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Can Tax-Loss Selling Explain the January Effect? A Note

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THE JANUARY EFFECT REMAINS an unexplained phenomenon in financial economics. That the month of January has consistently shown higher mean stock returns than other months has been documented by several researchers, beginning with Rozeff and Kinney [19]. Additional research by Keim [14] showed that approximately half the annual "size effect" is attributable to abnormal returns in January.[1] Roll [18] confirmed Keim's results, as has subsequent research such as Lakonishok and Smidt [15].[2]

The most often-advanced explanation for the January effect is tax-motivated transactions.[3] First proposed in 1942 by Wachtel [21], the tax-loss-selling hypothesis is frequently advocated in the popular press as a major influence on stock price behavior. Essentially, this hypothesis states that investors seek to reduce their taxes by realizing losses at year end, thereby depressing stock prices. Stocks return to equilibrium levels after year end, providing abnormally high returns in January.

Despite the popularity of the tax-loss-selling hypothesis, it is difficult to justify upon serious reflection. Justification requires acceptance of the proposition that speculators are prevented from purchasing tax-loss candidates before year end with the intention to sell them after year end. No theoretical justification has been presented for the hypothesis. Instead, using both theoretical and empirical approaches, Constantinides [5] argues that such action cannot explain the January effect if the market is efficient.[4] His results imply that empirical studies supporting the tax-loss-selling hypothesis provide evidence of market inefficiency.

As for the empirical evidence, Branch [3], Dyl [8], Roll [18], Reinganum [17], and Givoly and Ovadia [10] have provided support for the tax-loss-selling hypothesis.[5] Recently, Schultz [20] tested the tax explanation by examining the

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[1] More than half the January premium occurred during the first week of trading in the year, with the first trading day being particularly important.

[2] Roll found that the largest returns occurred on five consecutive days—the last trading day of December and the first four trading days of December.

[3] See Arbel [1] for a discussion of alternative explanations of the January effect.

[4] Constantinides' argument is that it is optimal for investors to realize losses as they occur, not at year end. Only irrational investors would regularly prefer to wait until the end of December to establish losses for tax purposes.

[5] However, both Reinganum [17] and Givoly and Ovadia [10], as well as others such as Lakonishok and Smidt [15], question whether this can be the only explanation of the January effect; furthermore, empirical evidence from other countries is mixed. Berges et al. [2] report that Canadian stocks showed a January effect even before capital gains were taxed, but they note that this could reflect the tax status of U.S. owners of Canadian stocks. Their overall conclusion is that their tests failed to support

returns of small firms prior to the 1917 tax act. He found no evidence of abnormal returns in January prior to 1917 but did find a January effect for the years 1918 through 1929. Schultz, therefore, "cannot reject a tax explanation for the 'January effect.'"

This note re-examines the tax-loss-selling explanation for the January effect by analyzing the longest pretax period available. We reconstruct and analyze a high-quality source of stock price data extending back to 1871, thereby allowing us to examine the behavior of stock prices for a much earlier and longer period than Schultz examined. Furthermore, unlike previous studies, we employ a parametric approach that takes into account the statistical problems that can arise when using certain stock market data. Finally, to determine whether the January effect exists for large firms, we analyze the returns to the Dow Jones Index. However, on the basis of our analysis of the early Dow Jones Indexes, we argue that previous research has casually linked these data together without examining the series, thereby possibly biasing the results. Our approach avoids this potential problem.

Contrary to Schultz, our results indicate that the January effect existed long before income taxes had an effective impact and that no significant change occurred in the January effect after income taxes were imposed. Our failure to support the tax-loss-selling hypothesis provides support for the position that some nontax factor influences stock price behavior in January. Our results also confirm the previous studies that indicate that the January effect is a small-firm effect.[6]

I. Modeling the January Effect

We use two alternative models to test for the January effect. Following Keim [14], Model 1 allows each month to have a different mean return:[7]

$$R_t = \beta_0 + \sum \beta_i M_{it} + \varepsilon_t, \tag{1}$$

where

R_t = stock returns in month t and

$M_{2t \cdots 12t} = 1$ if month is February $\cdots$ December,

0 otherwise.

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the tax-loss-selling hypothesis. Gultekin and Gultekin [11] find that most countries' data exhibited stock return seasonality consistent with their respective tax year conventions, although Australia was an exception. Brown et al. [4] also find that returns in Australia did not fit the tax-selling pattern.

[6] Pettengill [16], using data similar to ours, also finds evidence of a January effect prior to income taxation. However, our paper differs from his in several ways. Our paper takes into account the serial correlation and heteroscedasticity that exist in these data, while Pettengill fails to do so. In addition, we test formally whether the January effect differed across tax regimes and whether all non-January months have equal mean returns. Finally, Pettengill splices different Dow Jones indexes together in the early period. As discussed below, our analysis strongly suggests that this is inappropriate.

[7] In effect, Model 1 corresponds to an analysis of variance. The β_i ($i > 1$) measure the difference between the January returns and the average returns for each of the other months. The January return is captured by the intercept β_0.

If no seasonal pattern exists, the hypothesis that the β_i ($i > 1$) are all zero should not be rejected. The January effect would be supported if the β_i are negative ($i > 1$), jointly nonzero, and equal to one another.

The second model is contained within the first. If the maintained hypothesis is that all β_i ($i > 1$) are equal, then the appropriate model is

$$R_t = \beta_0 + \beta_1 JAN_t + \varepsilon_t, \tag{2}$$

where $JAN_t = 1$ if the month is January, 0 otherwise. In this model, the test for a January effect is simply the test that β_1 is significantly greater than zero.

II. Data and Estimation Technique

We use as our primary data the reconstructed industrial stock price series of the Cowles Commission as reported in *Common Stock Indexes* [7]. In constructing the stock price indexes in the 1930s, the Cowles Commission chose the value-weighted form used by the forerunner of Standard & Poor's organization; this index should be equivalent to the current value-weighted S&P index. The Commission constructed several indexes for the period 1871 through 1938 for all stocks quoted on the New York Stock Exchange based on meticulous records for each company. The industrial index increased to more than three hundred stocks in the 1930s and is the broadest available historical measure of industrial prices.[8]

The error terms in the models described above are likely to be both heteroscedastic and serially correlated. The Cowles stock price data are averages of the high and low for the month, and the rate-of-return series depends on the first difference of these averages. Working [24] noted that this should introduce some serial correlation into the return series, a hypothesis confirmed by Cowles [6]. In addition to this problem, heteroscedasticity is likely to be present. Tests for the equality of the variances of monthly returns indicate that the hypothesis of homoscedasticity should be rejected.[9]

While ordinary least squares (OLS) yields unbiased parameter estimates when the error term is heteroscedastic, serially correlated, or both, the OLS estimate of the variance-covariance matrix of the parameter estimates is biased and inconsistent, and hypothesis tests are invalid. Since the exact forms of the heteroscedasticity and serial correlation are uncertain, we estimate the parameters of Models 1 and 2 using OLS but estimate the related variance-covariance

[8] We have analyzed the complete Cowles monthly data series on stock prices and identified an apparent error in the original series for June 1884. We have corrected this error and made other adjustments as needed; for example, the stock market was closed between July 31, 1914, and December 11, 1914. See Wilson and Jones [23] for a complete description of the Cowles data and the revisions made in the data. It should be noted that we use the Second Edition of *Common Stock Indexes*, which updated the original data from the end of 1937 through 1938. This edition contains substantially revised data for the pre-1918 period, thereby raising questions about those studies that rely only on the First Edition.

[9] The test, also used by Keim [14], is described in Johnston [13, pp. 298–99]. For the period February 1871 through December 1917, the relevant χ^2 statistic is 22.0 (11 d.f.), significant at the five percent level. For the period January 1918 through December 1938, the variances are even less similar, with the relevant χ^2 statistic equal to 1347.1 (11 d.f.). Like all the formal tests for heteroscedasticity, this test assumes that the returns are serially uncorrelated and thus could be misleading for our data.

matrix using a method outlined in Gallant [9]. This method ensures a consistent estimate of the variance-covariance matrix regardless of the forms of heteroscedasticity and autocorrelation and hence allows us to make valid inferences from our hypothesis tests.[10]

Schultz [20] compared the returns on "small" stocks with the Dow Jones Industrial Average. To determine whether seasonality exists in this market measure, we applied the same tests for the pre- and post-tax period. During the period 1900 through 1929, examined here and by Schultz, the Dow underwent several important changes. For most of the period 1900 through September 1916, the Dow consisted of twelve stocks (DJ12), including two preferred, one utility, and one mining. On October 4, 1916, the number of firms increased to twenty (DJ20), and, on October 1, 1928, the coverage was increased to thirty stocks, with the adjustment for stock splits changed from a numerator to a denominator adjustment. Therefore, three different measures are spliced together for the period analyzed by Schultz [20], with no way to determine how the series were spliced together or what the effects were of doing so.

We have examined the twenty-three-month overlap between the DJ12 and the DJ20, created when the latter was extended back to the opening of the market after World War I (December 12, 1914). The monthly change in prices for this period had a correlation coefficient of only .67. The DJ12 increased 108.1 percent, and the DJ20 increased 88.4 percent. Furthermore, although we are unable to measure the impact of the shift that was made in 1928 from numerator adjustments to denominator adjustments for stock splits, it is well known that the two procedures lead to quite different temporal movements in the index.

To avoid these problems, we use the DJ12 to compute returns for the pretax period; that is, we use these data for the period January 1900 through September 1916. We use the DJ20 to compute returns for the post-tax period of 1918 through 1929. This procedure provides a cleaner measure of the Dow Jones Industrial Average while corresponding very closely to the pre- and post-tax periods used by Schultz.

III. Results

The models were estimated for two time periods: February 1871 through December 1938, the entire period for the Cowles Index, and January 1900 through December 1929, the period analyzed by Schultz [20]. Since the issue being examined is whether the income tax produced the January effect, separate estimates are made for the pre- and post-tax years for both the entire period and the Schultz period. Following Schultz, we assume that the effective post-tax period begins with 1918.[11] Table I gives the mean returns and variances for each month for the pre- and post-tax periods.

[10] The Gallant [9, pp. 137–139] procedure is a generalization of White's [22] heteroscedasticity-consistent estimator in that it also takes into account the serial correlation of the errors in computing estimates of the elements of the variance-covariance matrix of the OLS estimated coefficients. Hansen [12] derives a similar estimator in the context of generalized method of moments estimators.

[11] See Schultz [20, pp. 334–335] for a history of federal income taxes. Although an income tax existed for the years 1914 through 1917, marginal rates were extremely low and presumably could have had only minor effects on tax trading.

Table I

Summary Statistics for Cowles Industrial Data Divided into Two Periods[a]

Month	Feb. 1871 through Dec. 1917		Jan. 1918 through Dec. 1938	
	Mean Return	Variance	Mean Return	Variance
January	1.443	8.982	2.605	8.074
February	.363	11.081	.894	19.957
March	.328	11.687	.326	15.101
April	1.039	13.288	−.326	50.139
May	−.547	21.552	.283	103.140
June	−.276	19.077	−.606	41.068
July	−.517	14.902	3.215	30.063
August	1.119	16.500	2.385	105.110
September	.351	17.133	1.050	25.358
October	−.014	23.812	−2.516	58.903
November	.540	22.894	−.133	65.488
December	−.349	18.042	−.612	31.601

[a] Rates of return are calculated as monthly percentage changes.

Table II reports the estimates of Model 1 along with the F-statistics for the joint hypothesis that all β_i are zero $(i > 1)$ for each of the four periods described earlier. Both the t-statistics from the OLS estimates and the asymptotic t-statistics from the consistently estimated variance-covariance matrix are shown in Table II. Corresponding sets of F-statistics are also shown for each period.

The estimate for the pretax period of February 1871 through December 1917 (column 1 of Table II) yields negative coefficients for all non-January months. The asymptotic t-statistics from the consistently estimated variance-covariance matrix indicate that several months had significantly lower returns than January. The F-statistic calculated from the OLS variance-covariance matrix would lead one not to reject the joint hypothesis of no seasonality. However, the OLS variance-covariance matrix for the estimated coefficients is biased, resulting in an invalid hypothesis test. The F-statistic calculated from the consistent estimate of the variance-covariance matrix indicates that the joint hypothesis that the β_i $(i > 0)$ are all zero should be rejected. Furthermore, the test of equality of the β_i $(i > 1)$ using the corrected procedure indicates that this hypothesis cannot be rejected.[12] Thus, this evidence is consistent with the existence of a January effect prior to the imposition of income taxes.

When the pre-income tax period is shortened to 1900 through 1917 (column 2 of Table II), the joint hypothesis of no seasonality cannot be rejected at the five percent level but can be rejected at the ten percent level. All but one of the estimated β_i are negative. The hypothesis that the β_i are equal $(i > 1)$ cannot be rejected.[13]

Columns 3 and 4 of Table II report the estimates for the post-income tax periods (i.e., after 1918). As before, when the consistent estimate of the variance-

[12] The computed F-statistic is 1.27 (10,546 d.f.).

[13] The F-statistic is 1.03 (10,199 d.f.).

Table II

Coefficient Estimates for Model 1 for Two Pretax and Two Post-tax Periods, with Dual t-Statistics[a]

	Pretax		Post-tax	
	Feb. 1871 through Dec. 1917	Jan. 1900 through Dec. 1917	Jan. 1918 through Dec. 1938	Jan. 1918 through Dec. 1929
Constant	1.443	1.619	2.605	3.010
	(2.404)	(1.450)	(1.757)	(2.255)
	[2.182]	[2.183]	[4.305]	[3.378]
February	−1.079	−2.581	−1.711	−2.840
	(−1.279)	(−1.642)	(−0.816)	(−1.505)
	[−1.920]	[−2.821]	[−1.631]	[−2.298]
March	−1.115	−1.061	−2.279	−2.285
	(−1.321)	(−0.675)	(−1.087)	(−1.211)
	[−1.708]	[−0.841]	[−2.188]	[−1.521]
April	−0.404	−0.041	−2.931	−2.357
	(−0.478)	(−0.026)	(−1.398)	(−1.249)
	[−0.590]	[−0.031]	[−1.804]	[−1.678]
May	−1.990	−3.160	−2.322	−1.805
	(−2.357)	(−2.010)	(−1.108)	(−0.956)
	[−2.487]	[−2.649]	[−1.034]	[−1.179]
June	−1.718	−1.731	−3.211	−3.793
	(−2.036)	(−1.100)	(−1.531)	(−2.009)
	[−2.240]	[−1.284]	[−2.151]	[−2.502]
July	−1.960	−1.919	0.610	−1.421
	(−2.372)	(−1.221)	(0.291)	(−0.753)
	[−2.764]	[−1.516]	[0.464]	[−1.130]
August	−0.323	0.197	−0.220	−2.448
	(−0.381)	(0.124)	(−0.105)	(−1.297)
	[−0.440]	[0.137]	[−0.095]	[−1.686]
September	−1.092	−1.840	−1.555	−0.603
	(−1.286)	(−1.153)	(−0.741)	(−0.319)
	[−1.465]	[−1.383]	[−1.263]	[−0.545]
October	−1.457	−1.798	−5.121	−2.683
	(−1.716)	(−1.127)	(−2.442)	(−1.421)
	[−1.734]	[−1.103]	[−2.938]	[−1.718]
November	−0.903	−0.409	−2.738	−3.419
	(−1.064)	(−0.256)	(−1.306)	(−1.811)
	[−1.061]	[−0.254]	[−1.505]	[−1.274]
December	−1.792	−1.458	−3.217	−2.038
	(−2.112)	(−0.914)	(−1.534)	(−1.080)
	[−2.732]	[−1.287]	[−2.547]	[−1.362]
R^2	.024	.048	.051	.052
DW	1.352	1.325	1.327	1.403
$FOLS$[b]	1.230	0.910	1.180	0.660
	(11,546 d.f.)	(11,199 d.f.)	(11,240 d.f.)	(11,132 d.f.)
$FOLSG$[c]	2.090*	1.630	2.050*	1.400
	(11,546 d.f.)	(11,199 d.f.)	(11,240 d.f.)	(11,132 d.f.)

[a] Model: $R_t = \beta_0 + \beta_2 \text{ February}_t + \beta_3 \text{ March}_t + \cdots + \beta_{12} \text{ December}_t + \varepsilon_t$.
R_t = monthly rate of return.
February$_t$, March$_t$, etc. = monthly dummy variables.

Figures in parentheses are OLS t-statistics, and figures in brackets are asymptotic t-statistics from Gallant's estimator of the variance-covariance matrix.

[b] $FOLS$ = F-statistic from OLS estimates for the joint hypothesis that all $\beta_i = 0$ $(i > 1)$.

[c] $FOLSG$ = F-statistic from Gallant's estimator for the joint hypothesis that all $\beta_i = 0$ $(i > 1)$.

* F-statistic is significant at five percent level.

Table III

Coefficient Estimates for Model 2 for Two Pretax and Two Post-tax Periods[a]

	Pretax		Post-tax	
	Feb. 1871 through Dec. 1917	Jan. 1900 through Dec. 1917	Jan. 1918 through Dec. 1938	Jan. 1918 through Dec. 1929
Constant	.184	.173	.360	.674
	(1.021)	(.510)	(.804)	(1.709)**
	[.830]	[.433]	[.668]	[1.475]
January	1.259	1.446	2.245	2.336
	(2.011)**	(1.248)	(1.447)	(1.710)**
	[2.677]*	[1.787]**	[2.848]*	[2.354]*

[a] Model: $R_t = \beta_0 + \beta_1 \text{ January}_t + \varepsilon_t$.

Figures in parentheses are t-statistics from OLS, and figures in brackets are asymptotic t-statistics from Gallant's estimator of the variance-covariance matrix.

* Significant at one percent level for one-tailed test.

** Significant at five percent level for one-tailed test.

covariance matrix is used for the longer period, the hypothesis of no seasonality can be rejected at the five percent level. For the shorter period examined by Schultz, the hypothesis of no seasonality cannot be rejected at even the ten percent level. The individual β_i are all negative, with one exception. As in the pretax period, the hypothesis that the β_i are equal ($i > 1$) cannot be rejected.[14]

To test the hypothesis that the models have identical coefficients in the pretax and post-tax periods, a version of Model 1 was estimated for the entire period 1871 through 1938 using a dummy variable and interaction terms to allow each month to have a different mean return across tax regimes. Using the Gallant procedure to estimate the variance-covariance matrix, the joint hypothesis of identical coefficients in the two periods cannot be rejected.[15] Thus, on the basis of these results, there is evidence of a January effect that was similar in magnitude before and after the effective imposition of taxes.

Since the hypothesis of equal β_i cannot be rejected, this hypothesis is imposed to arrive at Model 2. While Model 1 allows for any departure from equal monthly mean returns, Model 2 is a direct test of the January effect since it allows only for January's mean return to differ. Table III reports estimates of Model 2 for the pre- and post-tax periods; as these estimates indicate, there is evidence of a statistically significant January effect in both the pre- and post-tax periods. As with Model 1, the results are stronger when the longer periods are used. Although the size of the January effect is larger after the imposition of the income tax, the hypothesis that the effects are equal across tax regimes cannot be rejected.[16]

[14] The relevant F-statistics are 1.33 (10,240 d.f.) for January 1918 through December 1938 and 1.05 (10,132 d.f.) for January 1918 through December 1929.

[15] The relevant F-statistics are 1.41 (12,786 d.f.) for February 1871 through December 1938 and 1.31 (12,331 d.f.) for January 1900 through December 1929.

[16] These tests also employ the corrected variance-covariance estimator. The relevant F-statistics are 1.23 (2,802 d.f.) for February 1871 through December 1938 and 1.01 (2,351 d.f.) for January 1900 through December 1929.

Table IV

Two Tests for January Effect Using the Dow Jones Industrial Average for the Pretax and Post-tax Periods

	Pretax (Jan. 1900 through Sept. 1916)	Post-tax (Jan. 1918 through Dec. 1929)
Tests that all months have the same mean return		
OLS	$F = .857$ (11,184 d.f.)[a]	$F = .644$ (11,132 d.f.)
OLSH[b]	$F = 1.315$	$F = 1.047$
Tests that January has higher mean return		
OLS	$t = -.688$[c]	$t = .347$
OLSH[b]	$t = -.849$	$t = .525$

[a] F-statistics are for the joint hypothesis that $\beta_i = 0$ $(i = 2,12)$ for Model 1: $R_t = \beta_0 + \beta_2 \text{ February}_t + \beta_3 \text{ March}_t + \cdots + \beta_{12} \text{ December}_t + \varepsilon_t$.

[b] OLSH indicates that variance-covariance matrix of estimated parameters was estimated using White's [22] procedure.

[c] t-Statistics are for the hypothesis that $\beta_1 = 0$ for Model 2: $R_t = \beta_0 + \beta_1 \text{ January}_t + \varepsilon_t$.

These results suggest that, when consistent estimates of the variance-covariance matrix are employed, there is strong evidence of a January effect in both the pre- and post-tax periods. To examine whether the January effect is absent from the returns on large firms, Models 1 and 2 also were estimated using the Dow Jones Industrial Average to compute returns. As discussed above, we feel it is inappropriate to link the alternative Dow Jones averages. Therefore, we use the DJ12 to compute returns for the pretax period and the DJ20 to compute returns for the post-tax period. Because the Dow Jones price measures employ end-of-the-month price data, we do not introduce serial correlation in computing the rate-of-return series. Therefore, the models can be estimated by OLS, and White's [22] procedure can be used to obtain a consistent estimate of the variance-covariance matrix in the event that heteroscedasticity is present.

Table IV presents the results for the Dow Jones data. No January effect is present in the returns from the DJ12 or DJ20 averages. In the pretax period, January has a smaller mean return than other months, but the difference is not statistically significant. For both models, the hypothesis that the coefficients are the same in the pre- and post-tax periods cannot be rejected.[17] The alternative estimation procedure produces very similar results indicating that the error terms display little heteroscedasticity.

IV. Conclusions

The tax-loss-selling rationale for the January effect in stock returns implies that the effect should be absent in the years prior to the effective imposition of income taxes. We test for the presence of this effect over roughly fifty years preceding the effective imposition of taxes and find that a January effect existed prior to income taxation; furthermore, there was not a statistically significant change in the effect after taxation was introduced. Our results are consistent with other studies that find the January effect closely related to the small-firm effect.

[17] The F-statistic is 1.41 (10,184 d.f.). The White [22] heteroscedasticity-consistent estimate of the variance-covariance matrix was used to compute the F-statistic.

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