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ABSTRACT

This study investigates the seasonal pattern of aggregate insider trading to help distinguish between two competing explanations for the seasonal pattern of security returns. The first potential explanation examined is that the January effect arises from predictable changes in turn-of-the-year demand for securities. The second potential explanation examined is that the January effect represents compensation for the higher risk of trading against informed traders at the turn of the year.

THIS STUDY EXAMINES AGGREGATE insider[1] trading activity to provide a new insight into the nature of the seasonal pattern of security returns reported by Keim [15] and others.[2] For the period from 1963 to 1979, Keim [15] documents unexpectedly large, positive returns to small firms during the first week of January.[3] The reason for the existence of seasonality in stock returns is not yet well understood, and hence, the January effect has become a paradox for models of equilibrium expected stock returns and the efficient-markets hypothesis.

This study examines two potential explanations for the January effect. One potential explanation is that the large positive return at the turn of the year arises due to price pressure from predictable seasonal changes in the demand for different securities. For instance, the proceeds from tax-loss selling or the payment of annual bonuses at year end result in an increase in cash for individual investors. Outside investors take most of their tax losses in December and then wait until January to buy the stock of other small firms with the proceeds of the sales. The increase in demand (albeit predictable) for small firms at the turn of the year results in an increase in the stock prices of small firms.[4]

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[1] The Securities and Exchange Act of 1934 defines insiders as officers, directors, and shareholders of ten percent or more of any equity class of securities.

[2] Banz [1] and Reinganum [18] document the relation between firm size and average stock returns. See also the special issue of the *Journal of Financial Economics*, 1983, volume 12, entitled *Symposium on Size and Stock Returns and Other Empirical Regularities*, for other papers. Schwert [26] provides an introduction and overview of the papers in this issue. Rozeff and Kinney [25] also provide evidence on stock market seasonality.

[3] Keim [15] finds that, over the period 1963 to 1979, the average difference between the stock returns of the smallest size firms and the stock returns of largest size firms equals fifteen percent in January. Of the fifteen percent difference, approximately eight percent is realized within the first five trading days in January. Blume and Stambaugh [2] and Roll [23] show that, with the returns on buy-and-hold portfolios, the size effect is restricted to January only.

[4] The evidence on the tax-loss selling is mixed. Studies by Branch [4], Reinganum [20], Givoly and Ovadia [11], Rozeff [24], and Ritter [21] provide evidence consistent with tax-loss selling. In contrast, Brown, Keim, Kleidon, and Marsh [5] conclude that the coincidence of the January effect with the end of the tax year in the U.S. is a casual, not a causal, association.

A second potential explanation is that the large positive returns in January for small firms represent compensation for the increased risk of trading against informed traders. Williams [32] provides a formal model where greater variability in outsiders' demand for stock provides greater opportunities for insiders to hide their information-motivated transactions. Glosten and Milgrom [12] conjecture that the positive return in January may represent compensation for the increased risk of trading against informed traders at the turn of the year. It is predicted that informed traders are more likely to possess nonpublic information at the turn of the year and thereby to impose greater losses on all opposing traders by either buying or selling stock.

Increased trading by informed traders is also expected to increase the bid-ask spread in the security. Hence, for the risk-premium hypothesis to be consistent with market efficiency, any profit opportunity from a strategy of buying stock in December and selling stock in January is expected to be less than the costs of trading, including the bid-ask spread cost.

The testable predictions of the two hypotheses are the following. If insiders in small firms respond to the large positive return in January, then increasing their net purchases in December would enable insiders to capture more of the price run-up in January. Hence, the price-pressure hypothesis predicts an increase in December in net insider purchases in small firms. Furthermore, December insider trading in small firms ought to be in contrast with December insider trading in large firms since no such increase is predicted for large firms. Alternatively, if insider trading is the source of the price pressure, then insiders would be expected to be net purchasers of stock in small firms in January. The risk-premium hypothesis predicts increased insider trading activity (either sales or purchases) in small firms in January. Moreover, the risk-premium hypothesis predicts a January seasonality in bid-ask spreads of small firms as well.

There are additional explanations (not examined here) for the January effect. One set of studies is concerned with econometric issues. These studies examine the stationarity of the January effect, stationarity of the risk measures, and deviations from normality. (See Christie and Hertzel [9], Brown, Kleidon, and Marsh [6], Chan and Chen [7], Hillion [13], and Reinganum [19].) Other studies examine the failure of asset-pricing models. (See Roll [22], Tinic and West [30], and Tinic and Rogalski [29].) Other types of risk exposure are also considered. (See Chan, Chen, and Hsieh [8].)

To observe the effects of the January effect in aggregate insider trading, it is not necessary that insiders actively purchase their own firms' stocks in December and sell their own firms' stocks in January. If insiders believe that the empirically documented price run-up in January in small firms represents a profit opportunity, then, in addition to any other transactions they might engage in, such as buying options or futures contracts on small firms in general, insiders would also be expected to adjust their planned transactions in their own firms. Seyhun [27] provides evidence of passive as well as active trading by insiders around firm-specific nonpublic information.[5] For instance, if an insider in a small firm already planned to purchase stock in his or her own firm, he or she would profit by

[5] Also see Seyhun [28] for information content of aggregate insider trading. For other studies of insider trading, see Finnerty [10], Jaffe [14], Pratt and DeVere [17], and Lorie and Niederhoffer [16].

accelerating and completing his or her purchase in December. Similarly, if an insider in a small firm planned to sell his or her stock, he or she would profit by delaying his or her sale until after the January effect were realized. As a consequence, insider sales in December would tend to be lower than the normal sales in other months, and insider purchases in December would tend to exceed normal purchases in other months. Hence, net insider purchases in December would be higher than otherwise.

The remainder of the paper discusses the data and presents the empirical findings and their interpretations. The paper is organized as follows. Section I contains the data sources and sample characteristics of the data. The empirical results of the study are in Section II, and the conclusions and implications are in Section III.

I. Data and Sample Characteristics

The insider trading data for this study come from a computer tape compiled by the Securities and Exchange Commission. The tape contains more than 1.5 million transactions by insiders in all publicly held firms from January 1975 to October 1981, for a total of eighty-two calendar months. This study analyzes a sample of insider transactions in 790 firms on the New York Stock Exchange (NYSE) and American Stock Exchange (AMEX) daily returns file of the Center for Research in Security Prices (CRSP).[6] Of the 790 firms, twenty-one did not report any insider trading between 1975 and 1981. Hence, the actual number of firms analyzed is 769.

Only open market sales and purchases by insiders during 1975 to 1981 are analyzed in this study. All other types of insiders' transactions, such as private sales and purchases, exercises of options, shares acquired through a plan, etc., are excluded since it is expected that insiders' open market sales and purchases are more likely to represent actions taken as a result of special insider information.[7] The distribution of insiders' open market sales and purchases by firm size are shown in Table I. Firm group 1 consists of firms with a market value of equity less than $25 million; group 2, between $25 million and $50 million; group 3, between $50 million and $250 million; group 4, between $250 million and $1 billion; and group 5, greater than $1 billion. This classification ensures that each size group contains more than 3000 insider transactions while maintaining a large diversity of firm sizes from less than $25 million to more than $1 billion.

The total sample contains 59,148 open market sales and purchases with a total dollar value of $11.1 billion. Of these, 24,371 are open market purchase transactions with a total dollar value of $6.8 billion, and 34,777 are open market sale transactions with a total dollar value of $4.3 billion. As expected, there are more insiders in larger firms and, therefore, a greater number of insider transactions.

[6] All 190 firms listed on option exchanges on January 1, 1977, are included in the sample. The remaining six hundred firms are selected using stratified random sampling based on the value of equity on January 1, 1977. This accounts for the heavier representation of the larger firms in the sample. The sample of firms analyzed in this study are identical to those in Seyhun [27, 28], which contain further characterizations of the sample.

[7] Numerous consistency checks on dates, prices, and shares were performed to eliminate approximately 1000 transactions containing apparent data errors out of about 60,000 transactions.

Table I

Distribution of the Number of Open Market Sales and Purchases, the Total Dollar Value of Sales and Purchases (in $ Millions), and the Number of Firms Grouped by the Average Month-End Market Value of Firms' Stock between January 1975 and October 1981

	The Average Market Value of Stock					
	Less than $25 Million (Group 1)	Between $25 and $50 Million (Group 2)	Between $50 and $250 Million (Group 3)	Between $250 Million and $1 Billion (Group 4)	More Than $1 Billion (Group 5)	All Firms
Number of Firms	104	68	173	267	157	769
Dollar Value of Sales	$51.3	$56.2	$550.0	$1431.7	$2235.7	$4324.7
Dollar Value of Purchases	$100.3	$126.2	$736.5	$1558.0	$4254.3	$6775.3
Number of Sales	1339	1325	5891	14,811	11,411	34,777
Number of Purchases	2802	1685	4661	8456	6767	24,371
Total Number of Trades	4141	3010	10,552	23,267	18,178	59,148

II. Empirical Results

A. *Seasonality of Stock Returns*

Tests presented in the empirical section examine the seasonal distribution of insiders' transactions to test the price-pressure and risk-premium hypotheses for the unusually high, positive stock returns to small firms in January. The price-pressure hypothesis states that the January effect arises from price pressure as a result of the seasonal changes in the demand for stocks of small firms. The risk-premium hypothesis states that the January effect represents compensation for the higher risk of trading against informed traders in January.

First, the seasonal pattern of stock returns documented by Keim [15] is replicated for the sample of 769 firms included in this study. It is necessary that firms included in this study exhibit a similar seasonality of stock returns if they are to have the potential of explaining the January effect. Second, the relation between the seasonality of insiders' transactions and the seasonality of security returns is examined.

Table II shows the seasonal pattern of the monthly returns from January 1975 to October 1981 for the 769 firms included in this study. In models (1) through (5), the dependent variable is the total monthly return to the equally weighted portfolio of firms in each of the five firm size groups. The independent variable is a dummy variable that takes a value of one for the month of January and zero for other months.

Model (1) shows that the returns to small firms (group 1) in January have been on average 12.9 percent higher than other months. This value is statistically significant and is twenty-five times higher than the average return for months

Table II

Regression of the Monthly Return to the Equally Weighted Portfolio of Firms Grouped by Size of Equity, RE_1 to RE_5, Against a Dummy Variable That Takes the Value of One in January and Zero Otherwise[a]

Model Number	Dependent Variable		Intercept		January Dummy	Adjusted R^2
(1) (Small Firms)	RE_1	=	0.005	+	0.129	0.177
			(0.6)		(4.3)[b]	
(2)	RE_2	=	0.013	+	0.081	0.099
			(1.8)		(3.2)[b]	
(3)	RE_3	=	0.014	+	0.066	0.071
			(2.0)[c]		(2.7)[b]	
(4)	RE_4	=	0.012	+	0.040	0.036
			(2.1)[c]		(2.0)[c]	
(5) (Large Firms)	RE_5	=	0.010	+	0.016	−0.001
			(2.0)[c] *		(1.0)	

[a] Sample period is from January 1975 to October 1981. The t-statistics for the estimated coefficients are shown in parentheses. Group 1 has average size of equity less than $25 million; group 2, between $25 and $50 million; group 3, between $50 and $250 million; group 4, between $250 million and $1 billion; and group 5, greater than $1 billion.

[b] Significant at the one percent level.

[c] Significant at the five percent level.

other than January. The returns for the month of January decline with increasing firm size. For large firms (group 5), the return in January is only 1.6 percent higher than other months, which is not statistically distinguishable from zero. This evidence indicates that the firms included in this study display a pattern of seasonality similar to that documented by Keim [15], Blume and Stambaugh [2], and others, and therefore, examination of the seasonal behavior of aggregate insider trading activity is warranted.

B. Insider Trading Variables

Two measures of insider trading activity are computed. First, net insider trading activity is measured by the aggregate net number of transactions by executives.[8] Define net number of transactions as follows:

$$NE_{i,t} = \textstyle\sum_{j=1}^{J_{it}} H_{t,j}, \quad t = 1, \cdots, 82, \text{ from January 1975 to October 1981,} \qquad (1)$$

where J_{it} denotes the total number of transactions in firm i and month t, and $H_{t,j}$ equals one if transaction j is a purchase or minus one if it is a sale. Aggregate net number of transactions in size group k and month t, $ANE_{k,t}$, is defined as follows:

[8] See Seyhun [27] for differences in information content of executives and large shareholders' transactions. Only the open market sales and purchase transactions of executives are included in the construction of *ANE*. Hence, transactions that might have an inherent seasonal component, such as acquisitions due to compensation plans, exercises of options, or conversion of securities, etc., are excluded.

$$ANE_{k,t} = \sum_{i=1}^{I_k} NE_{i,t}, \quad t = 1, \cdots, 82 \text{ and } k = 1, \cdots, 5, \tag{2}$$

where I_k denotes the number of firms in size group k. (See Table I for definition of size groups 1 through 5.) The aggregate net number of transactions by executives in group k and month t, $ANE_{k,t}$, is the sum of net number of transactions over the firms in size group k.

As a second measure of insider trading activity, the absolute value aggregate net number of transactions, $AANE_{k,t}$, is constructed. This variable measures insider trading activity without regard to the direction of transactions. Define

$$AANE_{k,t} = \sum_{i=1}^{I_k} | NE_{i,t} |, \quad t = 1, \cdots, 82 \text{ and } k = 1, \cdots, 5. \tag{3}$$

The statistical properties of insider trading activity as measured by the *ANE* and *AANE* variables are shown in Table III. In small firms, insiders have a greater number of purchases than sales, resulting in positive mean value of *ANE*. In large firms, insiders have a greater number of sales than purchases, resulting in a negative mean value of *ANE*. The differences in the standard deviations of

Table III

Statistical Properties of the Aggregate Net Number of Transactions by Executives, $ANE_1 \cdots ANE_5$, and Absolute Value of Aggregate Net Number of Transactions by Executives, $AANE_1 \cdots AANE_5$, from January 1975 to October 1981[a]

			Cross-Sectional Correlation				Serial Correlation		
Variable	Mean	Standard Deviation	ANE_2	ANE_3	ANE_4	ANE_5	r_1	r_2	r_3
ANE_1	6.84	14.7	0.29	0.40	0.14	0.01	0.28	0.11	−0.10
			(2.7)	(3.9)	(1.5)	(0.1)	(2.5)	(1.0)	(−1.0)
ANE_2	−0.96	10.9	—	0.51	0.44	0.30	0.37	0.11	−0.01
				(5.3)	(4.4)	(2.8)	(3.4)	(1.0)	(−0.1)
ANE_3	−25.90	38.6	—	—	0.66	0.53	0.34	0.13	0.18
					(8.6)	(5.6)	(3.1)	(1.2)	(1.6)
ANE_4	−105.13	75.6	—	—	—	0.82	0.38	0.07	0.07
						(12.7)	(3.5)	(0.6)	(0.6)
ANE_5	−83.88	50.8	—	—	—	—	0.30	0.04	0.10
							(2.7)	(0.4)	(0.9)

			Cross-Sectional Correlation				Serial Correlation		
Variable	Mean	Standard Deviation	$AANE_2$	$AANE_3$	$AANE_4$	$AANE_5$	r_1	r_2	r_3
$AANE_1$	11.18	3.8	0.17	0.24	0.28	0.30	0.15	0.05	0.05
			(1.6)	(2.2)	(2.6)	(2.8)	(1.4)	(0.5)	(0.5)
$AANE_2$	8.98	4.0	—	0.36	0.28	0.29	0.05	−0.05	0.04
				(3.5)	(2.6)	(2.7)	(0.5)	(−0.5)	(0.4)
$AANE_3$	42.77	13.1	—	—	0.75	0.67	0.37	0.24	0.21
					(10.3)	(8.1)	(3.4)	(2.2)	(1.9)
$AANE_4$	112.43	32.1	—	—	—	0.82	0.43	0.15	0.14
						(12.9)	(3.9)	(1.4)	(1.3)
$AANE_5$	86.46	23.1	—	—	—	—	0.25	0.00	0.10
							(2.3)	(0.0)	(0.9)

[a] The t-statistics are shown in parentheses. The serial correlation coefficients at lags 1 to 3 are denoted by r_1 to r_3.

ANE across firm size groups largely reflect the different number of firms in each group. The cross-sectional correlations of aggregate net number of executives' transactions are positive and generally greater in absolute value among firms of similar size. Hence, it appears that insiders in similar-size firms are more likely to buy or sell stock at the same time as compared with insiders in different-size firms. This is consistent with the interpretation that the executives in different firms react to the same economy-wide factors at the same time or that a similar seasonal pattern of trading occurs among insiders of different firms.

Table III also shows the time-series properties of the aggregate net number of transactions by executives. The insider trading variables generally have significantly positive first-order serial correlation coefficients, while the higher order serial correlation coefficients are insignificant. The significant first-order serial correlations are taken into account in the empirical tests presented in the next section.

The statistical properties of the absolute value of aggregate net number of transactions by executives (*AANE*) are also shown in Table III, and they are generally similar to the statistical properties of aggregate net number of transactions (*ANE*). The cross-sectional correlations of *AANE* are positive and generally increasing in magnitude with firm size. The serial correlations of *AANE* show a positive peak at first lag for medium and large size firms. For smaller size firms, the serial correlations are insignificant.

C. A Test of the Price-Pressure Hypothesis

To test the price-pressure hypothesis, seasonal distribution of aggregate insider trading activity is examined using regression analysis in Table IV, models (1) to (5). The dependent variable is the aggregate net number of transactions by executives (*ANE*) for firm sizes one through five. The independent variables are dummy variables that take the value of one for the months of December and January, respectively, and zero for other months.

The positive intercept coefficient in model (1) of Table IV indicates that executives in small firms generally tend to be net purchasers in their own firms. Furthermore, there is an additional significant increase in purchase activity for executives in small firms in the month of December. The estimated positive regression coefficient for the December dummy is significant at the one percent level.

For medium size firms in model (2), December trading is marginally negative, while, in model (3), December trading is not substantially different from other months. In model (2), if a first-order moving-average model is fitted for the residuals, then the negative December coefficient estimate is marginally significant at the ten percent level.

The negative intercept coefficients in models (4) and (5) of Table IV indicate that, on average, executives in large firms tend to be net sellers in their firms. Furthermore, for large firms, there is an additional increase in net insider sales in the month of December. The estimated regression coefficients for the December dummy variables are significantly negative at or about the one percent level.

Various tests are conducted to examine the sensitivity of the results in Table

Table IV

Regression of the Monthly Aggregate Net Number of Transactions by Executives in Five Firm Size Groups, $ANE_1 \cdots ANE_5$, against Dummy Variables That Take the Value One in December and January, Respectively, and Zero Otherwise[a]

Model Number	Dependent Variable	Intercept	December Dummy	January Dummy	Error Model	Serial Correlation of Residuals r_1	r_2	r_3
(1)	ANE_1	5.09	15.11	5.38	AR(1)	0.01	0.03	−0.13
		(2.21)[b]	(2.63)[c]	(1.52)				
(2)	ANE_2	−0.29	−8.05	1.01	AR(1)	0.02	−0.01	−0.08
		(−0.15)	(−1.95)[b]	(0.26)				
(3)	ANE_3	−25.79	2.04	2.06	AR(1)	0.00	−0.04	0.11
		(−3.98)[c]	(0.13)	(0.14)				
(4)	ANE_4	−97.16	−82.45	−8.33	AR(1)	0.01	−0.02	0.08
		(−7.51)[c]	(−2.97)[c]	(−0.32)				
(5)	ANE_5	−79.33	−49.60	−7.50	AR(1)	0.01	0.00	0.18
		(−9.87)[c]	(−2.50)[b]	(−0.40)				

[a] The sample period is from January 1975 to October 1981. The t-statistics for estimated coefficients are shown in parentheses. For error models, AR(1) denotes a first-order autoregressive process. Regressions are estimated using the ARIMA procedure in SAS. Under the null hypothesis of no serial correlation, the estimated serial correlation coefficients of regression residuals at lag k, r_k, have a standard error of about 0.11.

[b] Significant at the five percent level.

[c] Significant at the one percent level.

IV to empirical methodology. Table IV reports the results of including a first-order autoregressive model for the residuals using Box and Jenkins' [3] methods. Examination of the Box-Pierce Q-statistics at lags 6, 12, 18, and 24 indicates that, with first-order autoregressive error models, there is no remaining serial correlation of the residuals. Omitting the first-order autoregressive models for the residuals reduces the magnitude and statistical significance of the estimated coefficients but does not affect the interpretation of the estimated coefficients. Hence the findings are not produced by the inclusion of the error-term models. The heteroscedasticity tests of White [31] yield insignificant chi-square statistics for all five models. Hence, the t-statistics of the estimated coefficients do not appear to be biased due to heteroscedasticity. Finally, the Kolmogorov D-statistics indicate that residuals in all five models are approximately normally distributed. Hence, the usual interpretation of the significance levels of the estimated coefficients appears to be justified.[9]

Although not reported here, to test the sensitivity of the results in Table IV to the definition of the insider trading activity, standardized aggregate net number of transactions has also been used as the dependent variable. Standardization refers to netting out the sample mean and dividing by the sample standard

[9] A seemingly unrelated regression (Zellner [33]) for the five size groups would provide identical estimates as the ordinary least squares despite significant cross-equation correlations since the regressors (dummy variables) are the same across equations.

deviation of the net number of transactions for each firm. Hence, standardization ensures that each firm gets approximately the same weight in the aggregation. Using as the dependent variable the standardized aggregate net number of transactions results in less significant December coefficient estimates for the small firms. For larger firms, net insider trading in December is still significantly negative. Moreover, the difference between insider trading in large and small firms is also significantly negative in December.

The tests in Table IV have also been repeated using as dependent variables both the standardized and unstandardized aggregate dollar value of insider trading and aggregate proportion of the firms traded by insiders. These tests (not shown here) result in December coefficient estimates similar in signs to those in Table IV yet usually statistically insignificant at the conventional levels. Nevertheless, the difference between insider trading in large and small firms remains significantly negative in December. Moreover, the residuals in these regressions exhibit significant non-normality that makes the interpretation of the results troublesome. Overall, these findings suggest that passive trading by insiders tends to be more strongly associated with the number of transactions rather than the dollar volume of trading. Consequently, the findings in Table IV should be interpreted in a limited sense as capturing an average trading trend by insiders.

To interpret the findings so far, the evidence shown in Table IV suggests that the aggregate net insider trading in small firms is substantially different from the aggregate net insider trading in larger firms in the month of December. The positive December coefficient in small firms indicates that some insiders accelerate their planned stock purchases and delay their planned stock sales in December to take advantage of the positive return in January. In contrast, the negative December coefficients in large firms indicate a tendency to sell stock at year end. The evidence presented in Table IV is consistent with the joint hypothesis that some insiders in small firms are aware of the seasonal pattern of returns in their own firms and regard the positive return in January as a profit opportunity.

The evidence in Table IV also does not indicate any increased selling activity in January for small firms. Hence, insiders do not seem to be actively buying their own firms' stock in December and then selling the stock in January.[10] This finding is consistent with the interpretation that the seasonal pattern of aggregate insider trading is more likely to be due to passive adjustments by insiders rather than to actively buying and selling their own firms' stock to exploit the January effect.

To test whether the insider trading contributes to the January effect, the January coefficients are examined. Model (1) in Table IV indicates that insiders in small firms increase their net purchases of stock slightly in January. However, the regression coefficient of the January dummy variable is not statistically significant at the ten percent level. Furthermore, the results (not shown) that use, as the dependent variable, the standardized aggregate net number of transactions also do not indicate a significant increase in purchase activity by insiders

[10] Section 16(b) of the Securities and Exchange Act of 1934 requires that any profit obtained from a sale and a purchase transaction within six months of each other be returned to the corporation.

in the month of January. Hence, the available evidence suggests that the January price pressure in small firms does not come from insiders' transactions.

D. A Test of the Risk-Premium Hypothesis

To test the risk-premium hypothesis, Table V examines the insider trading activity in December and January without regard to the direction of transactions. As a measure of insider trading activity, the absolute value of the net number of insider transactions computed from equation (3) is used. This measure, denoted *AANE*, measures aggregate insider trading activity without netting out purchases in one firm against sales in another firm. If the January effect were to represent compensation for the expected losses to informed traders, then the month of January would be expected to be characterized by increased insider trading in either direction, which should be captured in a high value of the *AANE* variable.

In Table V, models (1) to (5) show that the month of December brings about increased insider trading in all firms, not just the small firms. In each model, the coefficient estimate of the December dummy variable is positive and statistically significant. A plausible interpretation of this evidence is that tax-related trading in December leads to substantially higher trading in all firms, not just small firms. In contrast with December, January is not characterized by a consistent pattern of increased insider trading activity. In models (1) to (5), the coefficient

Table V

Regression of Monthly Aggregate Absolute Value of Net Number of Transactions by Executives, $AANE_1 \cdots AANE_5$, Against Dummy Variables Equal to One in December and January, Respectively, and Zero Otherwise[a]

Model Number	Dependent Variable	Intercept	December Dummy	January Dummy	Error Model	Serial Correlation of Residuals r_1	r_2	r_3
(1)	$AANE_1$	10.62	6.21	1.23	—	0.00	0.08	0.02
		(25.00)[b]	(4.13)[b]	(0.88)				
(2)	$AANE_2$	8.62	4.88	0.09	—	0.00	0.01	−0.04
		(18.42)[b]	(2.95)[b]	(0.06)				
(3)	$AANE_3$	41.40	9.54	2.44	AR(1)	−0.01	0.14	0.04
		(18.90)[b]	(1.93)[c]	(0.53)				
(4)	$AANE_4$	107.14	38.20	8.97	AR(1)	0.08	−0.02	0.06
		(19.19)[b]	(3.49)[b]	(0.87)				
(5)	$AANE_5$	84.32	27.87	−4.89	AR(1)	0.04	0.01	0.12
		(24.83)[b]	(3.14)[b]	(−0.59)				

[a] The sample period is from January 1975 to October 1981. The t-statistics for estimated coefficients are shown in parentheses. For error models, AR(1) denotes a first-order autoregressive process. Regressions are estimated using the ARIMA procedure in SAS. Under the null hypothesis of no serial correlation, the estimated serial correlation coefficients of regression residuals at lag k, r_k, have a standard error of about 0.11.

[b] Significant at the one percent level.

[c] Significant at the ten percent level.

estimate of the January dummy variable is insignificant. Hence, on average, the level of insider trading activity in January is not distinguishable from the level of insider trading activity during any other month excluding December. This evidence suggests that, for an uninformed trader, the expected losses to corporate insiders ought not be higher in January than in other months.

The results in Table V also have been replicated by using, as the dependent variable, the standardized, aggregate absolute value of the net number of transactions. The results using the standardized dependent variable (not reported here) are similar to the results shown in Table V. This finding suggests that the results in Table V are not due to a few outlier firms. Furthermore, heteroscedasticity tests of White [31] produce insignificant chi-square statistics for all five models. Hence, heteroscedasticity does not appear to be a serious problem.

While not shown, further examination of the insider trading activity by calendar weeks shows that most of the increased insider trading activity in the month of December occurs during the third week of December. Furthermore, the first week of January is not characterized by an increased level of insider trading. Hence, the timing of the increased level of insider trading does not correspond with the timing of higher returns to small firms. In addition, insider trading in October through March is examined. Inclusion of these additional months also does not change any of the findings in Tables IV and V. For brevity, these results are not shown.

Assuming that the overall trading by all informed investors and the trading by corporate insiders are positively correlated, the evidence presented in Table V does not corroborate the hypothesis that the price run-up in January represents compensation for greater expected losses to informed traders. Instead, the evidence suggests that corporate insiders in all firms are more likely to trade in the month of December, presumably for tax reasons. Corroborating evidence is contained in Blume and Stambaugh [2], who report that the bid-ask spread bias in returns of small firms is not significantly higher in January than in other months. This finding still leaves open the possibility that a portion of the January effect can be a risk premium for other types of risk that have a January seasonal component.[11]

III. Conclusions and Implications

The evidence presented in this paper suggests that corporate insiders in small firms adjust their transactions at year end. Evidence indicates that some insiders in small firms tend to accelerate their planned stock purchases and postpone their stock sales in December. The increased number of net purchases in December in small firms contrasts with an opposite pattern of insider trading in larger firms. Becoming net purchasers of stock in December enables insiders in small firms to capture more of the positive return in January. Hence, this evidence is

[11] Chan, Chen, and Hsieh [8] report that a portion of the returns to small firms in January represents compensation for risk. One component of risk is measured by the difference between the return to low-grade bonds and the long-term government bonds, which themselves have a January seasonal component.

consistent with the interpretation that some insiders in small firms regard the price run-up in January as a profit opportunity. However, examination of the aggregate insider trading activity in January indicates that any price pressure on the prices of small firms in January does not come from additional increased insider purchases of stock since corporate insiders do not significantly increase their stock purchases in January.

The evidence presented in the paper also does not corroborate the hypothesis that the January price run-up in small firms represents compensation for increased expected losses to informed traders such as corporate insiders. The evidence shows that insider trading activity in small firms does not significantly increase in January. Hence, an uninformed trader in small firms is not more likely to trade against an informed insider in January than in other months. Consequently, the price run-up in small firms in January cannot be interpreted as compensation for greater expected losses against informed traders in January.

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