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**An empirical investigation of the relationship between tax-loss selling and the January effect**

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# AN EMPIRICAL INVESTIGATION OF THE RELATIONSHIP BETWEEN TAX-LOSS SELLING AND THE JANUARY EFFECT

A Dissertation

Presented to

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In Partial Fulfillment

Of the Requirements for the Degree

Doctor of Philosophy

by

Thomas M Dalton

May, 1992

**AN EMPIRICAL INVESTIGATION OF THE RELATIONSHIP BETWEEN TAX-LOSS SELLING AND THE JANUARY EFFECT**

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## ABSTRACT

Prior empirical studies have reached different conclusions regarding the influence of tax-loss selling on the January effect. Some studies present evidence supporting tax-loss selling by investors as a cause of the January effect while others demonstrate that the January effect remains strong even in the absence of a tax-loss selling motivation. Anomalies such as the January effect have been cited as evidence against the efficient market hypothesis.

The purpose of this study is to add to the body of knowledge regarding efficient markets by testing a popular explanation of the January effect, the tax-loss-selling hypothesis. This study tests the tax-loss-selling hypothesis by examining the price growth rate in commodity futures contracts before and after the enactment of major tax legislation affecting these contracts.

The tax-loss-selling hypothesis predicts the January effect to be present in commodity futures contracts with high tax-loss-selling potential prior to the Economic Recovery Tax Act of 1981 (ERTA), but not in commodity futures contracts with low tax-loss-selling potential. however, it predicts that after the enactment of ERTA the January effect should disappear or be greatly reduced in all commodity futures contracts. Using a pretest-posttest design with a comparison group, this study finds the pattern of price growth rates predicted by the tax-loss-selling hypothesis. Both OLS and WLS regression analysis is used to document the pattern.

# TABLE OF CONTENTS

CHAPTER 1 . . . . . . . . . . . . . . . . . . . . . . . . 1

INTRODUCTION . . . . . . . . . . . . . . . . . . . . 1

The Tax-Loss-Selling Hypothesis . . . . . . . 2

Description of the Research Setting . . . . . 3

Purpose of the Study . . . . . . . . . . . . 6

CHAPTER 2 . . . . . . . . . . . . . . . . . . . . . . . . 9

LITERATURE REVIEW . . . . . . . . . . . . . . . . . 9

Documentation of the January Effect . . . . . 9

Tests of Stocks with Potential Tax Losses . . 11

Tests of Stock Trading Volume . . . . . . . . 12

Tests of Stocks in Foreign Markets . . . . . 13

Analytical Evaluation of the Tax-Loss-Selling Hypothesis . . . . . . . . . . . . . . . . 15

Tests of the U.S. Pre-taxation Stock Market . 17

Tests of the Futures Market . . . . . . . . . 17

Summary of Literature Review . . . . . . . . 19

CHAPTER 3 . . . . . . . . . . . . . . . . . . . . . . . . 20

CONCEPTUAL FRAMEWORK . . . . . . . . . . . . . . . 20

Hypothesis Foundation . . . . . . . . . . . . 20

Hypothesis Rationale . . . . . . . . . . . . 20

Dependent Variable Surrogate . . . . . . . . 25

Independent Variable Surrogate . . . . . . . 29

CHAPTER 4 . . . . . . . . . . . . . . . . . . . . . . . . 33

RESEARCH DESIGN . . . . . . . . . . . . . . . . . . 33

General Design . . . . . . . . . . . . . . . . . 33

Test and Comparison Groups . . . . . . . . . . 33

Time Frame . . . . . . . . . . . . . . . . . . . 38

Data Source . . . . . . . . . . . . . . . . . . . 39

CHAPTER 5 . . . . . . . . . . . . . . . . . . . . . . . . 40

TESTS AND RESULTS . . . . . . . . . . . . . . . . . 40

Presentation of Rates of Growth . . . . . . . . 40

Regression Analysis of January Effect . . . . 56

Partial F-tests Between Groups . . . . . . . . 63

Test for Mid-month Anomalies . . . . . . . . . 70

Econometric Issues . . . . . . . . . . . . . . 75

CHAPTER 6 . . . . . . . . . . . . . . . . . . . . . . . . 80

ANALYSIS OF RESULTS . . . . . . . . . . . . . . . . 80

Mean Rates of Growth by Month . . . . . . . . . 80

Regression Analysis . . . . . . . . . . . . . . 81

Partial F-tests . . . . . . . . . . . . . . . . 82

Mid-month Rate of Growth Analysis . . . . . . 84

CHAPTER 7 . . . . . . . . . . . . . . . . . . . . . . . . 86

SUMMARY . . . . . . . . . . . . . . . . . . . . . . 86

Discussion . . . . . . . . . . . . . . . . . . . 86

Limitations . . . . . . . . . . . . . . . . . . . 87

Future Research . . . . . . . . . . . . . . . . 88

REFERENCES . . . . . . . . . . . . . . . . . . . . . . 90

# LIST OF TABLES
(with captions)

Table 1 Agricultural futures used in the study. All of these commodity futures are traded on one of the domestic exchanges shown and are listed in the Wall Street Journal. . . . . . . . . . . . . . . . . . . 39

Table 2 Comparison of PREHI means with PRELO means. The two highest and the two lowest PTS quartiles from all the trading days of each month are used in the calculations. . . . . . . . . . . . . . . . . . 42

Table 3 Comparison of PREHI means with PRELO means. Only the highest and the lowest PTS quartiles from all the trading days of each month are used in the calculations. . . . . . . . . . . . . . . . . . . . 44

Table 4 Comparison of PREHI means with PRELO means. Only the first four trading days of each month and the highest and the lowest PTS quartiles are used in the calculations. . . . . . . . . . . . . . . . . 47

Table 5 Comparison of POSTHI means with POSTLO means. The two highest and the two lowest PTS quartiles from all the trading days of each month are used in the calculations. . . . . . . . . . . . . . . . . 50

Table 6 Comparison of POSTHI means with POSTLO means. Only the highest and the lowest PTS quartiles from all the trading days of each month are used in the calculations. . . . . . . . . . . . . . . . . . . . 52

Table 7 Comparison of POSTHI means with POSTLO means. Only the highest and the lowest PTS quartiles from the first four trading days of each month are used in the calculations. . . . . . . . . . . . . . . . . 54

Table 8 Estimated regression coefficients and standard errors for the PREHI group using ordinary least squares. . . . . . . . . . . . . . . . . . . . . . 58

Table 9 Estimated regression coefficients and standard errors for the PREHI group using weighted least squares. . . . . . . . . . . . . . . . . . . . . . 59

Table 10 Estimated regression coefficients and standard errors for the PRELO group using ordinary least squares. . . . . . . . . . . . . . . . . . . . . . 60

Table 11 Estimated regression coefficients and standard errors for the PRELO group using weighted least squares. . . . . . . . . . . . . . . . . . . . . . 60

Table 12 Estimated regression coefficients and standard errors for the POSTLO group using ordinary least squares. . . . . . . . . . . . . . . . . . . . . 61

Table 13 Estimated regression coefficients and standard errors for the POSTLO group using weighted least squares. . . . . . . . . . . . . . . . . . . . . 61

Table 14 Estimated regression coefficients and standard errors for the POSTHI group using ordinary least squares. . . . . . . . . . . . . . . . . . . . . 62

Table 15 Estimated regression coefficients and standard errors for the POSTHI group using weighted least squares. . . . . . . . . . . . . . . . . . . . . 62

Table 16 Partial F-test of whether PREHI = PRELO using regression equations (2) and (3) and ordinary least squares. . . . . . . . . . . . . . . . . . . . . 65

Table 17 Partial F-test of whether PREHI = PRELO using regression equations (2) and (3) and weighted least squares. . . . . . . . . . . . . . . . . . . . . 65

Table 18 Partial F-test of whether PREHI = POSTHI using regression equations (4) and (5) and ordinary least squares. . . . . . . . . . . . . . . . . . . . . 67

Table 19 Partial F-test of whether PREHI = POSTHI using regression equations (4) and (5) and weighted least squares. . . . . . . . . . . . . . . . . . . . . 67

Table 20 Partial F-test of whether POSTHI = POSTLO using regression equations (6) and (7) and ordinary least squares. . . . . . . . . . . . . . . . . . . . 69

Table 21 Partial F-test of whether POSTHI = POSTLO using regression equations (6) and (7) and weighted least squares. . . . . . . . . . . . . . . . . . . . 69

Table 22 Mean rates of growth by month for the 11th through the 15th of each month over the period from September, 1978 through January, 1985. . . . . . . 71

Table 23 Estimated regression coefficients for the regression of mid-month rates of growth on monthly dummy variables using ordinary least squares. . . 73

Table 24 Estimated regression coefficients for the regression of mid-month rates of growth on monthly dummy variables using weighted least squares. . . 74

LIST OF FIGURES
(with captions)

Figure 1 Graph of the PRELO group from Table 2. The two lowest PTS quartiles from all trading days of each month are used in the calculations. . . . . . 43

Figure 2 Graph of the PREHI group from Table 2. The two highest PTS quartiles from all trading days of each month are used in the calculations. . . . . . 43

Figure 3 Graph of the PRELO group from Table 3. The lowest PTS quartile from all trading days of each month are used in the calculations. . . . . . . . 45

Figure 4 Graph of the PREHI group from Table 3. The highest PTS quartile from all trading days of each month are used in the calculations. . . . . . . . 45

Figure 5 Graph of the PRELO group from Table 4. The lowest PTS quartile from the first four trading days of each month are used in the calculations. . 48

Figure 6 Graph of the PREHI group from Table 4. The highest PTS quartile from the first four trading days of each month are used in the calculations. . 48

Figure 7 Graph of the POSTLO group from Table 5. The two lowest PTS quartiles from all trading days of each month are used in the calculations. . . . . . 5

Figure 8 Graph of the POSTHI group from Table 5. The two highest PTS quartiles from all trading days of each month are used in the calculations. . . . . . 51

Figure 9 Graph of the POSTLO group from Table 6. The lowest PTS quartile from all trading days of each month are used in the calculation. . . . . . . . . 53

Figure 10 Graph of the POSTHI group from Table 6. The highest PTS quartile from all trading days of each month are used in the calculations. . . . . . . . 53

Figure 11 Graph of the POSTLO group from Table 7. The lowest PTS quartile from the first four trading days of each month are used in the calculations. . 55

Figure 12 Graph of the POSTHI group from Table 7. The highest PTS quartile from the first four trading days of each month are used in the calculations. . 55

Figure 13 Graph of mid-month rates of growth from Table 22. . . . . . . . . . . . . . . . . . . . . . . . 72

# CHAPTER 1

## INTRODUCTION

The existence of "anomalies" in capital markets is puzzling. It is not clear how anomalies and their characteristic patterns of returns can exist in an efficient market. The existence of anomalies is well documented and publicized [Rozeff and Kinney 1976; Keim 1983; Reinganum 1983], and yet they have continued to exist over long periods of time.

One of the most extensively analyzed anomalies is the "turn-of-the-year" or "January" effect. This pattern of returns is characterized by unusually high returns during the first few days of the calendar year. Originally thought to be limited to relatively low-capitalized stocks, the pattern has been observed as well in stock indexes, debt instruments, and commodity futures indexes [Clark and Ziemba 1987; Wilson and Jones 1990; Gay and Kim 1987]. A number of explanations have been proposed for the January effect. These explanations include 1) December tax-loss-selling with an accompanying January rebound in prices, 2) an unknown risk factor in small stocks that makes small stocks only "appear" to have abnormal January returns, 3) information releases in early January that cause investors to bid up the price of small stocks, 4) the parking-the-proceeds hypothesis which conjectures that investors sell stocks in December, but delay reinvesting the sales proceeds until January, 5) financial statement window dressing by managers that cause investors to bid up the price of small stocks when the financial statements are released, 6) year-end employee bonuses invested in the stock market thereby temporarily increasing demand and stock prices in January, and 7) capital asset pricing model misspecification. None of these proposed explanations have completely explained the January effect. Although the data may partially support one hypothesis or another, uncertainty remains as to the true cause. This study attempts to reduce the uncertainty by examining one of the proposed causes of the January effect, the tax-loss-selling hypothesis.

## The Tax-Loss-Selling Hypothesis

The tax-loss-selling hypothesis holds that investors sell stocks with loss potential in December in order to recognize those losses and thereby offset them against capital gains. Tax-loss selling drives the price of these stocks to levels lower than they would be in the absence of tax selling pressure. In January, when the selling pressure disappears, investors buy back the stocks sold in December at prices that reflect the stocks' true, underlying values. This causes a regular pattern of positive returns during the first few days of the year.

A number of studies find that stocks with high tax-loss potential tend to have higher January returns than stocks with low tax-loss potential [Reinganum 1983; Givoly and Ovadia

1983; De Bondt and Thaler 1985, 1987]. Small stocks in particular have exhibited a net selling trend in December which abruptly changes to a net buying trend in January [Keim 1983]. The results of these studies are consistent with the tax-loss-selling hypothesis.

Other studies, however, find evidence against the tax-loss-selling hypothesis. For example, a January effect has been found in countries with non-December tax year-ends such as Australia [Brown et al. 1983] and Japan [Kato and Schallheim, 1985]. The January effect has been found in the United States as far back as 1871, some forty years before the United States had income taxes [Jones et al. 1987].

This study examines the tax-loss-selling hypothesis as an explanation of the January effect and attempts to gain insight into the anomaly by analyzing data from the commodity futures market before and after changes in the tax law affecting this market.

## Description of the Research Setting

A good setting for testing any research question allows a researcher to have direct control over the independent variables. It also allows random assignment of subjects to groups, [Kerlinger 1986, 349]. The independent variable in this study, tax law, cannot be manipulated nor can the subjects, security investors, be randomly assigned into different groups. Therefore, what Kerlinger [1986, 315] refers to as a "compromise design" is needed. The compromise design must allow for a pre-test and post-test of both the treatment group and a suitable comparison group. The proxy variables and the form of treatment must be selected so that the weaknesses of non-control over the independent variable and of non-random assignment can be mitigated to an acceptable degree. The Economic Recovery Tax Act of 1981 and the commodity futures contract market provide a suitable setting for these requirements.

The Economic Recovery Tax Act of 1981 (ERTA) changed the taxation of commodity futures contracts. Prior to ERTA, gain or loss from the sale of commodity futures contracts held for investment were taxed the same as any other capital asset. Capital gain or loss was recognized upon sale or exchange of the contract and was short-term or long-term depending on the holding period of the contract. For transactions of commodity futures contracts entered into after June 23, 1981, any gain or loss is treated as 60% long-term and 40% short-term without regard to the holding period of the contract. More importantly for this study, all commodity futures positions are "marked-to-market" at year-end for tax purposes and the holder of the position is forced to recognize any gain or loss even though the contract has not yet been sold. In effect, cash basis investors are placed on the accrual basis of accounting for any commodity futures contracts that they hold.

After June 23, 1981 there is no incentive for an investor to sell a commodity futures contract for the purpose of recognizing a tax loss. The tax loss will be recognized by the investor whether the contract is sold or not. Therefore, if the January effect existed in the commodity futures contract market prior to 1981, and if tax-loss selling were the motivation for the effect, then the effect should disappear after June 23, 1981.

A suitable comparison group should be unaffected by the actual treatment and yet be similar enough to the treatment group so that there is reasonable assurance it was only the treatment that caused the effect and not some unanticipated difference between the two groups. Under the tax-loss-selling hypothesis, it is only those assets with tax-loss potential that investors should want to sell in order to generate tax losses. These are assets that have declined in value since their purchase. Assets which have increased in value should result in a tax gain when sold and should not be suitable for generating tax losses at year-end. Futures contracts which have increased in value should not exhibit a January effect either before or after ERTA if the tax-loss-selling hypothesis is plausible. Investors would ignore these gain potential assets when seeking to generate year-end tax losses. Futures contracts which have not decreased in value, therefore, are likely candidates for a comparison group.

The commodity futures market and the enactment of ERTA provide a suitable research setting for testing the tax-loss-selling hypothesis. Commodity futures contracts that have experienced a decline in price since inception comprise a suitable test group. Further, results from this setting should be generalizable to other security markets, such as stocks, which have an observed January effect. Futures contracts are similar to stock securities in that they are exchanged on a regulated market by large numbers of investors.

Although similar in many ways to stocks as investments, commodity futures price patterns are not affected by the release of periodic financial statements. The absence of quarterly financial statements makes using commodity futures appropriate for isolating the tax-loss-selling explanation from other explanations such as the information release hypothesis.

ERTA provides an opportunity to examine these assets in the presence and in the absence of tax-loss selling motivation. The motivation for selling these assets at year-end to create tax losses exists prior to June 23, 1981 and effectively disappears in 1981 with the passage of ERTA. This makes ERTA a good tool for testing the tax-loss-selling hypothesis.

Commodity futures contracts that have not declined in price since inception make an acceptable comparison group. They are highly similar to contracts in the test group except for their tax-loss potential. Because the comparison group futures contracts will have less potential for losses, there should be little if any January effect observed in this group either before or after the passage of ERTA. A comparison of return patterns between these two groups provides useful insight into the tax-loss-selling hypothesis as an explanation of the January effect.

## Purpose of the Study

The general purpose of this research is to add to the body of knowledge regarding efficient markets. Current researchers such as Hand [1990], Harris and Ohlson [1990], Bernard and Thomas [1989] and Bernard [1989] are challenging the efficient market hypothesis. They provide evidence that investors to some extent may be functionally fixated on accounting values and therefore unable to unscramble true cash flow from accounting data. This challenge has important implications for the accounting profession since it suggests that managers can influence the market by manipulating accounting reports.

Market anomalies such as the January effect are often cited as evidence of inefficient markets. Whether anomalies actually provide such evidence depends upon the cause of the anomaly. To the extent that the tax-loss-selling hypothesis is a correct explanation of the January effect, it supports the argument that markets are inefficient. It is important, therefore, that an empirically verifiable explanation of the January effect is found.

As yet, the January effect as an anomaly has not been satisfactorily explained. The tax-loss-selling hypothesis is a popular explanation in the financial world [Jones et. al. 1987]. Some money managers have even suggested recently that tax selling is responsible for a new pattern of abnormally high November returns [Wall Street Journal 11/19/90]. These managers propose that the 1986 tax act gave mutual funds an incentive to sell declining stocks in October, thereby creating a rebound price effect in November similar to the historic January effect. This study attempts to provide additional evidence as to the cause of the January effect by testing the tax-loss-selling hypothesis.

The specific purpose of this study is to extend the methodology of analyzing return anomalies by testing return patterns of disaggregated commodity futures contracts before and after the enactment of ERTA. It does not appear that prior research has analyzed yearly return patterns using a sample of specific commodity futures contracts in a pre-test, post-test comparison group format over the ERTA time period. The marked-to-market rules enacted with ERTA provide an opportunity to gain insight into the tax motivated causes of the January effect. This study extends the current methodology by using ERTA to study the January effect.

This study also examines commodity futures prices for other yearly patterns such as a mid-April price depression and subsequent price rebound. If tax-loss selling can depress prices in December, it is possible that cash flow requirements to pay taxes in April can also cause a pattern of selling resulting in an April/May effect. As a part of an examination of the tax-loss-selling hypothesis, the data is studied for evidence of a mid-April tax motivated selling pattern.

# CHAPTER 2

## LITERATURE REVIEW

### Documentation of the January Effect

Stock return seasonality was first documented by Rozeff and Kinney [1976] using monthly rates of return on the New York Stock Exchange. They combined several indexes of stock returns over the period January, 1904 through December, 1974 and used the combined index to search for abnormal January returns.

The Rozeff and Kinney [1976] index reveals a higher mean of absolute returns in January as compared to other months. Over the entire test period, the average January mean is 3.48% while the average of each of the other 11 months ranges from 1.9% to -0.52%. These differences are generally statistically significant over the entire test period as well as several sub-periods.

Branch [1988] noted that Rozeff and Kinney's findings were highly sensitive to the index used. If Rozeff and Kinney had used a value weighted index rather than an equally weighted index, they would not have found a January effect. A value weighted index gives more weight to high capitalization stocks (large firms) than to low capitalization stocks (small firms). Subsequent researchers have found that it is small firms, not large firms, that demonstrate the

January effect [Keim 1983; Roll 1983; and Reinganum 1982, 1983].

Keim [1983] tied the small firm effect to the January effect and established the regression methodology commonly used in studies of this anomaly. Keim [1983] found that nearly 50% of the small firm effect occurs in January. Further, over 26% of January abnormal returns occur during the first week of trading and almost 11% of January returns are attributable to the first trading day of the year.

In Keim's methodology, firm returns are regressed on monthly dummy variables in each of 10 firm size portfolios. Firm size is determined by market value, computed as the number of common stock shares outstanding at year-end multiplied by the year-end price of the firm's common shares. The ranked firms are divided equally into 10 size portfolios ranging from the smallest firms to the largest. The regression equation used by Keim [1983] to identify the January effect is:

$$R_t = a_1 + a_2D_{2t} + a_3D_{3t} + \ldots + a_{12}D_{12t} + e_t$$

where $R_t$ is the average daily CRSP excess return for day t for the size portfolio under consideration, and where the dummy variables indicate the month of the year in which the excess return is observed ($D_{2t}$ = February, $D_{3t}$ = March, etc.). The excess return for January is measured by $a_1$, while $a_2$ through $a_{12}$ represent the differences between the excess return for January and the excess returns for the other months.

Although Keim's [1983] results do not specifically support the tax-loss-selling hypothesis for the January effect, Keim did suggest that tax-loss-selling might be a viable explanation for the January effect.

<u>Tests of Stocks with Potential Tax Losses</u>

Branch [1977] was the first to empirically test the tax-loss-selling hypothesis. Using the New York Stock Exchange composite index for the 1965-1974 period, he found no evidence of a general market trend before or after the first of the year. However, when he examined specific stocks that had experienced a decline in price during the year, the results were impressive. The price of these stocks were shown to rise briefly in the following year during the first week of January. Branch [1977] estimated that the average gain (neglecting commissions) which could have been made trading in these stocks ranges from 5% to 8%.

Roll [1983] concluded that evidence exists to support the tax-loss-selling hypothesis when he used negative stock returns during the year as a criteria for identifying potential tax-loss stocks. He found an inverse relationship between the yearly returns of these stocks (which are generally negative) and early January returns (which are generally positive). The tax-loss-selling hypothesis predicts this relationship. The results of Roll [1983] show this negative relationship for both a NYSE and AMEX sample of firms for every year between 1962 and 1979.

Reinganum [1983] tested the tax-loss-selling hypothesis more directly. He calculated a measure of potential tax-loss selling (PTS) for each firm as the ratio of a firm's price on the second to the last day of trading in the calendar year over the firm's highest price over the last six months of the calendar year excluding the final two trading days. For example, if the price on the second to the last trading day was 20, and the highest price over the last six months was 25, then the PTS is 0.80 (20/25). Using PTS to identify potential tax-loss stocks and using regression methodology similar to Keim [1983], Reinganum [1983] demonstrated that small firms with the largest tax-loss potential have the highest positive abnormal returns in January. This relationship is predicted by the tax-loss-selling hypothesis.

Other researchers have tested the tax-loss-selling hypothesis by examining stocks with potential tax losses and have generally found positive January returns associated with these stocks, [Givoly and Ovadia 1983; and De Bondt and Thaler 1985, 1987]. The findings of these researchers are generally consistent with the findings of Branch [1977], Roll [1983], and Reinganum [1983].

Tests of Stock Trading Volume

Dyl [1977] found abnormally low trading volume in December for common stocks that had appreciated during the year (winner stocks) and abnormally high trading volume in December for common stocks that had declined during the year (loser stocks). These results are consistent with the tax-loss-selling hypothesis. Dyl [1977] did not test for abnormal returns in January. He only examined trading volume and pointed out that the effect of trading abnormalities on year-end stock price are dependent on the availability of close substitutes for the stocks. In other words, if there are no close substitutes, and if there is a downward sloping demand curve, then the changing demand for tax-loss stocks should cause price movements in those stocks around the end of the year.

Lakonishok and Smidt [1986] also found evidence of increased trading activity of loser stocks in December. They found that winner stocks tend to have higher abnormal volumes than loser stocks until the end of the year. In December, the loser stocks have a higher abnormal volume than winner stocks. This observation supports the tax-loss-selling hypothesis since loser stocks should be subject to increased selling pressure in December, but not necessarily during the rest of the year.

Ritter [1988] examined the daily buy/sell ratios of the cash account customers of Merrill Lynch, Pierce, Fenner and Smith from 1971 to 1985. These cash account customers are non-institutional and are assumed to represent primarily individual investors. Ritter [1988] found that the buy/sell ratio is very low at the end of December and makes a very large increase at the beginning of the new year. This supports the hypothesis that individual investors sell heavily in December and purchase heavily in January, This pattern is generally consistent with a tax motivated January effect.

Tests of Stocks in Foreign Markets

Brown et al. [1983] applied regression methodology similar to that used in Keim [1983]. They obtained similar statistical results, but arrived at an opposite interpretation. Australian stock market returns, rather than

United States stock market returns, were regressed by Brown et al. [1983] and revealed positive January returns. However, Australia does not have a December tax year-end. It has a June tax year-end. With a June tax year-end, a pattern of July positive returns is predicted; however the only other pattern observed from the data was abnormally high August returns. The evidence, then, is inconsistent with the usual form of the tax-loss-selling hypothesis.

Kato and Schallheim [1985] examined the Japanese stock market and found a statistically significant January effect and June effect during the 1952 to 1980 period. At first this appears inconsistent with the tax-loss-selling hypothesis since there is no tax on capital gains for individual investors in Japan nor is there a benefit for capital losses. However, Kato and Schallheim [1985] noted that Japanese firms traditionally pay bonuses in June and December. After a purchase of traditional summer and winter gifts, the remaining bonus money may be usea for investment in capital markets thus stimulating stock prices twice yearly. Although the Kato and Schalheim [1985] results provide evidence for the "bonus" hypothesis rather than the tax-loss-selling hypothesis, the results do not necessarily rule out the tax-loss-selling hypothesis as an explanation for the January effect in the United States.

At least two research studies have taken advantage of the unique features of British tax law. British tax law provides for an April 5 rather than a December 31 tax year-end and, prior to April 6, 1965, did not provide for a capital gains tax. If the tax-loss-selling hypothesis is correct, abnormally positive returns should not appear until after April 6, 1965. Gultekin and Gultekin [1983] tested this prediction and found seasonality in the years before 1965 with the largest returns concentrated in April, but no seasonality after 1965. This result is exactly the opposite of what the tax-loss-selling hypothesis predicts. However Reinganum and Shapiro [1987] re-examined the data used in the Gultekin study and found inaccuracies. Reinganum and Shapiro [1987] corrected the inaccuracies and repeated the Gultekin and Gultekin [1983] test. Using the corrected data they found non-significant levels of seasonality before 1965 but significant seasonality after 1965 with peaks in January and April. The findings of Reinganum and Shapiro [1987] therefore support the tax-loss-selling hypothesis except for the January seasonal after 1965.

It is possible that integration of foreign markets and the American market is responsible for at least part of the observed January effect seen in foreign markets. These and other foreign market studies that reveal January effects in countries with non-December 31 tax year-ends or with no capital gains tax [also see Tinic et al. 1987] may simply reveal the level of integration with the U.S. market.

Analytical Evaluation of the Tax-Loss-Selling Hypothesis

Constantinides [1983, 1984] examined the effect of taxes on January returns from an analytical perspective. His studies visualize the presence of transaction costs as causing tax-loss-selling to increase gradually from January to

December and to suddenly cease in the first few days of January in the following year. Constantinides [1983, 1984] did not predict that tax-loss-selling will necessarily affect stock price. Only if investors are irrational and do not repurchase undervalued stocks will tax-loss-selling depress stock prices. In such an illiquid market, the end of tax-loss-selling at the end of December would cause a positive jump in returns during January. If investors are rational or are aware of seasonality, no effect on stock price would occur since investors would merely trade tax-loss-stocks among each other.

Simulation by Constantinides [1984] indicates that any tax-loss-selling effect would be increased by a distinction between short-term and long-term capital gain tax rates. Chan [1986] tested the analysis of Constantinides [1984] by empirically examining the relationship between January returns and the holding period of tax-loss stocks. Chan [1986] predicted that if the Constantinides analysis is correct regarding the tax-loss-selling hypothesis, then there should be a relationship between stock returns and the long-term or short-term character of potential tax-loss stocks. Chan [1986] found no evidence of the predicted relationship; therefore his results are inconsistent with the Constantinides [1984] optimal tax-loss trading strategy. Chan's [1986] results do not provide any direct evidence in support of the tax-loss-selling hypothesis.

Tests of the U.S. Pre-taxation Stock Market

Prior to the Revenue Act of 1913, the income tax for individuals did not exist. Even after 1913, marginal tax rates for investors were near zero until the Revenue Act of 1917. Schultz [1985] examined the returns of small firms from 1900 to 1929, a period which spans both of these tax acts. He calculated an index from Dow Jones Industrial stocks without adjusting for dividends or changes in the firms comprising the Dow. Using this index, Schultz [1985] found no January effect prior to the 1917 tax act. He found that the effect appears after the tax act in the period 1918 to 1929. This seems to be strong evidence for the tax-loss-selling hypothesis, however these results have not been confirmed by subsequent researchers.

The Jones et al. [1987] results, for example, contradict the Schultz [1985] findings. Jones et al. [1987] used a more consistent measure of returns than that used by Schultz [1985] and found that a January effect existed both before and after the imposition of income taxes. Although the Jones et al. [1987] results do not support the tax-loss-selling hypothesis, they do not necessarily contradict it. Their results do indicate that the January effect appears stronger when taxes are present. It is possible, therefore, that taxes contribute to the effect along with other causes.

Tests of the Futures Market

Research of futures market seasonality originally examined day-of-the-week patterns rather than the January effect [Chiang and Tapley 1983; Junkus 1986; and Phillips-

Patrick and Schneeweis 1988]. A current line of research is developing, however, that extends the study of futures markets to the January effect. These studies have observed a January effect in futures markets but have not specifically tested the tax-loss-selling hypothesis.

Gay and Kim [1987] found a trend of positive percentage price increases during the last trading day of December and the first four trading days of January in the Commodity Research Bureau futures price index for the 1956 to 1985 period. Their results imply a tax motivation because the trend decreased dramatically around the same time that the ERTA marked-to-market rules for futures contracts were enacted. ERTA removed any tax motivation for selling these contracts. The suggestion of a tax motivation is highly suspect, however, because it only derives from the observed trend in one index. Gay and Kim's observation was not intended as a rigorous test of the tax-loss-selling hypothesis. It can be easily argued that a number of influences other than ERTA caused the decrease in year-end price effect.

Other market studies have used indexes and specific commodities to demonstrate market inefficiencies but have not directly tested the tax-loss-selling hypothesis. Clark and Ziemba [1987] devised a strategy using Value Line and S&P indexes which takes advantage of the January effect. Cornett and Trevino [1989] examined futures contracts on corn, soybeans, and wheat and found intra-monthly patterns of returns. Kamara [1990] examined the effect of delivery method on market efficiency using soybean futures contracts. Intra-week market inefficiencies were examined by Chang and Kim [1988] using the Dow Jones Spot Commodity Price Index and by Ma [1986] using gold spot prices. Wilson and Jones [1990] found evidence of a January effect in both commercial paper and corporate bonds during various subperiods from 1857 through 1987.

## Summary of Literature Review

Prior research is inconclusive about the causes of the January effect. Intuitive explanations such as the tax-loss-selling hypothesis are not yet clearly supported by the evidence.

Research investigating the tax-loss-selling hypothesis has not yet taken full advantage of the opportunities available in the futures market. These opportunities are created by special tax provisions peculiar to this market. At best, prior research has used this market to find general trends consistent with the tax-loss-selling hypothesis. This study uses the futures market to specifically test the tax-loss-selling hypothesis.

CHAPTER 3

# CONCEPTUAL FRAMEWORK

## Hypothesis Foundation

The general hypothesis of this study involves the relationship between two unobservable constructs: (1) investor tax motivation and (2) predictable investor behavior in capital markets at year-end. The operational constructs are described in the Dependent Variable Surrogate and the Independent Variable Surrogate subsections of this chapter. The general hypothesis is examined by testing the five alternative hypotheses in Chapter 5, Tests and Results.

## Hypothesis Rationale

The rationale underlying the tax-loss-selling hypothesis has been articulated by several prominent researchers. This section presents the rationale as described by Brown et al. [1983] and Ritter [1988].

According to the Brown et al. [1983] description, tax laws encourage investors to sell securities which have experienced recent price declines so that investors can recognize the resulting short-term capital losses and then use these losses to offset capital gains. The selling motivation typically occurs at the end of the year when investors experience liquidity requirements and when the deadline for recognizing losses in the current year approaches. The selling pressure at year-end forces down prices even further than the declines that have occurred throughout the year. At the beginning of the new tax year, the tax selling pressure disappears and the prices rebound to equilibrium levels. The rebound in prices causes positive abnormal returns in the first few days of the year. These positive abnormal returns are characteristic of the January effect.

The January effect is evident in small firms rather than large firms, according to the Brown et al. [1983] description, because small firms typically have a greater variance in price fluctuation. This greater fluctuation increases the probability that small stocks will experience losses sufficient to motivate tax selling.

Brown et al. [1983] point out the obvious problems with the tax selling rationale. First, even if there is heavy tax related selling of a security at year-end, the selling does not necessarily imply a decline in price. A price decline from selling pressure requires a downward sloping demand curve characteristic of an inelastic product. Inelasticity occurs when there are no close substitutes for the product. Scholes [1972] points out, however, that securities with similar risk characteristics often serve as close substitutes. Since there are many small stocks that have similar risk, there may be many substitutes for tax-loss stocks and the demand curve for these stocks may be essentially horizontal. If the demand curve is horizontal, then there is no reason to expect a price decline from tax-loss selling.

Second, if price declines do occur, it is not clear why investors do not simply change the timing of tax-loss selling to avoid or exploit the decline. For example, an investor could sell tax-loss stocks a few days before the rush and avoid the additional temporary decline in price. Another investor could exploit the price decline by purchasing the tax-loss stocks during the temporary decline, and sell them in January when the price returned to normal. Either of these activities would eventually dissipate the temporary drop in price resulting from tax selling pressure.

Ritter [1988] proposes a variation of the common rationale and counters some of the objections to the tax-loss-selling hypothesis. His rationale, called the "parking-the-proceeds" hypothesis, conjectures that individual investors do indeed sell securities at year-end for tax purposes, but instead of immediately reinvesting the proceeds in similar securities, they "park" the proceeds until January. When the proceeds are reinvested, the buying pressure pushes up the price of small firm stocks in which individuals typically invest.

Ritter's [1988] rationale relies on three assumptions. First, individuals invest a greater share of their funds in small stocks than do institutional investors. This is important since it is individuals, not institutions, that "park" the proceeds of year-end stock sales. Second, buying and selling pressure affects small stock prices. This assumption requires a downward sloping demand curve for small stocks. Third, individuals delay reinvestment of December proceeds from small stock sales until January.

Ritter's [1988] rationale makes the argument for a downward sloping demand curve in small stocks more plausible. If individuals comprise a large proportion of small stock investors, and if they voluntarily remove large amounts of their funds from the small stock market in December, then the demand for small stocks at year-end should be downward sloping. A downward sloping demand curve would cause the price of tax-loss stocks to fall in December with increased selling pressure and to rebound in January with increased buying pressure.

Ritter [1988] presents evidence that individual investors behave according to his assumptions. Business Week, on April 18, 1986, reported the largest 1,000 firms in terms of market capitalization along with the percentage of each firm's stock held by institutions. Ritter [1988] finds a correlation between the percentage held and the log of firm market value, indicating that the proportion of individuals, as opposed to institutions, investing in a firm's stock increases as firm size decreases. He also examines the ratio of buying volume over selling volume of Merrill Lynch, Pierce, Fenner and Smith cash account customers (typically individuals) and finds that the ratio drops in December and abruptly rises in January. This pattern indicates that, as a group, individuals tend to sell stocks heavily in December, but defer replacing the stocks until January.

The rationale proposed by Ritter [1988] does not explain why individuals "park" their December proceeds. By failing to reinvest their proceeds in other small stocks immediately, individuals miss an opportunity to realize abnormal positive returns when prices rebound in January. The "parking-the-proceeds" hypothesis, therefore, contradicts the efficient market hypothesis.

If taxes affect the January seasonal price pattern, then it is reasonable to suspect that taxes may also affect security prices at other times of the year. One possibility is that investors are required to sell part of their portfolios in April to pay tax liabilities. If markets for the sold securities were illiquid, there would be a downward sloping demand for the securities and prices would be temporarily depressed. The depressed prices would be evidenced by negative returns near April 15 each year. This study considers other possible yearly patterns such as negative April returns.

In this study, no distinction is made between long-term and short-term holding period. Although Constantinides [1984] proposed that the tax-loss-selling effect would be increased by a distinction between short-term and long-term capital gain tax rates, this idea has not been supported by empirical evidence. This study assumes that any losses recognized from the sale of commodity futures, whether long-term capital losses, short-term capital losses, or ordinary losses, will contribute to the January effect. It is assumed that investors and dealers are motivated to recognize losses at year-end whatever the tax character.

Dependent Variable Surrogate

The dependent variable, year-end investor behavior in capital markets, is normally operationalized in January effect studies by the year-end observed patterns of stock returns. Instead of stock returns, this study uses observed price changes of commodity futures contracts.

Returns for individual commodity futures contracts are not readily available and must be calculated. Gay and Kim [1987] calculate a daily return from the futures price index published by the Commodity Research Bureau, Inc. as:

$$R_t = (I_t - I_{t-1}) / I_{t-1}$$

where $R_t$ is the return of the index on day t, $I_t$ is the index value on day t, and $I_{t-1}$ is the previous day's index value. This calculation represents a daily percentage change in price. According to Chang and Kim [1988], the role of margin deposits in futures markets makes it inappropriate to view the percentage change in a futures index as a true return on investment. Instead of the term "rate of return", they suggest "rate of growth". The term "rate of growth" will be used hereafter in this study.

Gay and Kim [1987] use index values to calculate rate of growth while searching for day-of-the-week effects. Their calculation is acceptable for this purpose because of the short time span required to study weekly seasonality. Differences in the make-up of contracts comprising the index are very small from the beginning of the week to the end of the week. If yearly seasonality is studied, however, the rate of growth calculated in January may come from an entirely different set of contracts than those used to calculate rate of growth during the other months. For example, a corn futures contract specifying delivery in March is different from a corn futures contract specifying delivery in May. Corn may normally sell for more in March than in May and the futures contracts will reflect this difference. This makes it impossible to use an index of contracts when analyzing yearly returns, even if the contracts all derive from the same underlying commodity. As the index moves through the year, the make-up of contracts in the index changes because the delivery dates change. Eventually the index reflects price changes between different products rather than price changes over time within the same product.

To avoid this problem, rate of growth in this study is calculated individually by commodity contract. An individual commodity contract is defined as a particular commodity future with a specified delivery month. For example, corn futures with a July delivery date is one contract while corn futures with a September delivery date is another contract.

Calculations are made by comparing changes in the closing price of each commodity contract from day to day. For example, when calculating rates of growth during January, contracts from 14 different commodities listed in the Wall Street Journal are chosen. Only the contracts within each commodity having 5, 6, 7, and 8 months remaining until delivery are chosen. The corn futures contracts chosen for January calculations are the May and the July delivery contracts since only these two corn futures contracts have delivery dates from 5 to 8 months away. These two contracts are used to calculate rate of growth throughout January for corn futures. The corn futures contracts chosen for February calculations are the July and the September delivery contracts since it is only these two contracts that have delivery dates from 5 to 8 months away. The corn futures contracts chosen for March calculations are also the July and the September delivery contracts since these two contracts are still the only corn contracts having 5 to 8 months until delivery. The same selection process is repeated for each of the remaining 13 commodities. The contracts chosen for any particular month are used in calculations throughout that month.

Rate of growth is not calculated between contracts having different delivery dates. For example, on the second trading day of March, one observation of the rate of growth for corn futures is the percentage growth in price of a July delivery contract from the first trading day of March to the second trading day of March. A second observation of the rate of growth of corn futures on the second trading day of March is the percentage growth in price of a September delivery contract from the first trading day of March to the second trading day of March. The July contract and the September contract are never compared to each other. The rate of growth on the first day of March for each contract is calculated as the percentage growth of that contract from the last day of February to the first day of March.

Rate of growth for each commodity future, on each trading day, is calculated as follows:

$$R_{it} = (P_{it} - P_{i(t-1)}) / P_{i(t-1)}$$

where $R_{it}$ is the rate of growth for commodity contract i on day t, $P_{it}$ is the price of the futures contract for commodity i on day t, and $P_{i(t-1)}$ is the price of commodity contract i on the previous day. Each calculation is considered a separate, independent observation.

Contracts having five to eight months until delivery are used in the calculations in order to exclude the beginning and ending months of a contract. This will exclude any unexpected influences on rates of growth that may occur due to a calculation being performed at the extreme points in the life of a contract.

Carrying costs (storage, insurance, and risk costs incurred before the delivery date) gradually diminish over the life of a contract and affect contract pricing. The change in carrying costs, however, is extremely small from day to day. Since carrying costs always grow smaller over time, they affect all contracts in the same direction. Since rates of growth are calculated as a day to day percentage change, carrying costs are not expected to have a significant effect on the calculations.

The futures prices selected for the rate of growth calculation is the daily closing price listed in the <u>Wall</u>

Street Journal. Using the closing price is consistent with similar commodity price research, [Chang and Kim 1988; Ma 1986; and Kamara 1990].

Independent Variable Surrogate

The independent variable, tax law, is operationalized by a change in the tax law that affected commodity futures contracts. The Economic Recovery Tax Act (ERTA), enacted in 1981, restructured the taxation of commodity futures trading. Prior to ERTA, gain or loss from holding a commodity futures contract was recognized for tax purposes only upon sale or exchange of the contract. If the contract was held as a capital asset, the investor recognized capital gain or loss upon disposition. The gain or loss was short-term or long-term depending on the holding period of the contract. In other words, investors were taxed on commodity futures contracts in the same fashion as other securities. If tax-loss-selling is associated with the January effect, then the January effect should be observed in commodity futures trading prior to the enactment of ERTA since the motivation for tax-loss-selling existed at that time.

The enactment of ERTA changed the taxation of commodity futures by requiring all commodity futures contracts to be "marked-to-market" at the end of the year for tax purposes. Under the marked-to-market rule, a commodity futures contract that is still held at the end of the year is treated as if it were sold for its fair market value on the last business day of the taxable year [I.R.C. Section 1256(a)(1)]. The holder of the contract must recognize for tax purposes any gain or loss on the contract, even though the contract has not been sold or exchanged.

All gain or loss from commodity futures contracts, whether from contracts terminated during the year, or from those held and marked-to-market at the end of the year, is capital gain or loss to an investor and is arbitrarily treated as 60% long term and 40% short term [I.R.C. Section 1256(a)(3)]. Actual holding period is irrelevant. Contracts that are held until making or taking delivery are treated the same way [I.R.C. Section 1256(c)(1)]. Treating all commodity futures contracts the same, whether held at year-end, terminated by offset during the year, or held until delivery, separates the gain or loss inherent in the commodity futures contract from the gain or loss in the underlying property. This tax treatment eliminates the motive for an investor to sell commodity futures contracts for tax purposes since the date and method of disposition is irrelevant to the investor's tax situation.

The ERTA provisions are effective in general for all regulated futures contracts entered into after June 23, 1981. On the June 23, 1981 effective date, regulated futures contracts were defined as futures contracts that were (1) traded on a U.S. board of trade designated as a contract market by the Commodity Futures Trading Commission, or a market determined by the Treasury to have rules adequate to carry out the purposes of I.R.C. Section 1256, (2) were marked-to-market daily with required margin deposits, and (3) required delivery of personal property.

On June 23, 1981, the term regulated futures contract was broad enough to include all agricultural futures traded on domestic commodity exchanges. Since that time, the definition has been expanded to include, interest rate futures, currency futures, stock market index futures, non-equity options, and stock option futures. These additional contracts were not considered regulated futures contracts when ERTA was enacted. For example, stock market index futures and Eurodollar futures contracts were not unambiguously considered regulated futures contracts until the Technical Corrections Act of 1982. Agricultural commodity futures, however, were clearly regulated futures contracts beginning in 1981, subject to the marked-to-market rules. This study only examines agricultural commodity futures so that investor uncertainty over the enactment date will not confound the results.

Dealers and traders (as opposed to investors) of regulated futures contracts are also required to follow the marked-to-market rules for most transactions. However, tax treatment differs from investors in two ways. First, if the contracts are not capital assets but are instead ordinary assets, all of the gain or loss is ordinary income or loss rather than capital gain or loss [I.R.C. Section 1256(f)(2)]. This only means that the 60/40 rule does not apply. The contracts are still marked-to-market at the end of the year as are investor's contracts.

Second, mixed straddles and qualified hedging transactions are both exempt from the marked-to-market rules [I.R.C. Section 1256(d)(4) and Section 1256(e)]. A mixed straddle consists of a regulated futures contract and an offsetting position in a non-regulated futures contract security. A qualified hedging transaction is one in which a dealer holds a regulated futures contract to hedge a position in the underlying property. Dealer positions that are exempt from the marked-to-market rules are susceptible to tax-loss selling even after the enactment of ERTA. Therefore, if evidence of the January effect remains after 1981, non-marked-to-market dealer trading may provide a partial explanation.

## CHAPTER 4

## RESEARCH DESIGN

### General Design

This research uses the pretest-posttest design with a comparison group. Kerlinger [1986, 315] calls this design a compromise design when it is used in a quasi-experimental setting since, in this setting, it is not possible to randomize nor is there direct control over the independent variable. However, this quasi-experimental design is the "best" counterpart of the completely randomized true experimental design [Abdel-Khalik 1979, 39]. It allows a before and after comparison between a group receiving the test treatment and a group not receiving the test treatment. The major threat to the internal validity of a quasi-experiment using this design is self-selection bias [Abdel-Khalik 1979, 39]. This threat is discussed in the econometric problems section of this chapter.

### Test and Comparison Groups

The test group is comprised of commodity futures contracts with a potential for tax-loss selling. Under the tax-loss-selling hypothesis, securities that can generate tax losses when sold should exhibit a greater degree of positive returns in January than other securities. The potential for generating tax losses can be measured by the relative decline in price of a security over a period of time. For example, a security that has declined in price over the year prior to December has a greater potential for tax-loss selling in December than a security that has gained in price over the year.

Reinganum [1983] uses this approach to examine the January effect in small stocks. In his study, he calculates a measure for each stock called "potential tax-loss selling" (PTS). This measure is the ratio of stock price on the second to the last trading day of the calendar year divided by the maximum stock price from the beginning of July through the second to the last trading day of the calendar year. The lower the calculated ratio, the higher the tax-loss selling potential and the more likely the stock will be sold at year-end for tax purposes. In other words, a low PTS value indicates a high potential for tax-loss selling and a high PTS value indicates a low potential for tax-loss selling. Reinganum [1983] only calculates PTS in December for each year of his study.

A variation of this method can be used to identify commodity futures contracts with tax-loss selling potential. The relative decline in price of a contract measures its PTS, or tax-loss selling potential. PTS in this study is calculated by comparing the highest price observed since the beginning of the contract (or 12 months, whichever is shorter) to the price observed on the second to the last trading day of the month of PTS calculation. Twelve months is used as a maximum time period for PTS calculation in order to remain consistent with prior studies using PTS as a variable. PTS is calculated in the month prior to the rate of growth calculation. For example, PTS of a corn futures contract in January will be the ratio of the contract's price on the second to the last trading day of December divided by the highest price of the contract observed prior to the second to the last trading day of December. As in Reinganum [1983], the higher the PTS value, the lower the tax-selling potential and the lower the PTS value, the higher the tax-selling potential.

The PTS calculation is made separately for each month of the year and for each separate contract. This is necessary for two reasons. First, contracts with different delivery dates will be used as the rate of growth is calculated throughout the year. For example, the January rate of growth calculations for corn futures is made with a May delivery corn futures contract and with a July delivery corn futures contract. The March rate of growth calculations for corn futures is made with a July delivery corn futures contract and a September delivery corn futures contract. It is not possible (or even desirable) to calculate a PTS for all of these contracts over the same standard interval of time as it is calculated in the stock studies. Most stock studies calculate PTS for all stocks over the June to December period. This is not possible for all commodity contracts since some of the contracts used to calculate rate of growth will not even have been in existence in the prior June to December period.

Second, calculating PTS only in the June to December period assumes that a pre-December decline in price is the only decline that can influence selling. For example, suppose investors tend to sell contracts with a low PTS value in whatever month they realize that the price decline had occurred (December or otherwise) and that the prices of these contracts tend to rebound in the following month. This would cause a contract which had declined in price in the pre-June period to experience a selling spree in June and a price rebound in July. Separating this contract into a high or low PTS value group based upon the pre-December PTS period would not reveal the June/July pattern. Instead, the test would be associating a selling pattern in June based upon a price decline that had occurred months before in December. It makes more sense to calculate PTS in the month preceding the rate of growth calculation.

The following is an example of PTS calculation for each month. One of the corn futures contracts used to calculate rates of growth during January, 1980 will be the July delivery corn futures contract. The PTS calculated in January, 1980 for the July delivery contract will be the closing price of the July delivery contract on the second to the last trading day of December, 1979 divided by the highest closing price of the July delivery contract observed on any trading day up to 12 months prior to the second to the last trading day of December, 1979. This PTS measure will be used to assign the contract either to the pre-ERTA high tax-loss-selling potential group (PREHI) or to the pre-ERTA low tax-loss-selling potential group (PRELO) for January, 1980. The PTS calculated in February, 1980 for the July delivery corn futures contract will be the price of the contract on the second to the last trading day of January, 1980 divided by the highest price of the contract observed on any trading day up to 12 months prior to the second to the last trading day of January, 1980. This PTS measure is be used to assign the contract either to the PREHI or to the PRELO groups for February, 1980. This procedure is used for each month throughout the testing period.

If the tax-loss-selling hypothesis is plausible, January rates of growth in the pre-ERTA period for the high tax-loss-selling potential group (PREHI) will be higher than either the pre-ERTA January rates of growth for the low tax-loss-selling potential group (PRELO) or the non-January rates of growth of either group.

Reinganum [1983] used PTS to separate his sample into quartiles of tax-loss selling potential. The same approach is used in this study. The pre-ERTA sample and the post-ERTA sample are each separated into quartiles. The eight groups (four in the pre-ERTA sample and four in the post-ERTA sample) created by the quartiles is further separated into test groups and comparison groups. The group with the highest tax-loss-selling potential in the pre-ERTA period is the pre-ERTA test group (PREHI) and the group with the lowest tax-loss-selling potential in the pre-ERTA period is the pre-ERTA comparison group (PRELO). The group with the highest tax-loss-selling potential in the post-ERTA period is the post-ERTA test group (POSTHI) and the group with the lowest tax-loss-selling potential in the post-ERTA period is the post-ERTA comparison group (POSTLO).

This study examines the change in January rate of growth before and after ERTA in the high tax-loss-selling potential commodity futures. It also examines the difference in January rate of growth between the high tax-loss-selling potential commodity futures and the low-tax-loss-selling potential commodity futures before and after ERTA. These comparisons will test the hypothesis that tax law directly contributes to the January effect.[1]

Time Frame

The data are collected over the 82 month period from April, 1978 through January, 1985. The pre-ERTA test period includes the 33 months from September, 1978 through June, 1981. The post-ERTA test period includes the 42 months from July, 1981 through January, 1985. Data collected prior to September, 1978 is used to calculate PTS. Because June, 23, 1981 is the effective date of the marked-to-market provisions, June, 1981 has been omitted from both testing periods.

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[1] The calculations can be illustrated by computing rate of growth and PTS on January 2, 1980 using the July delivery corn futures contract. The rate of growth calculation is made using the closing contract price on January 2, 1980 of $308.00 and the closing contract price on December 31, 1979 of $310.50. The rate of growth equals -0.008052, [($308.00 - $310.50) / $310.50].

The PTS calculation for the July delivery corn futures contract during January 1980 is made from the closing contract price on December 28, 1979 of $312.00 (the second to the last trading day of the previous month) and from the closing contract price on October 4, 1979 $318.50 (the highest contract price prior to December 28, 1979). The January 1980 PTS for the July delivery corn futures contract is 0.97959 [$312.00/$318.50].

Data Source

The commodity contract data were purchased from C.A.C.T. Seminars, a local firm specializing in tracking commodity prices listed in the Wall Street Journal. The data were received on 14 floppy discs in ASCII format. Fourteen commodity futures contracts are used. Each commodity future is divided into separate contracts based upon the delivery dates for the commodity. The commodity futures used in the study and their exchanges are listed on Table 1.

**Table 1** Agricultural futures used in the study. All of these commodity futures are traded on one of the domestic exchanges shown and are listed in the Wall Street Journal.

| Commodity Futures Contracts (exchange) | |
|---|---|
| Corn (CBT) | Soybean meal (CBT) |
| Oats (CBT) | Soybeans (CBT) |
| Live cattle (CME) | Cocoa (CSCE) |
| Live hogs (CME) | Pork bellies (CME) |
| Cotton (CTN) | Coffee (CSCE) |
| Orange juice (CTN) | Soybean oil (CBT) |
| Sugar (CSCE) | Wheat (CBT) |

| Commodity Exchange Key | |
|---|---|
| CBT | Chicago Board of Trade |
| CME | Chicago Mercantile Exchange |
| CSCE | Coffee, Sugar & Cocoa Exchange |
| CTN | New York Cotton Exchange |

# CHAPTER 5

## TESTS AND RESULTS

The research tests used in this study are divided into four parts. First, the calculated rates of growth by month are presented in both tabular and graphical form. Second, regression analysis is used to statistically test for the January effect in both the high tax-loss-selling commodity futures and the low tax-loss-selling commodity futures, before and after ERTA. Third, partial F-tests are used to make specific contrasts between the pre-ERTA high and low tax-loss-selling potential groups, between the pre-ERTA and post-ERTA high tax-loss-selling potential groups, and between the post-ERTA high and low tax-loss-selling potential groups. Fourth, calculated rates of growth (presented in tabular and graphical form) and regression analysis is used to test specifically for mid-month downward price pressure in April.

### Presentation of Rates of Growth

Mean rates of growth and their standard errors, calculated by month and by group, are presented in Tables 2 through 7. Table 2 compares rates of growth by month between the PREHI group (prior to June, 1981, high tax-loss-selling potential group) and the PRELO group (prior to June, 1981, low tax-loss-selling potential group). The two highest quartiles of tax-loss-selling potential and the two lowest quartiles of tax-loss-selling potential in the pre-ERTA period from all trading days of each month are used in the calculation of means. The mean growth rates shown on Table 2 are presented graphically in Figure 1 and Figure 2.

When all four quartiles of tax-loss-selling potential are used to calculate the means (Table 2), no January effect is evident. Therefore, rate of growth means are recalculated and presented in Table 3 using only the highest and the lowest quartiles of tax-loss-selling potential. In other words, the results on Table 3 show rate of growth means with the two middle quartiles deleted. The mean growth rates shown on Table 3 are presented graphically on Figure 3 and Figure 4.

Table 2 Comparison of PREHI means with PRELO means. The two highest and the two lowest PTS quartiles from all the trading days of each month are used in the calculations.

| MONTH | Descrip of values | PRELO | PREHI |
|---|---|---|---|
| Jan | Gr Rate | 0.000585 | -0.000062 |
| Jan | Std Dev | 0.012771 | 0.017036 |
| Feb | Gr Rate | 0.000512 | -0.000545 |
| Feb | Std Dev | 0.013296 | 0.014627 |
| Mar | Gr Rate | -0.002079 | -0.001222 |
| Mar | Std Dev | 0.013720 | 0.017007 |
| Apr | Gr Rate | -0.000044 | -0.000058 |
| Apr | Std Dev | 0.011623 | 0.016894 |
| May | Gr Rate | -0.001587 | -0.000407 |
| May | Std Dev | 0.041925 | 0.013682 |
| Jun | Gr Rate | 0.002391 | 0.001244 |
| Jun | Std Dev | 0.019341 | 0.014624 |
| Jul | Gr Rate | -0.000424 | 0.001897 |
| Jul | Std Dev | 0.018990 | 0.019856 |
| Aug | Gr Rate | 0.001235 | 0.002407 |
| Aug | Std Dev | 0.015431 | 0.016952 |
| Sep | Gr Rate | 0.001072 | 0.002105 |
| Sep | Std Dev | 0.013862 | 0.019343 |
| Oct | Gr Rate | 0.001687 | -0.000772 |
| Oct | Std Dev | 0.015855 | 0.015587 |
| Nov | Gr Rate | 0.000655 | 0.000032 |
| Nov | Std Dev | 0.016202 | 0.013626 |
| Dec | Gr Rate | -0.003096 | -0.000805 |
| Dec | Std Dev | 0.015735 | 0.013727 |

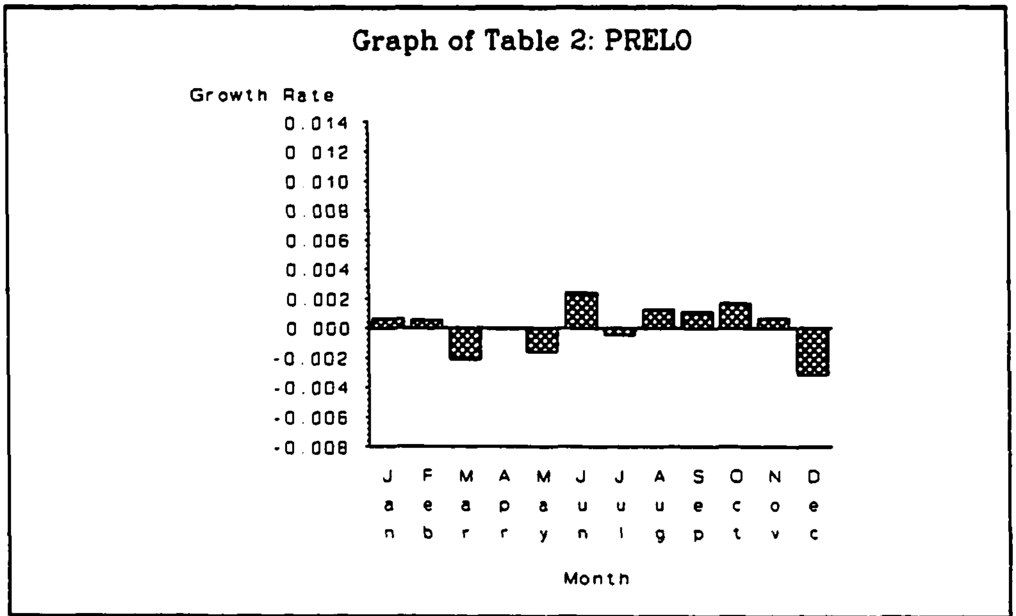


**Figure 1** Graph of the PRELO group from Table 2. The two lowest PTS quartiles from all trading days of each month are used in the calculations.

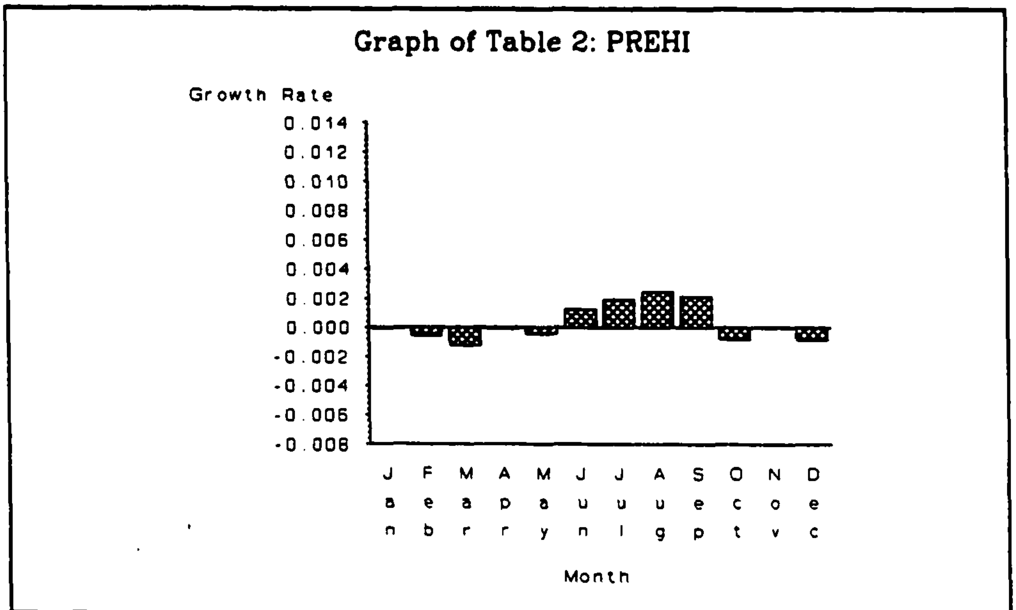


**Figure 2** Graph of the PREHI group from Table 2. The two highest PTS quartiles from all trading days of each month are used in the calculations.

**Table 3** Comparison of PREHI means with PRELO means. Only the highest and the lowest PTS quartiles from all the trading days of each month are used in the calculations.

| MONTH | Descrip of value | PRELO | PREHI |
|---|---|---|---|
| Jan | Gr Rate | 0.002321 | 0.000172 |
| Jan | Std Dev | 0.014794 | 0.020669 |
| Feb | Gr Rate | 0.001946 | -0.001157 |
| Feb | Std Dev | 0.014551 | 0.017047 |
| Mar | Gr Rate | -0.000939 | 0.000239 |
| Mar | Std Dev | 0.011325 | 0.019910 |
| Apr | Gr Rate | -0.000107 | 0.000024 |
| Apr | Std Dev | 0.010548 | 0.018786 |
| May | Gr Rate | 0.000290 | -0.000576 |
| May | Std Dev | 0.010484 | 0.014256 |
| Jun | Gr Rate | 0.001682 | 0.000986 |
| Jun | Std Dev | 0.020796 | 0.017421 |
| Jul | Gr Rate | -0.002947 | -0.000926 |
| Jul | Std Dev | 0.017991 | 0.018838 |
| Aug | Gr Rate | 0.001553 | 0.003412 |
| Aug | Std Dev | 0.013768 | 0.019101 |
| Sep | Gr Rate | 0.000142 | 0.001313 |
| Sep | Std Dev | 0.012900 | 0.022441 |
| Oct | Gr Rate | 0.002396 | 0.000645 |
| Oct | Std Dev | 0.015767 | 0.018444 |
| Nov | Gr Rate | 0.000233 | -0.000256 |
| Nov | Std Dev | 0.016486 | 0.014131 |
| Dec | Gr Rate | -0.004112 | -0.000930 |
| Dec | Std Dev | 0.016809 | 0.017593 |

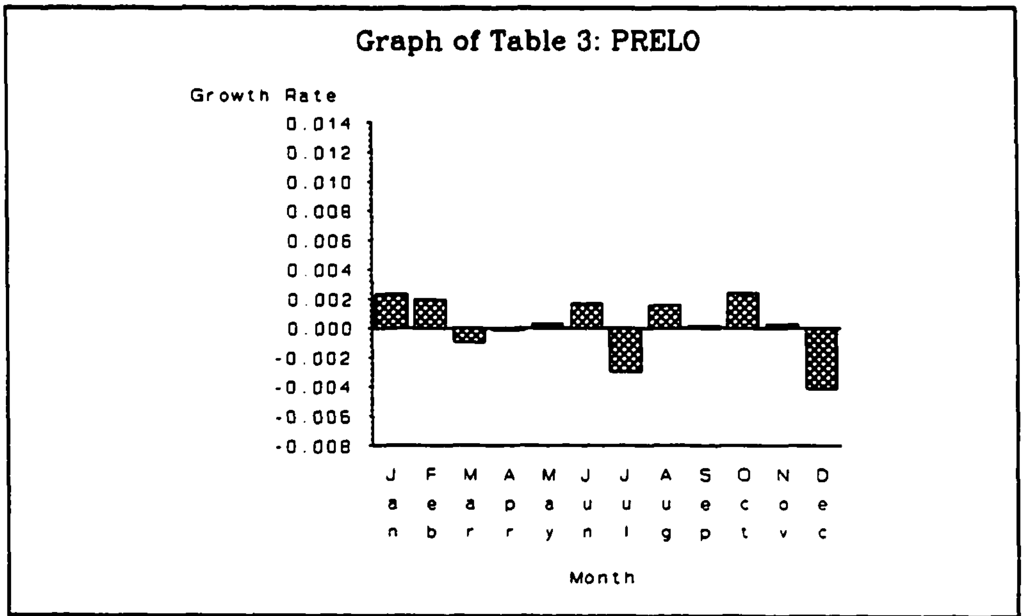


**Figure 3** Graph of the PRELO group from Table 3. The lowest PTS quartile from all trading days of each month are used in the calculations.

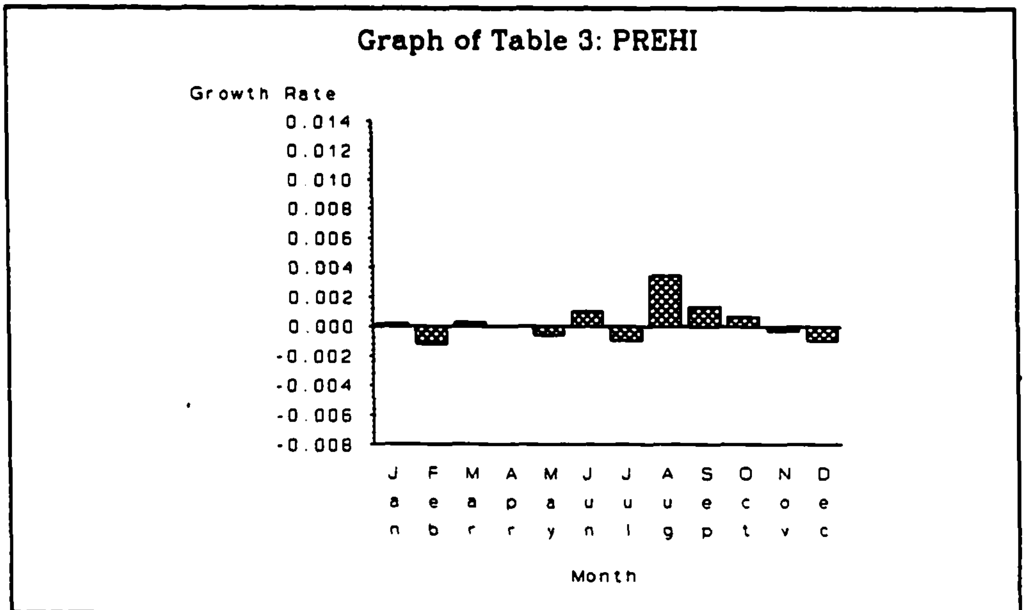


**Figure 4** Graph of the PREHI group from Table 3. The highest PTS quartile from all trading days of each month are used in the calculations.

No January effect is evident in Table 3, even though only the highest and lowest PTS quartiles were used. Because the January effect has generally been found in the first few days of the year [Keim 1983; Gay and Kim 1987], rate of growth means are recalculated a third time and presented in Table 4. The calculation uses only the first four trading days of each month and only the highest and the lowest quartiles of tax-loss-selling potential. The results of Keim [1983] and Gay and Kim [1987] suggest that this time period will reveal a January effect.

Positive January rates of growth are evident in Table 4 when calculated using the first four trading days of each month and using only the highest and lowest PTS quartiles. Table 4 also reveals positive rate of growth peaks in July and September. Gay and Kim [1987] found that mean daily returns in the Commodity Research Bureau futures price index peaked in July as well as January. They did not attempt to explain the July seasonal.

**Table 4** Comparison of PREHI means with PRELO means. Only the first four trading days of each month and the highest and the lowest PTS quartiles are used in the calculations.

| MONTH | Descrip of values | PRELO | PREHI |
|---|---|---|---|
| Jan | Gr Rate | -0.002516 | 0.004771 |
| Jan | Std Dev | 0.012451 | 0.019223 |
| Feb | Gr Rate | 0.003889 | 0.000528 |
| Feb | Std Dev | 0.015680 | 0.018518 |
| Mar | Gr Rate | -0.002343 | -0.004895 |
| Mar | Std Dev | 0.011837 | 0.022887 |
| Apr | Gr Rate | -0.002057 | 0.000443 |
| Apr | Std Dev | 0.011015 | 0.019607 |
| May | Gr Rate | 0.002791 | 0.000933 |
| May | Std Dev | 0.012395 | 0.016123 |
| Jun | Gr Rate | 0.002221 | -0.008036 |
| Jun | Std Dev | 0.014878 | 0.018654 |
| Jul | Gr Rate | 0.004417 | 0.011326 |
| Jul | Std Dev | 0.015824 | 0.015816 |
| Aug | Gr Rate | -0.002031 | 0.000846 |
| Aug | Std Dev | 0.013051 | 0.021527 |
| Sep | Gr Rate | 0.001162 | 0.004772 |
| Sep | Std Dev | 0.013181 | 0.020915 |
| Oct | Gr Rate | 0.002701 | -0.003079 |
| Oct | Std Dev | 0.014169 | 0.020584 |
| Nov | Gr Rate | -0.001455 | -0.005953 |
| Nov | Std Dev | 0.017059 | 0.014513 |
| Dec | Gr Rate | -0.007598 | -0.005949 |
| Dec | Std Dev | 0.012716 | 0.019213 |

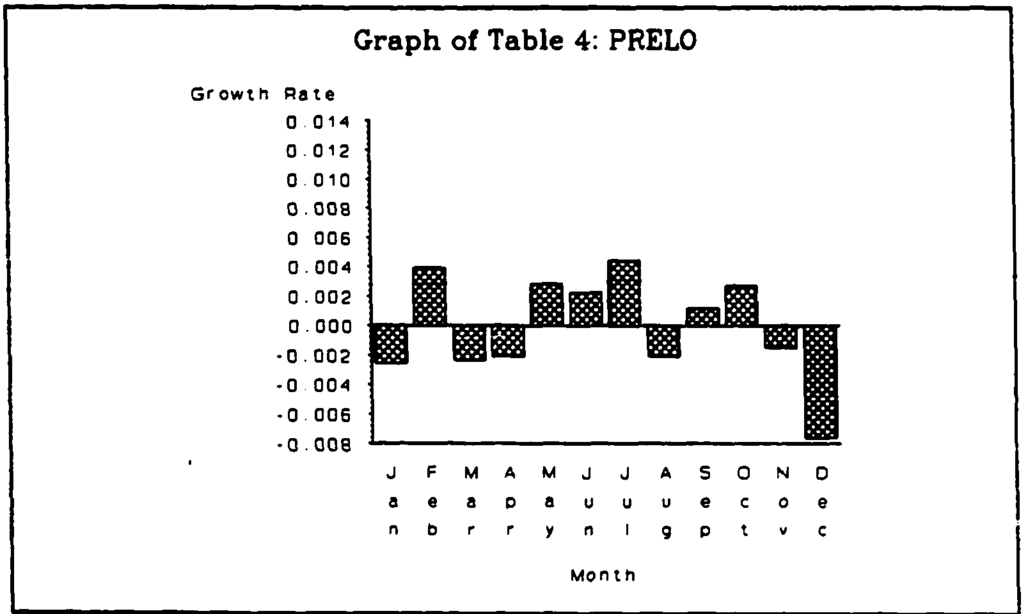


**Figure 5** Graph of the PRELO group from Table 4. The lowest PTS quartile from the first four trading days of each month are used in the calculations.

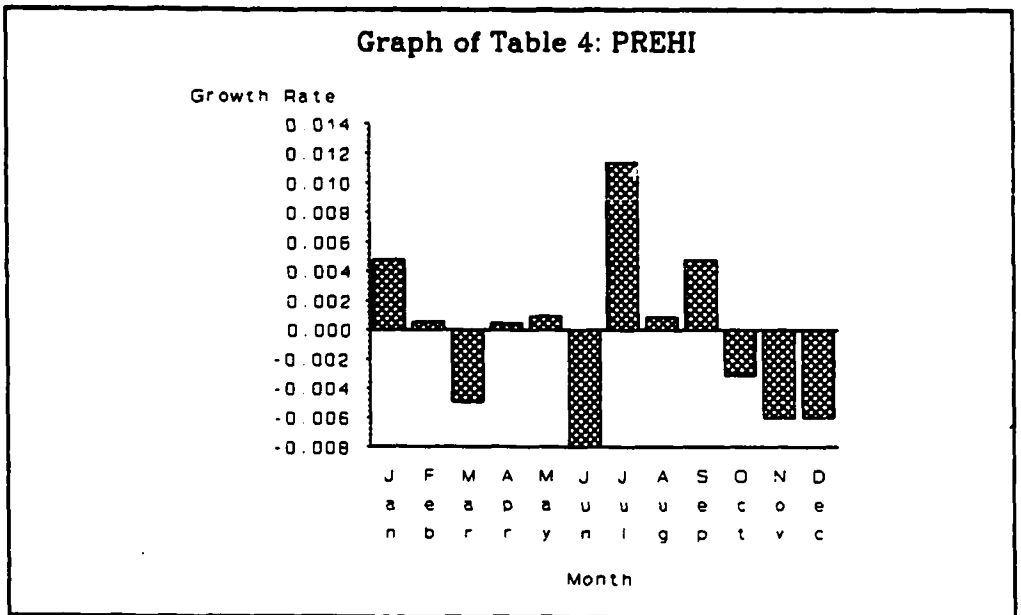


**Figure 6** Graph of the PREHI group from Table 4. The highest PTS quartile from the first four trading days of each month are used in the calculations.

Rates of growth by month are also examined in the post-ERTA period (subsequent to June, 1981). Tables 5 through 7 compare the POSTLO group (subsequent to June, 1981, low tax-loss-selling potential group) with the POSTHI group (subsequent to June, 1981, high tax-loss-selling potential group). As in the pre-ERTA comparisons, the means are calculated first using all trading days of the month and the two highest and the two lowest PTS quartiles, second using all trading days of the month and only the highest and the lowest PTS quartiles, and third using only the first four trading days of each month and only the highest and the lowest PTS quartiles.

None of the post-ERTA growth rate calculations indicate large positive January rates of growth relative to the other months. Disappearance of the January seasonal in the high tax-loss-selling group after the enactment of ERTA indicates that taxes may influence January rates of growth. This conclusion is unwarranted, though, without further testing since the PREHI January rate of growth mean is not shown to be significantly different from zero by any of the comparisons. However, the results do show a preliminary trend that can be supported by regression analysis in the next section. The mean growth rates shown on Tables 5 through 7 are shown graphically on Figures 7 through 12.

Table 5 Comparison of POSTHI means with POSTLO means. The two highest and the two lowest PTS quartiles from all the trading days of each month are used in the calculations.

| MONTH | Descrip of value | POSTLO | POSTHI |
|---|---|---|---|
| Jan | Gr Rate | -0.000126 | 0.001282 |
| Jan | Std Dev | 0.012926 | 0.015361 |
| Feb | Gr Rate | -0.001552 | -0.001396 |
| Feb | Std Dev | 0.012441 | 0.011517 |
| Mar | Gr Rate | 0.001748 | 0.001232 |
| Mar | Std Dev | 0.011352 | 0.012538 |
| Apr | Gr Rate | 0.000605 | -0.000646 |
| Apr | Std Dev | 0.011132 | 0.012010 |
| May | Gr Rate | -0.000967 | -0.000638 |
| May | Std Dev | 0.032283 | 0.012577 |
| Jun | Gr Rate | -0.000883 | -0.000879 |
| Jun | Std Dev | 0.015693 | 0.014142 |
| Jul | Gr Rate | -0.000901 | 0.001434 |
| Jul | Std Dev | 0.014591 | 0.015909 |
| Aug | Gr Rate | 0.001110 | -0.000914 |
| Aug | Std Dev | 0.016730 | 0.015837 |
| Sep | Gr Rate | -0.000968 | -0.001503 |
| Sep | Std Dev | 0.016165 | 0.012760 |
| Oct | Gr Rate | -0.000401 | 0.000665 |
| Oct | Std Dev | 0.013380 | 0.014917 |
| Nov | Gr Rate | -0.000287 | -0.000140 |
| Nov | Std Dev | 0.011195 | 0.013255 |
| Dec | Gr Rate | 0.000723 | -0.001409 |
| Dec | Std Dev | 0.010799 | 0.013832 |

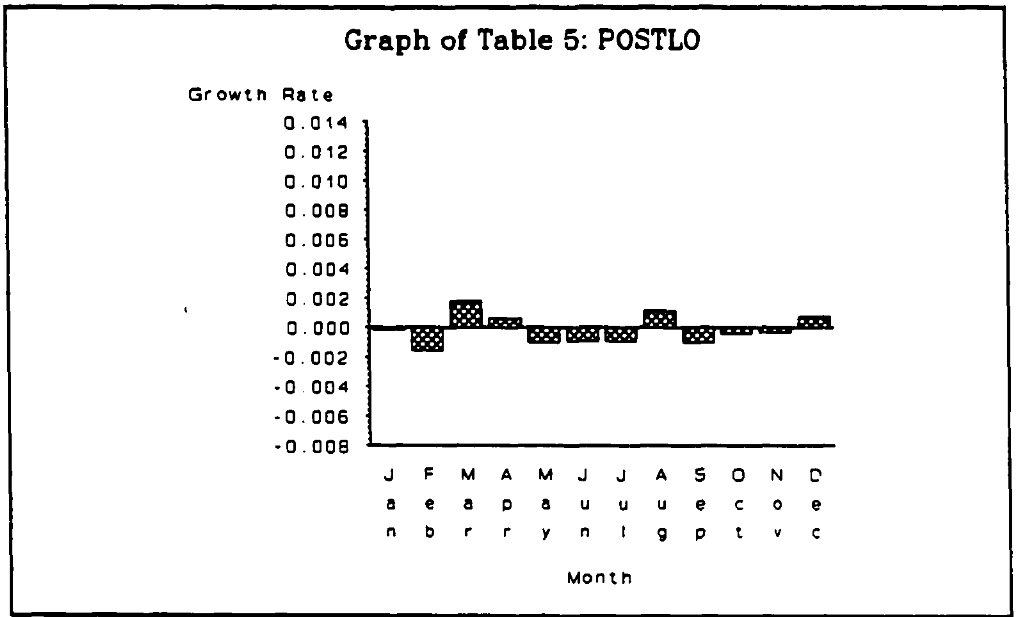


**Figure 7** Graph of the POSTLO group from Table 5. The two lowest PTS quartiles from all trading days of each month are used in the calculations.

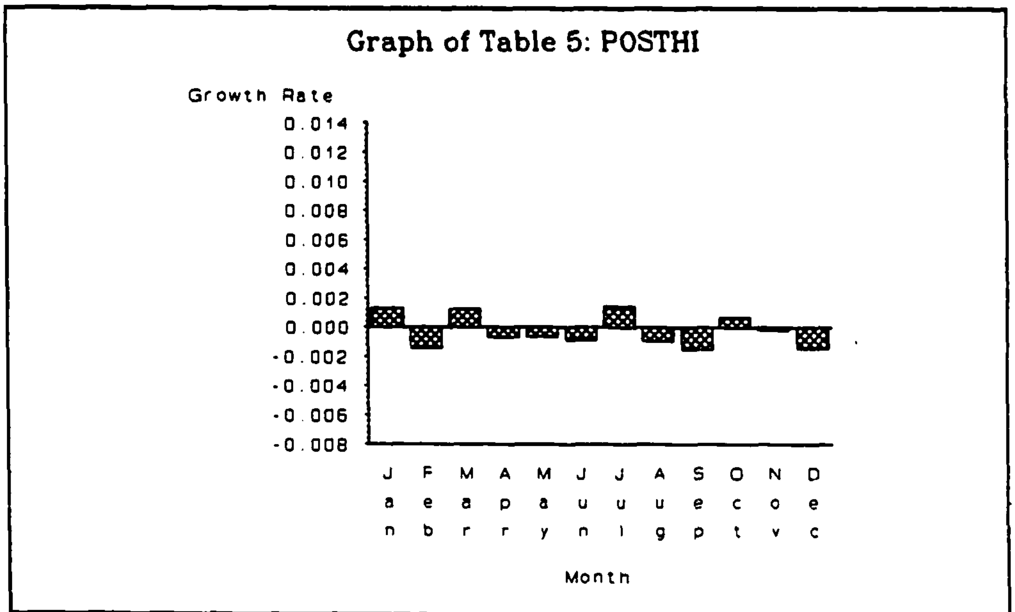


**Figure 8** Graph of the POSTHI group from Table 5. The two highest PTS quartiles from all trading days of each month are used in the calculations.

Table 6 Comparison of POSTHI means with POSTLO means. Only the highest and the lowest PTS quartiles from all the trading days of each month are used in the calculations.

| MONTH | Descrip of value | POSTLO | POSTHI |
|---|---|---|---|
| Jan | Gr Rate | -0.000130 | 0.001737 |
| Jan | Std Dev | 0.011995 | 0.016067 |
| Feb | Gr Rate | -0.001267 | -0.001978 |
| Feb | Std Dev | 0.012208 | 0.012422 |
| Mar | Gr Rate | 0.001192 | 0.000901 |
| Mar | Std Dev | 0.010281 | 0.012952 |
| Apr | Gr Rate | 0.000754 | -0.000232 |
| Apr | Std Dev | 0.009592 | 0.012977 |
| May | Gr Rate | -0.001384 | -0.000213 |
| May | Std Dev | 0.038753 | 0.015139 |
| Jun | Gr Rate | -0.001418 | -0.000651 |
| Jun | Std Dev | 0.017135 | 0.016780 |
| Jul | Gr Rate | 0.000071 | 0.002115 |
| Jul | Std Dev | 0.012317 | 0.017239 |
| Aug | Gr Rate | 0.003218 | -0.000701 |
| Aug | Std Dev | 0.018285 | 0.018148 |
| Sep | Gr Rate | -0.001189 | -0.001304 |
| Sep | Std Dev | 0.015375 | 0.014263 |
| Oct | Gr Rate | -0.001070 | 0.000507 |
| Oct | Std Dev | 0.012719 | 0.015483 |
| Nov | Gr Rate | 0.000430 | 0.000040 |
| Nov | Std Dev | 0.008740 | 0.013463 |
| Dec | Gr Rate | 0.000527 | -0.002174 |
| Dec | Std Dev | 0.009208 | 0.015230 |

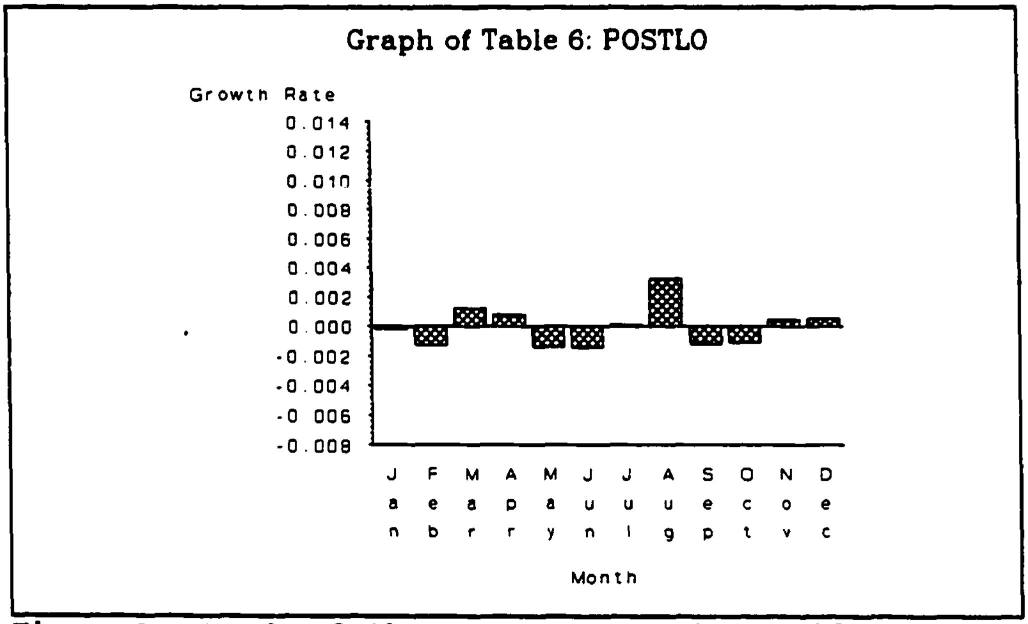


**Figure 9** Graph of the POSTLO group from Table 6. The lowest PTS quartile from all trading days of each month are used in the calculation.

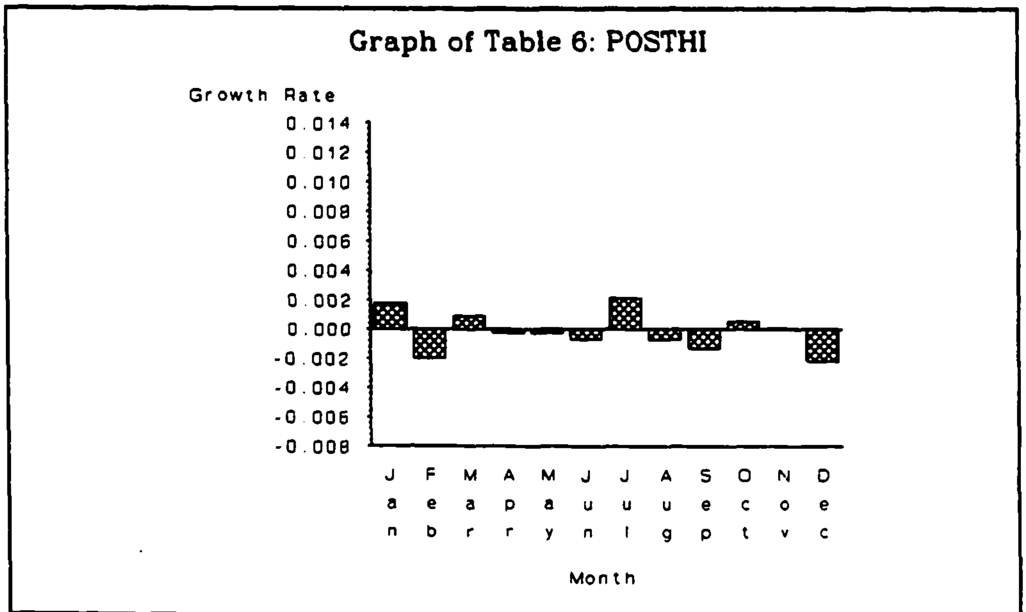


**Figure 10** Graph of the POSTHI group from Table 6. The highest PTS quartile from all trading days of each month are used in the calculations.

Table 7 Comparison of POSTHI means with POSTLO means. Only the highest and the lowest PTS quartiles from the first four trading days of each month are used in the calculations.

| MONTH | Descrip of values | POSTLO | POSTHI |
|---|---|---|---|
| Jan | Gr Rate | 0.000033 | 0.000560 |
| Jan | Std Dev | 0.012796 | 0.013553 |
| Feb | Gr Rate | -0.000700 | -0.001756 |
| Feb | Std Dev | 0.009711 | 0.009345 |
| Mar | Gr Rate | 0.002883 | 0.000896 |
| Mar | Std Dev | 0.006924 | 0.011281 |
| Apr | Gr Rate | 0.001996 | 0.001870 |
| Apr | Std Dev | 0.010728 | 0.012720 |
| May | Gr Rate | 0.001434 | -0.002577 |
| May | Std Dev | 0.007949 | 0.012484 |
| Jun | Gr Rate | -0.004530 | -0.005687 |
| Jun | Std Dev | 0.017079 | 0.016873 |
| Jul | Gr Rate | -0.004731 | 0.002925 |
| Jul | Std Dev | 0.012480 | 0.020719 |
| Aug | Gr Rate | 0.006330 | 0.001164 |
| Aug | Std Dev | 0.015952 | 0.016388 |
| Sep | Gr Rate | 0.001873 | -0.001595 |
| Sep | Std Dev | 0.013210 | 0.015381 |
| Oct | Gr Rate | -0.004288 | 0.003095 |
| Oct | Std Dev | 0.012006 | 0.015338 |
| Nov | Gr Rate | 0.001598 | 0.003013 |
| Nov | Std Dev | 0.008916 | 0.011081 |
| Dec | Gr Rate | -0.001268 | -0.002548 |
| Dec | Std Dev | 0.007778 | 0.013191 |

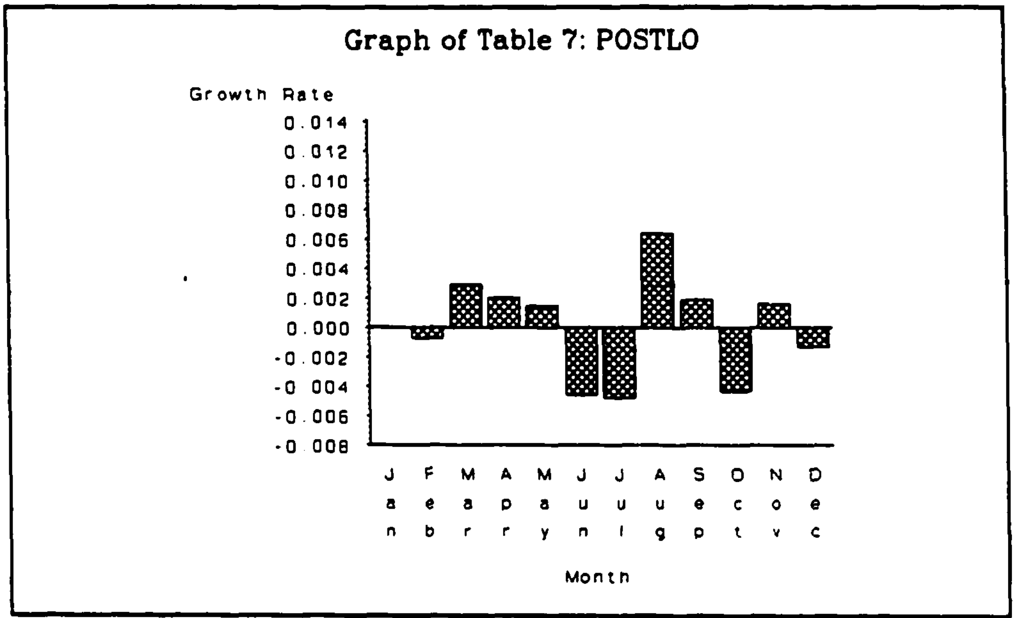


**Figure 11** Graph of the POSTLO group from Table 7. The lowest PTS quartile from the first four trading days of each month are used in the calculations.

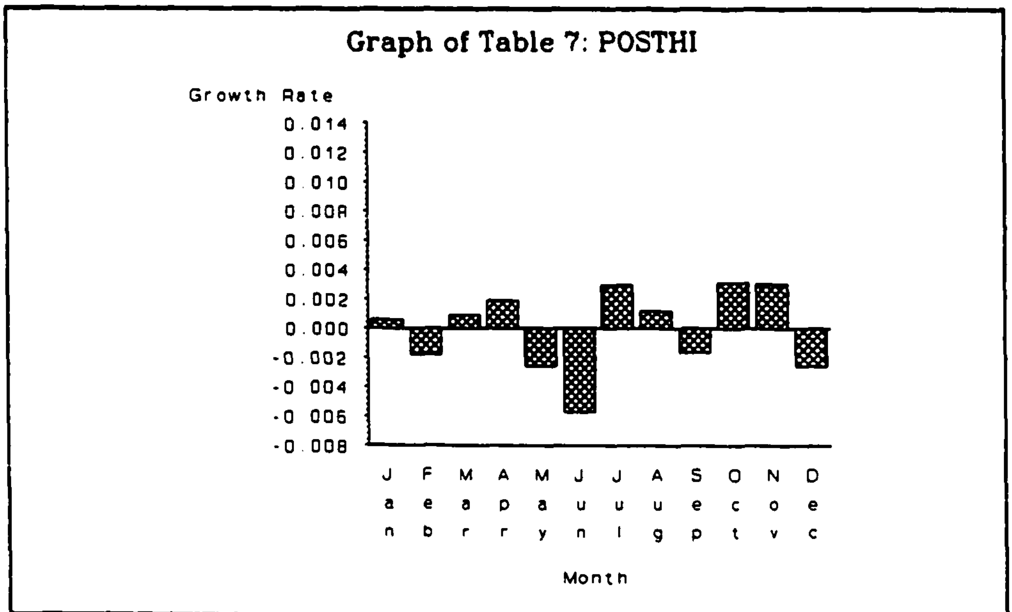


**Figure 12** Graph of the POSTHI group from Table 7. The highest PTS quartile from the first four trading days of each month are used in the calculations.

Regression Analysis of January Effect

The presence of a January effect is analyzed using regression analysis similar to that used by Keim [1983] and Reinganum [1983]. Rate of growth is regressed on 11 dummy variables, one representing each month. A separate regression is estimated for each of the following four groups: (1) PRELO - observations prior to June, 1981 in the first four trading days of each month in the highest tax-loss-selling quartile, (2) PREHI - observations prior to June, 1981 in the first four trading days of each month in the lowest tax-loss-selling quartile, (3) POSTLO - observations after June, 1981 in the first four trading days of each month in the lowest tax-loss-selling quartile, and (4) POSTHI - observations after June, 1981 in the first four trading days of each month in the highest tax-loss-selling quartile. The first four days of the month were chosen for sample because Gay and Kim [1987] had previously found a trend of positive percentage price increases over this period.

The regressions are estimated as:

$$R_{it} = b_1 + b_2D_{2t} + b_3D_{3t} + \ldots + b_{12}D_{12t} + e_{it} \qquad (1)$$

where $R_{it}$ is the rate of growth calculated for a commodity future on day t and contract i for the group under consideration, and where the dummy variables indicate the month of the year in which the rate of growth is observed ($D_{2t}$ = February, $D_{3t}$ = March, etc.). The excess rate of growth for January is measured by $b_1$, while $b_2$ through $b_{12}$ represent the differences between the excess rate of growth for January and the excess rates of growth for the other months. The first alternative hypothesis states the expected relationship between tax-loss-selling potential and positive January rates of growth in terms of the regression coefficients. If the tax-loss-selling hypothesis is correct then the PREHI group should demonstrate a statistically significant positive rate of growth in the first part of January.

> H1: The PREHI test group will demonstrate a positive rate of growth in the first four trading days of January compared to the first four trading days of other months, indicated by a significantly positive $b_1$ coefficient and significantly negative coefficients $b_2$ through $b_{12}$.

Table 8 presents the estimated regression coefficients and standard errors obtained using the PREHI test group.

A statistically significant January effect is evident in the PREHI group as well as a non-significant July seasonal. Since White's test indicates heteroskedasticity in the residuals, the standard errors of the coefficients are estimated a second time using weighted least squares. Each independent variable observation is weighted by the reciprocal of the estimated standard error for that observation. The estimated standard errors for each observation is calculated using the method described by Kmenta [1986, 291] and is discussed in Econometric Problems in this chapter. Table 9

Table 8 Estimated regression coefficients and standard errors for the PREHI group using ordinary least squares.

| Variable | Parameter Estimate | Standard Error | T for H0: Parameter=0 | Prob > |
|---|---|---|---|---|
| INTERCEP | 0.004771 | 0.00190006 | 2.511 | 0.0122 |
| FEB | -0.004242 | 0.00280657 | -1.512 | 0.1310 |
| MAR | -0.009665 | 0.00274250 | -3.524 | 0.0004 |
| APR | -0.004328 | 0.00266209 | -1.626 | 0.1044 |
| MAY | -0.003838 | 0.00277333 | -1.384 | 0.1668 |
| JUN | -0.012807 | 0.00391707 | -3.270 | 0.0011 |
| JUL | 0.006555 | 0.00473111 | 1.386 | 0.1663 |
| AUG | -0.003925 | 0.00292413 | -1.342 | 0.1799 |
| SEP | 0.000002 | 0.00374697 | 0.000 | 0.9996 |
| OCT | -0.007849 | 0.00348475 | -2.252 | 0.0246 |
| NOV | -0.010723 | 0.00329100 | -3.258 | 0.0012 |
| DEC | -0.010719 | 0.00338118 | -3.170 | 0.0016 |

Test of First and Second Moment Specification:
DF 11 Chisq Value 29.316 Prob>Chisq 0.0020

Durbin-Watson D 2.146 No. of obs. 796

presents the weighted least squares estimates. The results presented in Table 8 and Table 9 provide support for H1.

If the tax-loss-selling hypothesis is plausible, then the January effect will not be as strong in the post-ERTA period high tax-loss-selling potential group (POSTHI) or in either of the low tax-loss-potential groups (PRELO and POSTLO) as it is in the pre-ERTA test group (PREHI). A visual inspection of the regression coefficients indicates that this pattern exists. The estimated regression coefficients for the PRELO, POSTHI, and POSTLO groups are shown on Tables 10, 11, 12, 13, 14, and 15.

Neither autoregression or multicollinearity appear to be a problem in the regressions as discussed in <u>Econometric Problems</u> of this chapter. However because of significant heteroskedasticity, all regressions are estimated using weighted least squares as well as ordinary least squares.

**Table 9** Estimated regression coefficients and standard errors for the PREHI group using weighted least squares.

(Weighted Least Squares)

| Variable | Parameter Estimate | Standard Error | T for H0: Parameter=0 | Prob > |
|---|---|---|---|---|
| INTERCEP | 0.004771 | 0.00188580 | 2.530 | 0.0116 |
| FEB | -0.004242 | 0.00275771 | -1.538 | 0.1244 |
| MAR | -0.009665 | 0.00285363 | -3.387 | 0.0007 |
| APR | -0.004328 | 0.00265503 | -1.630 | 0.1035 |
| MAY | -0.003838 | 0.00263211 | -1.458 | 0.1452 |
| JUN | -0.012807 | 0.00384341 | -3.332 | 0.0009 |
| JUL | 0.006555 | 0.00433255 | 1.513 | 0.1307 |
| AUG | -0.003925 | 0.00300097 | -1.308 | 0.1913 |
| SEP | 0.000002 | 0.00383848 | 0.000 | 0.9996 |
| OCT | -0.007849 | 0.00354360 | -2.215 | 0.0270 |
| NOV | -0.010723 | 0.00298761 | -3.589 | 0.0004 |
| DEC | -0.010719 | 0.00335519 | -3.195 | 0.0015 |

**Table 10 Estimated regression coefficients and standard errors for the PRELO group using ordinary least squares.**

| Variable | Parameter Estimate | Standard Error | T for H0: Parameter=0 | Prob > |
|---|---|---|---|---|
| INTERCEP | -0.002516 | 0.00220895 | -1.139 | 0.2550 |
| FEB | 0.006405 | 0.00270540 | 2.368 | 0.0181 |
| MAR | 0.000173 | 0.00266410 | 0.065 | 0.9482 |
| APR | 0.000459 | 0.00299094 | 0.153 | 0.8781 |
| MAY | 0.005308 | 0.00286140 | 1.855 | 0.0640 |
| JUN | 0.004737 | 0.00293818 | 1.612 | 0.1073 |
| JUL | 0.006933 | 0.00459830 | 1.508 | 0.1320 |
| AUG | 0.000485 | 0.00344240 | 0.141 | 0.8880 |
| SEP | 0.003678 | 0.00305211 | 1.205 | 0.2285 |
| OCT | 0.005217 | 0.00258586 | 2.018 | 0.0440 |
| NOV | 0.001062 | 0.00253067 | 0.420 | 0.6749 |
| DEC | -0.005082 | 0.00249143 | -2.040 | 0.0417 |

Test of First and Second Moment Specification
DF 11 · Chisq Value 32.167 Prob>Chisq 0.0007

Durbin-Watson D 2.237 No. of obs. 834

**Table 11 Estimated regression coefficients and standard errors for the PRELO group using weighted least squares.**

(Weighted Least Squares)

| Variable | Parameter Estimate | Standard Error | T for H0: Parameter=0 | Prob > |
|---|---|---|---|---|
| INTERCEP | -0.002516 | 0.00207610 | -1.212 | 0.2258 |
| FEB | 0.006405 | 0.00265031 | 2.417 | 0.0159 |
| MAR | 0.000173 | 0.00248450 | 0.070 | 0.9445 |
| APR | 0.000459 | 0.00273640 | 0.168 | 0.8669 |
| MAY | 0.005308 | 0.00268687 | 1.975 | 0.0486 |
| JUN | 0.004737 | 0.00287610 | 1.647 | 0.0999 |
| JUL | 0.006933 | 0.00475078 | 1.459 | 0.1448 |
| AUG | 0.000485 | 0.00328092 | 0.148 | 0.8825 |
| SEP | 0.003678 | 0.00290829 | 1.265 | 0.2063 |
| OCT | 0.005217 | 0.00247525 | 2.108 | 0.0353 |
| NOV | 0.001062 | 0.00248106 | 0.428 | 0.6688 |
| DEC | -0.005082 | 0.00234691 | -2.165 | 0.0307 |

Table 12 Estimated regression coefficients and standard errors for the POSTLO group using ordinary least squares.

| Variable | Parameter Estimate | Standard Error | T for H0: Parameter=0 | Prob > |
|---|---|---|---|---|
| INTERCEP | 0.000032869 | 0.00107403 | 0.031 | 0.9756 |
| FEB | -0.000733 | 0.00153195 | -0.478 | 0.6325 |
| MAR | 0.002850 | 0.00169820 | 1.678 | 0.0936 |
| APR | 0.001963 | 0.00163039 | 1.204 | 0.2288 |
| MAY | 0.001401 | 0.00147356 | 0.951 | 0.3421 |
| JUN | -0.004563 | 0.00157625 | -2.895 | 0.0039 |
| JUL | -0.004764 | 0.00207354 | -2.297 | 0.0218 |
| AUG | 0.006297 | 0.00153195 | 4.110 | 0.0001 |
| SEP | 0.001840 | 0.00178584 | 1.030 | 0.3030 |
| OCT | -0.004321 | 0.00178584 | -2.420 | 0.0157 |
| NOV | 0.001565 | 0.00186028 | 0.842 | 0.4002 |
| DEC | -0.001301 | 0.00159305 | -0.817 | 0.4142 |

Test of First and Second Moment Specification
DF 11 Chisq Value 80.438 Prob>Chisq 0.0000

Durbin-Watson D 2.003 No. of obs. 1104

Table 13 Estimated regression coefficients and standard errors for the POSTLO group using weighted least squares.

(Weighted Least Squares)

| Variable | Parameter Estimate | Standard Error | T for H0: Parameter=0 | Prob > |
|---|---|---|---|---|
| INTERCEP | 0.000032869 | 0.00109827 | 0.030 | 0.9761 |
| FEB | -0.000733 | 0.00146737 | -0.499 | 0.6176 |
| MAR | 0.002850 | 0.00147824 | 1.928 | 0.0541 |
| APR | 0.001963 | 0.00158910 | 1.235 | 0.2169 |
| MAY | 0.001401 | 0.00136652 | 1.025 | 0.3056 |
| JUN | -0.004563 | 0.00175036 | -2.607 | 0.0093 |
| JUL | -0.004764 | 0.00210105 | -2.267 | 0.0236 |
| AUG | 0.006297 | 0.00166185 | 3.789 | 0.0002 |
| SEP | 0.001840 | 0.00184487 | 0.997 | 0.3188 |
| OCT | -0.004321 | 0.00178979 | -2.414 | 0.0159 |
| NOV | 0.001565 | 0.00169916 | 0.921 | 0.3571 |
| DEC | -0.001301 | 0.00144432 | -0.901 | 0.3678 |

**Table 14 Estimated regression coefficients and standard errors for the POSTHI group using ordinary least squares.**

| Variable | Parameter Estimate | Standard Error | T for H0: Parameter=0 | Prob > |
|---|---|---|---|---|
| INTERCEP | 0.000560 | 0.00117149 | 0.478 | 0.6325 |
| FEB | -0.002316 | 0.00209664 | -1.105 | 0.2695 |
| MAR | 0.000336 | 0.00178349 | 0.188 | 0.8506 |
| APR | 0.001310 | 0.00225691 | 0.580 | 0.5618 |
| MAY | -0.003138 | 0.00232459 | -1.350 | 0.1774 |
| JUN | -0.006247 | 0.00259760 | -2.405 | 0.0163 |
| JUL | 0.002365 | 0.00232459 | 1.017 | 0.3093 |
| AUG | 0.000603 | 0.00187682 | 0.321 | 0.7479 |
| SEP | -0.002155 | 0.00201493 | -1.070 | 0.2851 |
| OCT | 0.002535 | 0.00158877 | 1.595 | 0.1109 |
| NOV | 0.002452 | 0.00164808 | 1.488 | 0.1370 |
| DEC | -0.003109 | 0.00171257 | -1.815 | 0.0698 |

Test of First and Second Moment Specification
DF 11 Chisq Value 41.008 Prob>Chisq 0.000

Durbin-Watson D 1.967 No. of obs. 1094

**Table 15 Estimated regression coefficients and standard errors for the POSTHI group using weighted least squares.**

(Weighted Least Squares)

| Variable | Parameter Estimate | Standard Error | T for H0: Parameter=0 | Prob > |
|---|---|---|---|---|
| INTERCEP | 0.000560 | 0.00114667 | 0.489 | 0.6251 |
| FEB | -0.002316 | 0.00181996 | -1.273 | 0.2034 |
| MAR | 0.000336 | 0.00166044 | 0.202 | 0.8396 |
| APR | 0.001310 | 0.00215895 | 0.607 | 0.5442 |
| MAY | -0.003138 | 0.00220740 | -1.421 | 0.1555 |
| JUN | -0.006247 | 0.00277959 | -2.248 | 0.0248 |
| JUL | 0.002365 | 0.00268690 | 0.880 | 0.3790 |
| AUG | 0.000603 | 0.00195082 | 0.309 | 0.7572 |
| SEP | -0.002155 | 0.00205841 | -1.047 | 0.2954 |
| OCT | 0.002535 | 0.00160117 | 1.583 | 0.1137 |
| NOV | 0.002452 | 0.00153866 | 1.594 | 0.1113 |
| DEC | -0.003109 | 0.00166433 | -1.868 | 0.0621 |

The estimated regression coefficients from the PRELO, POSTHI, and POSTLO groups do not indicate the presence of a January effect. The PREHI group is the only group in which regression analysis reveals significant, positive January rates of growth.

Partial F-tests Between Groups

Partial F-tests between full and reduced regression models are used to statistically test the effect of the marked-to-market provisions and tax-loss selling potential on monthly rates of growth. Instead of monthly dummy variables as in the previous test, the dummy variables in these tests are:

PREHI - The pre-ERTA, high tax-loss selling potential group

PRELO - The pre-ERTA, low tax-loss selling potential group

POSTHI - The post-ERTA, high tax-loss selling potential group

POSTLO - The post-ERTA, low tax-loss selling potential group

The dependent variables are the rates of growth for the first four trading days of January. As in the previous regressions, only the highest and the lowest tax-loss-selling potential quartiles are used.

The tax-loss-selling hypothesis predicts that rates of growth for the PREHI group are higher than for the PRELO group, therefore the regression coefficient for the PREHI group should be significantly higher than the coefficient for PRELO.

A partial F-test between the following two regressions tests whether the PREHI mean rate of growth is significantly higher than the PRELO mean rate of growth:

Full $R_{it} = b_1 + b_2POSTHI + b_3PREHI + b_4PRELO + e_{it}$ (2)

Reduced $R_{it} = b_1 + b_2POSTHI + b_3(PREHI + PRELO) + e_{it}$ (3)

where $R_{it}$ is the rate of growth observation, $b_1$ is the constant term representing the effect of POSTLO, $b_2$ through $b_4$ are coefficients representing the difference in effect between POSTLO and the other groups, and $e_{it}$ is the error term for the observation. The reduced model assumes that there is no difference in the effect of PREHI and PRELO and therefore $b_3$ and $b_4$ are equal. The second alternative hypothesis states the expected relationship between the pre-ERTA high tax-loss-selling potential commodity futures and the pre-ERTA low tax-loss-selling potential commodity futures in terms of the regression coefficients.

> H2: Regression models (2) and (3) are statistically different, indicating that tax-loss selling potential in commodity futures contracts causes abnormal rates of growth in the first four trading days of January in the Pre-ERTA period.

The results of the partial F-test between regressions (2) and (3) are presented in Table 16 and Table 17. The mean rates of growth for the first four trading days of January between the highest and the lowest PTS quartiles are significantly different (p-value = 0.0101). The high tax-

**Table 16** Partial F-test of whether PREHI = PRELO using regression equations (2) and (3) and ordinary least squares.

Test of PREHI = PRELO
(Ordinary least squares)

| Variable | DF | Parameter Estimate | Standard Error | T for H0: Parameter=0 | Prob > |
|---|---|---|---|---|---|
| INTERCEP | 1 | 0.000560 | 0.00125680 | 0.446 | 0.6559 |
| POSTLO | 1 | -0.000528 | 0.00185351 | -0.285 | 0.7761 |
| PREHI | 1 | 0.004210 | 0.00192900 | 2.183 | 0.0296 |
| PRELO | 1 | -0.003077 | 0.00267347 | -1.151 | 0.2505 |

Test whether PREHI = PRELO

| | | | | | |
|---|---|---|---|---|---|
| Numerator: | 0.0015 | DF: | 1 | F value: | 6.8877 |
| Denominator: | 0.000223 | DF: | 401 | Prob>F: | 0.0090 |

loss-selling potential group (PREHI) has the higher mean rate of growth which is consistent with the tax-loss-selling hypothesis. The results shown of Tables 16 and 17 support H2.

**Table 17** Partial F-test of whether PREHI = PRELO using regression equations (2) and (3) and weighted least squares.

Test of PREHI = PRELO
(Weighted Least Squares)

| Variable | Parameter Estimate | Standard Error | T for H0: Parameter=0 | Prob > |
|---|---|---|---|---|
| INTERCEP | 0.000560 | 0.00118778 | 0.472 | 0.6373 |
| POSTLO | -0.000528 | 0.00172511 | -0.306 | 0.7599 |
| PREHI | 0.004210 | 0.00203071 | 2.073 | 0.0388 |
| PRELO | -0.003077 | 0.00244534 | -1.258 | 0.2090 |

| | | | | | |
|---|---|---|---|---|---|
| Numerator: | 0.1070 | DF: | 1 | F value: | 7.2921 |
| Denominator: | 0.014678 | DF: | 401 | Prob>F: | 0.0072 |

If the tax-loss-selling hypothesis is plausible, then the enactment of ERTA will eliminate incentive for tax-loss-selling in commodity futures and the January effect should disappear, even in the high tax-loss-selling potential group (POSTHI). The following two regressions test the difference in rate of growth means between the high tax-loss-selling groups before and after the enactment of ERTA:

Full $\quad R_{it} = b_1 + b_2PREHI + b_3POSTHI + b_4POSTLO + e_{it}$ (4)

Reduced $\quad R_{it} = b_1 + b_2(PREHI + POSTHI) + b_4POSTLO + e_{it}$ (5)

where $R_{it}$ is the rate of growth observation, $b_1$ is the constant term representing the effect of PRELO, $b_2$ through $b_4$ are coefficients representing the difference in effect between PRELO and the other groups, and $e_{it}$ is the error term for the observation. The reduced model assumes that there is no difference in the effect of PREHI and POSTHI and therefore $b_2$ and $b_3$ are equal. The third alternative hypothesis states the expected relationship between the high tax-loss-selling potential commodity futures before ERTA and the high tax-loss-selling commodity futures after ERTA in terms of the regression coefficients.

> H3: Regression models (4) and (5) are statistically different, indicating that the ERTA marked-to-market rules removed the incentive for tax-loss selling in commodity futures contracts after 1981.

The results of the partial F-test between regression models (4) and (5) are presented on Tables 18 and 19. The mean rates of growth for the first four days of January for

Table 18 Partial F-test of whether PREHI = POSTHI using regression equations (4) and (5) and ordinary least squares.

Test of PREHI = POSTHI
(Ordinary least squares)

| Variable | DF | Parameter Estimate | Standard Error | T for H0: Parameter=0 | Prob > |
|---|---|---|---|---|---|
| INTERCEP | 1 | -0.002516 | 0.00235964 | -1.066 | 0.2869 |
| PREHI | 1 | 0.007287 | 0.00277658 | 2.624 | 0.0090 |
| POSTHI | 1 | 0.003077 | 0.00267347 | 1.151 | 0.2505 |
| POSTLO | 1 | 0.002549 | 0.00272468 | 0.936 | 0.3500 |

Test whether PREHI = POSTHI

| | | | | | |
|---|---|---|---|---|---|
| Numerator: | 0.0011 | DF: | 1 | F value: | 4.7637 |
| Denominator: | 0.000223 | DF: | 401 | Prob>F: | 0.0296 |

high tax-loss-selling commodity futures is significantly higher before the enactment of ERTA (PREHI group) than after the enactment (POSTHI group). This indicates that the tax law change affected rates of growth and provides support for H3.

Table 19 Partial F-test of whether PREHI = POSTHI using regression equations (4) and (5) and weighted least squares.

Test of PREHI = POSTHI
(Weighted Least Squares)

| Variable | Parameter Estimate | Standard Error | T for H0: Parameter=0 | Prob > |
|---|---|---|---|---|
| INTERCEP | -0.002516 | 0.00213749 | -1.177 | 0.2398 |
| PREHI | 0.007287 | 0.00269849 | 2.700 | 0.0072 |
| POSTHI | 0.003077 | 0.00244534 | 1.258 | 0.2090 |
| POSTLO | 0.002549 | 0.00247670 | 1.029 | 0.3040 |

| | | | | | |
|---|---|---|---|---|---|
| Numerator: | 0.0631 | DF: | 1 | F value: | 4.2984 |
| Denominator: | 0.014678 | DF: | 401 | Prob>F: | 0.0388 |

If the tax-loss-selling hypothesis is plausible, then after the enactment of ERTA there should be no tax-loss-selling and January rates of growth should not be significantly different between high and low tax-loss-selling potential groups. The following two regressions compare January rates of growth between the test group (POSTHI) and the comparison group (POSTLO) in the post ERTA period:

Full $\quad R_{it} = b_1 + b_2PREHI + b_3POSTHI + b_4POSTLO + e_{it}$ (6)

Reduced $\quad R_{it} = b_1 + b_2PREHI + b_3(POSTHI + POSTLO) + e_{it}$ (7)

where $R_{it}$ is the rate of growth observation, $b_1$ is the constant term representing the effect of PRELO, $b_2$ through $b_4$ are coefficients representing the difference in effect between PRELO and the other groups, and $e_{it}$ is the error term for the observation. The reduced model assumes that there is no difference in the effect of POSTHI and POSTLO and therefore $b_3$ and $b_4$ are equal. The fourth alternative hypothesis states the expected relationship between the post-ERTA low tax-loss-selling commodity futures and the post-ERTA high tax-loss-selling commodity futures in terms of the regression coefficients.

> H4: Regression models (6) and (7) are not statistically different, indicating that it was the effect of the ERTA marked-to-market rules, rather than an extraneous variable, that removed the incentive for tax-loss selling in commodity futures contracts after 1981.

The results of the partial F-test between regression models (6) and (7) are presented on Tables 20 and 21. In the

Table 20 Partial F-test of whether POSTHI = POSTLO using regression equations (6) and (7) and ordinary least squares.

Test of POSTHI = POSTLO
(Ordinary least squares)

| Variable | DF | Parameter Estimate | Standard Error | T for H0: Parameter=0 | Prob > |
|---|---|---|---|---|---|
| INTERCEP | 1 | -0.002516 | 0.00235964 | -1.066 | 0.2869 |
| PREHI | 1 | 0.007287 | 0.00277658 | 2.624 | 0.0090 |
| POSTHI | 1 | 0.003077 | 0.00267347 | 1.151 | 0.2505 |
| POSTLO | 1 | 0.002549 | 0.00272468 | 0.936 | 0.3500 |

Test whether POSTHI = POSTLO

| | | | | | |
|---|---|---|---|---|---|
| Numerator: | 0.0000 | DF: 1 | | F value: | 0.0810 |
| Denominator: | 0.000223 | DF: 401 | | Prob>F: | 0.7761 |

Table 21 Partial F-test of whether POSTHI = POSTLO using regression equations (6) and (7) and weighted least squares.

Test of POSTHI = POSTLO
(Weighted Least Squares)

| Variable | Parameter Estimate | Standard Error | T for H0: Parameter=0 | Prob > |
|---|---|---|---|---|
| INTERCEP | -0.002516 | 0.00213749 | -1.177 | 0.2398 |
| PREHI | 0.007287 | 0.00269849 | 2.700 | 0.0072 |
| POSTHI | 0.003077 | 0.00244534 | 1.258 | 0.2090 |
| POSTLO | 0.002549 | 0.00247670 | 1.029 | 0.3040 |

| | | | | |
|---|---|---|---|---|
| Numerator: | 0.0014 | DF: 1 | F value: | 0.0935 |
| Denominator: | 0.014678 | DF: 401 | Prob>F: | 0.7599 |

post-ERTA period, the mean rates of growth for the first four days of January are not significantly different between commodities in the high tax-loss-selling potential and in the low tax-loss-selling potential. Since the high tax-loss-selling potential commodity futures had significantly higher January rates of growth than the low tax-loss-selling potential commodity futures in the pre-ERTA period, but not significantly higher rates of growth in the post-ERTA period, there is support for taxes as an explanation for the change in rate of growth.

Test for Mid-month Anomalies

Rates of growth in the mid-month period of each month are calculated to determine if rates of growth become negative during this period in months when tax payments are due. Mean rates of growth for the five day period from the 11th to the 15th of each month are calculated using all observations from September, 1978 through January, 1985. these means are presented in Table 22.

The means shown in Table 22 are graphed on Figure 13. An inspection of the rates of growth indicates a pattern of negative rates in April. Regression analysis is used to statistically test for the pattern. Rate of growth is regressed on monthly dummy variables as in equation (1) using a sample consisting of rate of growth observations during the period of the 11th through the 15th of each month. All months from September, 1978 through January, 1985 are combined for this regression. The regression coefficient $b_4$ (April) is expected to be negative in comparison to the other coefficients. The fifth alternative hypothesis states the expected relationship between mid-April rates of growth and rates of growth in other months in terms of regression coefficients.

H5: April rates of growth during the mid-month period will be negative compared to the rates of growth during the same period in all other months. This will be indicated by a significantly negative April coefficient compared to the coefficients for all other months.

Table 22 Mean rates of growth by month for the 11th through the 15th of each month over the period from September, 1978 through January, 1985.

Growth Rate Means: ALL YEARS, Days 11-15

| Month | Rate of Growth | Standard error |
|---|---|---|
| Jan | 0.0027073 | 0.014974 |
| Feb | -.0008538 | 0.012175 |
| Mar | -.0017564 | 0.012894 |
| Apr | -.0031382 | 0.011903 |
| May | -.0016674 | 0.011518 |
| Jun | 0.0008275 | 0.014056 |
| Jul | -.0002190 | 0.017552 |
| Aug | 0.0025033 | 0.017090 |
| Sep | -.0005332 | 0.015737 |
| Oct | 0.0007240 | 0.016332 |
| Nov | -.0003152 | 0.013851 |
| Dec | -.0014813 | 0.015380 |

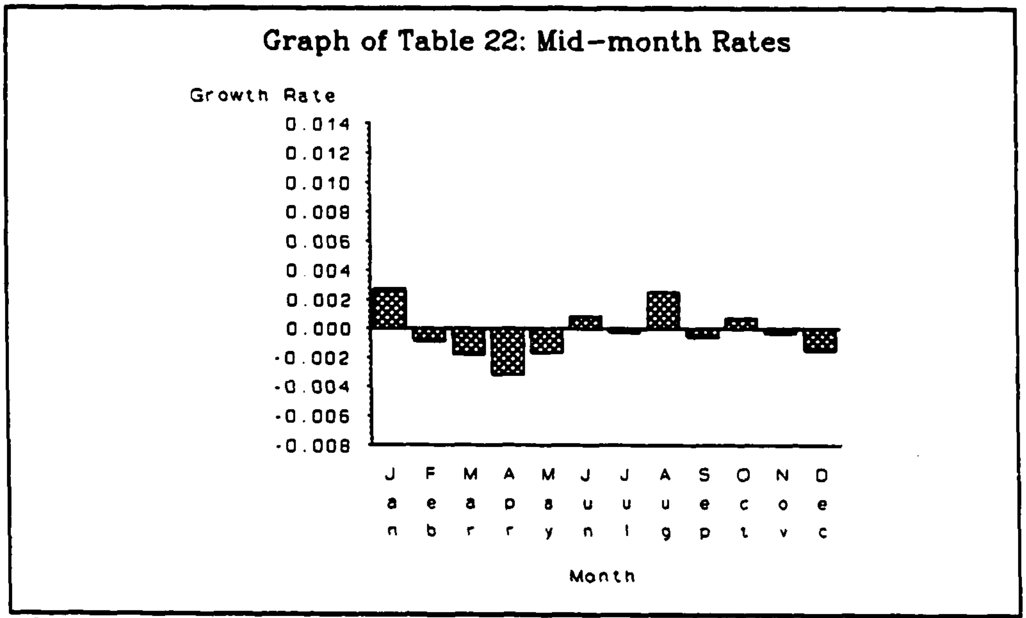


**Figure 13** **Graph of mid-month rates of growth from Table 22.**

**Table 23** Estimated regression coefficients for the regression of mid-month rates of growth on monthly dummy variables using ordinary least squares.

(Ordinary least squares)

| Variable | Parameter Estimate | Standard Error | T for H0: Parameter=0 | Prob > |
|---|---|---|---|---|
| INTERCEP | 0.002516 | 0.00051416 | 4.894 | 0.0001 |
| FEB | -0.003413 | 0.00082899 | -4.117 | 0.0001 |
| MAR | -0.003542 | 0.00078512 | -4.511 | 0.0001 |
| APR | -0.004911 | 0.00090319 | -5.437 | 0.0001 |
| MAY | -0.004521 | 0.00082614 | -5.473 | 0.0001 |
| JUN | -0.001162 | 0.00087551 | -1.327 | 0.1844 |
| JUL | -0.002720 | 0.00093052 | -2.923 | 0.0035 |
| AUG | -0.000404 | 0.00080802 | -0.500 | 0.6168 |
| SEP | -0.003972 | 0.00084947 | -4.676 | 0.0001 |
| OCT | -0.001446 | 0.00077307 | -1.871 | 0.0614 |
| NOV | -0.001404 | 0.00079086 | -1.775 | 0.0759 |
| DEC | -0.003081 | 0.00076776 | -4.012 | 0.0001 |

Test of First and Second Moment Specification
DF 11 Chisq Value: 185.763 Prob>Chisq: 0.0000

Durbin-Watson D 2.010 No. of Obs. 6851

**Table 24** Estimated regression coefficients for the regression of mid-month rates of growth on monthly dummy variables using weighted least squares.

(Weighted Least Squares)

| Variable | Parameter Estimate | Standard Error | T for H0: Parameter=0 | Prob > |
|---|---|---|---|---|
| INTERCEP | 0.002516 | 0.00052088 | 4.831 | 0.0001 |
| FEB | -0.003413 | 0.00077774 | -4.389 | 0.0001 |
| MAR | -0.003542 | 0.00075680 | -4.680 | 0.0001 |
| APR | -0.004911 | 0.00082702 | -5.938 | 0.0001 |
| MAY | -0.004521 | 0.00078430 | -5.764 | 0.0001 |
| JUN | -0.001162 | 0.00088162 | -1.318 | 0.1874 |
| JUL | -0.002720 | 0.00097652 | -2.785 | 0.0054 |
| AUG | -0.000404 | 0.00084943 | -0.476 | 0.6341 |
| SEP | -0.003972 | 0.00088482 | -4.489 | 0.0001 |
| OCT | -0.001446 | 0.00080672 | -1.793 | 0.0730 |
| NOV | -0.001404 | 0.00078068 | -1.798 | 0.0722 |
| DEC | -0.003081 | 0.00077166 | -3.992 | 0.0001 |

The estimated coefficients from the regression are presented on Table 23 and 24. The regressions show the presence of negative mid-month rates of growth for all months except January. Mid-April rates of growth do not appear to be significantly lower than any month other than January. However, even though non-significant, mid-April rates of growth appear to have a negative trend.

Econometric Issues

Autocorrelation and heteroskedasticity. Ordinary least squares (OLS) coefficient estimates are unbiased and consistent when the disturbance terms are autoregressive, however they are not efficient [Kmenta 1986, 308]. In the presence of autocorrelation, OLS estimators do not have the smallest variance among all unbiased estimators and correct inferences regarding the estimators cannot be made. The same is true in the presence of heteroskedasticity [Kmenta 1986, 275].

Autocorrelation and heteroskedasticity have been noted in prior regression studies of the January effect. For example, Keim [1983] notes a first order autoregressive scheme (AR1) in stock return residuals but does not adjust for the autocorrelation. He estimates that the results would remain significant even if the regression equation was calculated accounting for the autocorrelation. He also finds heteroskedasticity and mitigates the effect by using weighted least squares (WLS). Wilson and Jones [1990] find evidence of autocorrelation and heteroskedasticity in regression residuals when examining bond and commercial paper interest rate January effects. They estimate regression parameters using OLS but estimate the variance-covariance matrix using a method suggested by Gallant [1987].

In this research, the Durbin-Watson test is used to test for autocorrelation in each regression. The Durbin-Watson D results are shown for each regression on the tables presenting the ordinary least square estimated parameters.

White's test is used to test for heteroskedasticity in each regression. The Chi-square value and the p-value of each test is also shown for each regression on the tables presenting the ordinary least square estimated parameters.

The Durbin-Watson test reveals either insignificant or indeterminate autocorrelation in the regressions of the PRELO, PREHI, POSTLO, and POSTHI groups using monthly dummy variables. Significant first order autocorrelation is present in the full model used for the three partial regressions (-0.232), however it is a low level of autocorrelation. The presence of the January effect in the PREHI group and its absence in the PRELO, POSTLO, and POSTHI groups support the conclusions drawn from the partial regressions. Therefore, the conclusions drawn from the partial regressions are not expected to change because of the low level autocorrelation.

White's test indicates significant heteroskedasticity in all regressions. Weighted least squares is used to estimate the standard errors for parameter estimates in all regressions as well as the ordinary least squares estimates. The weighted least squares estimates are presented subsequent to the ordinary least squares regression estimates. In weighted least squares, the variance of the regression parameters are estimated using different weights for different independent variable values. The weights used in this study are the reciprocals of the standard errors, $\sigma_i$, estimated as $s_i$. The estimated variance $s_i^2$ is calculated using the following equations as suggested by Kmenta [1986, 291]:

$$s_i^2=\sum_{j=1}^{n_i} \frac{(Y_{ij}-\overline{Y}_i)^2}{(n_i-1)} \quad where$$
$$\overline{Y}_i=\frac{1}{n_i}\sum_{j=1}^{n_i} Y_{ij}$$

As the tabled results show, the conclusions are not changed by the weighted least squares regressions.

Multicollinearity. Multicollinearity is not mentioned as a problem in prior regression studies of the January effect using monthly dummy variables. Multicollinearity can bias the study against rejection of the null hypothesis by inflating the standard errors of the regression coefficients. In this research, variance inflation factors are calculated in each regression as a measure of multicollinearity. The variance inflation factor did not exceed 2.0 in any of the regressions. Therefore, multicollinearity is not considered to be a problem.

Omitted Variable Bias. Omitted variable bias is not mentioned as a problem in prior regression studies of the January effect using monthly dummy variables. If an independent variable is omitted from the regression equation that is correlated to both the dependent variable and an included independent variable, then the coefficient of the included independent variable will be biased. An omitted variable which is correlated both with rate of growth and the monthly dummy variables is not expected.

Self-selection bias. Self-selection bias is possible in non-experimental research because subjects cannot be randomly assigned to groups. Self-selection occurs when subjects are assigned to their groups partly because of unrecognized traits they possess that are related to variables in the research problem [Kerlinger 1986, 349]. These traits are hidden from the researcher, but cause subjects to "select themselves" into groups in ways that are correlated with the independent variable under examination.

In this study, there is a danger that commodity contracts are self-selected into PTS quartiles based upon some common factor that makes their prices move together and that this factor is correlated with a monthly pattern of returns.[1]

Self-selection is avoided in this research by using contracts for 14 different agricultural products. Many of

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[1] For example, suppose that corn and wheat were the only contracts that exhibited positive January rates of growth in the last 20 years, and that this was due to an unanticipated increase in January demand for corn and wheat in Russia each year until 1981. Also assume that corn and wheat contracts were self-selected into the pre-ERTA high tax-loss selling potential groups in January each year because they are both only grown in Kansas, and unanticipated good weather occurred in Kansas each year until 1981 creating an over-supply of corn and wheat along with a subsequent drop in price for those contracts over the year. The self-selection would lead to the erroneous conclusion that tax-loss selling caused the positive January rates of growth.

these products are substantially different such as corn and soybeans, cattle and oats, hogs and orange juice. It is expected that any unknown factors which influence both price activity through the year and the January pattern of growth rate will be randomized between different groups due to the differences in the products.

# CHAPTER 6

## ANALYSIS OF RESULTS

### Mean Rates of Growth by Month

The mean rates of growth by month presented in Table 2 through Table 7 demonstrate that the January effect is evident in commodity futures only in the first few days of January. Although Keim [1983] finds evidence of positive abnormal returns in small stocks throughout January, this evidence could not be found in commodity futures. Commodity futures, however, are more volatile than stocks. This volatility may cause January positive price trends in commodity futures to dissipate more rapidly than in stocks. The mean rates of growth presented in Table 2 through Table 7 are useful mainly in identifying the first few days of January as the time period in which the January effect is concentrated in commodity futures. Even in Table 7, however, where the sample is limited to the first four days of each month and to the highest and lowest tax-loss-selling potential quartiles, the January mean rate of growth in the PREHI group is not significantly different than zero. A significantly positive rate of growth is not required to show evidence of the January effect, only significantly higher rates in January compared to other months. The mean rates of growth by month presented in Table 7 show a trend in this direction but cannot provide statistical evidence of the trend. The regression analysis presented in Table 8 through Table 15 provides statistical evidence of the January effect.

Regression Analysis

The regression parameters presented in Table 10 and Table 11 clearly demonstrate a significant January effect in the PREHI group. The intercept parameter representing January is +0.004771 with a standard error of 0.001900 using ordinary least squares and 0.001886 using weighted least squares. Both methods indicate that the probability the results are due to chance is less than 2%. The remaining parameters are either negative or insignificantly positive. Although the July parameter estimate is a higher positive value than the intercept parameter (0.006555), it is not significantly different from zero.

The intercept (January) is significantly positive compared to the remaining coefficients when using the pre-ERTA, high tax-loss-selling potential group (PREHI), supports H1 and is consistent with the tax-loss-selling hypothesis. The tax-loss-selling hypothesis predicts that a January effect will be found in commodity futures before the enactment of ERTA in commodities that have experienced a declining price trend.

Regressions using the three remaining groups, PRELO, POSTHI, and POSTLO do not reveal a significant January effect under ordinary least squares or weighted least squares. The pre-ERTA group with low tax-loss-selling potential (PRELO) yields an insignificantly negative intercept coefficient

(-.002516) as presented in Table 8 and Table 9. The post-ERTA group with high tax-loss-selling potential (POSTHI) yields an insignificantly positive intercept coefficient (0.000560) as presented in Table 14 and Table 15. The post-ERTA group with low tax-loss-selling potential (POSTLO) yields an insignificantly positive intercept coefficient (0.000033) as presented in Table 12 and Table 13.

Finding a significant January effect in the PREHI group, but not the PRELO, POSTHI, or POSTLO groups provides support for the tax-loss-selling hypothesis. However partial F-tests comparing the January means between these groups are needed to provide stronger statistical support for the hypothesis.

Partial F-tests

Table 16 and Table 17 present results of the partial F-test comparing the mean January rate of growth in the PREHI group with the mean January rate of growth in the PRELO group. The partial F-test reveals that the PREHI group has a significantly higher January rate of growth compared to the PRELO group using both ordinary least squares (p-value = 0.0090) and weighted least squares (p-value = 0.0072). This test supports H2 and is consistent with the tax-loss-selling hypothesis. The tax-loss-selling hypothesis predicts that commodity futures in the pre-ERTA period with high tax-loss-selling potential (PREHI group) will have significantly higher rates of growth than commodity futures in the pre-ERTA period with low tax-loss-selling potential (PRELO group), and the results support this prediction.

Table 18 and Table 19 present the results of the partial

F-test comparing the mean January rate of growth in the PREHI group with the mean January rate of growth in the POSTHI group. The partial F-test reveals that the PREHI group has a significantly higher January rate of growth compared to the POSTHI group using both ordinary least squares (p-value = 0.0296) and weighted least squares (p-value = 0.0388). This test supports H3 and is consistent with the tax-loss-selling hypothesis. The tax-loss-selling hypothesis predicts the commodity futures in the pre-ERTA period with high tax-loss-selling potential (PREHI group) will have significantly higher rates of growth than commodity futures in the post-ERTA period with high tax-loss-selling potential (POSTHI group). The results presented on Table 18 and Table 19 indicate that the enactment of the marked-to-market provisions in ERTA and the removal of the incentive for tax-loss-selling in commodity futures after June, 1981 caused the January effect to disappear.

Table 20 and Table 21 present results of the partial F-test comparing the mean January rate of growth in the POSTHI group with the mean January rate of growth in the POSTLO group. The partial F-test reveals that the POSTHI group is not significantly different than the POSTLO group using either ordinary least squares (p-value = 0.7761) or weighted least squares (0.7599). This test supports H4 and is consistent with the tax-loss-selling hypothesis. The tax-loss-selling hypothesis predicts that commodity futures in the post-ERTA period with high tax-loss-selling potential (POSTHI group) will not have significantly higher rates of growth than commodity futures in the post-ERTA period with low tax-loss-selling potential (POSTLO group). Because the enactment of the marked-to-market provisions removed the incentive for tax-loss-selling, there is no reason after June, 1981 for the high tax-loss-selling group to have greater rates of growth than the low tax-loss-selling group. Comparison of the POSTHI group to the POSTLO group helps to rule out other causes besides the tax law change as causing the disappearance of the January effect in the post-ERTA period.

Mid-month Rate of Growth Analysis

The mean rates of growth calculated from the 11th through the 15th of each month presented on Table 22 indicate a pattern of negative rates of growth in April as compared to other months. Although none of the rates are significantly different from zero, the results suggest that mid-April rates of growth may be significantly lower than other months. Regression analysis is used to test this difference.

Regression parameters presented on Table 23 and Table 24 do not indicate that the mid-April rate of growth is significantly more negative than the other months other than January. In fact, all mid-month rates of growth are negative in comparison to January.

Because there is an indication that mid-April rates of growth are more negative than other mid-month rates of growth (although not significantly so in this study), it is possible that a significant pattern of mid-April negative rates of growth may appear in samples of other securities over different time periods. A possible extension of this study would be to analyze other securities for this pattern.

# CHAPTER 7

## SUMMARY

### Discussion

The results of this study provide support for the tax-loss-selling hypothesis. The research extends prior studies of the January effect by examining a sample of individual commodity futures contracts in a pre-test, post-test comparison group design built around the 1981 ERTA provisions. Prior studies have alluded to the possibility that ERTA may have an influence on the January effect, but have not specifically tested the influence of ERTA. This study uses the opportunity provided by ERTA to examine the influence of tax motivation on the January effect and finds statistical support for the tax-loss-selling hypothesis.

These findings imply market inefficiency since there is evidence that the market did not fully impound information about the yearly trend of tax-loss-selling when pricing commodity futures in the pre-ERTA period. If the commodity futures market is efficient, investors would be expected to anticipate the positive increase in prices at the first of the year and eliminate the increase through competitive bidding. A possible alternative explanation is that transaction costs prevented the market from eliminating the January seasonal. This study does not evaluate the influence of transaction costs.

The findings of this study imply that tax law changes directly affect market behavior. Specifically, the findings indicate that the ERTA marked-to-market provisions eliminated or greatly reduced tax-loss-selling in commodity futures contracts. The congressional purpose of the marked-to-market provisions was to eliminate tax motivated straddle transactions with commodity futures. Congress wished to eliminate transactions in which investors would sell the loss leg of a commodity future straddle at year-end and sell the gain leg the straddle in the following year. The results of this study indicate that tax motivated loss selling abated after the enactment of ERTA.

## Limitations

The study is limited by the effect of commodity futures contract carrying cost changes during the rate of growth calculation period. For example, part of the cost of a futures contract is the carrying cost of the underlying commodity such as storage costs and insurance costs. These costs decrease as the contract gets closer to maturity. With all other factors remaining constant, the decrease in carrying costs should cause a steady price decline and slightly decreasing rates of growth. The decreasing rates of growth caused by carrying cost changes interfere with detection of the January effect.

Since rates of growth are calculated as the percentage change in price from day to day, the time period for the carrying cost effect is relatively short. The carrying cost effect should not be correlated with high or low PTS or with any particular month so the effect should fall equally on all observations. Therefore, carrying costs are not expected to significantly affect the conclusions of the study.

Future Research

This study could be extended to examine other tax-related anomalies. Specifically, the trend of negative rates of growth observed in mid-April might be statistically verified using other securities or using other time periods.

The day-of-the-week anomaly, noted in the Pre-ERTA period in commodity futures by Gay and Kim [1987], has been observed to disappear in the post-ERTA period. A future study may seek establish a link between tax law changes and the day-of-the-week anomaly.

As mentioned in the introduction, some money managers have suggested recently that tax-loss-selling is responsible for a new pattern of abnormally high November returns in stocks [Wall Street Journal 11/19/90]. The managers conjecture that tax law changes brought about by the 1986 Tax Act affecting mutual funds created the conditions for the November seasonal. This hypothesis could be tested by comparing rates of return in high tax-loss-potential stocks in November to rates of return in other months. The comparison could be made before and after the tax law change and also with a comparison group of low tax-loss-selling potential stocks.

Finally, a further extension of this study could examine whether continuing trading patterns such as the November seasonal in stocks can be profitably traded when considering transaction costs.

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