Journal of Financial Economics 12 (1983) 13–32. North-Holland

SIZE-RELATED ANOMALIES AND STOCK RETURN SEASONALITY

Further Empirical Evidence

Donald B. KEIM*

University of Pennsylvania, Philadelphia, PA 19104, USA

Received June 1981, final version received June 1982

This study examines, month-by-month, the empirical relation between abnormal returns and market value of NYSE and AMEX common stocks. Evidence is provided that daily abnormal return distributions in January have large means relative to the remaining eleven months, and that the relation between abnormal returns and size is *always* negative and more pronounced in January than in any other month — even in years when, on average, large firms earn larger risk-adjusted returns than small firms. In particular, nearly fifty percent of the average magnitude of the 'size effect' over the period 1963–1979 is due to January abnormal returns. Further, more than fifty percent of the January premium is attributable to large abnormal returns during the first week of trading in the year, particularly on the first trading day.

1. Introduction

Recent empirical research in financial economics has revealed abnormal returns inconsistent with equilibrium in a market where the CAPM holds. Banz (1981) and Reinganum (1981) report a significant negative relation between abnormal returns and market value of common equity for samples of NYSE and NYSE–AMEX firms, respectively.[1] Whereas Banz and Reinganum implicitly assume that the negative relation between abnormal returns and size is stable over the periods examined, Brown, Kleidon and Marsh (1983) report a reversal of the size anomaly for certain years and reject the hypothesis of stationary year-to-year abnormal returns attributable to size.

*I would like to thank Susan Chaplinsky, Eugene Fama, Robert Holthausen, Michael Jensen, Allan Kleidon, Richard Leftwich, Merton Miller, Myron Scholes, and the referee Marc Reinganum for helpful comments. Financial assistance was provided by the Center for Research in Security Prices at the University of Chicago.

[1]Although it is not clear that the anomalous returns derive explicitly from failure of the CAPM to account for firm size, several studies have shown that anomalous return behavior associated with firm-specific variables is largely subsumed under the 'size effect'. For example, Reinganum (1981) finds that the relation between abnormal returns and *P/E* ratios reported by Basu (1977) appears to vanish after controlling for size. Keim (1980) and Stattman (1980) find a significant negative relation between abnormal returns and the degree to which market value of equity exceeds book value of equity, and also interpret this relation as a proxy for the size effect.

This study examines the month-to-month stability of the size anomaly over the period from 1963–1979. The evidence indicates that nearly fifty percent of the average magnitude of the risk-adjusted premium of small firms relative to large firms over this period is due to anomalous January abnormal returns. Further, more than twenty-six percent of the size premium is attributable to large abnormal returns during the first week of trading in the year and almost eleven percent is attributable to the first trading day. The data do not reveal significant seasonal behavior in any other month.

Hypotheses advanced by others to explain the size effect appear unable to explain the January effect. For example, Brown, Kleidon and Marsh argue that at least part of the size effect may be explained by an omitted risk factor in the pricing model. Even if part of the average size effect is due to an unspecified risk variable, however, the behavior observed in January cannot be due solely to this cause because risk alone cannot explain a return premium observed in the same month each year.[2] Stoll and Whaley (1983) contend that transaction costs can explain the size effect because such costs prevent arbitrageurs from eliminating the average return differential. However, only if transaction costs are seasonal in nature, implying some degree of market power for market makers, can such costs explain the January effect.

The results in this paper shed further light on the magnitude and nature of the size anomaly and also carry implications for empirical work using daily data. In particular, market efficiency studies relying on a model that does not account for non-stationary returns across months may be biased when investigating events concentrated in January if the event is unrelated to the true cause of the anomalous January abnormal returns.

1.1. Outline of the paper

The previously observed negative relation between abnormal returns and size is reproduced across a sample of NYSE and AMEX firms in section 2, and the relation is judged to be insensitive to misassessment of risk caused by infrequent trading of securities. In section 3 the effects of month-to-month stock return seasonality on the size effect are investigated, and evidence is presented that implies almost fifty percent of the average size anomaly is due to large January abnormal returns. Several possible explanations of the January effect are considered in section 4, and a brief summary is presented in section 5.

[2]Thus, it appears that we can separate the 'size effect' into two distinct components: a large premium *every* January and a much smaller and, on average, positive differential between risk-adjusted returns of small and large firms in every other month. A complete explanation of the 'size effect' requires two separate explanations for these very different phenomena.

2. Evidence on anomalous excess returns

In this section I investigate the anomalous negative relation between firm size, measured by total market value of common equity, and abnormal risk-adjusted returns for the sample of NYSE and AMEX firms used in this study. Careful attention is paid to Roll's (1981) conjecture that the apparent return premia of smaller firms may be at least partially attributable to an observed downward bias in the OLS betas for these portfolios. To avoid this bias, I employ beta estimates that adjust for non-synchronous trading and trading infrequency in the computation of abnormal returns. Although the adjustments result in a near monotone declining relation between beta and size, the portfolio abnormal returns computed with adjusted betas still exhibit a pronounced negative relation to firm size.

2.1. Data and portfolio selection

The data for this study are drawn from the CRSP daily stock files for the seventeen-year period from 1963 to 1979. The sample consists of firms which were listed on the NYSE or AMEX and had returns on the CRSP files during the entire calendar year under consideration. Thus, every year firms enter or leave the sample due to mergers, bankruptcies, delistings and new listings. The number of sample firms in a given year ranges from approximately 1,500 in the mid-1960's to 2,400 in the late 1970's.

Each year I rank all sample firms on the market value of their common equity. The market values, derived from the CRSP daily master file, are computed by multiplying the number of shares of common stock outstanding at year-end by the year-end price of the firm's common shares. I then divide the yearly distributions of market values equally into ten portfolios on the basis of size, portfolio one containing the smallest firms and portfolio ten containing the largest firms. Thus, each portfolio is updated annually and, on average, contains approximately two hundred firms.

2.2. Sensitivity of the size anomaly to trading infrequency

Roll (1981) conjectures that the size effect may be a statistical artifact of improperly measured risk. Scholes and Williams (1977) point out that non-synchronous trading of securities imparts a downward bias to the estimated beta when the underlying security trades infrequently. Dimson (1979) also argues that trading infrequency biases beta estimates and predicts a downward bias for infrequently traded shares and an upward bias for frequently traded shares. Roll maintains that since the shares of small firms are generally the most infrequently traded and the shares of large firms are the most frequently traded, the betas for small firms are downward biased while the betas of large firms are upward biased. Thus, estimation of abnormal returns using risk estimates that are not adjusted for trading infrequency may yield the observed size effect. In a recent paper, however, Reinganum (1982) reports that, while the direction of the bias in beta estimation is consistent with Roll's conjecture, portfolio excess returns computed with these adjusted betas will still exhibit a pronounced negative relation to firm size.

I have independently investigated the effects of improperly estimated betas on the size effect and the results corroborate Reinganum's (1982) findings. Estimates of OLS betas, Scholes–Williams betas and Dimson betas are presented, along with other statistics discussed below, in table 1 and three interesting results emerge. First, there is no distinguishable relation between the OLS estimates of beta and firm size measured by market value of equity. Of particular interest are the low levels of beta for the two smallest firm portfolios. Second, although the Scholes–Williams beta estimates for smaller firms are generally higher than the corresponding OLS estimates, there still is no distinct ordering of the betas according to firm size. Third, the Dimson beta estimate for the portfolio of smallest firms is significantly larger than the largest firm portfolio beta, and there is a near monotone declining relation between firm size and Dimson beta. However, not even the Dimson estimator results in the upward revision of small firm betas necessary to eliminate the excess returns for the small firm portfolios. Given the levels of average annual return for portfolio one (35.0%) and for the value-weighted market (7.0%) over the 1963–1979 period, extremely large average betas would have been necessary to ensure zero average excess returns.[3] Thus, the magnitude of the size anomaly does not appear to be sensitive to different estimators of beta.

2.3. The size anomaly: Some empirical results based on Scholes–Williams beta estimates

To test the relation between anomalous returns and size, I use security abnormal returns obtained from the CRSP daily excess return file. The CRSP daily excess return file contains control portfolios constructed by annually ranking the stocks in its file into ten portfolios of descending order of risk as measured by estimated Scholes–Williams betas. Security excess returns are computed as the security daily return less the equal-weighted daily return of the control portfolio into which the security is ranked.[4] I

[3]The Dimson estimator used in this study is defined in footnote e to table 1. The choice of ten lagged terms conforms with the findings of Schultz (1982) that even the smallest AMEX firms rarely go untraded for more than ten days. Nevertheless, Dimson estimates were also computed with twenty-one lagged, five leading, and contemporaneous value-weighted market returns and yielded small and large firm portfolio betas of 1.69 and 0.97, respectively. This larger beta for the small firm portfolio reflects positive but insignificant coefficients on the eleventh through twenty-first lagged market returns.

[4]See Center for Research in Security Prices (1980) for more details.

compute the average daily excess returns for the size deciles by weighting the CRSP excess returns for the securities in each decile equally.

Average daily portfolio excess returns are presented in table 1. The results are based on seventeen years of daily excess returns for each portfolio, beginning on the first trading day in January 1963.[5] That is, the daily excess returns for the twelve-month period following each portfolio update are stacked into one time series vector. Table 1 also contains the average median market value of equity for each portfolio and the first-order autocorrelation coefficient of the portfolio excess returns. The average excess returns are plotted in fig. 1 as a function of the decile of equity value. The plot suggests that, even after the Scholes–Williams adjustment of beta for non-synchronous trading, excess returns are a monotone decreasing function of firm size as measured by total market value of equity. The average return of the portfolio of smallest firms is about 20.7 percent per year (0.082 percent per day × 252 trading days per year) greater than the return implied by its beta risk. On the other hand, the portfolio of largest firms earned a return 9.6 percent per year

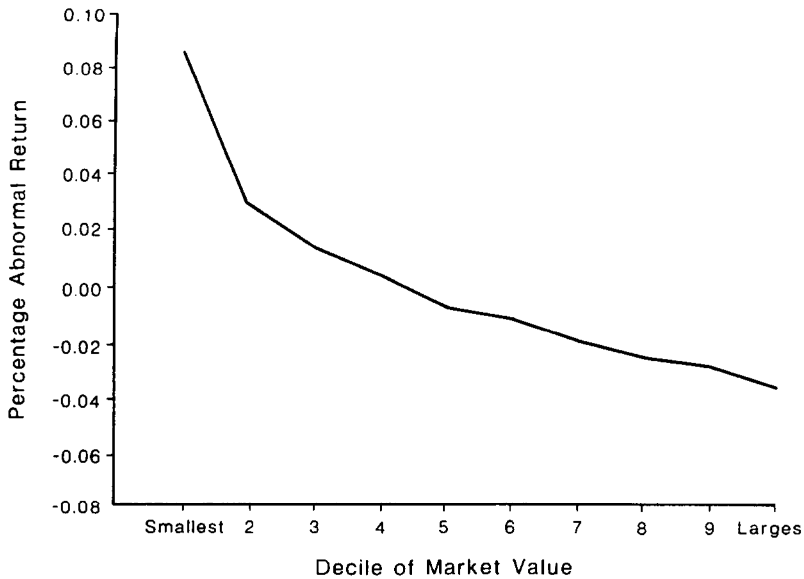

Fig. 1. Average daily abnormal returns (in percent) for ten market value portfolios constructed from firms on the NYSE and AMEX over the period 1963–1979. Abnormal returns are provided by CRSP.

[5]Think of the excess returns as being generated by the following strategy: Compute firm equity values on the last trading day of 1962. Construct portfolios on the first trading day in 1963 by ranking on the estimates of firm size and track the excess returns over the next twelve months. Update the firm equity values at the end of 1963, restructure the portfolios at the beginning of 1964 based on the updated information and again track the excess returns for one year. Repeat the process through December 1979.

Table 1
Average daily excess returns (in percent), size measured by market value of equity, beta estimates and autocorrelations of excess returns for ten portfolios constructed from firms on the NYSE and AMEX over the period 1963–1979.

Portfolio	Average excess return[a]	Market value of equity[b]	OLS beta[c]	Scholes–Williams beta[d]	Dimson beta[e]	1st order autocorrelation of excess return[a]
Smallest	0.082 (10.38)	$ 4.4	0.76	0.92	1.47	0.222
2	0.030 (5.83)	10.5	0.87	1.01	1.47	0.131
3	0.015 (3.88)	18.9	0.91	1.03	1.43	0.065
4	0.003 (0.97)	30.3	0.93	1.08	1.42	0.028
5	−0.007 (−2.24)	46.7	0.99	1.08	1.42	0.029
6	−0.014 (−4.82)	73.4	0.98	1.08	1.30	0.097
7	−0.020 (−6.40)	118.1	0.95	1.03	1.27	0.180
8	−0.024 (−6.74)	210.2	0.97	1.04	1.22	0.278
9	−0.029 (−7.20)	433.0	0.96	1.02	1.12	0.345
Largest	−0.038 (−7.19)	1092.1	0.96	0.97	0.98	0.351

[a]Excess returns are provided by CRSP; the excess return statistics and autorrelations are based on 4262 daily observations for each portfolio; t-statistics are in parentheses.

[b]Market value of equity is measured by the average, across all sample years, of the median market value of the particular size decile in each year; Market values are in millions of dollars.

[c]The OLS beta is obtained by regressing daily portfolio returns against the daily returns of the CRSP value-weighted index for the entire 1963–1979 period.

[d]The Scholes–Williams beta estimates are defined as

$$b_i = \sum_{k=-1}^{+1} B_{ik}/(1+2r), \quad i=1,10,$$

where r is the autocorrelation of the CRSP value-weighted daily market return and the B_{ik} are the slope coefficients from three separate OLS regressions,

$$R_{it} = a_{ik} + B_{ik} R_{m,t+k} + v_t, \quad k=-1,0,1, \quad i=1,10.$$

[e]The Dimson beta estimates are obtained by summing the slope coefficients on the ten lagged, five leading and the contemporaneous CRSP value-weighted daily market returns in the following OLS regression:

$$R_{it} = a_i + \sum_{k=-10}^{+5} b_{ik} R_{m,t+k} + u_{it}, \quad i=1,10.$$

(-0.032 percent $\times$ 252) less than that implied by its beta risk. The annual difference of 30.3 percent between the risk-adjusted returns of these two portfolios is comparable to the twenty-five percent annual difference reported by Reinganum (1981).[6]

Also of interest in table 1 is the pattern of autocorrelations across the various size portfolios. In particular, the average excess returns of the smallest *and* largest firms (e.g., portfolio one and portfolio ten) are highly autocorrelated while the intermediate portfolios exhibit little autorrelation. This appears counterintuitive given the evidence of Scholes and Williams (1977) and Dimson (1979) that only returns of infrequently traded shares (i.e., smaller firms) should display positive autocorrelation. Roll (1981) argues, however, that excessive autocorrelation in the extreme portfolios is consistent with the definition of excess returns as $r_t = R_{pt} - R_{ct}$, where p indicates the portfolio under study and c indicates an equal-weighted control portfolio. As Roll (1981, p. 882) points out:

> The serial covariance in excess returns, $\operatorname{cov}(r_t, r_{t-1})$, e.g., is composed of $\operatorname{cov}(R_{pt}, R_{p,t-1}) + \operatorname{cov}(R_{ct}, R_{c,t-1}) - \operatorname{cov}(R_{ct}, R_{p,t-1}) - \operatorname{cov}(R_{pt}, R_{c,t-1})$. When trading frequency is different for p and c, one of the first two dominates [implying positive autocorrelation]. Otherwise, the four mutually cancel.

Since the large firm portfolio excess returns are relative to an equal-weighted control portfolio (which is more heavily represented by smaller component firms) autocorrelation is induced into the large firm portfolio excess returns by the equal-weighted control portfolio.[7]

3. Size-related anomalies and stock return seasonality

In this section I investigate the month-to-month stability of the size anomaly. The evidence indicates that the magnitude of the anomaly depends on the month of the year and that nearly fifty percent of the anomaly is concentrated in the month of January. Further, more than twenty-seven percent of the size effect in an average year can be attributed to the first week of trading in January. I turn first, though, to the seasonality in stock returns.

[6]Reinganum (1981) finds the betas of the market value portfolios used in his study close to one and computes excess returns by subtracting the daily return of the equal-weighted NYSE–AMEX index from the daily portfolio return. I replicated Reinganum's risk-adjustment procedure using both equal- and value-weighted NYSE–AMEX control portfolios for the ten market value portfolios in this study with similar results.

[7]Small firm portfolio returns generally exhibit higher autocorrelation than large firm portfolio returns. As evidence, for the period 1963–1979, the daily equal-weighted NYSE–AMEX index has first-order autocorrelation of 0.408, while the daily value-weighted NYSE–AMEX index has first-order autocorrelation of 0.205.

3.1. Stock return seasonality

Much evidence has accumulated [see Fama (1965, 1970)] indicating that common stock prices follow a multiplicative random walk. Thus, the return on an equal-weighted portfolio of stocks conforms to the following process:

$$\tilde{R}_t = \mu + \tilde{e}_t, \tag{1}$$

where $\tilde{R}_t$ is the random portfolio return, $\mu = E(\tilde{R}_t | I_{t-1})$, I_{t-1} is the information set available at $t-1$, and $\tilde{e}_t$ is an i.i.d. random variable with zero mean.

While the random walk model implies portfolio return distributions are time-invariant, recent empirical evidence [French (1980), Gibbons and Hess (1981), Officer (1975), and Rozeff and Kinney (1976)] indicates that portfolio return distributions do indeed differ temporally. For example, Rozeff and Kinney test the seasonal model

$$\tilde{R}_{tm} = \mu + \lambda_m + \tilde{e}_t, \tag{2}$$

where m denotes the month of the year and $\tilde{e}_t$ is i.i.d. with zero mean. They examine monthly returns on an equal-weighted NYSE index over the period 1904–1974 and report large average monthly returns in January relative to the remaining eleven months. Rozeff and Kinney conclude that expected portfolio returns depend on the month of the year.

Keim (1982) presents evidence that the January seasonal in stock returns is more pronounced for portfolios of small firms than for portfolios of large firms. Employing the ten market value portfolios of NYSE–AMEX stocks used in this study, Keim tests the hypothesis of stable month-to-month average returns for each portfolio and finds that ability to reject the hypothesis declines as average firm size increases. In fact, the results indicate that one cannot reject the hypothesis for the portfolio of largest firms.

3.2. Seasonality and the size anomaly: A January effect

The finding that the magnitude of the January return seasonal is related to firm size provides some basis for suspecting month-to-month instability in the size effect. To address this suspicion, I plot in fig. 2 the negative relation between abnormal return and firm size separately for each month of the year during the period 1963–1979. The plots in fig. 2 dramatically display a difference in the size effect between January and the other eleven months. The January relation between abnormal returns and market value is steep and negative while the plots for the other months have, relative to January, only a slight negative slope and are tightly clustered around zero abnormal

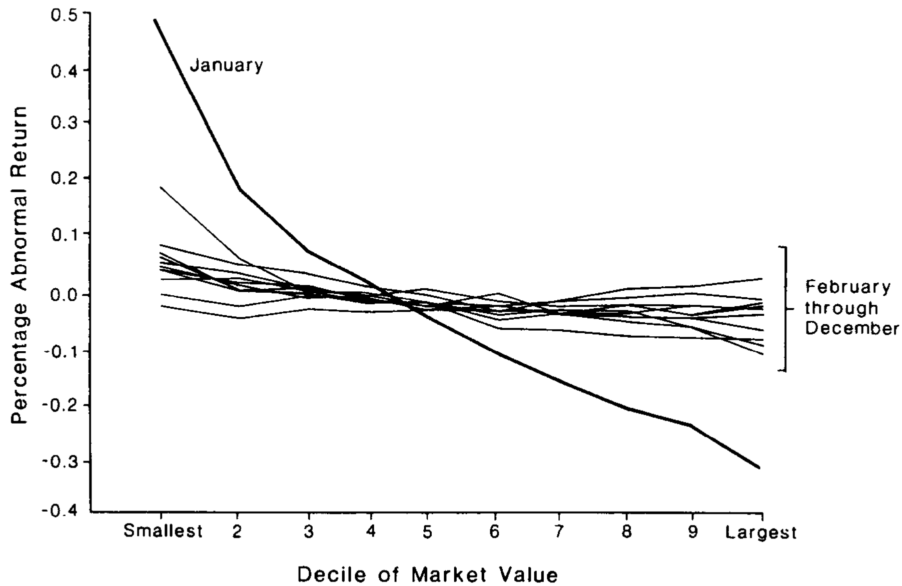

Fig. 2. The relation between average daily abnormal returns (in percent) and decile of market value *for each month* over the period 1963–1979. The ten market value portfolios (deciles) are constructed from firms on the NYSE and AMEX. Abnormal returns are provided by CRSP.

return. The figure shows clearly that the size effect is more pronounced in January than in the other months, and also that the anomaly has similar characteristics from February through December. (The non-January observation that stands apart from the cluster in the decile of smallest market value is the February observation.)

While fig. 2 suggests that the size effect is more pronounced in January than in any other month, further support for this conclusion can be obtained from examination of the month-to-month magnitude of the size effect measured by the difference in risk-adjusted returns between the smallest market value portfolio and the largest market value portfolio. Table 2 contains the differences in average daily CRSP excess returns between the smallest and largest market value portfolios for every month of every year during the period 1963–1979. Differences averaged across all months for each year (rightmost column) and averaged across all years for each month (bottom row) are also provided. The most striking feature of the data is the persistence, magnitude and statistical significance of the excess returns in January: a monthly size effect of 15.0% (average daily percentage return of 0.714 multiplied by twenty-one trading days per month) is implied in January in contrast to the implied monthly excess of return of 2.5% (0.121 × 21) averaged over all months and all years. No consistent pattern is apparent

Table 2

Average differences (t-statistics) between daily (CRSP) excess returns (in percent) of portfolios constructed from firms in the top and bottom decile of size (measured by market value of equity) on the NYSE and AMEX over the period 1963–1979.

	Jan.	Feb.	March	April	May	June	July	Aug.	Sept.	Oct.	Nov.	Dec.	Mean daily return over all months
1963	0.309	0.093	−0.085	−0.045	0.172	0.056	−0.026	0.000	0.040	−0.021	0.011	−0.123	0.032
	(2.26)	(1.23)	(−0.81)	(−0.55)	(2.56)	(0.73)	(−0.39)	(0.00)	(0.51)	(−0.26)	(0.08)	(−1.38)	(1.16)
1964	0.170	0.105	0.097	0.007	−0.084	(−0.037	0.121	0.104	0.077	0.136	−0.019	−0.111	0.048
	(1.52)	(1.30)	(1.37)	(0.12)	(−1.34)	(−0.56)	(1.66)	(1.42)	(1.08)	(2.03)	(−0.23)	(−1.64)	(2.14)
1965	0.288	0.228	0.204	0.133	0.025	−0.212	0.070	0.104	−0.023	0.314	0.343	0.202	0.137
	(2.34)	(2.68)	(2.63)	(2.25)	(0.38)	(−2.41)	(1.05)	(1.34)	(−0.44)	(4.16)	(3.74)	(1.79)	(5.29)
1966	0.388	0.448	0.183	0.192	−0.278	0.017	−0.009	−0.177	−0.025	−0.423	0.138	0.001	0.033
	(4.63)	(4.44)	(1.52)	(2.11)	(−1.82)	(0.24)	(−0.08)	(−1.78)	(−0.23)	(−2.31)	(1.47)	(0.01)	(0.92)
1967	0.765	0.413	0.142	0.149	0.240	0.599	0.403	0.235	0.512	0.268	−0.120	0.427	0.336
	(4.59)	(5.04)	(2.35)	(1.54)	(2.56)	(3.87)	(4.15)	(3.09)	(5.59)	(2.70)	(−0.64)	(4.40)	(9.34)
1968	0.834	−0.197	−0.079	0.427	0.727	0.096	0.222	0.348	0.345	−0.002	0.091	0.434	0.285
	(6.20)	(−1.25)	(−0.52)	(2.70)	(6.52)	(1.11)	(1.30)	(3.37)	(4.15)	(−0.03)	(0.88)	(4.12)	(6.78)
1969	0.128	−0.253	−0.059	−0.139	0.082	−0.265	−0.241	−0.073	−0.006	0.247	−0.148	−0.349	−0.085
	(1.00)	(−2.49)	(−0.62)	(−1.94)	(1.19)	(−2.47)	(−1.88)	(−0.61)	(−0.07)	(2.48)	(−2.16)	(−4.06)	(−2.76)
1970	0.612	−0.257	0.033	−0.213	−0.016	−0.082	−0.164	−0.074	0.315	−0.019	−0.434	0.136	−0.011
	(2.49)	(−1.58)	(0.31)	(−1.39)	(−0.07)	(−0.47)	(−0.80)	(−0.48)	(2.91)	(−0.16)	(−4.44)	(1.08)	(−0.22)

1971	0.886 (7.44)	0.432 (2.59)	0.159 (2.05)	−0.156 (−2.09)	−0.066 (−1.05)	−0.246 (−3.57)	−0.053 (−0.75)	0.004 (0.03)	0.095 (1.19)	−0.052 (−0.63)	−0.175 (−1.78)	0.130 (0.98)	0.074 (2.12)
1972	0.760 (5.32)	0.197 (2.44)	0.005 (0.06)	−0.054 (−0.68)	−0.181 (−2.87)	−0.061 (−0.69)	−0.022 (−0.26)	−0.204 (−2.63)	−0.098 (−1.03)	−0.200 (−2.19)	−0.065 (−0.68)	−0.154 (−1.19)	−0.008 (−0.26)
1973	0.368 (1.78)	−0.137 (−1.47)	−0.060 (−0.48)	−0.109 (−0.79)	−0.406 (−1.86)	−0.011 (−0.07)	0.367 (2.61)	−0.144 (−1.60)	−0.080 (−0.60)	0.169 (1.61)	−0.298 (−1.76)	−0.494 (−2.07)	−0.067 (−1.39)
1974	1.464 (3.67)	0.047 (0.25)	0.313 (2.12)	0.127 (0.86)	0.017 (0.12)	0.063 (0.31)	0.331 (1.57)	0.313 (1.09)	0.224 (0.65)	−0.380 (−1.20)	0.015 (0.06)	−0.187 (−0.48)	0.199 (2.39)
1975	2.068 (3.95)	0.278 (1.15)	0.929 (3.76)	−0.126 (−0.66)	0.293 (1.83)	0.426 (1.53)	0.489 (2.48)	−0.274 (−1.69)	−0.083 (−0.54)	−0.210 (−1.29)	0.123 (0.74)	0.220 (0.63)	0.345 (4.06)
1976	1.109 (4.02)	1.638 (4.55)	−0.062 (−0.50)	0.022 (0.18)	−0.089 (−0.75)	−0.158 (−1.17)	0.106 (1.04)	−0.125 (−0.93)	0.007 (−0.08)	−0.010 (−0.07)	0.225 (1.47)	0.534 (3.35)	0.250 (4.23)
1977	0.954 (4.60)	0.241 (2.29)	0.098 (0.98)	0.082 (0.64)	0.129 (1.11)	0.040 (0.41)	0.328 (2.44)	0.066 (0.74)	0.085 (0.78)	0.093 (1.16)	0.249 (1.73)	0.007 (0.06)	0.194 (5.05)
1978	0.444 (3.47)	0.343 (2.81)	0.350 (3.15)	0.268 (2.63)	0.414 (3.75)	0.213 (1.45)	−0.065 (−0.67)	0.755 (5.48)	0.328 (1.91)	−0.988 (−2.88)	0.230 (0.85)	−0.097 (−0.71)	0.186 (3.26)
1979	0.548 (4.13)	0.157 (1.30)	0.183 (1.34)	0.202 (2.07)	0.200 (1.57)	0.059 (0.61)	−0.039 (−0.32)	0.205 (2.80)	−0.104 (−0.85)	−0.092 (−0.44)	−0.049 (−0.32)	0.153 (1.20)	0.121 (3.05)
Mean daily return over all years	0.714 (11.81)	0.223 (5.07)	0.136 (4.34)	0.042 (1.42)	0.072 (2.13)	0.028 (0.80)	0.105 (3.11)	0.063 (1.88)	0.092 (2.74)	−0.070 (−1.65)	0.006 (0.16)	0.42 (0.93)	0.121 (10.31)

across the remaining eleven months; although a few display significant positive size premia (e.g., February, March and July) that are much lower than January, the month of October displays a size discount.

The studies by Banz (1981) and Reinganum (1981) implicitly assume that anomalous size-related excess returns are obtained continuously; i.e., month-by-month, year-by-year small firms earn larger returns than large firms after controlling for risk. The evidence in table 2 casts doubt on the month-to-month constancy of the anomalous size effect. In fact, a significant proportion of the size effect, averaged over 1963–1979, is due to return premia observed during January in *every* year. Exclusion of the large January abnormal returns reduces the overall magnitude of the anomaly by almost fifty percent: the average annual size premium of 30.3 percent declines to 15.4 percent when the January observations are removed.[8]

To test the null hypothesis of equal expected abnormal returns for each month of the year, I use the regression

$$R_t = a_1 + a_2 D_{2t} + a_3 D_{3t} + \ldots + a_{12} D_{12t} + e_t. \tag{3}$$

In the regression, R_t is the average daily CRSP excess return for day t for the size portfolio under consideration, and the dummy variables indicate the month of the year in which the excess return is observed (D_{2t} = February, D_{3t} = March, etc.). The excess return for January is measured by a_1, while a_2 through a_{12} represent the differences between the excess return for January and the excess returns for the other months. If the expected excess return is the same for each month of the year, the estimates of a_2 through a_{12} should be close to zero and the F-statistic measuring the joint significance of the dummy variables should be insignificant.

Due to heteroskedastic residuals in the OLS estimate of (3), I estimate eq. (3) as weighted least squares (WLS) where the weight applied to an observation in month t is the normalized value of the inverse of the standard error for month t of the residuals from the OLS estimate of (3).[9] WLS

[8]These abnormal returns are derived from table 2. The estimate of the average annual size premium equals the overall average daily difference in abnormal returns between the smallest and largest market value portfolios (0.121) multiplied by the number of trading days per year (252). The component of the size effect occurring in January equals the average daily difference between the smallest and largest market value portfolios (0.714) multiplied by the average number of trading days in January (21). The average January effect of 15% is 49.3% of the average magnitude of the size effect.

[9]Heteroskedasticity is induced into the residuals by way of the heteroskedasticity in the excess returns. The value of a generalized likelihood ratio test for equality of excess return variances across months [see Mood, Graybill and Boes (1974, p. 439)] ranges from 319.4 for portfolio 1 (smallest firms) to 27.2 for portfolio 5. These values are significant at the 1% level and one can therefore reject the hypothesis of equal excess return variance across months for any particular portfolio.

Although autocorrelated disturbances are present in the OLS estimates of (3), I do not adjust for the autocorrelation in the test of the month-to-month stability hypothesis. Vinod (1975) estimates of (3) are presented in table 3. In addition to tests of the hypothesis that the mean abnormal returns of the ten size portfolios are temporally constant, I also test for the month-to-month stationarity of the differences in mean abnormal returns between the smallest and largest market value portfolios. Three interesting results emerge. First, average excess returns for smaller firms appear disproportionately large in January relative to the remaining eleven months. For example, the F-statistic of 14.59 for the smallest firm portfolio is significant at any level and allows rejection of the null hypothesis. Second, and somewhat surprisingly, January abnormal returns for the larger firm portfolios are *negative* and lower than the mean excess returns in any other month. The large F-statistic of 17.63 for the largest firm portfolio also allows rejection of the hypothesis of temporal constancy of excess returns for large firms. Third, the estimates of (3) for differences in average excess returns between the smallest and largest market value portfolios indicates the observed size premium in January is positive and significantly larger than the average premium in any other month. The F-statistic of 18.9 permits rejection of the hypothesis of a stable month-to-month size effect.

3.3. A closer look at the January effect: The first five trading days

A major implication of the previous section is that a significant portion of the size effect, averaged over 1963–1979, is due to return premia observed in the month of January in every year. Closer examination of the abnormal returns within the month of January reveals a large portion of the size effect occurs during the first five trading days of the year. The magnitude of the size effect during the first week of trading in the year is shown in table 4 which contains the difference in average abnormal return between the smallest and largest market value portfolios for these five days. The first trading day's difference in abnormal returns between the smallest and largest market value portfolios averages 3.2 percent with a standard deviation of 2.0 percent for the 1963–1979 period. The first day's difference is positive in *every* year and the average difference is significant at any level. Further, the difference in abnormal return between the smallest and largest market value portfolios averaged 8.0 percent over the first five trading days in January.[10] Thus, 10.5 percent ($3.2 \div 30.4$) of the annual size effect for an average year

estimates bounds for the F-statistic from an OLS equation estimated in the presence of AR(1) errors. These bounds are used to assess the likelihood of making an incorrect test decision given that the residuals from the OLS estimate of (3) are approximately AR(1). The relevant bounds for the test here indicate that the F-statistic remains significant if (3) is estimated while directly accounting for the autocorrelation.

[10]The second through fifth trading days in January display 17, 16, 16 and 15 positive differences, respectively, out of a possible 17 over the 1963–1979 period.

Table 3

Test of the month-to-month stability of the size effect using CRSP daily excess returns (in percent) of portfolios constructed from firms in each decile of size (measured by market value of equity) on the NSYE–AMEX over the period 1963–1979.[a]

$R_t = a_1 + a_2D_{2t} + a_3D_{3t} + a_4D_{4t} + a_5D_{5t} + a_6D_{6t} + a_7D_{7t} + a_8D_{8t} + a_9D_{9t} + a_{10}D_{10t} + a_{11}D_{11t} + a_{12}D_{12t} + e_t$

Decile	a_1	a_2	a_3	a_4	a_5	a_6	a_7	a_8	a_9	a_{10}	a_{11}	a_{12}	R^2	F-statistic	Degrees of freedom
Smallest	0.43	−0.27	−0.35	−0.40	−0.36	−0.40	−0.37	−0.37	−0.40	−0.44	−0.43	−0.41	0.036	14.59	11;4250
	(13.28)	(−6.04)	(−8.62)	(−9.80)	(−8.78)	(−9.73)	(−9.04)	(−9.08)	(−9.68)	(−10.75)	(−10.33)	(−9.44)			
2	0.19	−0.13	−0.15	−0.17	−0.17	−0.18	−0.14	−0.18	−0.17	−0.21	−0.20	−0.16	0.020	8.03	11;4250
	(9.30)	(−4.76)	(−5.76)	(−6.65)	(−6.69)	(−6.69)	(−5.31)	(−7.10)	(−6.49)	(−8.05)	(−7.49)	(−5.85)			
3	0.09	−0.07	−0.05	−0.08	−0.08	−0.08	−0.07	−0.08	−0.09	−0.10	−0.07	−0.07	0.008	3.06	11;4250
	(5.92)	(−3.38)	(−2.80)	(−4.10)	(−4.30)	(−3.92)	(−3.41)	(−4.41)	(−4.57)	(−4.73)	(−3.73)	(−3.62)			
4	0.02	−0.01	−0.00	−0.03	−0.03	−0.03	−0.01	−0.01	−0.02	−0.04	−0.00	−0.01	0.004	1.39	11;4250
	(1.71)	(−0.53)	(−0.06)	(−1.73)	(−2.01)	(−1.62)	(−0.45)	(−0.83)	(−1.39)	(−2.56)	(−0.18)	(−0.75)			
5	−0.03	0.02	0.03	0.04	0.02	0.03	0.03	0.02	0.02	0.02	0.05	0.04	0.004	1.38	11;4250
	(−3.13)	(1.25)	(2.10)	(2.64)	(1.24)	(2.16)	(2.21)	(1.55)	(1.64)	(1.35)	(3.03)	(2.61)			
6	−0.09	0.06	0.07	0.07	0.07	0.07	0.09	0.09	0.09	0.07	0.09	0.07	0.014	5.54	11;4250
	(−7.84)	(3.90)	(4.61)	(5.02)	(4.80)	(4.47)	(6.01)	(6.46)	(5.98)	(4.98)	(6.16)	(4.89)			
7	−0.13	0.10	0.11	0.12	0.12	0.12	0.12	0.14	0.11	0.13	0.16	0.14	0.030	11.92	11;4250
	(−11.57)	(6.54)	(7.31)	(7.66)	(7.75)	(7.83)	(7.83)	(9.56)	(7.44)	(8.60)	(9.91)	(8.48)			
8	−0.19	0.15	0.17	0.18	0.17	0.19	0.17	0.19	0.16	0.21	0.21	0.19	0.044	17.85	11;4250
	(−14.00)	(7.96)	(9.72)	(10.36)	(9.72)	(10.61)	(9.52)	(11.30)	(9.40)	(11.89)	(11.44)	(10.17)			
9	−0.22	0.16	0.18	0.22	0.20	0.21	0.20	0.22	0.19	0.25	0.24	0.20	0.044	17.70	11;4250
	(−14.00)	(7.73)	(9.23)	(10.72)	(9.54)	(10.32)	(9.64)	(10.94)	(9.35)	(11.59)	(11.41)	(9.52)			
Largest	−0.28	0.22	0.23	0.28	0.28	0.29	0.24	0.28	0.23	0.34	0.28	0.26	0.044	17.63	11;4250
	(−13.83)	(8.20)	(8.86)	(10.79)	(10.60)	(10.86)	(8.94)	(10.67)	(8.54)	(11.81)	(9.97)	(9.67)			
Smallest −largest	0.71	−0.49	0.58	−0.67	−0.64	−0.69	−0.61	−0.65	−0.62	−0.78	−0.71	−0.67	0.047	18.90	11;4250
	(14.88)	(−7.68)	(−9.80)	(−11.47)	(−10.68)	(−11.39)	(−10.12)	(−10.90)	(−10.36)	(−12.55)	(−11.44)	(−10.58)			

[a]The dummy variables indicate in which month of the year each excess return is observed (D_{2t} = February, D_{3t} = March, etc.), and the estimated t-values are in parentheses. The F-statistic tests the hypothesis that a_2 through a_{12} are zero. Fractiles of the F-distribution: $F_{10,\infty}(95\%) = 1.83$, $F_{10,\infty}(99.9\%) = 2.96$.

Table 4

Differences in average daily (CRSP) excess returns (in percent) between portfolios constructed from firms in the top and bottom decile of size (measured by market value of equity) on the NYSE and AMEX for the first five trading days in each of the years 1963–1979.

Trading day	1963	1964	1965	1966	1967	1968
1	2.30	1.99	2.09	1.34	3.39	2.18
2	0.00	0.26	0.85	0.69	0.58	1.47
3	0.72	0.19	0.60	0.52	0.96	0.17
4	0.48	0.02	0.75	0.25	0.27	1.43
5	−0.69	0.38	0.36	0.43	1.08	1.55
Avg.	0.56	0.57	0.93	0.65	1.26	1.36

Trading day	1969	1970	1971	1972	1973	1974
1	1.25	3.98	1.66	2.36	3.68	6.46
2	0.65	3.68	0.88	1.56	1.59	5.72
3	−0.36	1.44	1.92	1.40	0.46	3.84
4	−1.63	1.19	1.59	1.63	0.84	3.19
5	−0.23	0.43	1.12	1.12	0.62	2.72
Avg.	−0.06	2.14	1.43	1.61	1.44	4.38

Trading day	1975	1976	1977	1978	1979	Avg. for 1963–1979
1	8.47	5.26	3.84	2.28	1.82	3.20
2	4.54	2.36	2.14	0.27	1.36	1.68
3	3.17	2.27	1.74	0.94	1.26	1.25
4	6.00	0.81	0.77	0.63	1.13	1.14
5	4.64	0.47	0.13	0.09	0.90	0.89
Avg.	5.36	2.23	1.72	0.84	1.29	1.63

can be attributed to the *first* trading day of the year and 26.3 percent ($8.0 \div 30.4$) can be attributed to the *first five* trading days. In contrast, if the small firm premium were spread uniformly throughout the year, then approximately 0.4 percent of the annual premium is expected on each day.

3.4. The January effect and the year-to-year stationarity of the size effect

Brown, Kleidon and Marsh (1983) report that the size effect is not stable from year to year during the 1967–1979 period, but identify two distinct subperiods when the relation between size and abnormal return is relatively stable. There is a stable positive relation from 1969 to 1973 and a stable negative relation from 1974 to 1979. The reversal of the size anomaly during the 1969–1973 period is apparent in the rightmost column of table 2. With the exception of 1971, large firms consistently outperform small firms over this period.[11]

The year-to-year instability of the size effect implies that estimation of (3) across all three subperiods, and without regard to important differences in the size effect within each subperiod, may be inappropriate. I therefore estimate eq. (3) within each subperiod identified above, and the results permit us to conclude that the January effect is insensitive to the year-to-year instability of the overall size effect. The average January differences in *monthly* percentage abnormal returns between the smallest and largest market value portfolios, and corresponding t-statistics in parentheses, for the three subperiods are: 1963–1968, 9.7(10.2); 1969–1973, 11.3(9.1); 1974–1979, 23.1(13.0). Two interesting observations are obtained from the subperiod evidence. First, the magnitude of the January effect increases through time in the 1963–1979 period. Second, and more importantly, even during the 1969–1973 subperiod when large firms have higher risk-adjusted returns than small firms, the size premium is significantly positive in January. An interesting implication is that after accounting for the (always) positive January effect, the year-to-year instability of the size anomaly remains.

3.5. The January anomaly and abnormal return autocorrelation

Roll (1981), following Scholes and Williams (1977) and Dimson (1979), argues that observed autocorrelations of abnormal returns for small firms may be due to infrequent trading. The evidence in the previous sections suggests an alternative explanation for the observed autocorrelations. In particular, non-stationary mean abnormal returns may induce autocorrelation in the time series calculations. If average daily abnormal returns in January are large relative to average daily abnormal returns during the remaining eleven months, then most daily abnormal returns will be less than the grand mean based on all days. This may result in positive computed autocorrelations, suggesting that autocorrelations be computed while controlling for a varying mean daily abnormal return. If the positive autocorrelations vanish, the test results imply that the source of the autocorrelation is a misspecified stochastic process and not infrequent trading.

To test the hypothesis that non-stationary mean abnormal returns may cause autocorrelation, I estimate eq. (3) with OLS. The residuals from (3) are mean-adjusted abnormal returns and, therefore, are used to compute autocorrelations that are free of the changing mean problem. The computed first-order autocorrelation of the residuals for the smallest market value

[11]The average differences in annualized percentage excess returns between the smallest and largest market value portfolios, and corresponding t-statistics in parentheses, for three subperiods are: 196–1968, 35.3(10.4); 1969–1973, $-5.0(-1.1)$; and 1974–1979, 55.4(8.5).

portfolio is 0.189 and for the largest market value portfolio is 0.320. These autocorrelations are not significantly different from those reported for the unadjusted excess return series in table 1.

Autocorrelations were also computed for another adjusted time series of abnormal returns, adjusted so that all January observations are excluded. Elimination of the January abnormal returns reduces the first-order autocorrelation to 0.187 for the smallest market value portfolio and 0.274 for the largest market value portfolio. These adjusted autocorrelations are still significantly different from zero for the extreme portfolios. It does not appear that the non-stationary mean of excess returns is the cause of the observed autocorrelation in excess returns.

4. Hypotheses regarding the January effect

Several hypotheses have been suggested to explain the January seasonal in stock returns. Most prominent are a tax loss selling hypothesis and an information hypothesis, although neither has been theoretically nor empirically linked to the return seasonal. The testable implications of these potential explanations may, nevertheless, yield insights into the nature of the January premia for small firms. It appears, however, that the ability of either hypothesis to explain the January effect is diminished either on grounds of plausibility or as a result of some preliminary testing.

4.1. *Tax loss selling hypothesis*

Wachtel (1942) and Branch (1977) formulate an explanation for disproportionately large January returns based on year-end tax loss selling of shares that have declined in value over the previous year.[12] Since size is measured here as total market value of equity, the smallest firm portfolios are biased toward inclusion of shares that have experienced large price declines and, therefore, are likely candidates for tax loss selling.[13] The possible association between tax loss selling and the January effect merits consideration.

The theoretical value of the tax loss selling hypothesis diminishes, however, with the existence of arbitrage possibilities in non-segmented markets with non-taxable investors.[14] In addition, the hypothesis is not clearly supported

[12] Dyl (1977) reports abnormally heavy trading volume at year-end for shares with previous twelve-month price declines and interprets the results as evidence of tax loss selling.

[13] Evidence presented in Keim (1982) indicates that the largest abnormal returns recorded in the first five trading days of January are associated with low-priced shares that have gravitated to the smallest market value portfolio. Most of these shares sell for less than two dollars.

[14] Roll (1982) argues that the annual pattern in small firm returns is strongly associated with tax loss selling, and conjectures that large transactions costs for smaller firm shares prevent arbitragers from eliminating the large abnormal returns in the first few days in January.

empirically. If the January effect is the result of year-end tax loss selling, then the magnitude and significance of the measured January effect should, *ceteris paribus*, vary with the level of personal income tax rates. For example, the January effect should be less significant in pre-World War II years when personal tax rates were relatively low. Keim (1982) reports, however, that the January effect is, on average, larger in the 1930's than in any subsequent subperiod.[15]

Of course, other things are not always equal. The methods available today for shielding income from taxes did not exist in the earlier years in this century. Thus, the marginal benefit of the capital loss offset may have been much greater in the 1930's than in later years, even though the tax rates in the former period were significantly lower. A more direct test of the tax loss selling hypothesis is possible though. The month-to-month behavior of abnormal returns across countries with tax codes similar to the United States' code but with differing tax year-ends (e.g., Great Britain has a May tax year-end) can be examined. If, in the month immediately following the tax year-end, abnormal returns of small firms in other countries are large relative to both other months and larger firms in that country, then the evidence is consistent with the tax loss selling hypothesis.[16]

4.2. *Information hypothesis*

Rozeff and Kinney (1976) note that 'January marks the beginning and ending of several potentially important financial and informational events... January is the start of the tax year for investors, and the beginning of the tax and accounting years for most firms, and preliminary announcements of the previous calendar year's accounting earnings and made'. Thus, at least for those firms with year-end fiscal closings, the month of January marks a period of increased uncertainty and anticipation due to the impending release of important information. In addition, the gradual dissemination of this information during January may have a greater impact on the prices of small firms relative to large firms for which the gathering and processing of information by investors is a less costly process. The recurrence of significant small firm premia at the same time each year due to inadequate adjustment of prices to information is, however, inconsistent with a rational expectations equilibrium in the market. Nonetheless, the information hypothesis is testable. One test involves aligning all firm excess returns in event time rather

[15]Branch (1977), Dyl (1977) and Roll (1982) examine tax loss selling in the post-1960 period only.

[16]Korajczyk (1982), using value-weighted stock market indices from eighteen countries with widely varying tax regimes and tax year-ends, finds evidence of a January seasonal in stock returns for each country (except Spain) over the period January 1973 to January 1982.

than in calendar time, where the event is the firm's fiscal year-end.[17] If the magnitude of the average excess returns immediately following the fiscal year-end in event time is greater than the measured excess returns in the month of January in calendar time, then the evidence would support the information hypothesis.

4.3. Other possible explanations

There remains the possibility that the measured January effect may not have an economic cause. That is, the effect may be due to spurious causes such as outliers, concentration of listings and delistings at year-end, or data base errors. Keim (1982) reports that the cross-sectional distribution of excess returns for the decile of smallest NYSE–AMEX firms is positively skewed for the first trading day in January. However, elimination of extreme observations (greater than three standard errors from the sample mean) does not significantly reduce the mean of the distribution. Roll (1982) investigates the other two 'non-exploitable' causes and dismisses them for lack of evidence.[18]

5. Summary and conclusions

Recent empirical investigations of the traditional CAPM report the existence of anomalous abnormal returns that appear to be negatively related to size. Evidence in section 3 indicates that daily abnormal return distributions in January have large means relative to the remaining eleven months, and that the relation between abnormal returns and size is always negative and more pronounced in January than in any other month — even in years when, on average, large firms earn larger risk-adjusted returns than small firms. Nearly fifty percent of the average magnitude of the size anomaly over the period 1963–1979 is due to January abnormal returns. Further, more than fifty percent of the January premium is attributable to large abnormal returns during the first week of trading in the year, particularly on the first trading day. Hypotheses advanced to explain the size effect appear unable to explain the January effect. Several alternative explanations with testable implications are discussed, but the tests are deferred for future research.

[17]Approximately sixty percent of the firms listed on the NYSE and AMEX have December 31 fiscal year-ends.

[18]The method used in this study for computing portfolio abnormal returns (1) requires restructuring portfolio composition at each year-end and (2) may induce a survival bias because of the requirement of one year of trading for all sample firms. Investigation of the January effect may be sensitive to these strict requirements. The results in section 3 were duplicated, however, with differences in returns between daily equal- and value-weighted NYSE–AMEX indices. Biases due to survival and year-end portfolio restructuring are eliminated when using those indices.

References

Banz, R.W., 1981, The relationship between return and market value of common stocks, Journal of Financial Economics 9, 3–18.

Basu, S., 1977, The investment performance of common stocks in relation to their price-earnings ratios: A test of the efficient market hypothesis, Journal of Finance 32, 663–682.

Branch, B., 1977, A tax loss trading rule, Journal of Business 50, 198–207.

Brown, P., A. Kleidon and T. Marsh, 1983, New evidence on the nature of size-related anomalies in stock prices, Journal of Financial Economics, this issue.

Center for Research in Security Prices, 1980, CRSP daily excess return file, Unpublished manuscript (Graduate School of Business, University of Chicago, Chicago, IL).

Dimson, E., 1979, Risk measurement when shares are subject to infrequent trading, Journal of Financial Economics 7, 197–216.

Dyl, E.A., 1977, Capital gains taxation and year-end stock market behavior, Journal of Finance 32, 165–175.

Fama, E.F., 1965, The behavior of stock market prices, Journal of Business 38, 34–105.

Fama, E.F., 1970, Efficient capital markets: A review of the theory and empirical work, Journal of Finance 25, 383–420.

Fama, E.F., 1976, Foundations of finance (Basic Books, New York).

French, K.R., 1980, Stock returns and the weekend effect, Journal of Financial Economics 8, 55–70.

Gibbons, M.R. and P.J. Hess, 1981, Day of the week effects and asset returns, Journal of Business 54, 579–596.

Keim, D.B., 1980, Asset pricing anomalies and capital market seasonality: Empirical evidence, Unpublished manuscript (University of Chicago, Chicago, IL).

Keim, D.B., 1982, Further evidence on size effects and yield effects: The implications of stock return seasonally, Unpublished manuscript (University of Chicago, Chicago, IL).

Korajczyk, R., 1982, Stock market seasonality: Some international evidence, Unpublished manuscript (University of Chicago, Chicago, IL).

Mood, A.M., F.A. Graybill and D.C. Boes, 1974, Introduction to the theory of statistics (McGraw-Hill, New York).

Officer, R.R., 1975, Seasonality in Australian capital markets: Market efficiency and empirical issues, Journal of Financial Economics 2, 29–52.

Reinganum, M.R., 1981, Misspecification of capital asset pricing: Empirical anomalies based on earnings yields and market values, Journal of Financial Economics 9, 19–46.

Reinganum, M.R., 1982, A direct test of Roll's conjecture on the firm size effect, Journal of Finance 37, 27–36.

Roll, R., 1981, A possible explanation of the small firm effect, Journal of Finance 36, 879–888.

Roll, R., 1982, The turn of the year effect and the return premia of small firms, Working paper (Graduate School of Management, University of California, Los Angeles, CA).

Rozeff, M.S. and W.R. Kinney, Jr., 1976, Capital market seasonality: The case of stock returns, Journal of Financial Economics 3, 379–402.

Scholes, M. and J. Williams, 1977, Estimating betas from nonsynchronous data, Journal of Financial Economics 5, 309–327.

Schultz, P., 1982, Some additional evidence on the small firm effect, Unpublished manuscript (University of Chicago, Chicago, IL).

Stattman, D., 1980, Book values and expected stock returns, Unpublished MBA honors paper (University of Chicago, Chicago, IL).

Stoll, H.R. and R.E. Whaley, 1983, Transaction costs and the small firm effect, Journal of Financial Economics, this issue.

Vinod, H.D., 1976, Effects of ARMA errors on the significance tests for regression coefficients, Journal of the American Statistical Association 71, 929–933.

Wachtel, S.B., 1942, Certain observations on seasonal movements in stock prices, Journal of Business 15, 184–193.