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THE JANUARY EFFECT ON A SMALL STOCK MARKET: LUMPY INFORMATION AND TAX-LOSS SELLING[+]

by

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ABSTRACT

A significant excess January return is reported for the Finnish (Helsinki) stock exchange for the years 1970-81. This excess return, however, appears only weakly linked to otherwise highly significant firm size premiums. Two propositions on the seasonal return pattern are tested. The first, suggesting that seasonality is caused by lumpy information, specifically by a bimodal distribution of stock split, dividend or issue announcements, is clearly refuted. The second--the tax-loss selling proposition--is strongly supported by the Finnish data. Apparently due to tax induced year-end selling pressure, January stock returns are found to depend positively on the market-wide demand for tax write-offs and negatively on the stock's individual past performance.

1. Introduction

For quite some time, the prospect of uncovering seasonal variation in the prices of perfectly storable commodities and assets has intrigued econometricians--the 1942 paper by Wachtel testifies to this fact. Although primarily confined to commodity markets, this search for seasonality appears to have become a hot topic in asset market research also. Here, interest has focused on a day-of-the-week effect (French 1980, and Lakonishok-Levi 1982) and recently especially on what has become known as the January or the turn of the (tax) year effect.

By that, reference is made to the abnormally high yields on securities observed on many, if not most, stock exchanges in the month of January. A vast amount of research on this effect has already been published, some of which is briefly surveyed below.

No single seminal paper on stock market seasonality appears to exist. The recent upsurge in the debate was preceded by the papers by Officer (1975) and Rozeff-Kinney (1976) on general seasonality and the Branch (1977) and Dyl (1977) papers on tax-loss trading. Branch was able to report excess returns to a simple tax-loss trading rule essentially consisting of purchases of 'losers'--stocks allegedly selling at prices depressed by heavy pressure from investors realizing their tax deductible capital losses--in December. Simultaneously, Dyl (1977), in a few simple regressions, showed that NYSE trading volume data also is consistent with a tax loss selling hypothesis. That hypothesis has since been examined in greater detail and additional empirical evidence has been produced by, among other, Gultekin and Gultekin (1982) and Brown, Keim, Kleidon and Marsh (1983).

The latest development seems to be the link between the small firm effect of Banz (1981), Reinganum (1981) and Reinganum (1982a) and the turn-of-the-year premium. Roll (1983) observes that for the 18 consecutive years covered by his data an equally weighted NYSE return exceeded the value weighted one. In an explicit test of a tax-related turn-of-the-year hypothesis, Reinganum (1983) concludes that tax loss selling might account for the abnormally high returns reported for early January, but that such behavior is unlikely to explain the large returns reported for small stocks later in that month.

The object of this paper is to examine the seasonal pattern in rates of return on a small European exchange--the Helsinki Stock Exchange (HESE) in Finland. Having in section 2 uncovered a pattern including a highly significant positive excess January return, it goes on to examine and test two major propositions on its emergence. According to the first proposition, the return seasonality reflects lumpiness, or more specifically, a seasonal pattern in the inflow of information to the stock market. In a general form, such a proposition is open to criticism based on rather obvious arbitrage arguments. In such a form it is also inevitably, for reasons associated with the immeasurability of information, hard to test.

The information-related test proposed in this paper is restricted to only one item of apparent informational value, namely announcements concerning stock dividends, splits, or equity rights issues. Since these are markedly seasonal and since their anticipation is known to produce positive price reactions, they may account for a significant portion of the seasonality contained in the market return data. This proposition is examined in greater detail in Section 3 of the paper.

The second proposition on the January excess returns is the more traditional tax-loss selling hypothesis. The fact that capital losses by Finnish legislation are deductible only from capital gains and not from other sources of income makes 'dumping of losers' profitable only in periods where investors have realized positive capital gains. Hence any tax-induced January effect in the Finnish market must be modeled a function not only of capital losses to specific assets, but of the overall market behavior also. Such a model will be developed and estimated in section 4 of this paper. The fifth and final section of the paper contains a summary and a discussion.

2. The evidence on Market Seasonality

The data for the present study consists of monthly, and where necessary daily, returns to the 50 stocks that have been listed on the HESE throughout the period from January 1970 to December 1981. In table 1 these stocks are arranged into ten portfolios in decreasing order of their market values on December last 1969.[1] For each portfolio mean returns in excess of the riskfree rate and the Dimson risk premium[2] are reported for each calendar month. The table also contains monthly excess returns to the value-weighted market index and Dimson beta estimates together with mean annual returns over the whole 1970-81 period for each portfolio.

Table 1 reveals a fairly strong January effect averaging roughly 2.5 percent. Positive excess returns to the market in the neighborhood of 1.6 percent are also reported for July and December. For May and the three fall months negative excess returns are reported almost throughout. The January effect appears to some extent linked to small firm effect. Mean excess January returns exceed 3 percent for portfolios 5-10 while falling below for remaining (large firm) portfolios. The coefficient of determination between mean excess January returns and portfolio capitalization is .39 and the t-statistic for the negative slope is--2.25.

To test the statistical significance of the seasonality, we employed a procedure of regressing monthly returns in excess of portfolio means on

Table 1. Mean monthly excess returns to the market portfolio and to 10 stock portfolios in decreasing order of market value at period's start.

Portfolio		Jan	Feb	Mar	Apr	May	Jun	Jul	Aug	Sep	Oct	Nov	Dec	Total	Dimson beta
Large firm	1	.0212	-.0109	.0221	-.0150	-.0266	-.0038	.0179	-.0106	-.0244	-.0133	-.0074	.0336	-.0216	.949
	2	.0154	.0021	.0041	-.0210	-.0211	.0030	.0206	-.0099	-.0252	-.0092	-.0000	.0259	-.0182	1.091
	3	.0173	.0053	.0004	.0120	-.0214	-.0051	.0208	-.0156	-.0182	-.0056	-.0115	.0160	-.0089	1.109
	4	.0266	.0027	-.0128	.0124	-.0207	.0099	.0306	-.0216	-.0082	-.0032	-.0169	.0108	.0071	1.119
	5	.0635	.0158	.0157	.0089	-.0172	.0098	.0015	.0035	-.0076	-.0182	-.0176	-.0081	.0442	1.023
	6	.0372	.0169	.0188	.0195	-.0271	.0120	.0146	.0007	-.0158	-.0287	-.0126	.0164	.0466	.941
	7	.0303	.0052	.0105	.0196	-.0286	.0167	.0043	-.0033	-.0274	-.0085	-.0125	.0177	.0260	1.004
	8	.0470	.0145	.0462	.0051	-.0200	.0007	.0012	.0195	-.0142	-.0195	-.0006	-.0084	.0686	.803
	9	.0740	.0574	.0090	.0216	-.0066	.0014	.0076	.0060	-.0021	-.0320	-.0319	-.0359	.0710	.987
Small firm	10	.0386	.0510	.0038	-.0041	-.0203	.0222	.0057	.0210	.0088	-.0112	-.0225	-.0019	.0870	.521
Market index		.0251	.0031	.0097	-.0026	-.0197	.0005	.0163	-.0036	-.0183	-.0115	-.0089	.0161		1.000

monthly dummies while suppressing the intercept. Note that by the deduction of portfolio means, i.e., the non-seasonal return component from monthly return data before regressing, we attempt falsification of the null hypothesis that no seasonality is present in the rates of return. This interpretation of the January effect differs from the one employed by Reinganum (1983) who, by regressing daily (total) returns on a set of dummies, only tests for departures from a zero return hypothesis.

The results set out in table 2 indicate that the January effect is indeed highly significant at the market level. It averages 3.2 percent which is significantly different from zero on the 1-percent level. The link between the small-firm and the January effect is, however, much less evident in ours than in Roll's or Reinganum's data. The January effect is significant (5-percent level) for portfolios 1 and 5 through 9, while exceeding 5 percent only for portfolios 5 and 9. In contrast, Reinganum (1983) found January effects that decreased almost monotonically with portfolio capitalizations.

Table 2 furthermore reveals a seasonal pattern in which the January effect is far from the only similarity between the American and the Finnish data. Although the fact that Reinganum's (1983) coefficient values are not actual seasonal effects but monthly means, makes a detailed comparison difficult; the seasonal pattern suggested by his data is strikingly similar to that contained in table 2 of this paper. Both data sets suggest that returns in the winter and spring months up until May are predominantly positive, May being distinctly negative, while the fall months, especially October, exhibit negative return effects. Obviously no explanation to these similarities can be offered.

The conclusion of this initial screening of the data must hence be that a significant January effect can be detected on the HESE. While an excessive

Table 2. Estimated coefficient values of monthly dummies for returns to ten stock portfolios composed in decreasing order of capitalization (1-largest, 10-smallest). Absolute t-statistics in parentheses.

	Portfolio	Jan	Feb	Mar	Apr	May	Jun	Jul	Aug	Sep	Oct	Nov	Dec	r^2
Large firm	1	2.43	-1.21	2.21	-1.01	-2.51	0.00	1.90	-1.09	-2.58	-1.25	-0.65	3.79	.217
		(2.14)	(1.07)	(1.94)	(.88)	(2.21)	(0.00)	(1.67)	(.96)	(2.27)	(1.10)	(.57)	(3.33)	
	2	1.63	0.40	0.60	-2.12	-1.87	0.29	2.53	-1.16	-2.14	-0.87	-0.12	2.62	.134
		(1.32)	(.33)	(.48)	(1.71)	(1.51)	(.24)	(2.05)	(.94)	(1.73)	(.70)	(.09)	(2.12)	
	3	1.93	0.50	0.28	0.55	-1.72	-0.38	1.92	-0.83	-1.85	-0.71	-1.58	1.83	.094
		(1.54)	(.40)	(.22)	(.43)	(1.37)	(.30)	(1.53)	(.66)	(1.47)	(.57)	(1.26)	(1.46)	
	4	2.09	0.63	-1.18	0.92	-1.97	0.52	2.54	-1.61	-0.63	-0.40	-1.72	0.86	.106
		(1.67)	(.50)	(.94)	(.73)	(1.58)	(.42)	(2.03)	(1.29)	(.50)	(.32)	(1.38)	(.68)	
	5	5.46	1.10	1.01	0.45	-2.09	0.11	-0.22	0.05	-1.27	-1.91	-1.96	-0.79	.160
		(4.02)	(.81)	(.74)	(.33)	(1.54)	(.08)	(.16)	(0.04)	(.93)	(1.40)	(1.44)	(.58)	
	6	3.34	1.51	1.31	1.20	-3.06	0.46	1.02	-0.10	-1.98	-3.16	-1.51	0.99	.178
		(2.68)	(1.21)	(1.05)	(.96)	(2.46)	(.37)	(.82)	(.08)	(1.59)	(2.54)	(1.21)	(.79)	
	7	2.91	0.06	0.73	1.24	-3.69	1.69	0.79	-0.16	-2.73	-0.84	-1.47	1.42	.144
		(2.17)	(.04)	(.54)	(.92)	(2.75)	(1.26)	(.59)	(.12)	(2.03)	(.63)	(1.09)	(1.06)	
	8	4.24	0.58	2.66	0.03	-2.47	-0.19	-0.65	1.80	-1.39	-2.39	-0.60	-1.65	.171
		(3.27)	(.44)	(2.05)	(.02)	(1.90)	(.15)	(.50)	(1.39)	(1.07)	(1.84)	(.46)	(1.27)	
	9	6.46	5.43	1.05	1.13	-0.19	-0.81	-0.09	0.36	-1.23	-4.05	-3.96	-4.11	.239
		(3.72)	(3.13)	(.61)	(.65)	(.11)	(.47)	(.05)	(.20)	(.71)	(2.33)	(2.28)	(2.37)	
Small firm	10	1.89	5.30	-0.74	-0.87	-2.40	1.16	0.15	0.66	-0.43	-1.50	-1.94	-1.24	.078
		(.91)	(2.55)	(.36)	(.42)	(1.16)	(.56)	(.07)	(.32)	(.20)	(.72)	(.93)	(.60)	
Equally weighted market		3.23	1.43	0.79	0.15	-2.20	0.28	0.99	-0.21	-1.62	-1.71	-1.53	0.37	.181
		(3.35)	(1.48)	(.82)	(.15)	(2.28)	(.29)	(1.02)	(.22)	(1.68)	(1.77)	(1.59)	(.38)	

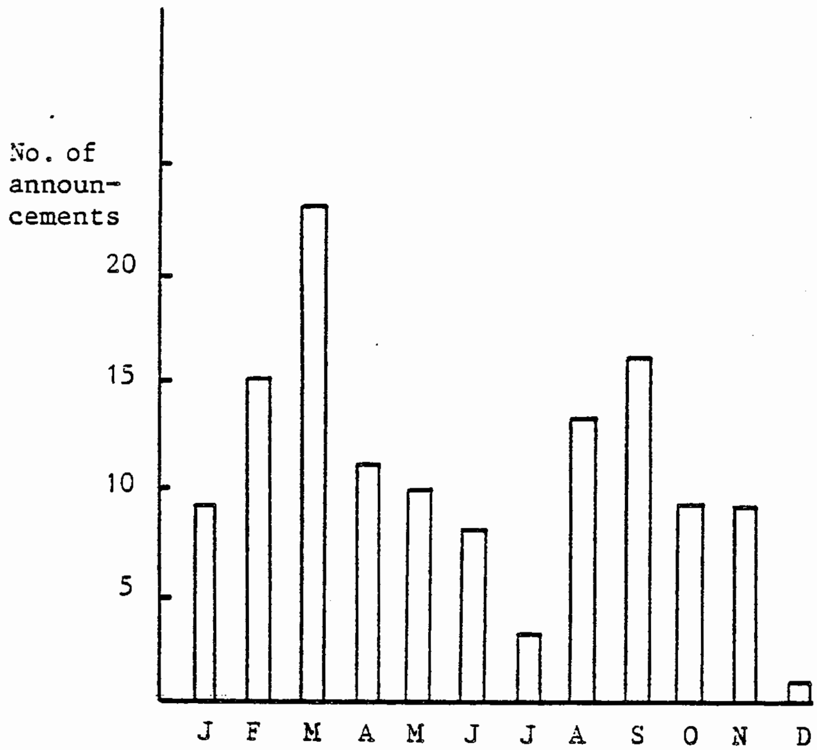
January return is somewhat more likely to occur in a small than a large firm stock, the link between the small firm and the January effects appears considerably weaker than in the American data. We would furthermore tend to regard even the American link as something of an open issue since, as long as a size premium is present in annual data--an effect for which a tentative but apparently empirically valid explanation has been offered in Berglund-Wahlroos (1983)--we would expect it to be present in any seasonal returns also. Hence, a monotonic relationship between January returns and portfolio capitalizations such as the one observed by Reinganum (1983, p. 18), does not imply causation, or for that matter, even correlation as long as the overall size effect has been left uncontrolled.

3. New Issues, Market Reactions and the January Effect

In the late sixties and early seventies a number of papers on stock price reactions to news concerning stock splits, stock dividends and equity rights issues were published.[3] The general indication was that prices reacted rather strongly to news concerning splits, stock dividends or issues, and that the initial reaction usually preceded the actual announcement by a period of two to eighteen months. Some of the papers also reported a minor reaction lag, i.e., residual price increases during a period of one or two months after announcement. The phenomenon (excepting the lag) was largely ascribed to signalling.

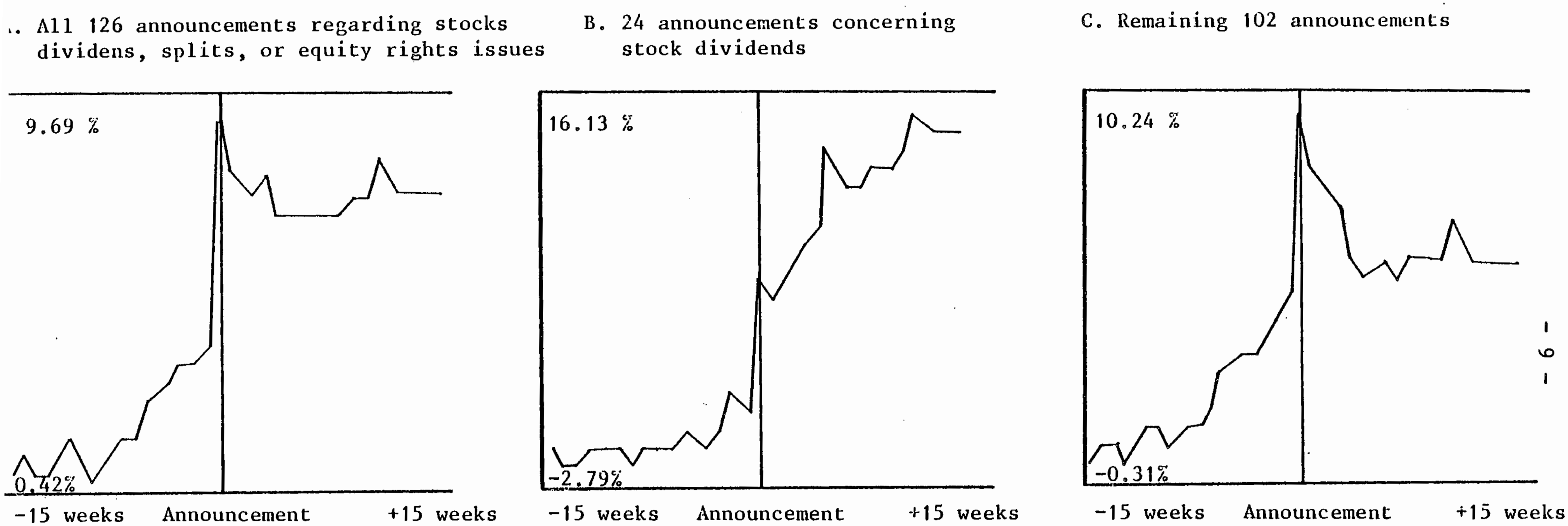
Announcements concerning stock dividends, splits or issues exhibit a distinct seasonal pattern on the HESE. Of the 126 announcements made in the period of September 1969 through December 1981, 39, or almost a third, occured in either March or September. The distribution of the announcements is given in figure 1 below.

Figure 1. The seasonal distribution of announcements concerning stock dividends, splits or equity rights issues on the HESE September 1969 through December 1981



Since announcements are unevenly distributed, the presence of a signalling effect with a sufficiently short lead time, i.e., a price increase preceding the announcement by one to four months, may account for a significant portion of, if not completely explain the seasonality in our data. To test this proposition, a modified Fama-Fisher-Jensen-Roll (1969) procedure was initially applied to weekly HESE data.[4] Residuals of the 'market' model were computed for each stock for a period of 15 weeks preceding and 15 weeks after each announcement.[5] Mean cumulative return residuals are graphed for all 126 announcements and separately for the 24 pure stock dividend and 102 remaining announcements contained in our 1969-1981 data in figure 2.

igure 2. Cumulative mean residuals from the market model for 126 announcements concerning stock dividends, splits and/or equity rights issues.



The cumulative residuals relating to all 126 announcements exhibit a fairly standard pattern of the type reported in several earlier studies. A strong positive trend in the residuals is detected from week 9 onwards up until the announcement week. This cumulative residual price increase amounts to nearly ten percent. A slight--probably insignificant--overshooting reported for the last week preceding the announcement is eliminated in 2-5 weeks time. This pattern can, however, be broken down into two somewhat different patterns, one related to the 24 stock dividend announcements, the other to the remaining 102 announcements. In the former case, no overshooting occurs. Instead, the cumulative mean residual, once the market has learned that a stock dividend has been announced, seems to take 3-5 weeks to adapt, levelling off at 15-16 percent. This outcome resembles the one reported by Marsh (1979) for the London exchange.

For the remaining 102 cases, a sharp drop of some 5 percentage units in cumulative residuals is reported for the weeks immediately following announcements. It is as if the market was able to foresee an upcoming announcement while still being unable to predict the contents of that announcement. Clearly then, the market must 'prefer' stock dividends to equity rights issues since a pure stock dividend announcement causes residual prices to increase by an average of 6-8 percent, whereas announcements containing equity rights issues tend to depress share prices from their pre-announcement high by some 5 percent. Any explanation for such market behavior would, it seems, have to be based on the different signalling properties of the two measures.[6]

For the purpose at hand, it is however sufficient to note that any such announcement, apparently because of its signal value, is associated with excess returns averaging roughly ten percent during the 5-8 weeks preceding it. Both the duration and the magnitude of these excess returns make them a likely candidate in explaining HESE seasonality. Since the modes of the announcement distribution occur in March and September, each preceded by heavy announcement activity in February and August, we would expect the stock market to yield excess returns in December-February and June-August, while otherwise performing less spectacularly. A glance at table 2 shows that, on an aggregate level at least, that might indeed be the case.

To test the proposition, we split our data into two sets, one consisting of return observations where an announcement is imminent or has recently been made, the other where it is not. If the proposition holds, we would expect the first set of observations (A) to have a significantly higher mean than the second set (B) and possibly contain a seasonal bias while we would expect the second (B) set to be virtually free of any seasonality.

Table 3 reports on regressions of returns in excess of sample means on a set of 12 monthly dummies with the intercept suppressed. The data consists of rates of returns to 49 stocks listed throughout the period 1970-81.[7] A total of 107 announcements concerning a split or an issue occured in these stocks. To avoid biasing the t-statistics in an upward direction by running regressions over the total 7056 observations which are highly unlikely to be independent of each other,[8] we construct an equally weighted index of all observations for a given month contained in each sample. This index is then regressed on the monthly dummies. In table 3, n denotes the total number of observations in each sample, whereas n' denotes the number of observations in the index.

The first row of table 3 lists the number of leads and lags (in months) around an announcement date that are contained in the A sample. The third row in turn lists the sample mean return which has been deducted from individual index return observations before running the regressions. As can be seen from that row, monthly mean returns increase rapidly with the approach of an announcement.

The results contained in table 3 appear to refute our proposition rather definitely, however. The actual outcome is pretty much the opposite of what was expected. Whereas sample A indeed, as was pointed out above, exhibits a much higher mean return, it contains no significant seasonality at leads shorter than 6 months; sample B, if anything, is more seasonally biased than was the total sample. While the January effect equalled 3.22 percent in the total sample, it equals 1.65-2.71 percent in sample A and 3.28-3.40 in sample B. Monthly coefficient signs are nevertheless more or less the same between the three samples.

The conclusion with respect to the announcement--seasonality proposition hence must obviously be negative. Since, however, stocks in which splits or issues occur tend to be 'winners', they are also unlikely to be good candidates for tax-loss selling. The fact that we are unable to register a January-effect for these stocks might then be taken as a highly tentative suggestion that that effect is produced by tax-loss selling.

4. Tax Loss Selling and the January Effect

For quite some time, capital gains taxation has been suggested to have undesirable effects on stock trading by giving investors an incentive to realize capital losses while deferring capital gains.[9] An investor subject to capital gains taxation, who during the tax year has realized some capital gains or earned other income from which capital losses may be deducted, will near the end of the tax year find it to his advantage to sell off his 'losers' while retaining his 'winners'. He thus creates a tax write-off which, in an

Table 3. Mean excess monthly returns (over sample averages) for stocks where announcements are imminent (A) and where they are not (B). t-statistics in parentheses.

Data set leads/lags months	Total market	A. Issue or split is imminent 6/1	3/1	1/1	-/-	B. No issue or split is imminent 6/1	3/1	1/1	-/1
n	7056	856	535	321	107	6200	6521	6735	6949
n'	144	134	124	110	64	144	144	144	144
mean	1.22	2.44	2.90	3.85	6.67	1.07	1.11	1.10	1.15
Jan	3.22	2.37	1.65	2.19	2.71	3.40	3.35	3.28	3.28
	(3.36)	(1.47)	(1.02)	(0.91)	(0.55)	(3.59)	(3.52)	(3.49)	(3.45)
Feb	1.46	0.97	2.10	2.98	2.93	1.54	1.39	1.41	1.41
	(1.52)	(0.63)	(1.30)	(1.31)	(1.03)	(1.62)	(1.46)	(1.49)	(1.48)
Mar	0.88	1.71	2.31	1.88	1.49	0.67	0.69	0.69	0.76
	(0.91)	(1.11)	(1.42)	(0.87)	(0.49)	(0.70)	(0.72)	(0.73)	(0.80)
Apr	0.13	2.03	0.64	0.10	-1.56	-0.11	0.04	0.03	0.12
	(0.14)	(1.26)	(0.37)	(0.04)	(0.48)	(0.11)	(0.05)	(0.03)	(0.12)
May	-2.03	-1.61	-1.70	-2.55	-2.59	-2.04	-2.08	-2.04	-2.04
	(2.12)	(0.95)	(1.00)	(1.06)	(0.61)	(2.16)	(2.19)	(2.17)	(2.15)
Jun	0.21	1.21	0.78	2.08	2.18	0.13	0.19	0.20	0.19
	(0.22)	(0.72)	(0.46)	(0.76)	(0.57)	(0.13)	(0.20)	(0.21)	(0.20)
Jul	0.94	1.16	0.72	-0.59	-4.34	1.04	1.04	1.07	1.02
	(0.98)	(0.69)	(0.40)	(0.24)	(0.72)	(1.10)	(1.09)	(1.14)	(1.08)
Aug	-0.25	-1.54	-1.24	0.78	-1.04	-0.23	-0.23	-0.29	-0.28
	(0.27)	(0.95)	(0.73)	(0.33)	(0.32)	(0.24)	(0.25)	(0.30)	(0.29)
Sep	-1.63	-1.76	-1.70	-2.38	-4.04	-1.60	-1.59	-1.59	-1.58
	(1.70)	(1.09)	(1.00)	(0.99)	(1.25)	(1.69)	(1.67)	(1.69)	(1.67)
Oct	-1.77	-3.14	-0.86	-2.38	-1.37	-1.73	-1.84	-1.78	-1.81
	(1.85)	(2.04)	(0.51)	(1.05)	(0.39)	(1.83)	(1.93)	(1.89)	(1.90)
Nov	-1.53	-0.38	-1.70	-2.04	2.79	-1.49	-1.41	-1.43	-1.53
	(1.60)	(0.25)	(1.05)	(0.85)	(0.80)	(1.58)	(1.48)	(1.52)	(1.62)
Dec	0.39	-0.75	-1.27	-0.12	-	0.45	0.48	0.45	0.47
	(0.40)	(0.49)	(0.78)	(0.05)	-	(0.47)	(0.51)	(0.47)	(0.50)

efficient market where expected certainty equivalent returns are the same for all stocks, will cost him only the brokerage fee plus any transaction tax. In an efficient market where demand for securities is perfectly elastic at the certainty equivalent market return, such tax-loss selling should, however, have only insignificant price effects due to temporary reshuffling of portfolios away from the perfectly diversified ones.

This prediction seems likely to be the true one. Even though the above argument, interpreted in terms of reservation prices, means that sellers would generally be willing to trade at prices significantly below prevailing equilibrium prices, the number of buyers with reservation prices equalling the equilibrium ones is likely to be sufficiently large to prevent any price decrease. The tax exempt investors--pension funds, charities and foundations --are by themselves a significant group likely to bid up prices to their equilibrium level. However, arbitrage operations are by no means restricted to this group since tax write-offs can be attained quite irrespectively of the tax status of, and at no cost to, the buying party. In the simplest of cases, two investors could gain the full tax credit by exchanging their 'losers' at prevailing equilibrium prices, while holding on to their 'winners'.[10]

Market power, interpreted as the power to extract some of the outside money available to the holder of a 'loser' by sheer control of demand, therefore seems an unlikely explanation for a tax effect on share prices around the turn of the tax-year. The group of buyers with reservation prices equal to or, if one allows for some diversity in the composition of individual portfolios due to, e.g., transaction cost (Mayshar 1981), approaching long run equilibrium ones, will mostly be large enough not to be able to collude efficiently.

Any tax loss selling argument concerning market price behavior towards the end of the tax year would therefore, it seems, have to rely on some inefficiency or a temporary disequilibrium in the market for 'losers' such that market clearing prices would actually fall below long run equilibrium ones while still exceeding the sellers' reservation prices. This inefficiency or disequilibrium is currently referred to as a selling pressure on the market. In the absence of trading friction such pressure could, given the fact that we purport to be dealing with a recurring phenomenon, be sustained only by highly implausible assumptions regarding the formation of price expectations.

Transaction costs might, however, offer the rudiments of an explanation. If the market is in equilibrium prior to the tax loss selling rally, and if the extent of, and the actual stocks that will be involved in the rally are hard to predict, the sellers might have to offer their 'losers' at a discount equalling at most twice the unit transaction cost. Just as in any transaction cost argument, this one neglects the investors already decided on a purchase that hence treat the transaction cost as sunk. The number of such investors may, however, be sufficiently small as compared to the number of tax loss sellers to permit a minor market price decrease.[11]

Empirical evidence, however, so far seems to point in the direction of tax loss selling as being responsible for at least some of the trading activity and price changes at the turn of the tax year. Quite consistently with a strict efficiency hypothesis, Dyl (1977) found that the tax-induced spread in reservation prices produces an abnormally low trading volume in 'winners' and a likewise abnormally high volume in 'losers' at year's end. The increase in trading volume for the 'losers' apparently has been induced by lower offer prices since Branch (1977) is able to report supernormal returns to a simple filter strategy designed to identify stocks subject to tax induced selling. Reinganum (1983) in turn reports significantly higher early January returns for 'losers' than for 'winners', although his conclusion is that tax loss selling cannot account for the entire January effect.

In what follows, a direct test of a tax loss selling hypothesis on the January effect is presented. Any such hypothesis is likely to contain institutional features. The essential institutional feature to be included in a test on Finnish data is that capital losses, according to local legislation, can be deducted only from capital gains. Hence, demand for tax write-offs will vary greatly, not with personal income, but with (realized) capital gains. The probability that a share of a given stock will be offered for sale towards the end of a tax year should thus increase with some aggregate measure of taxable capital gains realized during the tax year and decrease with its own past gains, i.e., the potential for write-offs offered by that stock.

At the onset of the analysis, let the capital market be in a CAPM-equilibrium:

$$(1) \qquad E(R_{i1t}) = R_F + (R_M - R_F)\beta_i$$

where R_{it1} denotes the return to asset i in the first month of year t, R_F and R_M denote the monthly riskfree and market returns respectively and β_i is the beta coefficient of asset i. For any asset <u>which can be the object of tax induced selling</u> in year t-1, the tax loss selling-hypothesis predicts a return premium at the onset of a new tax year. That premium is now assumed proportional to the probable monetary gain from a tax write-off proxied by the market return for the tax year t-1 (the total amount of capital gains realized in that tax year is assumed increasing with the market return) and inversely proportional to the book loss incurred by realizing a share of i. Obviously the latter measure must individually depend on the period during which that asset has been held, and aggregately on the distribution of holding periods over the investors. Since only return data was available to us, the most convenient proxy[12] for the tax-loss potential of stock i was its return in the preceding tax year t-1. Note that since, because of tax considerations, dividend yields to Finnish common stocks are generally very low (1-4% annually), the difference between return and price data is fairly small.

The expected rate of return to asset i in January of year t should thus be approximated by

$$(2) \qquad E(R_{it1}) = R_F + (R_{M1t}-R_F)\,\beta_i + \alpha_0 + D_i^I\,(\alpha_1 R_{Mt-1} - \alpha_2 R_{it-1}),$$

where D_i^I is a dummy variable receiving the value 1 when stock i has tax loss potential and the value zero otherwise, R_{Mt-1} is market return and R_{it-1} is the return to asset i in the previous tax year. We will not introduce an explicit definition of tax loss potential but instead run regressions of (2) for different values of I where I is defined as the cut-off point on R_{it-1} such that D_i^I receives the value 1 if $R_{it-1} < I$ and zero otherwise. Since we are not able to produce an equilibrium theory that would explicitly encompass the January effect, the introduction of a tax-loss selling effect into a CAPM-framework may seem doubtful. Consequently, the α-coefficients of equation (2) will also be estimated separately from the CAPM-induced coefficients.

Initial results are reported in table 4. Notice that in estimating (2) we have treated R_F as a constant to be estimated. The beta-estimates used are Dimson-betas (Dimson 1979) for the period 1970-81 with five leads and ten lags using the WI-index (Berglund-Wahlroos-Grandell 1983) as market portfolio. The expected value of R_{i1t} is proxied by its actual value. Since the regression is pooled, the number of observations equals the number of stocks (50) times the number of Januaries less one (11), i.e., 550.

The first three columns of table 4 give the I-value, the total number of observations and the number of observations for which $D_i^I = 1$ respectively. The columns under A give coefficient estimates for equation (2) based on the CAPM[13] whereas the columns under B give the coefficient estimates of

$$(3) \qquad R_{i1t} = \alpha_0 + D_i^I(\alpha_1 R_{Mt-1} - \alpha_2 R_{it-1}).$$

The initial impression is one of rather strong support for the tax-loss selling hypothesis. January returns are positively correlated with market returns over the previous tax year and (for $I < -.10$) negatively correlated with the stock's past individual performance. α_1 and α_2 estimates are more or less the same in cases A and B.

A negative sign on α_2 was not to be expected for low values of I because of transaction costs. An investor with a marginal tax rate of 50 percent will not benefit from realizing losses of less than twice the transaction costs associated with going out of and into a long position, i.e., 6 percent. As I decreases, the explanatory power of R_{it-1} increases sharply while the significance of past market returns drops. Notice, however, that the magnitude of the α_1-coefficient tends to increase with the tightness of the selection criterion I; its loss of significance is due to a heavy increase in coefficient variance. It may furthermore be interesting to note that γ_0 associated with the risk measure $(R_{M1t}\beta_i)$ also decreases with the absolute value of I. Hence, the CAPM tends to fit the data worse the fewer the observations that are assumed affected by a January premium.

The magnitude of the January effect as reflected in α_1 and α_2 is far from

Table 4. Coefficient estimates of equation (2) on the tax-loss selling effect in January prices. t-statistics in parentheses.

I	n	n'	A Intercept α_0+R_F	$\beta_i - R_F$	$R_{M1t}\beta_i$ γ_0	$D_i^I R_{Mt-1}$ α_1	$D_i^I R_{it-1}$ $-\alpha_2$	r^2	B Intercept α_0	$D_i^I R_{Mt-1}$ α_1	$D_i^I R_{it-1}$ $-\alpha_2$	r^2
∞	550	550	.024 (1.75)	-.015 (-.93)	.714 (2.55)	.135 (5.29)	.007 (.39)	.085	.028 (5.67)	.126 (4.98)	.006 (.31)	.074
.00	550	181	.041 (2.95)	-.005 (-.30)	.626 (2.16)	.181 (3.67)	.054 (1.33)	.039	.049 (10.20)	.162 (3.34)	.042 (1.05)	.030
-.10	550	112	.040 (2.86)	-.003 (-.18)	.582 (2.03)	.282 (4.24)	.011 (.25)	.044	.049 (10.35)	.263 (3.98)	.001 (.00)	.035
-.20	550	61	.041 (2.91)	-.003 (-.20)	.361 (1.26)	.210 (2.41)	-.004 (-.08)	.016	.046 (9.93)	.209 (2.41)	-.008 (-.18)	.013
-.30	550	33	.042 (2.95)	-.005 (-.31)	.314 (1.09)	.331 (2.05)	-.093 (-1.55)	.010	.044 (9.71)	.336 (2.08)	-.098 (-1.64)	.008
-.40	550	14	.043 (3.05)	-.007 (-.40)	.273 (.95)	.458 (1.68)	-.181 (-2.17)	.011	.043 (0.71)	.467 (1.72)	-.189 (-2.28)	.009
-.50	550	8	.042 (2.99)	-.006 (-.36)	.266 (.93)	.365 (1.19)	-.227 (-2.63)	.016	.043 (9.71)	.349 (1.25)	-.234 (-2.73)	.014

negligible. If, for instance, we are able to find a stock that has lost more than 40 percent during the previous year while the market as a whole has gone up by 10 percent, we would expect an excess January return on that stock of 11.8 percent or more! Likewise, a 30 percent loser would, under the same conditions, be expected to yield an excess return of 7.4 percent in the next January.

The model fit is far from perfect, however; r-squares average less than two percent. This is likely to be due to our rather strict operationalization of a 'loser'. The significance of α_2 increased strongly with the absolute value of I, but at the same time it seems evident that some of the 'minor losers', e.g., those that have lost less than 30 or 40 percent in the previous year, must provide some of the investors with tax loss selling opportunities. If that is so, we would--under the highly restrictive demand assumptions discussed earlier--expect the market return index of the previous tax year to explain something of the January return in virtually all stocks--not just the stocks we have singled out as primary candidates for tax induced selling. Hence, a better model formulation might be one where R_{Mt-1} is not multiplied by the dummy D_i^I. Put more explicitly, this latter model formulation presupposes that there are tax-loss selling opportunities in (virtually) all stocks for some of the investors and that consequently, there must exist a positive relationship between the need for tax write-offs as measured by our proxy for capital gains R_{Mt-1} and (virtually) all January returns (since tax-loss selling is by some mechanism assumed to depress the market) while the variable $D_i^I R_{it-1}$ serves to single out those stocks where the selling pressure is likely to be the heaviest.

The results contained in table 5 come out exactly as expected. α_1, the coefficient associated with the previous year's market return is highly

Table 5. Coefficient estimates of a modified equation (2) where the previous year's market return has not been multiplied by the dummy variable D. t-statistics in parentheses.

I	n	n'	A Intercept α_0+R_F	$\beta_i - R_F$	$R_{M1t}\beta_i$ γ_0	R_{Mt-1} α_1	$D_i^I R_{it-1}$ $-\alpha_2$	r^2	Intercept α_0	R_{Mt-1} α_1	$D_i^I R_{it1}$ $-\alpha_2$	r^2
∞	550	550	.024 (1.75)	-.015 (-.93)	.135 (5.92)	.135 (5.29)	.007 (.39)	.085	.028 (5.67)	.126 (.31)	.006 (.31)	.074
.00	550	181	.023 (1.67)	-.016 (-.98)	.695 (2.47)	.148 (6.47)	-.027 (-.64)	.085	.025 (3.99)	.141 (6.19)	-.037 (-.89)	.075
-.10	550	112	.023 (1.68)	-.016 (-.98)	.693 (2.46)	.148 (6.61)	-.029 (-.70)	.085	.025 (4.23)	.140 (6.32)	-.038 (-.94)	.075
-.20	550	61	.023 (1.67)	-.016 (-.99)	.704 (2.52)	.148 (6.95)	-.042 (-1.06)	.086	.025 (4.67)	.139 (6.59)	-.045 (-1.12)	.076
-.30	550	33	.023 (1.65)	-.017 (-1.05)	.698 (2.50)	.153 (7.32)	-.090 (-2.12)	.092	.024 (4.66)	.143 (6.96)	-.093 (-2.17)	.082
-.40	550	14	.022 (1.63)	-.014 (-.90)	.655 (2.34)	.150 (7.37)	-.139 (-2.64)	.096	.025 (4.98)	.141 (7.04)	-.148 (-2.82)	.087
-.50	550	8	.022 (1.64)	-.014 (-.91)	.666 (2.40)	.150 (7.43)	-.216 (-3.47)	.104	.025 (5.10)	.141 (7.09)	-.223 (-3.57)	.095

significant and constant in magnitude throughout the runs. An average of 14-15 percent of an annual (tax year) return is projected as an excess return into the next January. The magnitude and statistical significance of α_2 in turn increases rapidly with the absolute value of I. α_2-estimates for higher I-levels are pretty much the same as in the runs reported in the previous table; for I = -30% we receive estimates ranging from -.090 to -.098 while for I = -40% the range is -.139 to -.189. Notice however that the α_2-coefficient becomes statistically significant (5%) at the -30%-level in table 5 as opposed to the -40% level in the previous table. Note also that in table 5, $(R_{M1t}\beta_i)$ receives a rather stable and significant coefficient value (although one lower than the predicted $\gamma_0 = 1$).

Without any doubt, one can conclude that the January-effect is linked to market behavior closely resembling tax loss selling. The better the previous (tax) year's market--and hence presumably, the greater the need for tax write-offs to reduce capital gains taxes--the better the market performs in January. Returns are especially high on those stocks that have lost money in the previous year. The 'January loser premium' averages 9 percent while it approaches 15 percent for 40% plus losers. These premia are highly significant. All this evidence rather strongly conflicts with the notion of an efficient market since absent transaction costs and property taxes, the excess returns could be arbitraged away by any investor.

5. Conclusion

Several recent papers have produced evidence of supernormal returns to stock markets in January. The object of the present paper was two-fold: firstly to investigate whether a fresh data set, the 1970-81 rates of return on the Helsinki Stock Exchange in Finland contained seasonality, and secondly to try to 'explain' that seasonality by known properties of stock prices.

We started off by screening our data for seasonal effects. Rates of return were computed for every month in our 12-year data set for ten stock portfolios composed in decreasing order of stock capitalization. We found an excess return on the market portfolio equalling two and a half percent in January--by far the largest monthly premium. Although our data reveal a strong overall size effect--the portfolio with the lowest mean capitalization earned a risk-adjusted return exceeding the return on the highest mean capitalization portfolio by almost ten percent annually--the link between the January and size effect discussed by Roll (1982) and Reinganum (1983) is far less evident in our data than in theirs. Still, a significant correlation between January excess returns and portfolio mean capitalizations could be detected. This correlation was, however, no stronger--actually it was weaker --than the correlation found in annual returns, and may thus be ascribed to the very same cause as the annual size effect.

To be more precise, if it is that the size effect reflects high investor risk aversion combined with return distributions of varying skewness and kurtosis as argued in Berglund-Wahlroos (1983), we would expect to find a size premium in the month of January just as in any other month. Moreover, if the January effect is caused by the tax induced selling of losers, we should, based on the very same argument, expect a higher than proportional share of the effect to materialize in small stocks since these, subject as they are to infrequent trading, tend to have more leptocurtic return distributions than larger stocks. The stocks with highly leptocurtic (fat-tailed) return distributions would be more likely to be losers even in years of high overall market return and would hence offer the best (ex ante) opportunities for tax loss selling.

Next, monthly portfolio returns in excess of monthly equally weighted market averages were regressed on a set of monthly dummy variables. Dummy coefficients turned out to pass a 5 percent two-tailed significance test in January, and curiously enough, in May. The latter coefficient was significantly negative, however. The January premium thus computed was positive for all portfolios, varying between 1.6 and 6.5 percent while averaging 3.2 percent.

The first proposition tested was one explaining the market seasonality uncovered by the seasonal distribution of stock dividends, splits and equity rights issues on the HESE. Since these, apparently because of signalling, are associated with supernormal returns averaging ten percent over roughly a four-month period, they may well account for some of the market seasonality. To test this proposition, the observation matrix was split into two sets, one containing observations where a split or an issue was imminent, the other where it was not. It turned out that the data where an issue was imminent contained no significant January effect, whereas in the remaining data that effect, if anything, was reinforced by the omission. Splits and issues as such hence cannot explain the January effect. Since, however, splits and issues predominantly occur in 'winner' stocks, the results may be taken to suggest that the January effect is somehow linked to 'losers'.

The one proposition to contain such a feature is the tax loss selling hypothesis contained in work by Branch (1977), Gultekin-Gultekin (1982), Reinganum (1983) and others. The problem with that proposition is that, in order to work, it requires an assumption of market failure manifesting itself in a less than infinitely elastic demand for individual securities, or the presence of transaction costs of quite significant magnitude. Although any investor who has realized capital gains during the tax year can benefit from selling his 'losers', these should entail no discount since he could easily, via a financial intermediary, purchase them back himself at his previous selling price and still earn the full tax write-off.

Irrespectively of these (severe) theoretical misgivings about a tax-loss selling related explanation to the January effect, we devised a testable version of the proposition. It turned out to fit the data just as the tax loss proposition predicted. January returns are significantly positively related to market returns for the previous year--a proxy for capital gains and the consequent need for tax write-offs--and negatively related to the stock's own previous performance, i.e., the potential for tax write-offs offered by the stock.

This outcome, far from settling any January controversy, seems to present us with only a slightly smaller number of questions than did the original observation of seasonality. How is it that tax induced selling of securities affects the market valuation for those securities? We can only suggest that transaction costs, by themselves or in combination with property taxation, might play a role in the final argument. As was mentioned earlier, transaction costs may shield off some of the arbitrage created by the tax-loss sellers. Finnish brokerage fees and transaction taxes are admittedly rather high (1.5 percent on each purchase or scale) but--so it seems--not nearly high enough to explain the January effect in losers, often amounting to more than ten percent.

Property taxes do, however, constitute an additional transaction cost in cases where investors are required to maintain long positions in taxable assets over the turn of the tax year. Since property tax exempt assets like cash or savings deposits are readily available, an investor would have to consider property taxes a transaction cost item when speculating against tax-loss selling. Property taxes are not very high however, and what's more, a great number of potential investors are actually tax exempt due to their low net worth. One would therefore assume transaction cost effects arbitraged away by investors already decided on their future portfolios who adjust the timing of their purchases to profit from any available arbitrage, while property tax effects would seem likely to be eliminated by the tax exempt investors. Neither trading cost can thus, it seems, without additional assumptions account for the correlations reported in this paper.

We are hence left only with a number of ad hoc explanations, based either on the alleged presence of intangible and preferably fixed transaction costs arising from search or data processing, or on irrational investor behavior by itself or in combination with market inefficiencies. Any assumption of conspiracy, holding that brokers by buying and selling at depressed prices in their own or some insider investors account to create tax losses and decrease property taxation values, seems refutable by a reference to the fact that the equally weighted return index recorded a decrease in precisely six out of the twelve last December trading days included in our data set. The mean return for the last trading day of the year over the 12-year period was slightly positive.

In closing, it may nevertheless be worth noting that (implicit) fixed transaction costs would probably not have to be very high to leave us with the arbitrage we now seem to be observing. Since January premiums typically occur in the smaller stocks that are also typically thinly traded, the volume of the purchase needed to push up the price to its long run equilibrium level need not be very great.

FOOTNOTES

1. The use of fixed portfolios is elaborated upon in Berglund-Wahlroos (1983). The largest of the ten portfolios had an average capitalization of 203 million FIM in 1970 while the average capitalization of the smallest portfolio was only 1.8 million FIM. Notice also that the sampling procedure is virtually free from any survivorship bias. During the period under study, only three firms left the exchange, none because of bankruptcy.
2. The riskfree rate was set at the mean bank lending rate and the market portfolio return at the mean return to the value-weighted WI-index (Berglund-Wahlroos-Grandell 1983). Dimson betas were estimated from daily return data with five leads and twenty lags.
3. See, e.g., Fama-Fisher-Jensen-Roll (1969), Scholes (1972), and--for an analysis of Finnish data--Korhonen (1975).
4. This research is extensively reported on in Grandell (1983).
5. The market portfolio in this work was the value-weighted WI-index of Berglund-Wahlroos-Grandell (1983). Betas were estimated from weekly data and are unadjusted for any thin trading bias.
6. There is one such explanation. By Finnish law, new shares issued by a firm cannot be offered at a price lower than their nominal value. Hence, a stock dividend, effectively splitting up the firm's assets on a greater number of shares of the same nominal value as before, constitutes more of a 'bet' on a positive return development in the future than does a rights issue. As management recommends pulling down the market value of the share closer to its book value it increases the risk of being unable to issue new shares in the future.

7. Of the total of 50 securities, one, Kemi Oy, which was partly nationalized following a near-bankruptcy in 1978 is omitted since the three Kemi rights issues of 1978-79 were primarily an instrument to provide it with public funds.

8. Since betas are generally positive, it is obvious that running a pooled regression would greatly bias the t-statistics in an upward direction. The pooled regression would closely resemble expanding one's data set by replication of observations.

9. See, e.g., Haugen and Wichern (1973).

10. Were it not for its potential illegitimacy, an even simpler scheme could be suggested: An investor could gain the full write-off by selling his losers to himself. An obviously legal alternative strategy (although one subject to transaction costs if exercised on the exchange floor) would be to sell one's losers and purchase the closest substitutes available.

11. Raw empirical data suggests that there must be more to the January effect than this. The January excess return on the equally weighted portfolio of all stocks in table 2 already exceeds total transaction costs associated with going in to and out of a given stock (3%) by 0.23 percent. The mean January excess return for 'losers'--if they are indeed responsible for the effect--must be significantly higher than 3.2%.

12. A perfect measure of the tax loss potential of a share seems almost impossible to devise. Branch (1977) defined a 'loser' as a stock reaching its yearly low in the last whole trading week of the year, while Reinganum (1982b) used the ratio of the stock's price on the second to the last trading day of the year to its high in the preceding six months.

13. Note that we do not restrict the value of the coefficient γ_0 associated with $(R_{M1t}\beta_i)$ to unity.

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