209; 0.2; 79	40; 17 1209; 0.2; 85	41; 24 597; 0.4; 56
PCE[f]	14.7; 11.3 95.2; 0.5; 13.8	14.1; 11.1 95.2; 0.5; 13.2	16.6; 12.2 78.3; 0.5; 15.4
SHR[g]	11.3; 7.2 96.6; 0.0; 13.3	14.5; 11.3 96.6; 0.5; 13.4	2.0; 0.0 52.9; 0.0; 7.5
SEN[h]	-8.7; -6.6 -0.1; -53.7; 8.1	-8.5; -6.3 -0.2; -53.7; 8.2	-9.3; -7.1 -0.1; -37.6; 8.0
LEV[i]	-7.4; -5.1 0.0; -53.7; 8.0	-8.5; -6.3 -0.2; -53.7; 8.2	-4.2; -1.0 0.0; -37.4; 6.5
PCH[j]	90.9; 93.3 100.0; 50.9; 9.2	88.4; 89.8 99.6; 50.9; 8.8	98.5; 100.0 100.0; 65.4; 5.4

[a] The greatest group consists of common-for-nonconvertible debt transactions, while the least group consists of those transactions that are not common-for-nonconvertible debt transactions

[b] Percentage of column total given in parentheses

[c] Includes values for equity and debt. Equity is the market value of common stock and the liquidation value of preferred stock. Debt is the book value of long-term debt and current liabilities. Values taken from sources nearest, yet prior, to announcement date

[d] Mean and median are reported in the first row. Maximum, minimum, and standard deviation are reported in the second row. Values are in millions of dollars for the first two variables and percentages for the last five variables

[e] For common stock offerings, the value (if not given by sources) is calculated by multiplying the price the day before the announcement times the planned number of new or primary shares

[f] Planned value of primary offering divided by equity value

[g] Planned percentage change in outstanding common shares

[h] Par value of the planned senior reduction divided by firm value

[i] Par value of the planned net reduction in debt divided by firm value

[j] Planned percentage of total common shares expected to be owned by insiders after the announcement divided by their percentage of total common shares owned before the offering

Relative size variables are crucial for our regression tests when discriminating between an adverse selection effect and various signaling effects. These five particular variables represent the variety of relative size variables reported previously (Cornett and Travlos, 1989; Finnerty, 1985; and Hull and Moellenberndt, 1994). Similar to others who have analyzed combinations offerings (Hull and Moellenberndt, 1994; Masulis and Korwar, 1986), the secondary offering component (because it does not impact the capital structure) is excluded when figuring values for relative size variables. Like Cornett and Travlos (1989), we calculate values for PCH under the assumption that insiders do not participate in the primary common stock offering.

Panel C shows that statistics for the greatest and least groups are similar for firm value, planned offering value, PCE, and SEN. Statistics for SHR, LEV, and PCH are less similar between groups. The dissimilarity for SHR and PCH reflects the less likely change in common stock for least transactions (e.g., SHR is zero percent and PCH is 100 percent when common stock is not the junior security). The dissimilarity for LEV reflects the less likely net change in debt for least transactions (e.g., LEV is zero if preferred stock is being retired or if the transaction is convertible debt for nonconvertible debt).

METHODOLOGY AND STATISTICAL TESTS

We use the ordinary least squares (OLS) market model procedure described by Brown and Warner (1985) to test if a sample's announcement period excess return (ER) or cumulative excess return (CER) is significantly different from zero. We calculate alphas and betas using equally weighted CRSP NASDAQ and CRSP NYSE/AMEX market indices for respective OTC and NYSE/AMEX firms. Our estimation period consists of days +41 to +240 after the announcement date (e.g., after event day 0). We choose a post-announcement estimation period because stock issues often occur during (or toward the end) of bull markets.[8]

When testing the null hypothesis that the mean excess return for the greatest group is less negative or equal to the mean return for the least group, we calculate a standard parametric one tailed t statistic for testing the equality of the means of two nonpaired samples. The research hypothesis is that the greatest group will have a greater negative return (e.g., the test statistic will be negative). Thus, a negative t statistic significant at the conventional 5 percent level supports adverse selection theory. When calculating t statistics, variances are assumed unequal if F values reject the hypothesis that portfolio variances are equal. We also give nonparametric Wilcoxon rank-sum z statistics to ascertain if parametric results are driven by outliers.

We use the OLS regression methodology to determine if an observation's location within either the greatest and least group can exercise a significant impact on stock returns. When regressing explanatory variables against CERs, we test the null hypothesis that an explanatory variable's coefficient is equal to zero. OLS robust t statistics, F values, R^2 values (adjusted and unadjusted), and variance inflation factors are calculated.[9]

[8] The empirical results reported in our study are robust to the following methodological variations: alphas and betas using the Scholes and Williams (1977) procedure; a 200 day estimation period consisting of days -220 to -21 before the initial announcement; the CRSP NASDAQ and CRSP NYSE/AMEX value-weighted market indices; and the mean-adjusted and market-adjusted models.

[9] Variance inflation factors are calculated to ascertain multicollinearity possibilities. See Belsley, Kuh, and Welsch (1980) for more details on the relationship between variance inflation factors and multicollinearity.

Robust t statistics are corrected using the White (1980) asymptotically consistent standard error estimates. We figure one tailed t statistics for explanatory variables because each represents a capital structure theory having a definite prediction concerning the sign of the coefficient.

EXCESS RETURN RESULTS

Panel A in Table 2 reports daily excess stock return (ER) results for five event days from days -2 through +2.[10] There is a -2.59 percent mean cumulative excess return (CER) for the total sample for these five days. Most of this negative five day CER occurs on the announcement day where the ER is -1.39 percent. We also find noticeable negative activity on days -1 and +1 where the ERs are -0.38 percent and -0.64 percent. The former can be explained by leakage (or late reporting), while the latter reflects the fact that for some observations the market can be closed at the time the announcement occurs.

The last three columns show that the greatest group has a more negative ER for each of the five days. For day 0 the mean ER for the greatest group is -0.85 percent more negative than the mean ER for the least group. This -0.85 percent difference is significant at the 1 percent level for both the parametric and nonparametric tests ($t = -3.72$ and $z = -3.47$). The similarity of these t and z statistics demonstrates that outliers do not explain the ER differences between groups.

Panel B gives three day CER results when using event days -1, 0, and +1. The CERs for the total sample, greatest group, and least group are all statistically significant from zero at the 1 percent level for both the parametric and nonparametric tests.[11] There is a -1.07 percent difference when subtracting the mean three day CER for the greatest group from that for the least group. Parametric and nonparametric statistics reveal that the CER difference between groups is significant at the 1 percent level ($t = -3.06$ and $z = -3.15$).

We also examine the smaller negative CER for the least group by looking at the CER results for each of its eight junior-for-senior classes. The mean three day CERs and two tailed t statistics (in parentheses) when testing if the CER for each class is equal to zero are given below.

27 common-for-convertible debt	-0.08% (-0.09)
6 common-for-nonconvertible preferred	-1.31% (-1.67)
6 common-for-convertible preferred	-0.45% (-0.34)
71 convertible preferred-for-nonconvertible debt	-1.57% (-2.77)
7 convertible preferred-for-convertible debt	-1.85% (-1.39)
26 nonconvertible preferred-for-nonconvertible debt	-0.47% (-0.63)
4 nonconvertible preferred-for-convertible debt	-8.08% (-1.89)
99 convertible debt-for-nonconvertible debt	-2.15% (-4.25)

The sample size for four of the eight junior-for-senior classes in the least group are too small to make meaningful individual class inferences. Nonetheless, a general inference can be made: there appears to be a more negative market response when the class is convertible preferred-for-nonconvertible debt or convertible debt-for-nonconvertible debt. The

[10] Extending the event period for days immediately beyond the five days reported in Table 2 adds nothing noteworthy to our analysis.

[11] Although we use three day CERs in reporting our results, similar findings occur if two day CERs for days 0 and +1 are used.

Table 2—Excess Stock Return Results for 971 Junior-for-Senior Transactions, 1970-1989

Event Day(s)	Total Sample (n = 971)	Greatest[a] Group (n = 725)	Least[a] Group (n = 246)	Excess Return Comparison Between Groups
Panel A: Daily Excess Return Results				
-2	-0.12%; -1.51[b] 45%; -2.92***	-0.18%; -1.90*[b] 43%; -3.53***	0.04%; 0.25[b] 51%; -0.26	-0.22%; -1.17; 968[c] -1.18
-1	-0.38%; -4.36*** 43%; -4.33***	-0.40%; -4.21*** 43%; -3.97***	-0.31%; -1.56 44%; -1.79*	-0.09%; -0.47; 372 -0.54
0[d]	-1.39%; -14.03*** 30%; -12.36***	-1.61%; -13.94*** 28%; -11.70***	-0.76%; -4.02*** 36%; -4.46***	-0.85%; -3.72***; 967 -3.47***
1	-0.64%; -6.90*** 40%; -6.26***	-0.67%; -6.03*** 40%; -5.39***	-0.54%; -3.37*** 40%; -3.19***	-0.13%; -0.69; 502 -0.31
2	-0.06%; -0.81 46%; -2.79***	-0.10%; -1.06 45%; -2.56***	0.03%; 0.19 46%; -1.15	-0.13%; -0.70; 969 -0.36
Panel B: Three Day Cumulative Excess Return Results				
-1, 0, 1	-2.41%; -15.71*** 28%; -13.90***	-2.68%; -15.25*** 26%; -12.89***	-1.61%; -5.23*** 33%; -5.48***	-1.07%; -3.06***; 969 -3.15***

*** Significant at the 1 percent level
** Significant at the 5 percent level
* Significant at the 10 percent level

[a] The greatest group consists of common-for-nonconvertible debt transactions, while the least group consists of those transactions that are not common-for-nonconvertible debt transactions

[b] The first row reports the mean excess return followed by the traditional two tailed t statistic (when testing if the mean excess return equals zero). The second row gives the percent of the sample that has a positive excess return followed by the two tailed binomial z statistic (when testing if the percent equals 50 percent)

[c] The first row reports the following: the difference when subtracting the mean excess return for the least group from the mean excess return for the greatest group; the nonpaired one tailed t statistic (when testing the null hypothesis that the mean excess return for the greatest group is less negative or equal to the mean excess return for the least group); and the degrees of freedom. The third row reports the z statistic for the one tailed nonparametric Wilcoxon rank-sum test

[d] Day 0 is the announcement day

CERs for these two classes are statistically significant at the 1 percent level—no other class is significant at the 5 percent level (or even the 10 percent level). Assuming eventual conversion into common stock for these two convertible junior issues, both classes can be viewed as consisting of delayed common-for-nonconvertible debt transactions. In this sense, it is not surprising that the statistical CER results for these two classes resemble those for the common-for-nonconvertible debt class (which form the greatest group).

Our results suggest that the market response is more negative when capital structure changes involve common-for-nonconvertible debt transactions. This finding is predicted by adverse selection theory. Due to the nature of our sample characteristics (as described earlier and reported in Table 1), however, an argument can be made that the difference in

CERs between groups is also consistent with the signaling theories of Leland and Pyle (1977), Ross (1977), and Fama (1985).[12]

First, the difference in CERs is consistent with Leland and Pyle (1977) who predict a more negative return if the market perceives insiders are decreasing their common shares. Panel A in Table 1 shows that junior-for-senior transactions in the least group are rarely a combination offering where principal owners (such as insiders) can sell their holdings through a registered secondary offering. Additionally, transactions in the least group seldom involve issuing common stock. Thus, insiders in this group are less equipped to lower their ownership proportions (by not participating) in the new stock issue.

Second, the more negative mean CER for the greatest group is supported by the debt level signaling model of Ross (1977) because (as reported in Panel C in Table 1) this group retires much larger amounts of debt. Finally, the signaling effect hypothesized by Fama (1985) may be a factor because (as seen in Panel A in Table 1) the least group is less likely to reduce nonbank debt. As argued by Hull and Moellenberndt (1994), the market will suspect that bank debt reductions are prompted by negative inside information obtained by bankers in the lending process.

Another explanation for return differences between the greatest and least groups can arise from the fact that (as seen in Panel A in Table 1) the least group consists of relatively fewer OTC-listed firms. OTC observations involve smaller firms associated with less public and private information. For such firms, differential information theory (Verrechia, 1980; Bhushan, 1989) predicts that their announcements will be more unexpected. It follows that a negative event (such as an equity or equity-like offering) will be more surprising and thus more negative for OTC firms. Any differential information effect caused by the prevalence of smaller OTC firms in the least group, however, may be offset because this group also contains relatively fewer private swaps which are larger firms.

To determine if potential sample biases influence the results reported in Table 2, we conduct tests comparing CERs between the greatest and least groups when observations having characteristics that are unevenly distributed between the two groups are eliminated. The last column in Table 3 reveals that the test eliminating combination offerings (test 1) and the test deleting bank debt reductions (test 2) cause test statistics with magnitudes ($t = -2.03$ and $z = -2.16$ for test 1; $t = -2.05$ and $z = -2.06$ for test 2) that are less than those reported in Panel B in Table 2 for the total sample ($t = -3.06$ and $z = -3.15$). Additionally, the 5 percent significance level reported for tests 1 and 2 is less than the 1 percent level given for the total sample.

[12] Although not considered as likely, there are other theories that potentially can explain the different CERs found between groups. For example, optimal capital structure models [e.g., agency models based in Jensen and Meckling (1976) or tax shield/financial distress models such as Kim (1978)] predict a negative effect for common-for-nonconvertible debt transactions if these transactions cause the firm to move further from its optimal debt-to-equity ratio. Additionally, the corporate tax model of Modigliani and Miller (1963) hypothesizes a negative CER if these greatest transactions cause a greater net decrease in debt. (This theory implies that bankruptcy and agency costs do not impact firm value.) Also, the issuance expenses model of Hull and Fortin (1993/1994) predicts greater negative CERs if issue costs are greater for common-for-nonconvertible debt transaction. The negative response increases as the cost per new share and the relative size of the offering increases. Finally, wealth redistribution theory (Galai and Masulis, 1976) predicts a greater negative CERs if greatest transactions cause greater shifts in wealth from junior claimants to senior claimants.

Table 3—Announcement Period Stock Return Results for Various Samples of Junior Security Offerings that Reduce Outstanding Senior Forms of Borrowing

Test 1 deletes the 138 combination offerings. Test 2 eliminates the 300 bank debt reductions. Test 3 removes the 212 OTC-listed firms. Test 4 omits the 222 private swaps. Test 5 deletes the 676 observations with at least one of the following characteristics: a combination offering, a bank debt reduction, an OTC listing, or a private swap

Test	Total Sample (n = 971)	Greatest[a] Group (n = 725)	Least[a] Group (n = 246)	Excess Return Comparison Between Groups
Test 1	(n = 833) -2.16%; -13.56***[b] 29%; -12.23***	(n = 594) -2.36%; -12.84***[b] 27%; -11.00***	(n = 239) -1.65%; -5.29***[b] 32%; -5.50***	-0.71%; -2.03**; 831[c] -2.16**
Test 2	(n = 671) -2.06%; -11.51*** 30%; -10.31***	(n = 485) -2.29%; -10.87*** 29%; -9.40***	(n = 186) -1.47%; -4.35*** 34%; -4.40***	-0.82%; -2.05**; 669 -2.06**
Test 3	(n = 759) -2.22%; -14.02*** 27%; -12.45***	(n = 546) -2.45%; -13.47*** 26%; -11.21***	(n = 213) -1.61%; -5.15*** 31%; -5.55**	-0.84%; -2.41***; 757 -2.30**
Test 4	(n = 749) -2.62%; -14.34*** 28%; -12.31***	(n = 533) -3.01%; -14.06*** 25%; -11.48***	(n = 216) -1.64%; -4.83*** 33%; -4.90***	-1.37%; -3.45***; 747 -3.68***
Test 5	(n = 295) -1.97%; -7.91*** 31%; -6.70***	(n = 163) -2.38%; -7.22*** 28%; -5.56***	(n = 132) -1.46%; -3.88*** 33%; -3.83***	-0.92%; -1.85**; 293 -1.74**

*** Significant at the 1 percent level
** Significant at the 5 percent level
* Significant at the 10 percent level

[a] The greatest group consists of common-for-nonconvertible debt transactions, while the least group consists of those transactions that are not common-for-nonconvertible debt transactions

[b] The first row reports the mean three day CER followed by the traditional two tailed t statistic (when testing if the mean three day CER equals zero). The second row gives the percent of the sample that has a positive mean three day CER followed by the two tailed binomial z statistic (when testing if the percent equals 50 percent)

[c] The first row reports the following: the difference when subtracting the mean three day CER for the least group from the mean three day CER for the greatest group; the nonpaired one tailed t statistic (when testing the null hypothesis that the mean three day CER for the greatest group is less negative or equal to the mean three day CER for the least group); and the degrees of freedom. The third row reports the z statistic for the one tailed nonparametric Wilcoxon rank-sum test

The last column of Table 3 shows that omitting OTC firms (test 3) or private swaps (test 4) gives statistics significant at the 1 percent level. The decrease in the magnitude of statistics for test 3 and the increase for test 4 are consistent with a differential information effect related to the size of the announcing firm. For example, differential information theory suggests that announcements by smaller OTC firms are more revealing (and thus more negative); similarly, the announcements by larger firms, e.g., firms undergoing private swaps, are less revealing (and thus less negative).

Test 5 simultaneously eliminates all observations (combination offerings, bank debt reductions, OTC firms, and private swaps) that can bias our tests. Omitting these observations produces a mean CER difference between groups of -0.92 percent. This difference

is associated with test statistics (t = -1.85 and z = -1.74) that are significant at the 5 percent level. The -0.92 percent difference is similar to the -1.07 percent difference reported for the total sample in Panel B in Table 2. Thus, deleting all observations that cause potential sample biases not only produces significant statistics, but also does not substantially alter the CER difference between groups. In conclusion, we find that potential biases cannot overturn our support for an adverse selection effect.

REGRESSION RESULTS

We use the OLS regression methodology to further examine the implication (in terms of market response) of an observation's location within the two groups. Our general regression model is:

$$CER = b_0 + b_1TYP + b_2RSZ + b_3COM + b_4BAN + b_5LIS + b_6SWA$$

where:

CER	=	Three day cumulative excess return for event days -1, 0, and +1;
TYP	=	1 if least transaction (e.g., not common-for-nonconvertible debt), else 0;
RSZ	=	Relative size variable (see Panel C in Table 1) that may be included in the test;
COM	=	1 if combination offering, else 0;
BAN	=	1 if senior reduction not identified as a bank debt reduction, else 0.
LIS	=	1 if listed on CRSP NYSE/AMEX return file, else 0; and
SWA	=	1 if a private swap, else 0.

The dummy variable TYP is tested to capture the impact of an observation's location within the greatest and least groups. Both adverse selection theory and signaling theory (predicated on decreases in fractional insider shareholdings or in relative debt levels) predict a positive coefficient for TYP. This prediction holds even for tests that delete observations suspected of causing the greatest group to have greater signaling effects due to suspected sample biases.[13]

Leland and Pyle (1977) predict a negative coefficient for COM because insiders are suspected of unloading their shares through the secondary component. This is particularly true for the 138 combination offerings in our sample where the secondary sales average about half of primary sales. Fama (1985) hypothesizes a positive coefficient for BAN as bank debt reductions are likely to convey negative news by bankers.[14] Positive coefficients are expected for LIS and SWA because NYSE/AMEX and private swap observa-

[13] We also test the dummy variable TYP, when TYP = 0 includes not only common-for-nonconvertible debt but also those two classes that have more negative CERs (e.g., the convertible preferred-for-nonconvertible debt and convertible debt-for-nonconvertible debt). We also tried a series of dummy variables representing various groups of classes (as opposed to the two groups for which results are reported) for which differences in CERs exist. The results of these tests produce weaker statistical results than are reported for the simple division into the two groups that we use.

[14] COM is identical to the variable used by Hull and Moellenberndt (1994) and similar to the variable used by Masulis and Korwar (1986). Both studies report significant negative coefficients. BAN is identical to the variable used by Hull and Pinches (1994/1995) and similar to the variable used by Hull and Moellenberndt (1994). Both studies find significant positive coefficients.

tions represent firms that are relatively larger in size. Because more information exists for these firms, their junior-for-senior announcements should be less revealing (and thus less negative).

Because greater absolute magnitudes for the RSZ variables are expected to be associated with greater negative signaling (as hypothesized by Leland and Pyle, 1977 and Ross, 1977), a negative coefficient is predicted when RSZ is PCE or SHR. Similarly, a positive coefficient is hypothesized when RSZ is SEN, LEV, or PCH.[15] Replacing TYP with RSZ variables is important to help ascertain if any significant coefficients for TYP can be attributed to the fact observations in the greatest group undergo greater capital structure changes, particularly those related to changes in residual insider ownership and outstanding debt. It is also important to test RSZ variables without TYP included as these variables can be significantly correlated with TYP. This significant correlation is especially true when RSZ is SHR, LEV, or PCH. (These are the three variables that focus on changes in common stock or debt.) For the total sample Pearson and Spearman rhos between TYP and any of the latter three RSZ variables are all greater than 0.22 and significant at the 1 percent level. Finally, RSZ variables should not be used together as rhos between any two of these five variables are all greater than 0.45.

Table 4 reports the robust OLS regression results.[16] Panel A gives results when the total sample is tested and an RSZ variable is not used. Although the adjusted and unadjusted R^2 values are modest, the F value of 6.61 is significant at the 1 percent level. Variance inflation factors are below values conventionally accepted as signifying multicollinearity. All coefficients have their predicted signs, with TYP and COM significant at the 1 percent level and BAN significant at the 5 percent level.

The significant positive coefficient for TYP supports the adverse selection notion that the forms of borrowing being issued and retired (as represented by our classification into either the greatest group or the least group) influences announcement period returns. The results for TYP also support the signaling models of Leland and Pyle (1977) and Ross (1977). The results for COM and BAN are consistent with the signaling models of Leland and Pyle (1977) and Fama (1985). The coefficient signs for LIS and SWA are consistent with a differential information effect related to firm size. LIS is not significant, however, while SWA is only marginally significant at the 10 percent level.

The next three panels demonstrate that the introduction of RSZ variables does not substantially alter any of the statistical results given in Panel A. When RSZ variables are used with TYP, the first rows for Panels B, C, and D report that coefficients for PCE, LEV, and PCH are all insignificant. When TYP is deleted, the second rows for Panels B, C, and D show that PCE and LEV remain insignificant, while PCH is marginally signif-

[15] RSZ variables traditionally are used in regression tests by pure leverage decrease studies (Cornett and Travlos, 1989; Finnerty, 1985; and Hull and Moellenberndt, 1994) as well as other security offering studies (Masulis and Korwar, 1985; and Mikkelson and Partch, 1986) where the purpose of the offering involve changes in either the asset structure or the capital structure (or both). For both sets of studies RSZ variables are hypothesized to capture some sort of negative signaling effect (or effects). It is possible for RSZ variables also to capture agency effects. For our pure leverage decrease tests, however, the nature of the agency effect is hard to predict, e.g., a negative agency effect from insiders lowering their residual ownership proportions can be offset by a positive agency effect from reducing onerous debt covenants.

[16] The reported regression results are similar if we use other methodological variations. These variations include standardized CERs, two day CERs consisting of event days 0 and +1, standard OLS t statistics, and weighted least squares. We find no evidence that outliers (e.g., values that are three standard deviations removed from the average value) for relative size variables alter our findings.

Table 4—Regression Results for 971 Junior-for-Senior Transactions, 1970-1989

The general regression model is:

$$CER = b_0 + b_1TYP + b_2RSZ + b_3COM + b_4BAN + b_5LIS + b_6SWA$$

where:

CER = The three day cumulative excess return for event days -1, 0, and +1;
TYP = 1 if least transaction (e.g., not common-for-nonconvertible debt), else 0;
RSZ = The relative size variable (see Panel C in Table 1) that may be included in the test;
COM = 1 if combination offering, else 0;
BAN = 1 if senior reduction not identified as a bank debt reduction, else 0;
LIS = 1 if listed on CRSP NYSE/AMEX return file, else 0; and
SWA = 1 if a private swap, else 0.

Constant b_0	TYP b_1	RSZ b_2	COM b_3	BAN b_4	LIS b_5	SWA b_6	R^2 (Adj R^2)	F
Panel A: Without RSZ								
-0.033	0.009		-0.013	0.008	0.003	0.004	0.033	6.61***
-7.24***	2.28***		-2.54***	2.10**	0.59	1.26*	(0.028)	
0.00	1.12		1.14	1.19	1.11	1.30		
Panel B: RSZ = PCE								
-0.032	0.009	-0.004	-0.013	0.008	0.002	0.004	0.033	5.52***
-5.56***	2.32***	-0.28	-2.56***	2.09**	0.48	1.01	(0.027)	
0.00	1.12	1.31	1.14	1.19	1.20	1.45		
-0.029		0.014	-0.015	0.009	0.003	0.001	0.028	5.62***
-5.46***		0.70	-3.02***	2.43***	0.51	0.34	(0.023)	
0.00		1.32	1.09	1.17	1.22	1.36		
Panel C: RSZ = LEV								
-0.034	0.009	-0.005	-0.013	0.008	0.003	0.005	0.033	5.51***
-6.01***	2.27***	-0.21	-2.54***	2.10**	0.61	1.27*	(0.027)	
0.00	1.19	1.35	1.14	1.19	1.22	1.41		
-0.030		-0.008	-0.015	0.008	0.003	0.002	0.028	5.56***
-5.58***		-0.37	-2.97***	2.42***	0.64	0.51	(0.023)	
0.00		1.26	1.09	1.17	1.22	1.28		
Panel D: RSZ = PCH								
-0.039	0.008	0.008	-0.013	0.008	0.002	0.004	0.033	5.52***
-2.16**	1.89**	0.38	-2.56***	2.09**	0.50	1.00	(0.027)	
0.00	1.56	1.73	1.14	1.19	1.18	1.52		
-0.055		0.023	-0.015	0.009	0.002	0.001	0.030	5.56***
-3.24***		1.44*	-2.88***	2.29***	0.44	0.22	(0.025)	
0.00		1.23	1.09	1.18	1.17	1.28		

*** Significant at the 1 percent level
** Significant at the 5 percent level
* Significant at the 10 percent level

In the first seven columns the first row reports estimated coefficients; the second row gives robust t-statistics using the White (1980) heteroskedasticity adjustment procedure; and the third row records variance inflation factors. Except for the constant term (CONSTANT), the t test is one tailed. The next to the last column reports unadjusted and adjusted R^2 values (the latter in parentheses), while the final column gives F values

icant at the 10 percent level. The latter finding offers some evidence that TYP may capture changes in ownership proportions of insiders. We would expect this, given that insiders for firms in the least group are not typically positioned to use the junior issue to change their common stock ownership proportions. Although not reported in Table 4, the results for SHR and SEN (like those for PCE and LEV) are also insignificant with small magnitudes for their respective coefficients. This is true when tested with or without TYP present. In conclusion, the results for the relative size tests indicate that the significant support for TYP is due to adverse selection theory (as opposed to signaling theory associated with changes in insider holdings or in debt levels).[17]

We conduct other regression tests. For example, we test the sample used for test 5 reported in Table 3 (where combination offerings, bank debt reductions, OTC firms, and private swaps are deleted). Similar results are found for TYP and RSZ variables with the exception that PCH is no longer significant. We perform other deletion tests. For example, the deletion of both combination offerings and bank debt reductions (the two deletions mostly likely to cause unequal signaling) leaves our results generally unchanged, e.g., TYP retains its significance while RSZ variables (including PCH), LIS, and SWA are insignificant. The results of these tests call into question our earlier (weak) inference that TYP may capture changes in ownership proportions of insiders.

The relative size results are somewhat surprising given the significant correlation that exists between TYP and RSZ variables. Because relative size variables (which are traditionally hypothesized to capture signaling effects) are not an important explanatory factor, it appears the forms of borrowing being issued and retired exercise some sort of combined effect that the market perceives as more important than the sheer relative magnitude of the capital structure change.[18] In a nutshell, our results suggest (with or without potential sample biases removed) a dummy variable representing the greatest and least groups in itself influences the market's response to pure leverage decrease announcements. Thus, we conclude that the wealth effect suggested by adverse selection theory is a significant and important determinant of our announcement period stock price behavior.

[17] The results for TYP are potentially consistent with the free cash flow theory of Jensen (1986). This theory predicts a positive coefficient because a wealth effect decreases in the looseness of the commitment bonding the payment of future cash flows. Replacing nonconvertible debt with common stock is the junior-for-senior class most apt to cause managers to squander excess cash on poor investment opportunities. The focus of the free cash flow theory is the magnitude of a firm's undistributed cash flow beyond that required to fund projects with positive net present value. The magnitude of this undistributed or free cash flow is hypothesized to be positively related with managerial inefficiency. Thus, the free cash flow theory predicts that firms with greater relative magnitudes of free cash flow will have greater negative CERs for leverage decrease announcements. When testing free cash flow variables suggested by extant research (Lehn and Poulsen, 1989; Gupta and Rosenthal, 1991), we find insignificant results. Consequently, we conclude that our results for TYP are explained by adverse selection theory as opposed to free cash flow theory.

[18] RSZ variables are also generally insignificant when we use our general model to separately test the greatest and least groups. For these tests RSZ variables exhibit significance when combinations of other explanatory variables (especially COM and BAN) are deleted from tests. The latter result is generally true for similar tests conducted on our total sample. Thus, a significant coefficient finding for an RSZ variable (unlike TYP) appears to hinge upon its capacity to proxy for effects captured by other variables. If so, researchers should be aware of the potential error-in-variables problem when using relative size variables.

SUMMARY AND CONCLUSIONS

We examine 971 pure leverage decreases consisting of junior-for-senior transactions. Our sample is unique in its size and the variety of classes of junior-for-senior transactions included. Our sample includes five major forms of borrowing ranked in terms of priority of claims. Two groups are formed based upon our rankings. The greatest group comprises the 725 common stock offerings that reduce nonconvertible debt. The least group includes the remaining 246 offerings that are not common-for-nonconvertible transactions.

As suggested by adverse selection theory, we hypothesize that announcement period stock returns for the greatest group will convey more negative information. Testing this hypothesis may be biased unless we control for sample characteristics that are unevenly distributed between the two groups. We find that the mean announcement period stock return for the greatest group is significantly more negative than the return for the least group. We show that the difference in returns between groups can fall somewhat when tests are corrected for suspected sample biases. In particular, deleting combination offerings and bank debt reductions decreases the unequal signaling between groups. Nonetheless, the potential sample biases appear to explain only a small part of the return difference between groups.

Regression tests offer further evidence that an observation's classification within one of the two groups influences the market response to junior-for-senior announcements. The magnitude of the relative size of the junior-for-senior transaction does not appear to explain excess stock returns. This is true even when correcting for possible sample biases. We conclude that the forms of borrowing being issued and retired convey information that is valuable. Because this information goes beyond that hypothesized by the signaling models of Leland and Pyle (1977) and Ross (1977), we are left to infer that an adverse selection effect is highly instrumental in explaining the different returns between the two groups.

REFERENCES

1. Belsley, D., E. Kuh, and R. Welsch, *Regression Diagnostics* (New York: Wiley 1980).

2. Bhushan, R., "Firm Characteristics and Analyst Following," *Journal of Accounting and Economics*, 11 (1989), pp. 255-274.

3. Brown, S., and J. Warner, "Using Daily Stock Returns: The Case of Event Studies," *Journal of Financial Economics*, 14 (March 1985), pp. 3-31.

4. Choe, H., R. Masulis, and V. Nanda, "Common Stock Offerings Across the Business Cycle," *Journal of Empirical Finance*, 1 (February 1993), pp. 3-31.

5. Copeland, T., and W. Lee, "Exchange Offers and Stock Swaps—New Evidence," *Financial Management*, 20 (Autumn 1991), pp. 34-48.

6. Cornett, M., and N. Travlos, "Information Effects Associated with Debt-for-Equity and Equity-for-Debt Exchange Offers," *Journal of Finance*, 44, no. 2 (June 1989), pp. 451-468.

7. Fama, E., "What's Different about Banks?" *Journal of Monetary Economics*, 15 (January 1985), pp. 29-39.

8. Finnerty, J., "Stock-for-Debt Swaps and Shareholder Returns," *Financial Management*, 14 (Autumn 1985), pp. 5-17.

9. Galai, D., and R. Masulis, "The Option Pricing Model and the Risk Factor of Stock," *Journal of Financial Economics*, 3 (January-March 1976), pp. 53-81.

10. Gupta, A., and L. Rosenthal, "Ownership Structure, Leverage, and Firm Value: The Case of Leveraged Recapitalizations," *Financial Management*, 20 (1991), pp. 69-83.

11. Hull, R., "Stock Price Behavior of Pure Capital Structure Issuance and Cancellation Announcements," *Journal of Financial Research*, 42 (Fall 1994), pp. 439-448.

12. Hull, R., and R. Moellenberndt, "Bank Debt Reduction Announcements and Negative Signaling," *Financial Management*, 23, no. 2 (Summer 1994), pp. 21-30.

13. Hull, R., and R. Fortin, "Issuance Expenses and Common Stock Offerings for Over-the-Counter Firms," *Journal of Small Business Finance*, 3, no. 1 (1993/1994), pp. 1-17.

14. Hull, R., and G. Pinches, "Firm Size and the Information Content of Over-the-Counter Stock Offerings," *Journal of Small Business Finance*, 4, no. 1 (1994/1995), pp. 31-56.

15. Israel, R., A. Ofer, and D. Siegel, "The Information Content of Equity-for-Debt Swaps," *Journal of Financial Economics*, 25 (December 1989), pp. 249-370.

16. Jensen, M., "Agency Costs of Free Cash Flow, Corporate Finance, and Takeovers," *American Economic Review* 76 (May 1986), pp. 323-329.

17. Jensen, M., and W. Meckling, "Theory of the Firm: Managerial Behavior, Agency Costs and Ownership Structure," *Journal of Financial Economics*, 3, (October 1976), pp. 305-360.

18. Kim, E. Han, "A Mean-Variance Theory of Optimal Capital Structure and Corporate Debt Capacity," *Journal of Finance*, 33, no. 1 (March 1978), pp. 45-63.

19. Lehn, K., and A. Poulsen, "Free Cash Flow and Stockholder Gains in Going Private Transactions," *Journal of Finance*, 44, no. 3 (July 1989) pp. 771-787.

20. Leland, H., and D. Pyle, "Informational Asymmetries, Financial Structure, and Financial Intermediation," *Journal of Finance*, 32, (May 1977), pp. 371-388.

21. Lucas, D., and R. McDonald, "Equity Issues and Stock Price Dynamics," *Journal of Finance* 45, no. 4 (September 1990), pp. 1019-1044.

22. Masulis, R., "The Impact of Capital Structure Change on Firm Value: Some Estimates," *Journal of Finance*, 38 (March 1983), pp. 107-126.

23. Masulis, R., and A. Korwar, "Seasoned Equity Offerings: An Empirical Investigation," *Journal of Financial Economics*, 15 (January-February 1986), pp. 91-118.

24. Mikkelson, W., and M. Partch, "Valuation Effects of Security Offerings and the Issuance Process," *Journal of Financial Economics*, 15 (January-February 1986), pp. 31-60.

25. Modigliani, F., and M. Miller, "Corporate Income Taxes and the Cost of Capital: A Correction," *American Economics Review*, 53 (June 1963), pp. 433-443.

26. Myers, S., "The Capital Structure Puzzle," *Journal of Finance*, 39, no. 3 (July 1984), pp. 575-592.

27. Myers, S., and N. Majluf, "Corporate Financing and Investment Decisions When Firms Have Information that Investors do not Have," *Journal of Financial Economics*, 13 (June 1984), pp. 187-221.

28. Narayanan, M., "Debt Versus Equity Under Asymmetric Information," *Journal of Financial and Quantitative Analysis*, 23 (1988), pp. 39-51.

29. Noe, T., "Capital Structure and Signaling Game Equilibria," *Review of Financial Studies*, 1 (1989), pp. 331-355.

30. Pinegar, J., and R. Lease, "The Impact of Preferred-for-Common Exchange Offers on Firm Value," *Journal of Finance*, 41, no. 4 (September 1986), pp. 795-814.

31. Rogers, R., and J. Owers, "Equity for Debt Exchanges and Shareholder Wealth," *Financial Management*, 14 (Autumn 1985), pp. 18-26.

32. Ross, S., "The Determination of Financial Structure: The Incentive-Signaling Approach," *Bell Journal of Economics*, 8 (Spring 1977), pp. 23-40.

33. Scholes, M., and J. Williams, "Estimating Betas from Nonsynchronous Data," *Journal of Financial Economics*, 5 (December 1977), pp. 309-328.

34. Verrechia, R., "The Rapidity of Price Adjustments to Information," *Journal of Accounting and Economics*, 2 (March 1980), pp. 63-92.

35. White, H., "A Heteroskedasticity-Consistent Covariance Matrix Estimator and a Direct Test for Heteroskedasticity," *Econometrica*, 48 (May 1980), pp. 817-838.