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New Evidence on The January Effect Before Personal Income Taxes

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ABSTRACT

We examine the returns of stocks in Cowles Industrial Index before and after the introduction of personal income taxes in 1917. This is distinct from earlier studies because we cross-sectionally analyze the relationship between the returns of the individual stocks and measures of tax-loss selling potential and size. We find that excess returns at the turn-of-the-year and for the month of January were not significant until after 1917. These results provide strong support for the tax-loss selling hypothesis as an explanation for the January seasonal in the returns of small firms.

DESPITE YEARS OF CONCERTED research, the January effect, first detected by Rozeff and Kinney (1976), remains an unexplained anomaly of the finance literature. Keim (1983) reports that risk-adjusted January returns are especially pronounced for small firms, and Roll (1983) shows that fifty percent of the excess return occurs during the first four trading days of January and the last trading day in December (or the turn-of-the-year). Recently, Chan and Chen (1988) and Handa, Kothari, and Wasley (1989) show that the size effect is insignificant when betas are estimated over long horizons. This suggests that the January effect is evidence of return seasonality rather than a misspecification of the Capital Asset Pricing Model.

One explanation for the turn-of-the-year seasonality is the tax-loss selling hypothesis. This hypothesis states that investors wanting to realize capital losses in the current tax year create downward price pressure, at year end, on securities that have previously experienced negative returns. Subsequently, at the beginning of the new tax year, this selling pressure is relieved and the affected securities earn excess returns as their prices rebound. The portfolio window-dressing hypothesis of Haugen and Lakonishok (1988) represents an alternative, but not necessarily mutually exclusive, explanation for the January effect. However, the two hypotheses are difficult to differentiate because they both rely on year-end selling pressure in losing stocks, and both predict a January effect concentrated at the turn-of-the-year. One difference is that the tax-loss selling hypothesis implies no return seasonality prior to the introduction of personal income taxes.

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Recent studies by Schultz (1985) and Jones, Pearce, and Wilson (1987) test the tax-loss selling hypothesis by examining whether a return seasonal existed before the introduction of personal income taxes with the War Revenue Act of 1917. Schultz analyzes the turn-of-the-year returns of small firm portfolios and does not find an effect prior to the War Revenue Act. On the other hand, Jones, Pearce, and Wilson (hereafter referred to as JPW) use the monthly Cowles Commission Industrial Index and detect a January effect before 1917. These conflicting results present a dilemma regarding this issue.

In this paper, our primary objective is to make use of a more comprehensive data set in order to determine the unsettled issue of whether the January and turn-of-the-year effects existed before the introduction of personal income taxes. A secondary objective is the reconciliation of Schultz with JPW. To perform the analysis, we disaggregate the Cowles Industrial Index, used by JPW, and collect price data for the individual stocks in the index. This approach allows us to perform cross-sectional analysis as well as investigate whether the conflicting evidence presented by Schultz and JPW is sample specific or the result of methodological differences. The paper extends the investigation of seasonal returns by recognizing the well established post-tax cross-sectional relationships between excess January returns (or turn-of-the-year returns) and measures of tax-loss selling potential and firm size (e.g., Berges, McConnell, and Schlarbaum (1984), Reinganum (1983), and Roll (1983)). If personal taxes are unrelated to the January effect, then we expect these post-tax cross-sectional relationships to hold in the pre-tax period.

The remainder of the paper is organized as follows. Section I reviews previous studies of a January effect prior to and after the passage of the War Revenue Act. Section II describes the data used in this study. The empirical analysis is presented in Section III. We find no evidence of systematically abnormal turn-of-the-year and January returns in the pretax period. On the other hand, the evidence from the post-tax period is consistent with the seasonal effects documented in earlier studies. Section IV provides a summary and concludes that the change in return behavior in 1917 provides strong support for the tax-loss selling hypothesis as an explanation of the turn-of-the-year and January effects of our post-tax period.

I. Prior Studies

Schultz (1985) examines the turn-of-the-year (the next to last trading day in December through the eighth trading day in January) returns of a small firm portfolio and the Dow Jones Industrials Average (DJIA) for each year from 1900 to 1929. The small firm portfolio includes all stocks listed in the *Wall Street Journal*, from all exchanges, with a bid price of five dollars per share or less. In addition, Schultz computes nine-day returns on the DJIA and a small firm portfolio for several non-January periods. His analysis indicates that prior to the War Revenue Act of 1917, the turn-of-the-year returns of the small firm portfolios were not significantly different from the

DJIA benchmark. Furthermore, he shows that before 1917, the turn-of-the-year returns of the small firm portfolios were not different from the returns of the non-January periods. However, he finds that after 1917, the turn-of-the-year returns of the small firm portfolios were significantly higher than the corresponding benchmark returns. Consequently, Schultz is unable to reject tax-loss selling as a viable explanation for the turn-of-the-year effect.

JPW (1987) re-examine the issue by using the monthly price series of the Cowles Commission Industrial Index. The index price is computed by value-weighting the average of the monthly high and low price of each stock in the index.[1] JPW detect a January seasonal in the monthly returns of the index for the periods before and after the War Revenue Act.[2] Consequently, they conclude that the simultaneous occurrence of the January effect and the turn-of-the-tax-year is a mere coincidence.

Although the results of JPW are impressive, they should not necessarily cause us to reject the equally impressive findings of Schultz. The source of the conflicting results appears to be the different data sets employed in the papers. In the next section of this paper, we describe a data set, with cross-sectional and time-series observations, that allows us to further the investigation of this issue.

II. Data

The primary data sources used in this paper are the *Wall Street Journal* and the Cowles Commission's *Common-Stock Indexes* (1938). In the 1930s, the Cowles Commission for Research in Economics reconstructed several value-weighted common stock indices for the period from 1871 through 1938. These indexes include stocks traded on the NYSE and, subsequent to 1925, stocks traded on a few other exchanges, such as the New York Curb. To accurately measure the actual returns of the month of January and the turn-of-the-year, we collected closing bid and ask prices, as quoted in the *Wall Street Journal*, for each of the stocks included in the Cowles Industrial Index.[3] Returns are calculated for each stock as follows:

$$R_{it} = \frac{(\text{Bid}_{2it} + \text{Ask}_{2it})/2 - (\text{Bid}_{1it} + \text{Ask}_{1it})/2}{(\text{Bid}_{1it} + \text{Ask}_{1it})/2}, \tag{1}$$

where the subscripts $2it$ and $1it$ refer to the closing prices for stock i at the end and beginning, respectively, of an interval from year t.[4] Note that

[1] JPW require Gallant's (1987) procedure for computing t-statistics because a return series calculated from these averages is likely to be serially correlated and heteroscedastic.

[2] JPW test for a January seasonal over two pre-tax/post-tax periods. The first period covers the entire series of the Cowles index from 1871 to 1938. A second period from 1900 to 1929 is analyzed to address the earlier work of Schultz. JPW's conclusions are basically the same for both periods.

[3] A list of every stock included, at any time, in the Cowles Commission Indexes is provided in *Common Stock Indexes* (1939 second edition) on pages 454–475. On the basis of this list, we constructed a database with 1178 records. From this database, we determined the stocks that are included in the Cowles index for the appropriate periods.

[4] Prices were adjusted for stock splits. Dividends were not included. Schultz states the exclusion of dividends has no affect on his results.

because we compute returns with the midpoints of the closing bid and ask prices, our returns are not subject to the bias suggested in Keim (1989).

Four return intervals are measured relative to each January from 1900 through 1929. Following Schultz (1985), one interval consists of the last trading day in December and the first eight trading days in January. The rationale for this interval is that, according to Keim (1983) and Roll (1983), most of the January effect occurs during the first four trading days in January and the last trading day in December. A second nine-day control interval covers the last trading day in June and the first eight trading days in July. The third interval includes the last trading day in December and all of January. A fourth interval covers the last trading day in June and all of July. Thus, the monthly returns subsume the nine-day intervals.[5]

The month of July return was chosen for comparison because JPW found July to have the statistically lowest return, relative to January, in their pre-tax period and the highest return, relative to January, in their post-tax period. Consequently, we are reasonably certain that the average July return for the pre-tax period is not unusually high nor is the corresponding return for the post-tax period unusually low.

The *Wall Street Journal* did not supply daily price quotations for all stocks in the Cowles index. The percentage of Cowles index stocks for which we were able to find prices averaged 85% across all intervals and ranged from 62% in January 1912 to 100% in July 1915. Equally weighted portfolio returns ($\overline{R}_t$) were calculated for each interval from the returns (R_{it}) of the stocks, in the Cowles index, for which prices were found.

In addition, we made use of two other data sources. Yearly high and low prices for the stocks of the turn-of-the-year and January portfolios were collected from the *Commercial and Financial Chronicle*, and Dow Jones Industrial Averages (DJIA) were collected from *The Dow Jones Averages: 1885–1985*, edited by Pierce (1986). The yearly highs and lows are used to compute tax-loss selling and risk measures. The returns on the DJIA are employed as benchmarks for calculating excess returns.

III. Empirical Analysis

In this section, we compare the pre- and post-tax returns of the stocks in the Cowles Industrial Index. Later, we use the cross-sectional observations of the individual stocks of the Cowles index to test the relationship between returns and measures of risk and tax-loss selling potential.

[5] Two further points concerning these return intervals deserve comment. First, since microfilmed copies of the *Wall Street Journal* were not available for every day of the intervals we took adjacent days in two cases. Thus, the June-July interval of 1900 and the turn-of-the-year interval of 1901 actually have ten days. Occasionally stock prices were not readable or missing. In such cases, we took the stock price from an adjacent day. Second, for almost every interval, we examined, the morning edition of the *Wall Street Journal*, where it was available. However, in 1922 only the evening edition of the *Journal* was microfilmed for the sixth to eighth trading days in July. Since some stock prices which were usually stated in the morning edition were not stated in the evening edition, we shifted the July 1922 interval three days backward.

A. Subperiods for Analysis

Although a U.S. federal personal income tax was first implemented in 1913, losses were not deductible against gains until The Revenue Act of 1916, and rates were not raised to substantial levels until the War Revenue Act of October 1917. In addition, December 1917 was the first date the *Wall Street Journal* contained any mention of tax-motivated selling.[6] Therefore, to investigate whether the introduction of personal income taxes had a significant influence on return behavior, we divide the data into two subperiods: a pre-tax period covering the turn-of-the-years from 1899–1900 through 1916–1917, and a post-tax period, which covers 1917–1918 through 1928–1929. These subperiods are consistent with those chosen by Schultz and JPW.

B. Analysis of the DJIA-adjusted Returns of Our Equally Weighted Cowles Industrial Index

Table I shows the return ($\overline{R}_t$), on an equally weighted portfolio, for each turn-of-the-year and each June-July turn-of-the-month, net of the return on the DJIA. This approach is intended to illuminate any size effect that may exist in our equally weighted Cowles index. Recall that Keim (1983) shows that the turn-of-the-year and January anomalies are concentrated in small firms.

During the pretax period (Panel A), the mean DJIA-adjusted returns for the turn-of-the-year and the June-July turn-of-the-month are 0.24% and −0.02%, respectively. Note that these means are calculated as simple averages across the years rather than weighting by the number of stocks in each year of the index. Comparative return differences, *t*-statistics, and nonparametric *z*-score equivalents are presented in Panel C. The comparative statistics indicate that the pre-tax turn-of-the-year returns are not significantly greater than the DJIA returns nor the June-July returns. Given the small number of portfolios, the nonparametric statistics may be appropriate because no assumptions are made about the underlying distributions. The details of the test-statistics are reported in the footnotes to Table I.[7]

The average net returns of the post-tax period (Panel B) for the turn-of-the-year and June-July are 1.63% and 0.08%, respectively. The test-statistics, in Panel C, indicate that the turn-of-the-year returns are significantly greater than the DJIA returns and the June-July returns. In addition, the last

[6] Note that the Revenue Act of 1918 allowed for the deduction of losses, above the level of gains, against ordinary income. Also, in 1921 the rate on capital gains (held for two years or more) was lowered to 12 1/2%. This was about one-sixth the highest marginal rate on income. The 1924 act reduced the tax shield of long-term capital losses to the same 12 1/2% level that applied to long-term gains. This later legislation may have heightened the incentives for year-end selling.

[7] Because the Welch test assumes small samples and unknown variances, it may be more appropriate than the test-statistic shown in footnote c. The Welch test results in a more conservative approach because the degrees of freedom are reduced from $N_x + N_y - 2$. However, the fewer degrees of freedom does not change our interpretation of the results.

comparison in Panel C indicates that the average post-tax turn-of-the-year returns are significantly greater than the corresponding pre-tax returns.

The above results clearly support Schultz's contention that the turn-of-the-year effect, as measured against the DJIA, is not detectable prior to the

Table I

Cowles Index Returns, Net of the Returns on the DJIA, at the Turn-of-the-Year (T-O-T-Y) and the June-July Turn-of-the-Month (T-O-T-M)

The Cowles index returns ($\overline{R}_t$) are computed by equally weighting the returns on the stocks in the Cowles index during each return interval. The mean returns for the pre-tax period and the post-tax period are simple averages across the net returns ($\overline{R}_t$ − $DJIA_t$) of each interval in the associated subperiod.

T-O-T-Y		June-July T-O-T-M	
Year	$\overline{R}_t - DJIA_t$	Year	$\overline{R}_t - DJIA_t$
Panel A: Pre-tax period			
1899-1900	0.0383	1900	0.0165
1900-1901	−0.0014	1901	−0.0004
1901-1902	−0.0051	1902	−0.0216
1902-1903	0.0054	1903	−0.0177
1903-1904	0.0053	1904	0.0016
1904-1905	−0.0181	1905	−0.0057
1905-1906	−0.0039	1906	−0.0075
1906-1907	−0.0012	1907	0.0222
1907-1908	0.0280	1908	−0.0028
1908-1909	0.0014	1909	−0.0087
1909-1910	0.0055	1910	0.0233
1910-1911	−0.0038	1911	−0.0025
1911-1912	−0.0012	1912	−0.0049
1912-1913	−0.0089	1913	−0.0101
1913-1914	0.0062	1914	0.0028
1914-1915	0.0044	1915	0.0008
1915-1916	0.0108	1916	−0.0027
1916-1917	−0.0165	1917	0.0187
Simple average	0.0024		−0.0002
Panel B: Post-tax period			
1917-1918	−0.0010	1918	0.0016
1918-1919	0.0184	1919	0.0090
1919-1920	0.0120	1920	0.0071
1920-1921	0.0709	1921	−0.0021
1921-1922	0.0037	1922	−0.0162
1922-1923	0.0193	1923	0.0092
1923-1924	0.0262	1924	0.0026
1924-1925	0.0033	1925	0.0051
1925-1926	0.0020	1926	−0.0032
1926-1927	0.0136	1927	−0.0133
1927-1928	0.0195	1928	0.0100
1928-1929	0.0072	1929	0.0005
Simple average	0.0163		0.0008

Table I—*Continued*

Panel C: Comparative Statistics

	Diff.	t-value	z-equiv.
Pre-tax T-O-T-Y average vs. zero:	0.0024	0.784[a]	0.675[b]
Pre-tax T-O-T-Y vs. T-O-T-M averages:	0.0026	0.625[c]	0.933[d]
Post-tax T-O-T-Y average vs. zero:	0.0163	2.938[a]	2.940[b]
Post-tax T-O-T-Y vs. T-O-T-M averages:	0.0155	2.544[c]	2.569[d]
Post-tax vs. pre-tax T-O-T-Y averages:	0.0139	2.302[c]	2.519[d]

[a] This t-statistic is computed as $t_{(N-1)} = (\sqrt{N}\,\overline{X})/\sigma_x$, where $\overline{X}$ and σ_x are the simple average and standard deviation of the $X_t = (R_t - \text{DJIA})$ time-series over the N years in the subperiod.

[b] z-score equivalent of Wilcoxon signed-rank statistic.

[c] This t-statistic is computed as $t_{(N_x+N_y-2)} = (\overline{X} - \overline{Y})/\sqrt{\sigma_x^2/N_x + \sigma_y^2/N_y}$, where $\overline{X}$, $\overline{Y}$, σ_x, and σ_y are as above, but X_t and Y_t represent the comparable series.

[d] z-score equivalent of the Wilcoxon-Mann-Whitney statistic.

introduction of personal income taxes. The most striking result in Table I is that the turn-of-the-year returns in the pre-tax period are insignificantly different from the returns on the DJIA, and the average is only one-seventh the magnitude of the corresponding post-tax return.

C. Reconciling the Conflicting Results

The above findings are based on an equally weighted version of the same index used by JPW. Consequently, the source of the conflicting results can be narrowed to the following:

(1) Our return observations are calculated from daily data over nine-day intervals, whereas JPW compute returns by differencing the Cowles index of average monthly high and low prices.
(2) Our results (in Table I) are net of the return on the DJIA, whereas JPW analyze gross returns.
(3) Our Cowles Industrial Index is equally weighted.
(4) Our data set is missing, on average, about 15% of the stocks in the Cowles index.

The first potential source of the conflicting results is the difference in return measures. We use a nine-day interval, as does Schultz, in order to concentrate on tax-induced effects. However, there may be some nontax-related effect that occurs later in January. On the other hand, the JPW method of computing returns (from the averages of highs and lows) may be introducing some bias in their results. To investigate further these potential sources of conflict, we analyze the returns of the full month (including the last trading day of the previous month).

Table II shows the January returns, net of the DJIA, for the pre-tax and post-tax periods plus the net July returns for the pre-tax period. In the pre-tax period (Panel A), the mean DJIA-adjusted returns for January and

July are 0.86% and 0.16%, respectively. The t-statistic, shown in Panel C, indicates that the average January return is borderline significantly greater than zero; however, the nonparametric test is insignificant. In addition, the net January returns are not significantly greater than the net July returns.

Table II

Cowles Index Returns, Net of the Returns on the DJIA, in the Month of January and the Month of July

The Cowles index returns ($\bar{R}_t$) are computed by equally weighting the returns on the stocks in the Cowles index during each return interval. The mean returns for the pre-tax period and the post-tax period are simple averages across the net returns ($\bar{R}_t$ − DJIA $_t$) of each interval in the associated subperiod.

January		July	
Year	$\bar{R}_t$ − DJIA$_t$	Year	$\bar{R}_t$ − DJIA$_t$
Panel A: Pre-tax period			
1899-1900	0.0367	1900	0.0137
1900-1901	−0.0200	1901	0.0072
1901-1902	0.0084	1902	−0.0047
1902-1903	0.0072	1903	0.0047
1903-1904	0.0213	1904	0.0006
1904-1905	−0.0180	1905	0.0057
1905-1906	0.0185	1906	−0.0054
1906-1907	−0.0219	1907	0.0071
1907-1908	0.0456	1908	−0.0061
1908-1909	0.0053	1909	−0.0087
1909-1910	−0.0288	1910	−0.0143
1910-1911	−0.0037	1911	−0.0041
1911-1912	−0.0079	1912	0.0048
1912-1913	0.0113	1913	0.0005
1913-1914	0.0851	1914	0.0150
1914-1915	0.0167	1915	0.0012
1915-1916	0.0172	1916	0.0063
1916-1917	−0.0187	1917	0.0048
Simple average	0.0086		0.0016
Panel B: Post-tax period			
1917-1918	0.0219		
1918-1919	0.0195		
1919-1920	−0.0026		
1920-1921	0.0698		
1921-1922	0.0173		
1922-1923	0.0145		
1923-1924	0.0007		
1924-1925	−0.0102		
1925-1926	0.0094		
1926-1927	0.0107		
1927-1928	0.0273		
1928-1929	−0.0245		
Simple average	0.0128		

Table II—*Continued*

Panel C: Comparative statistics

	Diff.	t-value	z-equiv.
Pretax January average vs. zero:	0.0086	1.293[a]	0.830[b]
Pretax January vs. July averages:	0.0070	1.020[c]	0.981[d]
Post-tax January average vs. zero:	0.0128	1.912[a]	1.765[b]
Post-tax vs. Pre-tax January averages:	0.0042	0.436[c]	0.699[d]

[a] This t-statistic is computed as $t_{(N-1)} = (\sqrt{N}\,\overline{X})/\sigma_x$, where $\overline{X}$ and σ_x are the simple average and standard deviation of the $X_t = (R_t - \text{DJIA})$ time-series over the N years in the subperiod.
[b] z-score equivalent of Wilcoxon signed-rank statistic.
[c] This t-statistic is computed as $t_{(N_x+N_y-2)} = (\overline{X} - \overline{Y})/\sqrt{\sigma_x^2/N_x + \sigma_y^2/N_y}$, where $\overline{X}$, $\overline{Y}$, σ_x, and σ_y are as above, but X_t and Y_t represent the comparable series.
[d] z-score equivalent of the Wilcoxon-Mann-Whitney statistic.

The mean DJIA-adjusted January return in the post-tax period (Panel B) of 1.28% is significantly greater than zero but, as shown in Panel C, insignificantly larger than the corresponding pre-tax return.

The mean DJIA-adjusted pre-tax January return of 0.86% is more than three times the magnitude of the associated average from Table I (0.24%), and only in January of 1914 is the full month return (8.51%) considerably larger than the corresponding turn-of-the-year return (0.62%).[8] This does not suggest a pattern of excessive returns in the latter two-thirds of January. The last comparison of Panel C is insignificantly different because the average net January return of 1.28% in the post-tax period is actually less than the associated turn-of-the-year return (1.63%) of Table I. Consequently no unusually large returns are being earned in the last two-thirds of January during the post-tax period. This throws additional doubt on the borderline significance of the pre-tax January return.

The above results indicate that any tax-induced January effect in the post-tax period is concentrated at the turn-of-the-year and dissipates in the remainder of the month. Consequently, it is clear that use of full month versus turn-of-the-year returns at least partially explains the discrepancy between JPW and Schultz.

JPW are critical of Schultz for his use of the DJIA as a benchmark because the average changes in composition from twelve to twenty stocks in 1916 and then to thirty stocks in 1928. Instead, JPW report gross January returns for the pre- and post-tax periods, and they use a dummy variable regression model to test if the January returns are larger than those of each of the other

[8] The commentary of the *Wall Street Journal*, in mid-January of 1914, indicates that the market rally of the previous week was motivated by lower interest rates, and that the lower rates made the high dividend yields of some low-price stocks very attractive. If these low-price stocks were among the smaller firms of the Cowles index, then this could partially explain the very large net return for the month.

months. Their post-tax January return of 2.61% is larger than ten of eleven months and significantly greater than eight of those ten (see Table II of JPW, p. 458). Also, their pretax January return of 1.62% is larger than ten of eleven months; however, our gross pre-tax January return (not shown in the tables) is only 0.90%. Given, the equal weighting of our index, it seems likely that we would have enhanced any pre-tax effect. Consequently, this raises the question of whether JPW's method of calculating returns (from the averages of highs and lows) may have arbitrarily inflated their January returns. We do not directly test this assertion because of the burden associated with collecting price data from all the stocks of the Cowles index in the months other than January and July.

One could argue that the missing observations of our sample may include many of the small stocks of the Cowles index and, therefore, we may have understated the pretax returns. In the next subsection, we address this and the possibility that the DJIA benchmark may have biased our results.

D. Cross-sectional Analysis of Returns When Stocks Are Segmented by Share Price

To avoid reliance on the DJIA, we would, ideally, like to compare the returns of the large and small firms of our portfolios. See Roll (1983) for an explanation of why a tax-induced seasonality would be concentrated in small firms. Although we do not have data on market values, Stoll and Whaley (1983) have shown share price and firm size (market value) to be highly positively correlated. Therefore, we use bid price as a proxy for firm size and split the portfolios into two segments with equal number of stocks. The results are presented in Table III.

In the pre-tax period (Panel A), no consistent pattern is detectable for the difference in turn-of-the-year returns, between the low-price stocks (LBP) and the high-price stocks (HBP) of our sample. The sign is positive in only eight of eighteen years, and the null hypothesis, that the difference is less than or equal to zero, is rejected only once by the cross-sectional Wilcoxon-Mann-Whitney test. However, in the post-tax period (Panel B), low-price stocks outperform high-price stocks in eleven of twelve years and by a statistically significant amount in eight of those eleven years. Although not shown here, the results are similar for the January (full months) returns.

Panel C shows that the pre-tax average return comparisons, for the turn-of-the-year versus zero and the June-July turn-of-the-month, are both insignificant. On the other hand, the associated post-tax comparisons are both significantly positive. The last two comparisons in Panel C show that only the post-tax turn-of-the-year average is significantly greater than the respective pre-tax return. These results indicate the potential for a January effect in Tables I or II was not diminished by the omission of some small stocks (from the Cowles index) or the use of the DJIA as a benchmark. In addition, the distributions of share price in the years of the pretax period are approximately equivalent to those of the post-tax period. Consequently, the procedure of subtracting the average return of the one-half the stocks with the highest prices from that of the half with the lowest prices does not dilute any pre-tax January effect in Table III. This again leads us to conclude that the discrepancy between our result and JPW is due to their method of computing returns.

A somewhat troublesome result is the large fluctuation of the turn-of-the-year returns in the post-tax period of Table III (as well as Table I). This is especially evident in the very large return for the 1920–1921 interval that tends to make one suspicious of the associated t-statistics. Commentary in the *Wall Street Journal* from late December of 1920 suggested that the volume of tax-loss selling was much higher than in previous years because of the large number of losses incurred by investors since the beginning of a bear market in October of 1919. Further investigation revealed this was the only turn-of-the-year in the post-tax period that followed a sustained bear market. This is similar to Roll's (1983) finding that the January effect is more prevalent when the market has declined in the previous year.[9]

E. Cross-sectional Analysis of Returns When Stocks are Segmented by Tax-Loss Selling Potential

The substantial change in return behavior from the pre- to the post-tax period suggests that the seasonality in the size effect may be tax-induced. As an additional test of taxes, as the source of the change in return behavior, we segment stocks by the following tax-loss selling measure (TLSM): $\text{TLSM}_{it} = \text{Bid}_{1it}/\text{High}_{it}$, where Bid_{1it} is the closing bid price for stock i on the next to last trading day of year t and High_{it} is the highest closing price for stock i in year t. Berges, McConnell, and Schlarbaum (1984), Reinganum (1983), and others document that in the modern era this tax-loss selling proxy is comingled with the January returns of small firms. If this correlation is unrelated to income taxes, then it should hold in the pre-tax period.

Table IV shows the mean turn-of-the-year and January returns for both the pre-tax and the post-tax period. During the pre-tax period (Panel A), the difference in mean returns between the half of the stocks with the highest tax-loss selling measures (HTLSM) and the half with the lowest tax-loss selling measures (LTLSM) displays no consistent pattern. The null hypothesis, that the difference is less than or equal to zero is rejected only once at the turn-of-the-year and twice for the full month. On the other hand, in the post-tax period (Panel B), stocks with high tax-loss selling potential consistently outperform stocks with low tax-loss selling potential. All signs are positive at the turn-of-the-year, and there are only two exceptions for the full month returns. Panel C indicates that the average net returns of the post-tax period are significantly greater than zero, but only the post-tax turn-of-the-year is larger than the respective pre-tax return.

[9] Additional analysis, not shown here, indicates that this negative relationship, between the turn-of-the-year return and the prior year return of the DJIA, holds throughout the post-tax period but is not present in the pre-tax period.

The distribution of the tax-loss selling measures of the sample indicates that the stocks of the pre-tax period have relatively higher potential than those of the post-tax period. This is not really surprising given the bull market of the mid to late 1920s. These findings provide additional support for

Table III

Difference in Mean Returns between the Low Bid Price (LBP) and the High Bid Price (HBP) Segments, of the Cowles Index, at the Turn-of-the-Year (T-O-T-Y) and the June-July Turn-of-the-Month (T-O-T-M)

The mean returns for the LBP and the HBP segments are computed by equally weighting the net returns ($\overline{R}_t$ − DJIA$_t$) of the one-half of the Cowles index stock with the lowest and highest prices, respectively, in each interval. Asterisks indicate the level of significance in cross-sectional nonparametric tests (see footnotes below). The mean returns for the pre-tax period and the post-tax are simple averages across the returns of each interval in the associated subperiod.

T-O-T-Y		June-July T-O-T-M	
Year	LBP-HBP	Year	LBP-HBP
Panel A: Pre-tax period			
1899-1900	0.0111	1900	0.0229
1900-1901	−0.0168	1901	0.0018
1901-1902	−0.0210	1902	−0.0400
1902-1903	0.0034	1903	−0.0447
1903-1904	−0.0112	1904	0.0086
1904-1905	−0.0136	1905	0.0206
1905-1906	0.0170	1906	−0.0065
1906-1907	0.0015	1907	0.0304**
1907-1908	0.0212	1908	0.0070
1908-1909	0.0047	1909	−0.0197
1909-1910	−0.0261	1910	−0.0500**
1910-1911	−0.0059	1911	−0.0051
1911-1912	−0.0165	1912	−0.0257
1912-1913	−0.0172	1913	−0.0118
1913-1914	0.0070	1914	−0.0004
1914-1915	0.0556***	1915	−0.0391
1915-1916	−0.0184	1916	−0.0156
1916-1917	−0.0058	1917	0.0199
Simple average	−0.0017		−0.0026
Panel B: Post-tax period			
1917-1918	0.0199**	1918	0.0086
1918-1919	0.0196*	1919	0.0181
1919-1920	0.0333***	1920	0.0173***
1920-1921	0.1121***	1921	0.0083
1921-1922	−0.0016	1922	0.0031
1922-1923	0.0223**	1923	0.0048
1923-1924	0.0557***	1924	0.0156
1924-1925	0.0270	1925	0.0014
1925-1926	0.0174**	1926	0.0030
1926-1927	0.0260***	1927	−0.0035
1927-1928	0.0094	1928	0.0055
1928-1929	0.0075	1929	−0.0085
Simple average	0.0291		0.0061

Table III—*Continued*

Panel C: Comparative statistics

	Diff.	t-value	z-equiv.
Pre-tax T-O-T-Y average vs. zero:	−0.0017	−0.369[a]	−0.872[b]
Pre-tax T-O-T-Y vs. T-O-T-M averages:	0.0009	0.100[c]	−0.095[d]
Post-tax T-O-T-Y average vs. zero:	0.0291	3.374[a]	2.940[b]
Post-tax T-O-T-Y vs. T-O-T-M averages:	0.0230	2.569[c]	3.031[d]
Post-tax vs. pre-tax T-O-T-Y averages:	0.0308	3.406[c]	3.450[d]
Post-tax vs. pre-tax T-O-T-M averages:	0.0087	1.130[c]	1.031[d]

[a] This t-statistic is computed as $t_{(N\text{-}1)} = (\sqrt{N}\,\bar{X})/\sigma_x$, where $\bar{X}$ and σ_x are the simple average and standard deviation of the $X_t = (R_t - \text{DJIA})$ time-series over the N years in the subperiod.
[b] z-score equivalent of Wilcoxon signed-rank statistic.
[c] This t-statistic is computed as $t_{(N_x+N_y-2)} = (\bar{X} - \bar{Y})/\sqrt{\sigma_x^2/N_x + \sigma_y^2/N_y}$, where $\bar{X}$, $\bar{Y}$, σ_x, and σ_y are as above, but X_t and Y_t represent the comparable series.
[d] z-score and equivalent of the Wilcoxon-Mann-Whitney statistic.
*, **, and *** denote in that year the distribution of LBP is significantly different from the distribution of HBP by the cross-sectional Wilcoxon-Mann-Whitney test, at the 10%, 5%, and 1% levels, respectively.

the tax-loss selling hypothesis as an explanation for the seasonality in returns of small firms.

IV. Summary and Conclusions

The primary objective of this paper is to use an equally weighted version of the Cowles Industrial Index, that we have constructed, to determine if a turn-of-the-year or January effect existed before the introduction of personal income taxes in 1917. The analysis indicates that, in the pre-tax period, the turn-of-the-year returns of our index are not significantly greater than the returns for the DJIA or the June-July turn-of-the-month. However, these differences are highly significant in the post-tax period. Additional evidence indicates that only in the post-tax period did stocks with relatively high tax-loss selling potential earn significantly larger turn-of-the-year and January returns than stocks with lower tax-loss selling potential. Similar results are found for low-price stocks versus high-price stocks. Also, regardless of the benchmark, the net turn-of-the-year returns of the post-tax period are significantly greater than those of the pre-tax period. Note that these results do not provide a solution to the size effect; however, they do indicate that the turn-of-the-year and January seasonals in small firms were induced by the War Revenue Act of 1917.

Many earlier papers find mixed results regarding the tax-loss selling hypothesis. It is possible that additional influences, such as portfolio window-dressing, may further magnify or complicate return anomalies. However, we believe that the evidence of a dramatic change in return behavior, beginning in 1917, indicates that the tax-loss selling hypothesis plays the major role in explaining the seasonal returns of small firms.

Table IV

Difference in Mean Returns between the High Tax-Loss Selling Measure (HTLSM) and the Low Tax-Loss Selling Measure (LTLSM) Segments, of the Cowles Index, at the Turn-of-the-Year (T-O-T-Y) and in the Month of January

The mean returns for the HTLSM and the LTLSM segments are computed by equally weighting the net returns ($\bar{R}_t$ − DJIA_t) of the one-half of the Cowles index stocks with the lowest and highest tax-loss selling potential, respectively, in each interval. Asterisks indicate the level of significance in cross-sectional nonparametric tests (see footnotes below). The mean returns for the pre-tax period and the post-tax are simple averages across the returns of each interval in the associated subperiod.

	T-O-T-Y	January
Year	HTLSM-LTLSM	HTLSM-LTLSM
Panel A: Pre-tax period		
1899-1900	0.0131	−0.0041
1900-1901	−0.0528	−0.0808
1901-1902	0.0003	0.0447
1902-1903	0.0017	0.0093
1903-1904	−0.0390	0.0456
1904-1905	−0.0146	−0.0301
1905-1906	−0.0005	0.0250
1906-1907	0.0032	−0.0336
1907-1908	0.0274	0.0336
1908-1909	−0.0125	−0.0121
1909-1910	−0.0169	−0.0508
1910-1911	0.0034	0.0137
1911-1912	−0.0047	0.0015
1912-1913	−0.0096	0.0271
1913-1914	−0.0001	0.1174***
1914-1915	0.0327*	0.0298
1915-1916	−0.0391	−0.0364
1916-1917	−0.0331	0.0128*
Simple average	−0.0078	0.0063
Panel B: Post-tax period		
1917-1918	0.0353**	0.0307**
1918-1919	0.0180**	0.0031
1919-1920	0.0435***	0.0114
1920-1921	0.1277***	0.1129***
1921-1922	0.0088	0.0297**
1922-1923	0.0268***	0.0272**
1923-1924	0.0608***	0.0192
1924-1925	0.0268*	0.0201
1925-1926	0.0130	0.0092
1926-1927	0.0092	−0.0147
1927-1928	0.0152**	0.0087
1928-1929	0.0084	−0.0174
Simple average	0.0330	0.0200

Table IV—*Continued*

Panel C: Comparative statistics	Diff.	t-value	z-equiv.
Pretax T-O-T-Y average vs. zero:	−0.0078	−1.478[a]	−1.350[b]
Pretax January average vs. zero:	0.0063	0.597[a]	0.500[b]
Post-tax T-O-T-Y average vs. zero:	0.0330	3.350[a]	3.001[b]
Post-tax January average vs. zero:	0.0200	2.040[a]	2.158[b]
Post-tax vs. pre-tax T-O-T-Y averages:	0.0408	3.959[c]	3.704[d]
Post-tax vs. pre-tax January averages:	0.0137	0.911[c]	0.572[d]

[a] This t-statistic is computed as $t_{(N-1)} = (\sqrt{N}\,\overline{X})/\sigma_x$, where $\overline{X}$ and σ_x are the simple average and standard deviation of the $X_t = (R_t - \text{DJIA})$ time-series over the N years in the subperiod.

[b] z-score equivalent of Wilcoxon signed-rank statistic.

[c] This t-statistic is computed as $t_{(N_x+N_y-2)} = (\overline{X} - \overline{Y})/\sqrt{\sigma_x^2/N_x + \sigma_y^2/N_y}$, where $\overline{X}$, $\overline{Y}$, σ_x, and σ_y are as above, but X_t and Y_t represent the comparable series.

[d] z-score equivalent of the Wilcoxon-Mann-Whitney statistic.

*, **, and *** denote in that year the distribution of HTLSM is significantly different from the distribution of LTLSM by the cross-sectional Wilcoxon-Mann-Whitney test, at the 10%, 5%, and 1% levels, respectively.

A secondary objective of this paper is to reconcile the findings of Schultz (1985) and JPW (1987). We find that their conflicting conclusions are primarily due to the different methods of computing returns. The Cowles index used by JPW is computed by value-weighting the average of the monthly high and low prices of each stock in the index. Analysis of our equally weighted version of the index (using beginning and end of period prices) indicates that the method may have arbitrarily inflated their pre-tax January returns.

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