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ANOMALIES AND EQUILIBRIUM RETURNS
IN A SMALL STOCK MARKET

by

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Abstract

Size and turn of the year related stock pricing anomalies are documented on the Helsinki Stock Exchange, Finland. Further inspection reveals that the magnitude of the size premium is roughly the same as the turn of the year seasonal, and that March through December returns are completely free of size premiums. A tax loss selling hypothesis on the formation of January returns, incorporating the institutional restrictions that capital losses can only be deducted from capital gains, cannot be rejected. While the January seasonal may be explained in terms tax-loss selling and transaction costs, the fact that its expectation is not arbitraged away cannot. It is argued that persistent long-run excess risk adjusted returns to small stocks must reflect an asset pricing model misspecification, specifically the omission of higher than second order moments from the CAPM. A four-moment parameter preference model combined with a proxy for estimation risk is found to reduce residual size premiums to less than half of those reported for the CAPM.

1. Introduction

The empirical evidence that excess returns can be earned over time by ranking securities over certain variables by now seems almost exhaustive. Banz (1981) and Reinganum (1981) detect a size effect, Basu (1977) and Reinganum (1981) report on a price-earnings ratio effect, while Barry and Brown (1983) uncover a period-of-listing effect. Simultaneously, the number of papers reporting seasonality in stock returns has grown rapidly with Branch (1977), Reinganum (1983), Roll (1983), Givoly and Ovadia (1983), Keim (1983), Brown, Keim, Kleidon, and Marsh (1983), and Tinic and Barone-Adesi (1983). Several of these papers also report links between such anomalies. Reinganum (1983) and Keim (1983) find that the positive January seasonal is more pronounced in smaller firms; up to half of the total firm size effect may be attributed to the month of January.[1] Similarly, Basu (1983) documents linkage between the P/E and the size effects.

Most of this work is based on American data. The only non-American studies mentioned above, Brown, Keim, Kleidon and Marsh (1983) and Tinic and Barone-Adesi (1983), both analyze the tax-loss selling hypothesis on the turn-of-the-year effect under tax regimes different from the American.

Although the quality of non-American data in terms of length of the available time-series and the number of securities covered mostly is vastly inferior to American, closer analysis of such fresh data sets seems warranted for a number of reasons. Most important among these, at least where several European exchanges are concerned, is the fact that the data sets are free from some commonly referred to shortcomings of the American data, such as the problems associated with the bid-asked spread and its treatment in the CRSP data base. Since several European exchanges like the Helsinki exchange studied in this paper are silent, with separate price fixing and free trading phases, prices recorded will always be actual trading, rather than bid or asked prices.[2] Typically, one and, at most, two such trading prices will be recorded. Virtually all trading will take place at these prices.[3]

There are, however, at least two additional reasons for turning to exotic data sets for answers regarding anomalies. The first, the differences in tax systems and other institutional features encountered, actually appears to have been the motivation for both Brown, Keim, Kleidon and Marsh (1983) and Tinic and Barone-Adesi (1983). Ideally, institutional differences, such as in the Australian and Canadian cases, permit more powerful testing, especially of tax related and other institutional hypotheses.

Furthermore, many if not most explanations to both size and calendar related anomalies rely on minor market or data imperfections such as transaction costs (Stoll and Whaley (1983), Schultz (1983)), thin trading (Roll (1981), Reinganum (1982)) or deficient information (Barry and Brown (1983)). In an international comparison such imperfections are not prominently featured on the NYSE or even the AMEX. Rather, they seem to decrease in magnitude with the size of the exchange (Berglund-Wahlroos-Ornmark (1983)). The standard explanation to the failures of the CAPM and some other simple equilibrium models in explaining European data (e.g., Hawawini, Michel and Viallet (1983)), is that the data is infected by numerous imperfections and that market volumes are generally too low to permit arbitrage arguments from working. If that is the case, European exchanges--smaller ones in particular--should be ideal for studies of pricing anomalies.

The present paper constitutes an analysis of size, calendar, and information related anomalies and their interdependence in Finnish security prices. More specifically, we start out by screening the data for size and January effects, the latter of which have been discussed in one previous paper also.[4] This paper then goes on to show that most of the size premium occurs in the month of January and that the remaining part, apparently because of extremely thin trading in very small stocks, occurs in early February. The remaining ten months are completely free from a size effect.

The January seasonal is then shown to be produced by market behavior that closely resembles tax loss selling. Small firm stock, it turns out, is more likely to be subject to such selling pressure due to higher return variability. The fact that the size premium may be tax induced and occurs only in January-February returns does not, however, resolve the long run equilibrium stock price issues contained in the debate. To the extent these turn-of-the-year premia are foreseeable, they should be eliminated by the arbitrage process. The remaining part of the paper, therefore, addresses the question of equilibrium pricing of securities in small, taxed stock markets. It is argued that because of the seasonality and informational as well as transactional imperfections, stock return distributions in small markets will tend to be strongly non-normal. In the presence of investors possessing utility functions of higher order then the quadratic, higher moments or co-moments of such return distributions will be priced. Such a pricing theory, expecially when combined with an effort to account for parameter estimation risk, turns out to fit the data reasonably well, yielding correct signs and mostly significant market prices for third and fourth moment risk. However, prices will still contain a (statistically insignificant) premium on the stocks' capitalization. The market hence does not appear to fully foresee the January price shocks in the long run. Removal of return observations potentially tainted by tax loss selling improves regression estimates slightly.

The paper is organized as follows: in section 2 the size and the turn-of-the-year effects and their interdependence is reported upon. Section 3 deals with the tax loss selling explanation to the January seasonal. In section 4, a multimoment pricing model is surveyed and tested.

2. Some Anomalous Regularities in the Data

Capitalizations and Expected Returns.

A total of 66 stocks were traded on the Helsinki Stock Exchange on December 3, 1981. During the 12-year period under study, starting January 1, 1970, the number of securities traded varied between 53 and 66. Only three securities were removed during that period, none because of bankruptcy.[5] This study covers the 50 stocks quoted throughout the period. Holding period returns are computed by taking logarithmic first differences of each security's index points in the WI-index. Mean returns are hence consistently computed using buy and hold estimators.

The WI-index measures the return to a position in a given stock under the assumption that dividends are reinvested. Details concerning the treatment of splits, stock dividends, and issues are given in Berglund-Wahlroos-Grandell (1983).

In Table 1 the 50 stocks are grouped into ten portfolios in decreasing order of their market capitalization on December 31, 1969.[6] For each of the portfolios mean capitalizations and returns over the 12-year period are computed. Mean returns are also reported for three four-year subperiods. Three separate sets of portfolio betas are then computed using the value-weighted WI-market index to measure returns to the market portfolio. Ordinary stock betas are estimated both from monthly and from daily return data. In line with Roll's (1981) argument and subsequent practice, Dimson (1979) betas with five leads and 20 lags are also estimated:

$$(1) \qquad \beta_j^D = \sum_{t=-5}^{20} \beta_{jt},$$

where

$$R_{jT} = \alpha_j + \sum_{t=T-5}^{T+20} \beta_{jt} R_{Mt},$$

and where β_j^D is the Dimson beta for stock j, R_{jt} and R_{Mt} are the returns to stock j and the market portfolio, respectively, and where α_j and β_{jt} are regression coefficients.

In computing the excess returns to each portfolio the mean return government bonds with a mean maturity of three years was used as he risk free rate of return[7] and the mean return to the WI-index as the market return. For the HESE it turns out that the difference between ordinary betas estimated from monthly data and Dimson betas is minute. Excess returns have, nevertheless, been computed using Dimson betas.

INSERT TABLE 1 HERE

The presence of a (long run) small firm effect in the data is evident. Mean excess returns decrease monotonically with stock capitalizations with just one exception--portfolio 7.[8] This effect appears to be a long, rather than a short term phenomenon in that its presence in the return data for the full 12-year period is more obvious than in any of the 4-year subperiods. The magnitude of the small firm effect, while apparently somewhat smaller than the corresponding effect uncovered by, for example, Reinganum (1981 and 1982), is still significant in magnitude: a spread of more than ten percent in annual excess returns on the portfolios is recorded.

Rather than running a "pooled" regression (Reinganum (1982)), a pure cross sectional regression is used to test the significance of the size effect. The reason for this is that a pooled regression is highly volatile to departures from normality in stock return distributions.[9] The cross section selected covers individual stocks rather than portfolios. This design is chosen because the number of securities does not permit a Kraus-Litzenberger (1976) or Barry-Brown (1983) partitioning based on several stock characteristics such as higher moment risk, size and/or information. To select portfolios based on just one characteristic--second moment systematic risk (beta) as in Fama-Macbeth (1973) would, in turn, bias a test in favor of mean variance related hypotheses since, independently of underlying stock returns, portfolio returns would tend to converge to normality.

To avoid problems with non-normally distributed prediction errors, all preferred stocks of firms whose common stock is also listed were removed.[10] Average annual (1970-81) returns ($\bar{R}_j$) to the remaining 42 stocks were then regressed on betas (β_{ij}) and capitalizations (S_j)

$$(2) \qquad \bar{R}_j = \gamma_0 + \gamma_1\beta_{ij} + \gamma_2 + S_j + \mu_j.$$

Table 2 reports on regression coefficients using both monthly and Dimson betas. Average returns ($\bar{R}_j$) are measured both by logarithmic first differences and, to reduce sampling dispersion assuming that the underlying distributions are symmetric stable with a characteristic exponent $1 < \alpha < 2$, by (.28 - .72) truncated means. While the explanatory power of the betas is negligible, the relationship between mean returns and capitalization is strongly negative. The size discount elasticity equals roughly one and five-tenths of one percent, and is significant on the one percent level. Its presence is slightly more evident in the truncated than in the ordinary means. In a linear model, the monthly size discount equals roughly $.3 \times 10^{-4}$ on each million of stock value and is also significant on the one percent level.

INSERT TABLE 2 HERE

The Seasonality of Stock Returns

Table 3 lists monthly portfolio returns in excess of the risk free rate and a Dimson risk premium[11] for each calendar month over the 1970-81 period for the ten market value portfolios defined earlier. The table also contains monthly excess returns to the value-weighted market index and beta estimates together with mean annual returns over the whole period for each portfolio.

INSERT TABLE 3 HERE

A fairly strong January effect averaging 2.5 percent is reported. Positive excess returns to the market in the neighborhood of 1.6 percent are also reported for July and December. The January returns are linked to firm capitalizations: large firm portfolios 1-4 earn an average excess January reurn of 1.5 to 2.5 percent, while small firm portfolios 5-10 all exhibit average excess January returns exceeding 3 percent. The coefficient of determination between mean excess January returns and mean portfolio capitalization is .3 and the t-statistic for the negative slope is -2.25.

Tests for significance are performed by regressing monthly returns in excess of portfolio means on monthly dummies while suppressing the intercept. By the deduction of portfolio means--i.e., the non-seasonal return component from monthly return data before regressing, we attempt falsification of the hypothesis that the return data contain no seasonality.[12]

INSERT TABLE 4 HERE

The results of Table 4 indicate that significant seasonality, particularly in the form of a strongly positive January seasonal, is contained in the data. The January seasonal in the market index averages 3.2 percent and is significantly different from zero on the .5 percent level. The seasonal is also significant (5 percent level) for portfolios 5 and 9. By comparison, Reinganum (1983) found January returns that decreased almost monotonically with capitalizations. Much of the difference is, however, accounted for by the apparently thin trading induced excess February returns for the very smallest stocks.

The Seasonality of the Small Firm Premium

Table 4 strongly suggests that most of the size related premium in stock returns occurs in the month of January. For very thinly traded (small) firms, some of this turn-of-the-year premium may, it seems, be projected into February.[13] Table 5 lists overall mean monthly returns for the ten market value portfolios[14] together with the corresponding mean January, February, February through December, and March through December monthly returns. Table 5 further corroborates the view that small firm premia are limited to the turn of the calendar year. Where overall portfolio returns decrease almost monotonically with capitalization, February through December returns exhibit a much more random pattern. The size related pattern disappears completely when

February returns are excluded: the mean March through December return is exactly the same (.774%) for the top five as for the bottom five market value portfolios.

INSERT TABLE 5 HERE

The significance test for the size premium reported on in Table 2 is now repeated on the subsets of February-December and March-December means, respectively. The equation estimated is (2), both in its original version and augmented by a variable measuring the time in months the security has been listed on the HESE prior to the period under study. This period of listing variable has previously (Barry-Brown (1983)) been shown to be consistent with a variety of models of differential information[15] as well as highly significant in explaining NYSE stock returns.

As expected, the size premium is no longer statistically significant for February-December returns and actually weakly negative for March-December returns. The regression results confirm the outcome of a visual inspection of the magnitudes of the overall size effect relative to the January-February small firm seasonal in Table 1 and 3. The difference in average annual returns between the large (1-2) and the small (9-10) firm portfolios was no bigger than the turn-of-the-year spread, i.e., 8-11 percent. Hence, our results are similar to American ones based on buy-and-hold returns (Blume and Stambaugh (1983)).

In close correspondence to the NYSE results reported by Barry and Brown (1983) a period of listing premium appears to be present in HESE returns. Contrary to these NYSE results, introduction of the variable to measure the amount of publicly available return information did not, however, eliminate the size premium. Furthermore, it appears that in the present data set, most of the period of listing effect occurs at the onset of a new year, although not necessarily only in January.

3. Seasonality and Tax-Loss Selling

Capital gains taxation has been suggested to have undesirable effects on stock trading by giving investors an incentive to realize capital losses while deferring capital gains.[16] An investor subject to capital gains taxation, who during the tax year has realized some capital gains or earned other income from which capital losses may be deducted, may near the end of the tax year find it to his advantage to sell off his "losers" while retaining his "winners." He thus creates a tax write-off which, in an efficient market where expected certainty equivalent returns are the same for all stocks, will cost him only the brokerage fee plus any transaction tax.[17] In an efficient market where demand for securities is perfectly elastic at the certainty equivalent market return, such tax-loss selling should, however, have only insignificant price effects due to temporary reshuffling of portfolios away from the perfectly diversified ones.

This prediction seems likely to be the true one. Even though the above argument, interpreted in terms of reservation prices, means that sellers would generally be willing to trade at prices significantly below prevailing equilibrium prices, the number of buyers with reservation prices equalling the equilibrium ones is likely to be sufficiently large to prevent any price decrease. The tax exempt investors--pension funds, charities, and foundations --are by themselves a significant group likely to bid up prices to their equilibrium level. However, arbitrage operations are not restricted to this group since tax write-offs can be attained quite irrespectively of the tax status of, and at no cost to, the buying party. In the simplest of cases, two investors could gain the full tax credit by exchanging their "losers" at prevailing equilibrium prices, while holding on to their "winners".[18]

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

Transaction costs might, however, offer the rudiments of an explanation. If the market is in equilibrium prior to the tax loss selling rally, and if the extent of, and the actual stocks that will be involved in the rally are hard to predict, the sellers might have to offer their "losers" at a discount equalling at most twice the unit transaction cost plus any property tax liability incurred by holding the taxable security over the turn of the tax year. Just as any transaction cost argument, this one neglects both the investors already decided on a purchase that hence treat the transaction cost as sunk and the investors whose low net worth make them exempt from property taxation. The number of investors who belong to both categories may, however, be sufficiently small as compared to the number of tax loss sellers to permit a minor market price decrease.[19]

The Finnish market provides an interesting testing ground for a tax loss selling hypothesis of January excess returns. This is due to the circumstance that capital losses, according to Finnish tax legislation, can be deducted only from capital gains. Hence, demand for tax write-offs will vary greatly, not with personal income but with (realized) capital gains. The probability that a share of some given stock will be offered for sale towards the end of a tax year should thus increase with some aggregate measure of taxable capital gains realized during the tax year and decrease with its own past gains, i.e., the probability that it, depending on its holding period in the individual portfolio, will provide its holder with a tax write-off if sold.

Stocks are now arranged into groups depending on their performance in each calendar year contained in the present data set. By doing this we attempt a partitioning of the annual and subsequent January return observations into groups with different tax loss sales potential.[20] Assuming similar holding period distributions for all stocks, the better a given stock's performance over the past 12 months, the smaller is the expected number of investors who are able to realize capital losses by sale of that stock by the end of the year. Figures 1(a) through (k) graph mean returns to stocks by their previous year's performance 240 trading days (roughly one calendar year) prior to and 50 trading days (roughly the months of January and February) past the turn of the year. The vertical axis measures the average natural logarithm of the stock price relative to its price on January 1. Note that since the scale of the vertical axis is adjusted to the price variability within each group, a constant angle at the turn of the year would indicate a January effect which is roughly (inversely) proportional to the stock's previous performance.

INSERT FIGURE 1 HERE

The averages of groups of stocks that have lost ground in the previous year all reach their lows at about one trading week before the end of the year. January returns are roughly 16% in the -50% loser group, 5-6% in the -30% or more loser group, and roughly 2% for stocks that had lost 10% or more in the previous year. For winner stocks and for minor losers (-10%-0%) a slight negative deviation from the price "trend" is usually discernible just prior to the year's end. The better the stock's past year's performance, the closer the January returns are to the long-run average returns for the group.

It may be interesting to note also that cross-sectional standard deviations of stock returns exhibit a drastic drop in late December for virtually all stock categories. For the major losers, -40% and -50% or more, the cross-sectional standard deviation drops from a typical 12%-18% to just 4% at the turn of the year. For minor losers the drop is somewhat smaller, from 9%-13% to 4%, while winners exhibit standard deviation drops from typical 6%-16% to just 2%-4% at year's end.

In an effort to reject the tax loss selling hypothesis as stated, (3) is estimated

$$E(R_{it1}) = R_F + (R_{Mt1} - R_F)\beta_i + \alpha_0 + \alpha_1 R_{Mt-1} + \alpha_2 D_i^I R_{it-1}, \tag{3}$$

where $E(R_{it1})$ is the expected return to asset i in January of year t, R_F and R_{Mt1} are the risk free and the market rate or return, respectively, R_{Mt-1} and R_{it-1} are the returns to the market and to asset i in the previous year, respectively, and D_i^I is a dummy variable such that $D_i^I = 1$ if $R_{it-1} < I$ and $D_i^I = 0$ otherwise. Since we have not yet produced an equilibrium theory that would explicitly encompass the January effect, the introduction of a tax-loss selling effect into a CAPM framework may seem doubtful. Consequently, the

α-coefficients of equation (3) will also be estimated separately from the CAPM induced coefficients.[21]

Initial results are reported in Table 7. Notice that in estimating (2) we have treated R_F as a constant to be estimated.[22] The beta estimates used are Dimson betas for the period 1970-81 with five leads and ten lags using the value-weighted WI-index (Berglund-Wahlroos-Grandell (1983)) as market portfolio. The expected value of R_{it1} is proxied by its actual value. Since the regression is pooled, the number of observations equals the number of stocks quoted throughout the period (50) times the number of Januaries less one (11), i.e., 550.

INSERT TABLE 7 HERE

The first three columns of Table 7 give the I-value, the total number of observations and the number of observations for which $D_i^I = 1$, respectively. Columns under A give coefficient estimates for the unabridged equation (3) whereas columns under B given corresponding estimates of a version of the same equation where CAPM induced variables have been omitted. The tax-loss selling hypothesis predicts $\alpha_1 > 0$ and $\alpha_2 < 1$ for I less than the negative of the marginal tax rate times round trip transaction costs.

The tax-loss selling hypothesis cannot be rejected. α_1 is highly significantly positive throughout while α_2 passes a 5 percent test of significance at I = -30% and less. Remaining coefficients receive correct signs and feasible magnitudes: the risk free non-seasonal January return component is estimated at .014-.017 and γ_0 is not significantly different from unity in any except the first regression where I is set at infinity. Estimates for α_1 and α_2 are remarkably similar in the two versions of the regression, and their magnitude is far from irrelevant. The predicted January return to, for example, a >30% loser in a year when the market gains 20% would exceed 8%.

As pointed out earlier, the January seasonal is primarily confined to smaller stocks. This is also apparent from figures 2(a) through (e) where all stocks traded on the HESE during 1970-81 have been arranged into five portfolios in decreasing order of initial capitalization. These figures present standardized stock prices 50 trading days prior to and past year's end. In addition to suggesting that big losers typically are smaller stocks, Figure 2 clearly suggests a thin trading phenomenon in the turn of the year rally.

INSERT FIGURE 2 HERE

The post low rally in the largest stocks appears to start in mid-December and is almost over by the turn of the year. In portfolios 2 and 3, prices begin a distinct upward climb in the last trading week of the year to be followed by subsequent rapid price increases in early January. In the smallest stocks, the post low rally starts after the turn of the year and continues into February.

This data screening suggests that the temporal location of the turn-of-the-year premium probably is influenced by thin trading biases in stock prices. It does not, however, offer any explanation to why the turn of the year effect predominantly is a small stock phenomenon. Table 8 below, however, offers at least one tentative explanation. Small stock returns are generally more volatile than are large stock returns. Most of this volatility, as evidenced by the low betas reported for portfolios 8 through 10 in Table 1, is unsystematic and hence presumably unpriced. In a taxed asset market the simple variance of stock return, however, gains new significance, not as a direct determinant of prices but as a determinant of the probability that a stock can be the subject of tax loss selling. Given a positive mean return, the higher the price variance, the higher the probability of a negative finite holding period price change.

INSERT TABLE 8 HERE

This is not an equilibrium argument, however. If market behavior towards losers at the turn of the year is predictable, as it appears to be, one would expect long run market prices to react to the expectation of tax induced turn-of-the-year premia on high variance stocks, resulting in an elimination of the overall size effect. The notions that the small firm effect occurs at the start of a new year and that this seasonality appears to be taxation induced, therefore, do not explain the long run size premium in stock prices.

4. Long Run Equilibrium Asset Prices

Introduction.

The size premium indicates misspecification of the underlying asset pricing model.[23] The misspecification is likely to be associated with one of three limiting assumptions contained in the CAPM: (i) the single period time horizon, (ii) the single "pervasive force" or factor regime, or (iii) the exclusion of distribution characteristics of higher than second order. Extensive work on an explanation related to the second restriction imposed on the CAPM equilibrium has been performed by Chan, Chen and Hsieh (1983) who conclude that more than four-fifths of the size premium can be attributed to pervasive forces other than the market portfolio excluded from the CAPM.[24]

In this section attention is focused on the third restriction, effectively requiring either stock return distributions to be completely specified by two parameters or investor utility functions to be quadratic.[25] Samuelson (1970) has demonstrated that for compact distributions, the quadratic (utility) solution to the portfolio problem is asymptotically correct as risk (time) approaches zero, but that it can always be improved upon by inclusion of third and higher moments. Most recent work on empirical return distributions tends to reject normality. Return distributions are typically positively skewed and leptocurtic.[26] It should be noted that if this indeed is the case, APT models, based as they are on biparametric distributions, are unlikely to provide us with the final resolution to the controversy.

The non-normality of the return distributions may be induced by one of several processes linked to firm capitalizations. Fixed trading costs are by themselves sufficient to induce kurtosis in thinly traded stocks through their restraining impact on continuous arbitraging.

Similarly, monitoring costs may serve to add to the tails of the return distribution of small stocks. Since large scale attempts to take advantage of fresh information rapidly eliminates any profit opportunity in the case of small stocks, and since monitoring costs are essentially fixed and roughly the same for all securities, monitoring will be more intense for large firms. Hence, while the price of large firm stocks will tend to react even on information of minor importance, small firm stock is likely to react only on more significant new information. Finally, if treated as a shock effect imposed on an otherwise Gaussian return distribution generated by a normally distributed inflow of information, tax related turn-of-the-year seasonality, predominantly occuring in small stocks, may serve as an explanation.[27]

Model and Measurement

The general parameter preference theorem of asset valuation by Rubinstein (1973) states that for any single period maximizers of expected utility of future wealth, reservation expected returns to any asset j are made up of the expected return to some asset k and the sum of the products of the measures of the individual's aversion to m^{th} moment risk and the m^{th} comoments of $(R_j - R_k)$ and the individual's future wealth. Assuming the existence of a risk free asset F, homogeneous expectations and portfolio separation, equilibrium returns are given by

$$(4) \qquad E(R_j) = R_F + \sum_{m=2}^{\infty} \theta_m \sigma_m(R_j, \tilde{W})$$

where θ_m is a measure of market aversion for m^{th} moment risk, σ_m denotes the m^{th} comoment, and $\tilde{W}$ measures the return to the mutual fund of risky assets.

If portfolio separation does not obtain, a number of cases emerge. In all of these, as in the two-moment model of Mayshar (1981), market prices will reflect the varying composition of individual portfolios and hence include a premium on risk. In terms of the empirical unique model this requires the return equation to contain not only the comoments but also measures of second and higher moments of return distribution as well. Hence, in a four-moment world the most general empirical model imaginable would contain six independent variables measuring the second, third and fourth central moments and comoments of the return distributions, respectively.

For an investor exhibiting strict consistency in his direction of preference of moments, consistent risk aversion and positive marginal utility of wealth, Scott and Horvath (1980) have proved that, interpreted as a measure of individual risk aversion,

(5) $\theta_m < 0$ if m is odd, and

$\theta_m > 0$ if m is even.

Investors will have positive preference for odd central moments and negative preference for even central moments. Similar theorems covering the third moment have previously independently been proven by Arditti (1967) and Kraus and Litzenberger (1976).[28]

The explicit empirical model to be used in this paper is an extension of Kraus and Litzenberger's (1976) and Friend and Westerfield's (1980) to include the fourth central moment:

(6) $$\bar{R}_i = \gamma_1 + \sum_{m=2}^{4} \gamma_m \hat{\sigma}_{mi}(R_i, R_m) + \mu_i$$

where

$$\hat{\sigma}_{mi}(R_i, R_M) = \frac{\sum_t^n (R_{it} - \bar{R}_i)(R_{Mt} - \bar{R}_M)^{m-1}}{\sum_t^n (R_{it} - \bar{R}_i)^m}.$$

In line with (5) our hypothesis is that $\gamma_m > 0$ for even m and $\gamma_m < 0$ for odd m.

Data and Testing Procedure

The theory surveyed above relates investors' ex ante reservation returns to subjective estimates of ex ante comoments of the distribution of returns. Since these are not observable, an extension of the theory that suggests hypotheses relating these expectations to observable entities is required.

A Fama-Macbeth (1973) type expectations hypothesis is effectively excluded from consideration because of data limitations.[29] Instead, the Kraus-Litzenberger (1976) hypothesis that ex post return distribution moments estimated from a subset of the available return data are unbiased estimates of the correspondingly expectational variables is adopted. Note that the unavailability of a short term risk free asset in regulated Finnish markets requires us to state the hypothesis in total rather than excess return form.

The small number of stocks in the Finnish market rules out a complete cross classification over the three independent variables as a basis for portfolio formation. Rather than inducing normality by basing formation only on second comoments, tests are performed using single asset parameter estimates. Since this increases (asymptotic) bias in parameter estimation as compared to a portfolio test, it biases the test in favor of rejection of H_1.

To avoid the risk of a spurious correlation in regressions of $\bar{R}_i$ on the set of σ_{mi}, the data is partitioned into two separate sets for the purpose of parameter estimation (1 pass) and testing (2 pass), respectively. While this procedure is designed to handle the spurious correlation problem, it also serves to increase the power of the test by permitting two independent regressions to be run.[30] The partioning is based on sequential order. Even numbered months are allocated to one set, odd numbered to the other set. To avoid allocation of all particular monthly seasonals to one set, the selection criterion is reversed annually.

Accordingly, in the first pass the three σ_{mi} parameters (m=2,3,4) of the model are estimated from a subset of 72 monthly return observations for each of the 42 stocks. In the second pass, the time series means over the remaining 72 observations are regressed on the σ_{mi}'s cross sectionally. To test for size and period of listing effects, variables measuring the natural logarithm of the firm's capitalization at period's start and the number of monthly listings available are also entered into the regression.

Results

Table 9 reports the OLS-estimates of the expanded equation (6) for two slightly different data sets. Set A (rows 1-4) contains all 144 observations partitioned into first and second pass sets of equal size. Two independent regressions on this data are reported upon, one where the first pass was made through subsample 1 and the second pass subsequently in subsample 2, and the second where this order was reversed.

Row (1) contains the coefficient estimates of the parameter preference pricing model as stated in (6). Coefficient signs on third and fourth comoments are correct and, excluding the third comoment in second pass regression on subsample 1, significant at the 5% level. F-statistics for the regressions do not, however, permit us to reject the null hypothesis. Coefficient signs for second comoments (betas) are, moreover, erroneous. Augmented by a period of listing variable, the pricing model passes a 5%-level F-test in both subsamples. Correctly signed comoment prices increase with the inclusion of the informational proxy. 21%-31% of cross sectional return variation is explained by this model.

INSERT TABLE 9 HERE

The premium associated with firm size is smaller, and indeed, statistically insignificant in this pricing model as compared to CAPM as reported in Table 2. The inclusion of the informational proxy, however, appears crucial to the reduction of the size premium as indicated by the regressions reported in row (4).

These results would seem to indicate that a general parameter preference model performs significantly better than does the CAPM in explaining stock prices. The fact that the size premium is lower indicates that at least some of the size related turn-of-the-year seasonals are indeed foreseen by investors. To provide additional tentative evidence relating to this question, some potential January price shocks were removed from the data. This was done by eliminating from the means all January return observations preceeded by a negative annual return, and hence, in line with the tax loss selling argument, potentially affected by the tax related selling pressure on the market. The number of monthly return observations thus removed equalled 151 out of a total of 6048. Equation (6) was then reestimated. Results are in section B (rows 5-8) of Table 9.

In both subsamples model fit is slightly improved. In second pass subsample 1 the improvement in r^2 is 2.5%-6% and 1%-3% in second pass subsample 2. Given that the four-moment parameter preference model is the true one, this result indicates that a portion of the apparently tax related January seasonals are unexpected.

As a final test of the size anomaly related performance of the several stock pricing models surveyed in this paper, the residuals of these models were regressed on firm capitalizations and their natural logarithms. Results are reported in Table 10.

INSERT TABLE 10 HERE

The residual monthly size elasticity in the CAPM averages -1.53 which translates into -.35% per each $100 million of capitalization. When the informational proxy measuring months of listing on the HESE is included, the residual size effect is reduced by roughly one-fourth in the log-linear model and a third in the linear one. Corresponding reductions following the introduction of the third and fourth comoments into the pricing model are slightly less than 20%. The residuals of the listing augmented four moment parameter preference model contain less than half of the original size related premium.

It is thus indicated that the long run size effect in stock prices--virtually all of which takes the form of apparently tax related turn-of-the-year premia--is only partly made up of unexpected return shocks. To a large extent the size premium appears to reflect a misspecification of the underlying valuation model, or rather the use of an oversimplified model. A model generalization designed to allow for departures from quadratic restrictions and assumptions of zero estimation risk produces a significant reduction in the residual size premium.

5. Summary and Conclusions

This paper investigated the presence and magnitude of some size and calendar related pricing anomalies on a small security market, the Helsinki Stock Exchange. The descriptive analysis indicated several departures from the outcomes expected in an efficient market. The mean annual return on a portfolio of the smallest stocks on the HESE amounted to 19.8% in the period 1970-81 while, for the same period, the portfolio of the largest stocks earned an average of only 10.6 percent. The risk adjusted excess return on the small firm portfolio was 8.7% a year while it was negative (-2.2%) on the large firm portfolio. Significance tests in the Reinganum (1982) tradition clearly showed that the phenomenon is unlikely to be random. In a cross sectional regression of mean stock returns, size variables received coefficients which were significantly negative on the .005 level.

Stock returns were also markedly seasonal. The equally weighted market index contains a January seasonal of 3.2% which is significantly positive on the .005 level. This seasonal is more pronounced in small stocks. January excess returns average 2%-6.5% in the small firm portfolios while they amount to only 1.5%-2.5% in the large firm ones. The rate of return records for the smallest stocks furthermore appear to contain a fairly strong thin trading bias causing some of the turn-of-the-year premia to show up in early February returns. January-February excess returns average 1.5%-2.5% for the largest and 5%-11% for the smallest stocks.

Since the turn of the year return spread among firms of different size is roughly the same as the overall size effect, it is not surprising to find that February through December returns contain only a weak, and March through December returns contain no size effect whatsoever. Regression of average monthly returns for March through December on capitalization yields an insignificantly positive slope coefficient. This fact in no way explains away the premium itself. In an efficient capital market the long excess run returns earned by holders of small stocks should attract other investors, quite independently of whether they tend to occur at the turn of the year or during the summer.

Moreover, as is seen from table 4, January is not the only month in which stock returns seem related to firm sizes. February, August and December returns are also statistically significantly related to firm capitalizations.

Even more puzzling is the fact that of these, the December returns actually contain a size discount, i.e. the returns are highest for the largest firms in that month. This discount is fairly clearly visible in figure 2 also. The presence of what appears to be a highly complicated seasonal pattern in monthly returns suggests extreme caution in the interpretation of our results as they relate to any link between size- and seasonal premia. Indeed, the main argument of this paper is that seasonals cannot provide us with a causal explanation to persistent return differences between stocks of different capitalization.

Noting the fact that Finnish tax legislation permits deductions on capital losses only from capital gains, the paper goes on to suggest a testable version of the familiar tax-loss selling hypothesis of Branch (1977), Reinganum (1983) and others. That hypothesis cannot be rejected. It is found that January returns depend negatively on the stock's previous year's performance and positively on the market's previous year's performance, the latter of which simply measures aggregate capital gains and consequent need for tax write-offs. If tax-loss selling is actually responsible for the turn-of-the-year seasonal, the link to the size effect appears to be the simple fact that return variance is significantly higher in small than in large stocks. Since mean returns are positive, the probability that a small stock will end up a loser over some finite holding period is much higher than the corresponding probability for a large stock.

In its final section, the paper addresses the question of how long run equilibrium prices are determined in such a seasonal market. It argues that over long periods of time and consequent changes in wealth an assumption of quadratic utility is likely to be a poor approximation. Since empirical stock return distributions are known to exhibit non-Gaussian properties, the mean-variance framework, requiring either quadratic utility functions or return distributions completely characaterized by two parameters, is likely to be insufficient in generality to yield accurate stock return predictions.

In line with several previous papers by Rubinstein (1973), Kraus and Litzenberger (1976), Scott and Horvath (1980), and Friend and Westerfield (1980), the effect of nonnegative third and higher return distribution moments on the required rate of return is examined. Estimation of a four-moment parameter preference model yields correct and mostly significant signs on measures of systematic skewness and kurtosis. The model fails to pass a 5% F-test, however. When estimation risk (Bawa, Brown and Klein, 1979) is accounted for by a period of listing proxy, the model cannot be rejected.

In the framework produced by the model, the residual size premium is reduced to less than half of its original (CAPM) size. The remaining size premium is unpriced, apparently because it was unexpected. The fact that the four-moment parameter preference model fits the data from which potential January shocks--January returns that had been preceded by a year of negative returns--have been omitted somewhat better than the total data, indicates that those January returns still contain a large unexpected component.

Notes

[1]Other results indicate that even more of the size premium is attributable to the month of January. After correcting for biases contained in daily closing prices by computing returns for buy and hold portfolios, Blume and Stambaugh (1983) find that the size effect is roughly half of what was reported in Banz (1981) and Reinganum (1981) and that all of it is due to January.

[2]See, e.g., Kent (1973), pp. 106-106.

[3]Rarely a large block transaction will occur at prices different from the ones fixed during the calling phase. Such transactions are by nature best compared to secondary distributions on the NYSE.

[4]Wahlroos-Berglund (1983).

[5]Of the three stocks removed, one was removed because of a merger, the second following nationalization, and the third because of the introduction of more frequently trading preferred stock in the same corporation.

[6]Note that this paper deals with long run equilibrium stock prices. The size effect hypothesis to be rejected hence is also a long run one, relating excess returns over extended periods of time to historic capitalizations.

[7]Since the Finnish bond market is regulated, excess demand or supply may prevail in the primary market. Bonds of shorter maturity than five years are seldom issued. The difference between the internal rate of return on a bond portfolio purchased in the secondary market and rolled over at maturity of 2.5 to 3.5 years and the nominal rate on newly issued bonds has averaged .68% in 1977-81. Adjusting the excess returns for this difference would not affect their order.

[8]This portfolio contains the pulp and woodworking firm of Kemi, a controlling interest in which, at the verge of its bankruptcy in 1978, was purchased by the government.

[9]Evidence in Berglund and Liljeblom (1983) indicates that Finnish stock return distributions may converge only at very long differencing intervals. Weekly, bi-weekly and monthly return distributions are typically leptocurtic and positively skewed. For some U.S. evidence on the same phenomenon see Fama (1965) and Upton and Shannon (1979).

[10]In addition to seven such stocks, the Metsälitton Teollisuus preferred was also removed. The price of this security is administered by the central co-operative bank to behave like a bond in order to make it usable as an

instrument of payment in lumber deals. Since 1977 the only significant price changes around the nominal value of the share have occurred in the two months preceding annual divident payments.

[11]Dimson betas were estimated from daily return data with five leads and twenty lags, MPR was set at the difference of the WI-index and the risk free rule of return.

[12]This interpretation of the seasonality hypothesis appears to differ from the one employed by Reinganum (1983) who, by regressing daily (total) returns on a set of dummies, tests for departures from a zero return hypothesis.

[13]This thin trading effect is contained in Reinganum's (1983) results as well. Analyzing daily data he uncovers supernormal returns for small stocks past the first seven trading days of January. Considering the thinness of the Helsinki market as compared to the NYSE it is hardly surprising that thin trading biases may linger on for two to three times as long. Related results reported in Berglund-Wahlroos-Örnmark (1983) indicate that the serial correlation of stock price changes, apparenetly induced by the dissemination of information produces a lag structure which on the HESE measures up to two trading weeks. A similar lag structure of a considerably shorter duration is contained in the NYSE results by Fama (1965).

[14]To enhance compatibility with the results reported in Table 6, preferred stocks have been removed. The number of stocks contained in the ten portfolios is hence 42.

[15]See Bawa, Brown and Klein (1979) and the appendix to Barry and Brown (1983).

[16]See, e.g., Haugen and Wichern (1973).

[17]Note that the mode of trading on the HESE briefly discussed earlier, specifically the use of a separate price-fixing stage of trading in most cases make simultaneous (the same day) purchases and sales possible. Hence no bid-asked spread is contained in total transaction costs.

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

[19]Raw empirical data suggests that there must be more to the January effect than this. The January excess return on the equally weighted portfolio of all stocks in Table 2 already exceeds total transaction costs associated with going into and out of a given stock (3%) by 0.23%. The mean January

excess return for "losers"--if they are indeed responsible for the effect--must be significantly higher than 3.2%.

[20]A perfect measure of the tax loss potential of a share seems almost impossible to devise since it is dependent on the distribution of individual holding periods of which there is no record. Branch (1977) defined a "loser" as a stock reaching its yearly low in the last whole trading week of the year, while Reinganum (1983) used the ratio of the stock's price on the second to the last trading day of the year to its high in the preceeding six months.

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

[22]The main reason for doing this is the absence of a short-term bond market mentioned early, and hence of reliable records of short term risk free interest rates.

[23]The alternative hypotheses (i) that the premium is produced by data shortcomings, or (ii) erroneously estimated second order risk, have been dealt with in Blume and Stambaugh (1983), Roll (1981), Roll (1982), and Reinganum (1982).

[24]The macrovariable primarily responsible for the performance of the Chan, Chen and Hsieh (1983) pricing model was the difference in return of "under BAA" and "AAA" bond portfolios typically closely related to the return spread between large and small firms.

[25]Note that the CAPM can be generalized to allow any set of distributions belonging to the symmetric stable class with characteristic exponent $1 < \alpha \leqslant 2$, provided all assets have the same characteristic exponent (Fama (1971)).

[26]Beedles (1979), Simkowitz and Beedles (1980), Fielitz and Rozelle (1983) and, for results relating to the data base underlying the present paper, Berglund and Liljeblom (1983). In addition to providing general evidence on departures from normality, these studies, especially Beedles (1979) and Berglund-Liljeblom (1983) suggest that several return distribution moments may be linked to firm size.

[27]The seasonality based explanation is the most elegant one in that it ties together the several arguments presented in this paper. Note, however, that non-normality does not have to follow from seasonality for the parameter preference model to explain the size anomaly, previously shown to be just a seasonal phenomenon.

[28]Kraus and Litzenberger (1976) furthermore claim to be able to prove a preference theorem for any higher moment but find no need to do so. Additional work on three-moment asset pricing models and their application is

contained in Friend-Westerfield (1980) and Kane (1982). Kraus-Litzenberger and Friend-Westerfield were generally able to reject the hypothesis that skewness does not enter the pricing equation. Friend and Westerfield, however, point out that the conclusion with respect to the paramter preference model as a whole is highly sensitive to the market indices used.

[29]Available are 12 years of data. The instability and slow convergence of return distributions restrict us to monthly data and consequent 144 observations on each stock. The fact that higher return distribution moments require fairly long estimation periods rules out a test of the multiple-overlapping period type.

[30]In this sense the methodology bears some resemblance to the use of a post-sample period prediction set common in time series analyses.

[31]Negative relationships between betas and rates of return are reported by Hawawini, Michel and Viallet (1983) for the French stock market. In a reproduction of Fama-Macbeth (1973) very similar to theirs, we received negative coefficients for betas on the Helsinki exchange also.

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Table 1. Portfolio statistics 1970-81, Helsinki Stock Exchange. Returns are computed from the difference between starting and ending data points of a return index based on the KOP-data base, described in Berglund-Wahlroos-Grandell (1983).

Portfolio	Mean Capitalization Mill.FIM 1970	Daily Beta	Monthly Beta	Dimson Beta	Mean Annual Return 1970-81 %	Mean Annual Return 1970-73	Mean Annual Return 1974-77	Mean Annual Return 1978-81	Mean Excess Return %
1	203.2	.921	.998	.949	10.64	26.87	-9.59	14.67	-2.16
2	139.0	1.013	1.074	1.091	11.54	36.13	-17.98	16.46	-1.82
3	77.8	.927	1.102	1.109	12.55	30.25	-16.63	24.03	-0.89
4	46.7	.800	1.140	1.119	14.19	36.59	-7.96	13.93	0.71
5	35.8	1.459	.912	1.023	17.51	41.24	-4.14	15.42	4.42
6	31.7	.704	1.010	.941	17.42	36.85	-3.97	19.37	4.66
7	19.3	.648	1.079	1.004	15.62	37.21	-6.87	16.50	2.60
8	6.5	.444	.727	.803	19.07	27.72	0.39	29.11	6.86
9	3.9	1.044	.875	.987	20.05	35.39	-0.31	25.05	7.10
10	1.8	.363	.499	.521	19.78	17.29	23.09	18.95	8.70

Table 2. The small firm effect in HESE data; Coefficients of regression of security return means on betas and capitalizations. 42 stocks 1970-81, monthly data, t-statistics in parentheses. Coefficients multiplied by 1000 for convenience.

	β Monthly Data	Dimson β Daily Data	ln Size	Intercept	r^2	F
Means	-.03 (.01)		-1.55 (3.05)	17.2 (7.37)	.219	5.48
		-1.80 (.65)	-1.37 (2.54)	18.2 (8.15)	.228	5.75
Truncated Means	.17 (.08)		-1.67 (4.10)	13.5 (7.23)	.333	9.72
		-1.28 (.58)	-1.54 (3.53)	14.3 (8.01)	.338	9.96

Table 3. Mean monthly excess returns to the market portfolio and to 10 stock portfolios in decreasing order of market value at period's start.

Portfolio		Jan	Feb	Mar	Apr	May	Jun	Jul	Aug	Sep	Oct	Nov	Dec	Total	Dimson beta
Large firm	1	.0212	-.0109	.0221	-.0150	-.0266	-.0038	.0179	-.0106	-.0244	-.0133	-.0074	.0336	-.0216	.949
	2	.0154	.0021	.0041	-.0210	-.0211	.0030	.0206	-.0099	-.0252	-.0092	-.0000	.0259	-.0182	1.091
	3	.0173	.0053	.0004	.0120	-.0214	-.0051	.0208	-.0156	-.0182	-.0056	-.0115	.0160	-.0089	1.109
	4	.0266	.0027	-.0128	.0124	-.0207	.0099	.0306	-.0216	-.0082	-.0032	-.0169	.0108	.0071	1.119
	5	.0635	.0158	.0157	.0089	-.0172	.0098	.0015	.0035	-.0076	-.0182	-.0176	-.0081	.0442	1.023
	6	.0372	.0169	.0188	.0195	-.0271	.0120	.0146	.0007	-.0158	-.0287	-.0126	.0164	.0466	.941
	7	.0303	.0052	.0105	.0196	-.0286	.0167	.0043	-.0033	-.0274	-.0085	-.0125	.0177	.0260	1.004
	8	.0470	.0145	.0462	.0051	-.0200	.0007	.0012	.0195	-.0142	-.0195	-.0006	-.0084	.0686	.803
	9	.0740	.0574	.0090	.0216	-.0066	.0014	.0076	.0060	-.0021	-.0320	-.0319	-.0359	.0710	.987
Small Firm	10	.0386	.0510	.0038	-.0041	-.0203	.0222	.0057	.0210	.0088	-.0112	-.0225	-.0019	.0870	.521
Market Index		.0251	.0031	.0097	-.0026	-.0197	.005	.0163	-.0036	-.0183	-.0115	-.0089	.0161		1.000

Table 4 Estimated coefficient values of monthly dummies for returns to ten stock portfolios composed in decreasing order of capitalization (1=largest, 10=smallest). Absolute t-statistics in parentheses.

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

Table 5. Some seasonal return characteristics for 10 market value portfolios of Finnish stocks.

Firm Size	Portfolio	Mean Monthly Returns (percent) Overall	January	February	February-December	March-December
Large	1	.88	3.21	-.13	.67	.75
	2	.91	2.04	1.66	.81	.72
	3	.81	2.76	1.31	.63	.56
	4	1.24	3.10	2.20	1.07	.96
	5	1.43	6.40	1.99	.98	.88
	6	1.41	4.77	2.94	1.11	.92
	7	1.05	3.63	.98	.81	.80
	8	1.45	5.52	2.04	1.08	.98
	9	1.46	7.82	6.83	.88	.28
Large	10	1.59	4.14	6.06	1.36	.89

Table 6. Cross-sectional regression estimates of the coefficients of the augmented equation (2) for mean monthly returns; (a) all months, (b) excluding Januaries, and (c) excluding Januaries and Februaries. 42 Finnish stocks monthly and daily data 1970-81. Coefficients multiplied by 1000; t-values in parentheses.

Means of Monthly Returns	Intercept	β Monthly Data	Dimson β Daily Data	ln Size	Months Listed	r^2	F
(a) January-December							
(1)	17.19	-.035		-1.55		.219	5.48
	(7.37)	(.01)		(3.05)			
(2)	18.64	.449		-1.22	-.0067	.303	5.51
	(7.99)	(.18)		(2.39)	(2.14)		
(3)	18.91		.026	-1.19	-.0066	.303	5.49
	(8.69)		(.01)	(2.25)	(2.02)		
(b) February-December							
(4)	14.00	-2.22		-.86		.117	2.57
	(5.81)	(.84)		(1.63)			
(5)	15.46	-1.73		-.53	-.0066	.205	3.27
	(6.38)	(.68)		(.99)	(2.06)		
(6)	15.39		-2.22	-.47	-.0060	.208	3.32
	(6.82)		(.76)	(.86)	(1.77)		
(c) March-December							
(7)	10.00	-3.65		.28		.035	.72
	(3.59)	(1.20)		(.46)			
(8)	11.01	-3.31		.51	-.0046	.071	.96
	(3.80)	(1.09)		(.80)	(1.20)		
(9)	9.85		-1.99	.43	-.0043	.050	.66
	(3.61)		(.57)	(.65)	(1.04)		

Table 7. Coefficient estimates of equation (3) on the tax-loss selling effect on January stock prices. t-statistics in parentheses.

			A						B			
			Intercept	β_i	$R_{Mt}I\beta_i$	R_{Mt-1}	$D_i^I R_{it-1}$		Intercept	R_{Mt-1}	$D_i^I R_{it-1}$	
I	n	n'	α_0+R_F	$-R_F$	γ_0	α_1	α_2	r^2	α_0	α_1	α_2	r^2
∞	550	550	.024	-.015	.135	.135	.007	.085	.028	.126	.006	.074
			(1.75)	(-.93)	(5.92)	(5.29)	(.39)		(5.67)	(4.98)	(.31)	
.00	550	181	.023	-.016	.695	.148	-.027	.085	.025	.141	-.037	.075
			(1.67)	(-.98)	(2.46)	(6.61)	(-.64)		(3.99)	(6.19)	(-.89)	
-.10	550	112	.023	-.016	.693	.148	-.029	.085	.025	.140	-.038	.075
			(1.68)	(-.98)	(2.46)	(6.61)	(-.70)		(4.23)	(6.32)	(-.94)	
-.20	550	61	.023	-.016	.704	.148	.042	.086	.025	.139	-.045	.076
			(1.67)	(-.99)	(2.52)	(6.95)	(-1.06)		(4.67)	(6.59)	(-1.12)	
-.30	550	33	.023	-.017	.698	.153	-.090	.092	.024	.143	-.093	.082
			(1.65)	(-1.05)	(2.50)	(7.32)	(-2.12)		(4.66)	(6.96)	(-2.17)	
-.40	550	14	.022	-.014	.655	.150	-.139	.096	.025	.141	-.148	.087
			(1.63)	(-.90)	2.34)	(7.37)	(-2.64)		(4.98)	(7.04)	(-2.82)	
-.50	550	8	.022	-.014	.666	.150	-.216	.104	.025	.141	-.223	.095
			(1.64)	(-.91)	(2.40)	(7.43)	(-3.47)		(5.10)	(7.09)	(-3.57)	

Table 8. Mean variances and higher moments of monthly returns to stocks arranged in 10 portfolios in decreasing order of capitalization at period's start.

		Means of (Coefficients of)		
Portfolio		Variance	Skewness	Kurtosis
Large	1	.0033	.4252	2.503
	2	.0037	.1338	2.670
	3	.0046	.4497	2.107
	4	.0042	.9975	3.490
	5	.0054	.8488	5.920
	6	.0047	.3899	3.206
	7	.0058	-.2055	4.000
	8	.0078	.5245	9.240
	9	.0110	.5409	6.124
Small	10	.0128	.0922	3.958

Table 9. Estimated (OLS) coefficients of equation (6) explaining long run stock returns, augmented by variables measuring firm size and listing period. All regression coefficients are multiplied by 1000 for convenience; t-values are in parentheses.

Subsamples:																
First Pass	1								2							
Second Pass	2								1							
Variable	γ_1 Intercept	γ_2 σ_2	γ_3 σ_3	γ_4 σ_4	ln Size	Listing Period	r^2	F	γ_1 Intercept	γ_2 σ_2	γ_3 σ_3	γ_4 σ_4	ln Size	Listing Period	r^2	F
A. All Observations																
(1)	7.70	-5.64	-5.08	11.75			.140	2.06	16.80	-14.34	-.41	13.13			.111	1.57
	(2.60)	(1.12)	(2.48)	(2.07)					(4.89)	(1.97)	(.54)	(2.17)				
(2)	10.50	-5.25	-5.84	13.21		-.0087	.212	2.49	20.73	-12.94	-.36	14.14		-.0143	.310	4.16
	(3.23)	(1.08)	(2.88)	(2.37)		(1.84)			(6.30)	(1.99)	(.53)	(2.61)		(3.28)		
(3)	11.89	-1.32	-5.39	10.32	-1.14	-.0064	.245	2.34	22.13	-8.82	-.46	11.45	-1.15	-.0119	.351	3.89
	(3.49)	(.23)	(2.64)	(1.72)	(1.26)	(1.27)			(6.57)	(1.26)	(.70)	(2.04)	(1.50)	(2.62)		
(4)	10.62	-.14	-4.74	8.33	-1.55		.212	2.48	20.05	-7.39	-.56	9.10	-1.84		.227	2.72
	(3.24)	(.03)	(2.38)	(1.43)	(1.83)				(5.69)	(.99)	(.79)	(1.53)	(2.36)			
B. Excluding Potential January Shocks																
(5)	6.70	-4.79	-5.82	12.32			.169	2.57	16.20	-15.50	-.61	15.14			.136	1.99
	(2.17)	(.91)	(2.72)	(2.08)					(4.58)	(2.07)	(.80)	(2.43)				
(6)	9.24	-4.43	-6.51	13.65		-.0079	.221	2.63	20.64	-13.92	-.55	16.28		-.0161	.370	5.44
	(2.70)	(.86)	(3.04)	(2.32)		(1.59)			(6.29)	(2.14)	(.83)	(3.02)		(3.71)		
(7)	10.22	-1.69	-6.20	11.63	-.79	-.0063	.236	2.22	22.54	-8.29	-.69	12.61	-1.57	-.0130	.439	5.64
	(2.81)	(.28)	(2.84)	(1.82)	(.82)	(1.17)			(6.90)	(1.23)	(1.08)	(2.31)	(2.11)	(2.93)		
(8)	8.96	-.52	-5.56	9.67	-1.20		.207	2.41	20.29	-6.74	-.80	10.06	-2.32		.306	4.07
	(2.56)	(.09)	(2.61)	(1.56)	(1.33)				(5.82)	(.91)	(1.14)	(1.70)	(3.01)			

Table 10. Slope coefficients for residuals of different cross-sectionally estimated pricing models regressed on capitalizations or logs of capitalizations. Standard deviations in parentheses.

Second-Pass Sample:	2		1		Means	
Size Variable:	S ($\times 10^5$)	ln S ($\times 10^3$)	S ($\times 10^5$)	ln S ($\times 10^3$)	S ($\times 10^5$)	ln S ($\times 10^3$)
Pricing Model:						
2-moment parameter preference (CAPM)	-3.34 (1.80)	-1.18 (.75)	-3.66 (1.75)	-1.88 (.69)	-3.50	-1.53
CAPM + Period of Listing	-2.55 (1.80)	-.90 (.74)	-2.02 (1.64)	-1.21 (.65)	-2.29	-1.11
4-moment parameter preference	-3.02 (1.67)	-.90 (.69)	-2.75 (1.68)	-1.44 (.67)	-2.89	-1.27
4-moment parameter preference + period of listing	-1.87 (1.64)	-.69 (.67)	-.92 (1.52)	-.80 (.61)	-1.40	-.75

Figure 1. Average logs of stock prices relative to January 1 prices 240 trading days prior and 50 trading days past the turn of the year, grouped according to performance in the previous year.

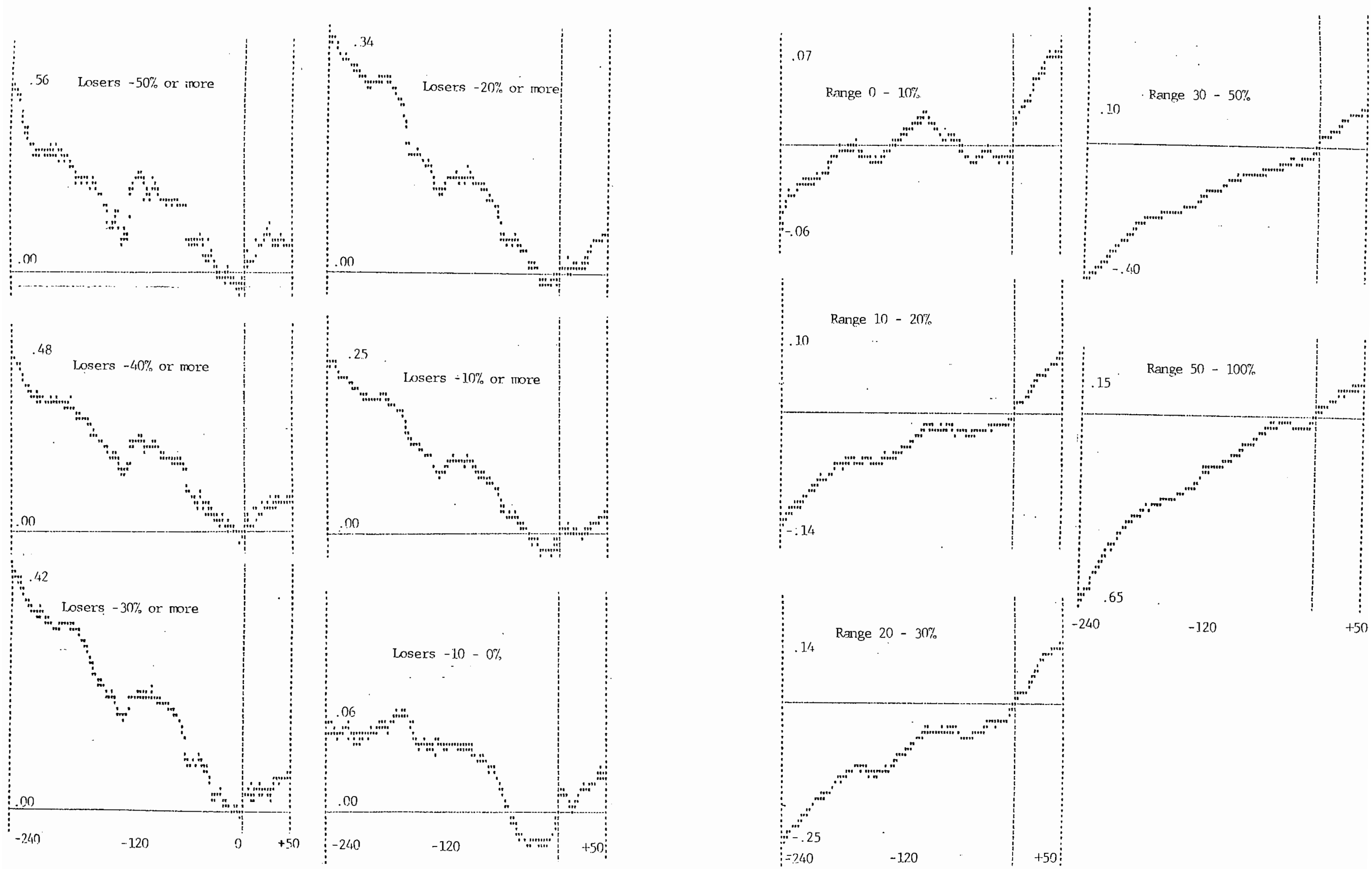

Figure 2. Average logs of stock prices relative to January 1 prices 50 days prior and 50 days past the turn of the year. Stocks are arranged into five market value portfolios in decreasing order of initial capitalization.