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STOCK RETURN SEASONALITIES AND THE TAX-LOSS SELLING HYPOTHESIS

Analysis of the Arguments and Australian Evidence*

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A 'tax-loss selling' hypothesis has frequently been advanced to explain the 'January effect' reported in this issue by Keim. This paper concludes that U.S. tax laws do not unambiguously predict such an effect. Since Australia has similar tax laws but a July–June tax year, the hypothesis predicts a small-firm July premium. Australian returns show pronounced December–January and July–August seasonals, and a premium for the smallest-firm decile of about four percent per month across all months. This contrasts with the U.S. data in which the small-firm premium is concentrated in January. We conclude that the relation between the U.S. tax year and the January seasonal may be more correlation than causation.

1. Introduction

Recent empirical studies by Banz (1981) and Reinganum (1981) report a significant negative relation between abnormal returns and market value of common equity for samples of NYSE and NYSE–AMEX firms respectively. Brown, Kleidon and Marsh (1983) find that the apparent relation is actually reversed over certain intervals of time, while Keim (1983) shows that the

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small firm premium is always positive in January from 1963 to 1979. Keim reports that nearly fifty percent of the average annual size effect can be attributed to the month of January, and more than half of the January effect occurs during the first week of trading. A complete explanation of the apparent size effect requires, therefore, separate explanations for both the sizable small firm premium every January and the smaller and non-stationary premiums in the other months.

Subsequent studies of the January effect have examined a tax-loss selling hypothesis [discussed in Wachtel (1942), Branch (1977) and Keim (1983)] that the small firm premium in the first few days of the year is a reaction to tax selling pressure at the end of the tax year for the shares of these firms. These studies have taken two different empirical approaches. First, Reinganum (1983) and Roll (1983) examine U.S. stock return data and report that January returns are greater for those firms with larger price declines over the previous six- and twelve-month period respectively, which they interpret as support for the tax-loss hypothesis. However, neither study controls for other variables that may affect January returns or investigates whether the phenomenon is unique to January.

A second method of obtaining evidence on the tax-loss hypothesis is to examine stock return patterns on stock markets in countries with different tax year-ends, on the assumption that international arbitrage is inhibited. Gultekin and Gultekin (1982) and Korajczyk (1982) both examine the monthly returns of value-weighted stock market indices in countries with widely differing tax laws and tax year-ends. Both studies find evidence of a persistent but generally less significant (than the U.S.) January effect in most of the countries, with Gultekin and Gultekin interpreting the evidence as support for the tax-loss hypothesis and Korajczyk interpreting it as not supportive. One drawback to these studies is that a single market index does not allow discrimination between small and large firms. Further, analysis of a value-weighted index, which is more heavily influenced by large firms, does not permit detection of the more important seasonal patterns in *small* firm stock returns that are allegedly the result of country-specific tax laws.

In this study, we extend the discussion of the tax-loss selling hypothesis and also examine the month-to-month small firm return premium for a sample of Australian stocks for the period 1958 to 1981. Although the basic idea behind the tax-loss selling hypothesis seems straightforward, a number of factors mitigate any impact that tax-related selling may have on stock prices. We argue that, at best, the tax-loss hypothesis leads to ambiguous predictions.

In Australia, the tax year-end is June 30, and the tax treatment of capital gains/losses is such that the tax-loss selling hypothesis predicts a July seasonal in returns for small stocks. We find that the raw returns for most Australian stocks exhibit pronounced December–January and July–August seasonals, with the largest (and roughly equal) effects in January and July. The persistence of these seasonals across almost all size categories results in an Australian 'size effect' — the smallest 10% of the firms earn an average monthly premium relative to the other nine deciles of at least 4% — with a month-by-month behavior that is very different from that observed in the U.S. Unlike the U.S. effect, the Australian size *premium* for small firms, whether measured with raw returns or with abnormal returns relative to the Sharpe–Lintner CAPM, appears fairly constant across all months.

It might be possible to construe some of our empirical results as more indicative of the degree of integration of world capital markets than as inconsistent with the tax-loss selling hypothesis. However, although this may 'explain' the Australian January seasonal, it leaves unexplained the August and December seasonals. The theoretical ambiguity of tax effects and the Australian evidence presented here tend to support the conclusion that the relation between the tax year and the U.S. January seasonal may be more correlation than causation.

Sections 2 and 3 discuss the tax-loss hypothesis with respect to the U.S. and Australia, and sections 4 and 5 report the data and results. Section 6 briefly discusses our results in an integrated capital markets context, and contains concluding remarks.

2. Tax-loss selling

The tax-loss selling hypothesis has been advanced most often to explain the January effect. The hypothesis maintains that tax laws influence investors' portfolio decisions by encouraging the sale of securities that have experienced recent price declines so that the (short-term) capital loss can be offset against taxable income.[1] Small firm stocks are likely candidates for tax-loss selling since these stocks typically have higher variances of price changes and, therefore, larger probabilities of large price declines. Importantly, the tax-loss argument relies on the assumption that investors wait until the tax year-end to sell their common stock 'losers'. For example, in the U.S., a combination of liquidity requirements and eagerness to realize capital losses before the new tax year may dictate sale of such securities at year-end. The heavy selling pressure during this period supposedly depresses the prices of small firm stocks. After the tax year-end, the price pressure disappears and prices

[1] Thus, the rationale for a year-end tax effect arises from the ability of capital losses to directly offset ordinary income. In the U.S., realized capital gains and losses are short-term if the asset has been held for one year or less, and long-term otherwise. Short-term gains or losses are taxed as ordinary unearned income, while long-term gains and losses are taxed at forty percent of the investor's marginal tax rate on ordinary income since sixty percent of long-term gains or losses are excluded from taxable income. Net short-term losses and net long-term gains incurred in the same year, however, offset each other. Also, net short-term losses and fifty percent of net long-term losses are limited to a deduction level of $3,000 in any year. Unused capital losses may be carried forward indefinitely.

rebound to equilibrium levels. Hence, small firm stocks display large returns in the beginning of the new tax year.[2]

There are, as Roll (1983) and others recognize, obvious problems with the above argument. The major problem is that even if it were established that there is heavy tax-related selling of a particular security at year-end, this does not necessarily imply any price decline. Realizations of capital gains and losses need not be associated with a shift in the demand functions for stocks. Although the price pressure hypothesis suggests that a firm's common stock — being unique — has a downward sloping demand curve, it is generally recognized that securities with similar risk characteristics will serve as close substitutes for inclusion in an investor's portfolio [see Scholes (1972)]. In this case, the demand curve is essentially horizontal and tax-related selling should have no effect on security prices.

Further, the mechanics of this price pressure argument are unclear. If sales are made purely for tax purposes (as opposed to a desire to liquidate a portfolio), then the proceeds must be reinvested somewhere. Portfolio requirements would suggest that, all else equal, it would be desirable to replace the security by one with similar characteristics. Although 'wash sales' are precluded by tax laws, investment brokers publish lists of similar stocks to aid tax exchanges [Sharpe (1981, p. 205)] that would weaken any price pressure effect. Sharpe concludes (1981, p. 205):

> End-of-year sales and purchases motivated by tax considerations are fairly common. Volume in securities that experienced substantial price changes during the year tends to be high as holders sell to realize gains or losses. *However, no major fall in prices appears to result from this pressure.* Buyers apparently recognize that the sellers are motivated by knowledge of the tax laws, and not some previously unrecognized news of disastrous developments affecting the companies in question. (Emphasis added.)

If the price pressure argument were true, we should observe a general decline in prices of small firm stocks in December even if no unusually detrimental information about them was filtering into the market. Keim (1982) finds evidence of general price increases for these shares in December.

Finally, there appear to be relatively simple ways to avoid or exploit any 'low' price, if it did in fact exist. Those who sold a little before the rush would avoid the large price pressure loss.[3] Anyone not forced to sell an underpriced

[2]The same line of reasoning implies the potential for a late April effect as investors experience liquidity drains during the income tax filing period, although no such seasonal has been found. Constantinides (1982, p. 38) discusses a variant of this hypothesis, under which selling pressure throughout the year is 'suddenly relieved at the beginning of January'. Our primary arguments apply to this variant also.

[3]This is especially the case since they are free to redefine their personal tax year to differ from the usual December 31 year-end.

security for tax purposes would stand to gain large excess returns by purchase of those securities.

Even the case for abnormal tax-loss selling of small firm stocks at year-end is not clearcut. It is easy to see why it would be preferable to time tax-loss sales at year-end as opposed to (say) a few days *later* (in the following year) to avoid waiting another twelve months for the benefit of the tax loss. However, why not recognize capital losses as they occur *throughout* the year? Given the existing tax structure and wash sale rules, Constantinides (1982, p. 35) argues that 'investors have an incentive to realize capital losses immediately'. This would result in only a fraction of total tax-loss selling being affected by realizations in December rather than in January. Reinganum (1982, p. 9) argues that if investors do not follow a year-end policy, they 'might accumulate short-term losses in excess of the (legal) amount which can be deducted from income', but this argument does not explain why the investor would not move his or her tax-loss sales back slightly to avoid the rush. Although Reinganum (1983) finds a negative relation between January returns and their percent change in the immediately preceding months, Keim (1982) finds this relation in *all* calendar months, not just January.

There may also be tax-related selling of large as well as small firm stocks. If *marginal* transactions costs are low or zero, it is not obvious why losses on the stocks of larger firms would not also be realized. In fact, costs associated with realizing a given dollar loss may be higher for small firm securities. This is particularly important since the limit of $3,000 on deductions for capital losses may be binding, especially for larger portfolios. Although losses can be carried forward, an investor may not want to sell only big losers (small stocks) since that would ignore possible diversification losses, and reduce the present value of the losses. In addition, it is not clear that small firms will necessarily provide the big losses, since even though they tend to have high return variance, they also typically have high *expected* returns.

Finally, Constantinides (1982) shows that, under certain conditions, some taxable investors will also find it optimal to realize capital gains on small firm stocks as well as, or instead of, capital losses, indicating that the emphasis on tax-*loss* selling is misplaced. In short, the U.S. tax position is somewhat ambiguous, and certainly does not necessarily imply the year-end price pressure on small firm stock prices relied on in the tax-loss hypothesis. Further, such pressure on prices in December does not show up in the U.S. data.

3. Australian taxes

The Australian tax year is from July 1 to June 30 for all but a few Australian taxpayers. Capital gains/losses are either treated as ordinary income/deductions and taxed at the ordinary rate (in general, if the taxpayer is classified as a 'share trader'), or are not subject to taxes at all (in general, if the taxpayer is an 'investor').

All tax-paying financial institutions [e.g., insurance companies, banks, and non-exempt superannuation (pension) funds] are automatically classed as share traders. Share traders pay taxes at normal rates on all realized gains and deduct all realized losses, regardless of the length of holding period. Other taxpayers may choose to nominate all or part of a portfolio as being held for trading purposes, although this does not bind the Taxation Commissioner.

The Commissioner is more likely to be persuaded that shares are held for investment purposes if they pay dividends and are not turned over frequently. Investors are not taxed on gains, nor can they deduct losses, on shares held for a year or more. Gains on shares held for less than a year are taxed, and a recent court decision[4] denies a deduction for losses on shares held less than a year. Isolated 'speculative' transactions by investors, for example, purchases of shares in non-dividend-paying oil exploration companies, attract tax on any gains. Losses normally will be deductible if the investor promptly notifies the taxation authorities of the purchase.

Since the rationale for a tax effect comes from the ability of capital losses to directly offset ordinary income, 'traders' in Australia appear to have similar incentives for June tax-loss selling as exist in the U.S. in December. In fact, the incentive may be stronger since they are not limited to the $3,000 deduction.

Certainly there appears to be no reason to predict a January seasonal based on Australian tax law. An alternative possibility is that if shares are accessible to both Australian and U.S. investors, then a January seasonal in the U.S. may show up in Australia to prevent arbitrage. However, this would create another arbitrage possbility for those not subject to U.S. taxes, which extends the class of those able to profit from any tax-selling induced price changes. Further, the securities of main interest, those of small firms, are unlikely to be of primary interest to foreign investors. The same arguments would apply to Australian listed stocks held by U.S. investors.

4. Data

4.1. General

The data base used is a merged version of three monthly data files. In all cases, only the most senior, ordinary share is included in the merged data base of 1924 shares. The first file, Brown's '$N=909$' file, covers the period

[4]F.C.T. v. Werchon and Anor. (1982) 82 A.T.C. 4332.

January 1958 to December 1973, and comprises all 'industrial' shares listed on an Australian stock exchange, with par value not less than $1 million (Aust.). The second 'AGSM/CRA' file covers January 1958 through February 1979, and comprises all listed mining and oil shares where the company's main operations are in Australia. The third file, an early version of the AGSM Share Data File, begins January 1974, is current, and comprises all listed Australian shares. The merged file is used to construct a value-weighted market index, with the weights based on market values of equity in the previous month. A file of riskless rates of return is constructed from Australian Treasury Bills.[5]

Ten size-ranked portfolios have been constructed, with updating each month t based on market value of equity deciles in month $t-2$. Consequently, the first rate of return available for these portfolios is in March 1958. The rates of return used in this paper are based on discrete compounding,[6] and returns are equally-weighted within each of the ten size-ranked portfolios.

Not all shares were listed for the full period and fewer were traded in every month. The number of shares for which returns are available ranges from a minimum of 281 (February 1958) to a maximum of 937 (April 1974). Two events that increased the number of available securities are capitalization changes that accompanied decimalization of currency in Feburary 1966, and the introduction of Australian Associated Stock Exchange uniform listing in February 1972.

Until January 1974 the smallest shares on the merged file are dominated by mining and oil shares, because of the $A1 million minimum per value imposed on the first industrial file. However, since then the inclusion of all industrial shares adds another set of small stocks. Time series plots of rates of return for the decile of smallest firms reveal that although there are outliers around 1974, the small firm portfolio does not show a readily apparent general increase in variance after 1974, which may indicate that small industrials are not significantly smaller than small mining and oil stocks.

In any event, there is an increase in the number of relatively small stocks in 1974. Further, the data error rate is likely to be higher in the early version of the AGSM Share Data File because it had not been validated to the same

[5]Details of file construction and composition for all files are available from The Australian Graduate School of Management, University of New South Wales, P.O. Box 1, Kensington, N.S.W. 2033, Australia.

[6]We repeated all tests based on continuously compounded returns. For raw returns, the results are essentially unchanged from those reported, although the size premium appears larger in discrete returns. For abnormal returns we used an equally-weighted market index, which removed more of the seasonal effect from small firm portfolios than did the value-weighted index used in reported results.

extent as the other two files.[7] To assess the sensitivity of our results to the structural change in our data base, analysis was conducted on subperiods before and after January 1974, and results are reported where pertinent. Also, robust regression techniques have been used to gauge the sensitivity of our results to extreme returns.[8]

4.2. *Average returns and autocorrelation*

The average rate of return, OLS beta and autocorrelation functions for the value-weighted market index and all ten portfolios are given in table 1 for the overall period March 1958 to June 1981. This table is meant to summarize the Australian data, and allows comparison with the U.S. size effect studies [e.g., Banz (1981) and Reinganum (1981)] and also with the earlier work of Officer (1975) who examined seasonalities in Australian industrial stocks from 1958 to 1970.

The cross-sectional pattern[9] of higher average returns for small firms in table 1 mirrors that reported by Banz (1981) and Reinganum (1981) for the U.S. The size effect, whatever its origin, shows up in the Australian data.[10] A significant feature relative to the statistics for U.S. monthly returns, is the high first-order serial correlation ranging from 0.16 for the largest firms to 0.33 for portfolio 8. These estimates are consistent with Officer (1975, p. 38) who attributes this to more severe non-trading in the Australian market than in the U.S. In contrast to the U.S. data, the autocorrelation is less pronounced at both ends of the size spectrum than for the medium-size firms.[11] The value-weighted market index, which puts primary emphasis on the largest firms, shows similar autocorrelations to portfolio 10.

Officer (1975) finds persistence over different subperiods of significant positive autocorrelation at lags 6 and 9, and significant negative autocorrelation at lags 13 and 14. The latter were regarded as particularly surprising, but he reports that they did not appear to have any predictive

[7]Extreme outliers (price relatives less than 0.2 or greater than 5.0) were verified. There were 87, of which 42 proved to be correct. The rest were deleted.

[8]The robust methods used follow Huber (1964, 1977) and Krasker–Welsch (1979). Outliers are defined relative to both regression and explanatory variable space. All computations have been made using the BIF and BIFMOD programs in the TROLL package [see Peters, Samarov and Welsch (1982)]. The results were not significantly changed.

[9]Note that 'the pattern' does not consist of ten independent observations (since the returns are cross-sectionally dependent), and it must be interpreted with this in mind.

[10]Preliminary tests [Brown, Kleidon and Marsh (1982)] indicate that the stochastic nature of the U.S. size effect (after allowing for a deterministic January seasonal) reported by Brown, Kleidon and Marsh (1983) does not seem to describe the Australian size-ranked portfolios. Any change in the size premium seems more akin to a steady, deterministic drift, although the stochastic component may be obscured by high volatility of Australian returns.

[11]Aside from sampling variation, a possible explanation for this pattern lies in the popularity and, therefore, heavier trading volume of small firm 'penny stocks' relative to (say) the stocks of small NYSE–AMEX firms.

Table 1

Average monthly rates of return, ordinary least squares beta estimates,[a] and correlograms of monthly rates of return for ten portfolios constructed by equally weighting monthly rates of return on all Australian industrial, mining and oil stocks[b] in each decile of size (measured by market value of equity),[c] and for the value-weighted market index of all Australian industrial, mining and oil stocks, over the period March 1958 to June 1981.

Size decile	Average[d] rate of return (%)	OLS beta	Autocorrelation estimates for the monthly rate of return[e]							
			$\hat{\rho}_1$	$\hat{\rho}_2$	$\hat{\rho}_3$	$\hat{\rho}_6$	$\hat{\rho}_9$	$\hat{\rho}_{12}$	$\hat{\rho}_{13}$	$\hat{\rho}_{14}$
Smallest	6.754 (0.637)	1.040	0.18	0.07	0.08	−0.03	0.14	−0.04	−0.12	−0.10
2	2.231 (0.393)	0.868	0.19	0.01	0.03	0.07	0.19	0.06	−0.02	−0.14
3	1.743 (0.305)	0.657	0.18	−0.01	0.14	0.06	0.17	0.02	−0.15	−0.14
4	1.319 (0.271)	0.661	0.28	0.06	0.04	0.07	0.23	0.04	−0.10	−0.20
5	1.476 (0.240)	0.609	0.30	0.11	0.07	0.07	0.23	0.01	−0.14	−0.23
6	1.267 (0.237)	0.613	0.28	0.05	0.07	0.05	0.21	0.07	−0.12	−0.16
7	1.150 (0.239)	0.659	0.31	0.08	0.04	0.10	0.20	0.02	−0.10	−0.18
8	1.221 (0.239)	0.678	0.33	0.07	0.06	0.04	0.21	0.01	−0.13	−0.20
9	1.181 (0.247)	0.759	0.24	0.07	0.13	0.03	0.21	0.03	−0.14	−0.17
Largest	1.023 (0.286)	0.950	0.16	0.01	0.08	−0.03	0.19	0.03	−0.09	−0.22
Index	1.073 (0.291)		0.16	0.03	0.05	0.02	0.22	0.02	−0.08	−0.20

[a]The OLS beta is computed by regressing monthly portfolio rates of return on the monthly rates of return of a value-weighted (market) portfolio over the period March 1958 to June 1981.

[b]The industrial stocks comprise all those listed on an Australian stock exchange, but prior to December 1973, include only those with a par value of at least \$A1 million. The mining and oil stocks comprise all those which are listed and have their main operations in Australia.

[c]Market values of equity used for weighting rates of return in month t are computed two months prior to t, i.e., at the end of month $t-2$, $t=$ March 1958 to June 1981.

[d]The average returns in percent are based on 280 observations for each equally-weighted portfolio of stocks in each size decile over the period March 1958 to June 1981. The standard errors given in parentheses will be biased downward because of the autocorrelation in the rates of return.

[e]The standard error of these autocorrelation estimates is approximately 0.06.

ability. Over the subperiod from January 1974 to June 1981, which is outside Officer's sample period, we find only weak evidence of negative autocorrelation at these lags. Similarly, autocorrelation estimates at lags 6 and 9 vary widely in sign and magnitude across subperiods in our sample.

In short, there appears reliable evidence (and explanation) for positive first-order serial correlation and this shows up in subsequent tests. Although some odd apparent correlation shows up at other lags, it is much less reliable and may be due to sampling variation. Table 1 also contains the sample market model beta estimates for the ten portfolios. The beta estimates are downward-biased due to the non-trading just described, and we report results using Dimson (1979) betas later.

5. Results

5.1. Raw data

Praetz (1973), using spectral analysis, reports apparent seasonalities in the Sydney All Ordinaries Index with peaks in January–February and July–August and troughs in March–April and November–December. Using time series analysis, Officer (1975, p. 46) concludes that 'the results do not give a clear indication of any particular type of seasonality', although there is some evidence of correlation between the returns of March and September. However, his results (1975, table 5, p. 44) may show some effects of a January seasonal, since for his sample index the largest lag 12 serial correlation occurs for the month of January.

To test for seasonality in Australian mean monthly stock returns, we estimate the following dummy variable OLS regression:

$$R_{pt} = a_{p1} + \sum_{i=2}^{12} a_{pi} D_{it} + \varepsilon_{pt}, \qquad p = 1, \ldots, 10, \tag{1}$$

where

R_{pt} = return on portfolio p in month t,
D_{it} = seasonal dummy for calendar month i,
i = February, ..., December.

The intercept a_{p1} indicates average returns for January, and the dummy coefficients a_{pi} indicate the average differences in return between January and each respective month. In table 2, we present estimates of eq. (1) for the overall period March 1958 to June 1981. The results of table 2 have been graphed in fig. 1 for visual assessment. Fig. 3 contains an analogous plot of the returns for the five market value portfolios of NYSE stocks used in Keim (1982) for comparison with U.S. results.

Table 2 and fig. 1 contain several important results. First, the smallest firm portfolio 1 shows higher average returns across all months than do the other portfolios. The results in table 1 indicate an average monthly return of 6.75% for portfolio 1 and 2.23% for portfolio 2. Table 2 and fig. 1 show that this average monthly premium of about 4% seems fairly constant across all months. Although the statistical significance of these results is somewhat difficult to judge, given the high first-order autocorrelation in the residuals shown in table 2 and the cross-sectional dependence in the returns, it seems that the size-related premium for the smallest firms exists in all months. Note also the extra information obtained by the analysis of the size-ranked portfolios vis-à-vis the market index, which looks most like the large-firm portfolio 10.

Second, for all portfolios, the dummy coefficients in table 2 are negative for almost all non-January months, indicating that January does earn apparently higher returns than most other months. The main exceptions are the July coefficients for the four smallest portfolios.

Third, there appears to be more than just a January premium. For portfolio 1, the smallest firms, there is considerable sampling variation across months, so that although (say) January and July appear to earn higher returns than (say) June, the F-statistic reported in table 2 indicates that the

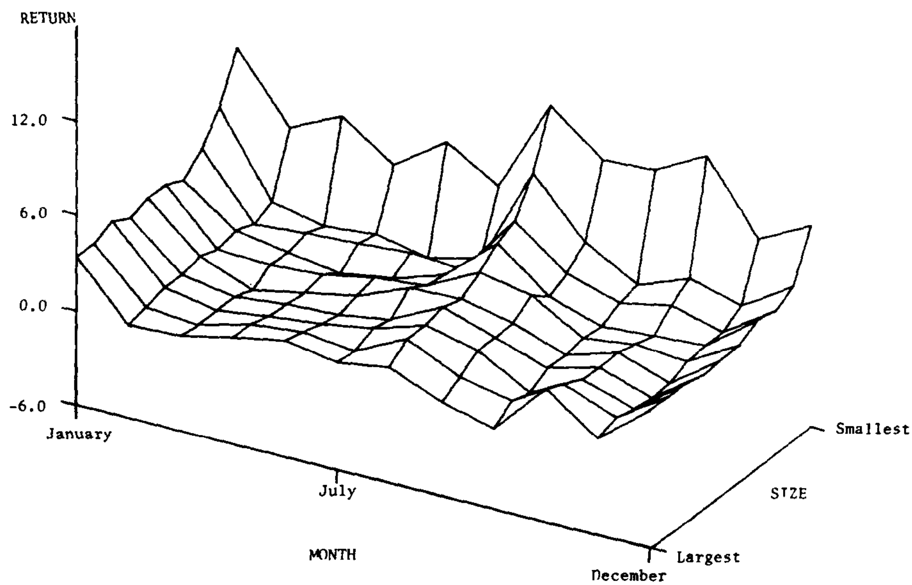

Fig. 1. Plot of the average monthly rates of return (in percent) for ten market value portfolios constructed from Australian industrial and mining stocks for each month over the period March 1958 to June 1981.

Table 2

Estimates of variation in month-by-month average monthly (%) rates of return (t-statistics in parentheses) on ten portfolios constructed by equally weighting monthly rates of return on all Australian industrial, mining and oil stocks[a] in each decile of size (measured by market value of equity),[b] and for the value-weighted market index of all Australian industrial, mining and oil stocks, over the period March 1958 to June 1981.

$$R_{pt} = a_{p1} + \sum_{i=2}^{12} a_{pi} D_{it} + \varepsilon_{pt}, \qquad p = 1, \ldots, 10, \quad t = 1, \ldots, T^{c}$$

Portfolio	Jan.	Feb.	March	April	May	June	July	Aug.	Sept.	Oct.	Nov.	Dec.
Smallest	8.86	−4.18	−2.64	−4.90	−2.61	−4.61	1.34	−1.25	−1.03	0.68	−3.77	−2.06
	(3.67)	(−1.22)	(−0.78)	(−1.45)	(−0.77)	(−1.36)	(0.39)	(−0.37)	(−0.30)	(0.20)	(−1.10)	(−0.60)
2	6.01	−5.17	−5.88	−5.60	−6.28	−5.52	0.77	−2.77	−4.63	−3.40	−4.27	−2.21
	(4.55)	(−2.77)	(−3.18)	(−3.03)	(−3.40)	(−2.98)	(0.41)	(−1.48)	(−2.48)	(−1.82)	(−2.28)	(−1.18)
3	4.27	−3.99	−3.72	−3.58	−3.31	−3.57	0.25	−2.16	−3.17	−2.46	−3.33	−1.16
	(4.08)	(−2.70)	(−2.54)	(−2.45)	(−2.26)	(−2.44)	(0.24)	(−1.46)	(−2.15)	(−1.66)	(−2.25)	(−0.79)
4	3.11	−2.36	−2.96	−3.45	−2.94	−1.87	0.93	−1.60	−2.66	−2.43	−2.19	0.28
	(3.37)	(−1.81)	(−2.29)	(−2.68)	(−2.28)	(−1.45)	(0.72)	(−1.23)	(−2.04)	(−1.86)	(−1.68)	(0.21)
5	3.68	−3.29	−3.41	−3.24	−2.59	−2.27	−0.62	−1.20	−3.10	−2.63	−2.93	−1.05
	(4.48)	(−2.83)	(−2.97)	(−2.82)	(−2.25)	(−1.98)	(−0.53)	(−1.03)	(−2.66)	(−2.27)	(−2.52)	(−0.91)
6	3.70	−2.89	−4.21	−3.54	−2.90	−1.59	−1.46	−2.15	−3.34	−2.33	−3.50	−1.16
	(4.57)	(−2.53)	(−3.71)	(−3.13)	(−2.56)	(−1.41)	(−1.28)	(−1.88)	(−2.92)	(−2.03)	(−3.06)	(−1.01)
7	3.30	−2.76	−3.22	−3.01	−2.82	−2.31	−0.68	−1.80	−2.92	−2.06	−3.14	−0.86
	(4.01)	(−2.37)	(−2.88)	(−2.62)	(−2.45)	(−2.01)	(−0.58)	(−1.55)	(−2.51)	(−1.77)	(−2.70)	(−0.74)
8	3.96	−3.38	−3.99	−3.54	−3.23	−2.69	−1.67	−2.45	−4.28	−2.63	−3.65	−1.24
	(4.87)	(−2.94)	(−3.51)	(−3.11)	(−2.84)	(−2.36)	(−1.45)	(−2.13)	(−3.72)	(−2.29)	(−3.17)	(−1.08)
9	3.39	−3.18	−3.47	−2.61	−2.30	−2.70	−1.12	−2.69	−3.38	−1.47	−2.81	−0.63
	(3.98)	(−2.64)	(−2.92)	(−2.19)	(−1.94)	(−2.27)	(−0.93)	(−2.24)	(−2.81)	(−1.22)	(−2.33)	(−0.53)
Largest	3.39	−3.47	−3.31	−2.58	−1.97	−2.56	−2.03	−3.42	−4.29	−0.91	−3.24	−0.61
	(3.44)	(−2.49)	(−2.41)	(−1.87)	(−1.43)	(−1.86)	(−1.46)	(−2.45)	(−3.08)	(−0.65)	(−2.33)	(−0.44)
Market	3.14	−3.22	−2.97	−1.90	−2.16	−2.19	−1.90	−2.80	−4.21	−0.91	−2.86	0.31
	(3.14)	(−2.27)	(−2.12)	(−1.35)	(−1.54)	(−1.56)	(−1.34)	(−1.97)	(−2.97)	(−0.64)	(−2.02)	(0.22)

	Autocorrelations of $\hat{\varepsilon}_{pt}$[d]								
	$\hat{\rho}_1$	$\hat{\rho}_6$	$\hat{\rho}_9$	$\hat{\rho}_{12}$	$\hat{\rho}_{13}$	$\hat{\rho}_{14}$	DW	F-stat.[e]	Prob$>F$[f]
Smallest	0.22	−0.03	0.16	−0.08	−0.13	−0.12	1.57	0.75	0.686
2	0.20	0.01	0.26	−0.05	−0.04	−0.12	1.60	3.09	0.001
3	0.20	0.02	0.21	−0.06	−0.17	−0.13	1.61	2.02	0.027
4	0.29	0.02	0.33	−0.06	−0.14	−0.19	1.42	2.39	0.008
5	0.33	0.03	0.31	−0.07	−0.18	−0.23	1.34	2.05	0.025
6	0.30	−0.01	0.30	−0.02	−0.16	−0.14	1.40	2.26	0.012
7	0.33	0.06	0.26	−0.06	−0.13	−0.16	1.34	1.78	0.053
8	0.36	−0.01	0.28	−0.09	−0.17	−0.19	1.28	2.36	0.009
9	0.27	−0.01	0.24	−0.05	−0.16	−0.16	1.46	1.77	0.059
Largest	0.18	−0.06	0.21	−0.04	−0.10	−0.23	1.65	1.77	0.059
Market	0.18	−0.00	0.25	−0.06	−0.09	−0.21	1.64	1.75	0.064

[a]The industrial stocks comprise all those listed on an Australian stock exchange, but prior to December 1973, include only those with a par value of at least $A1 million. The mining and oil stocks comprise all those which are listed and have their main operations in Australia.

[b]Size rankings are up-dated monthly and are based on market values of equity two months earlier, i.e., in month t, portfolios are formed using equity values at the end of month $t-2$, $t=1,\ldots,T$.

[c]For each portfolio p, the intercept coefficient $\hat{a}_{p1}$ is an estimate of that portfolio's average rate of return in January. The t-statistic for $\hat{a}_{p1}$, in parentheses under that coefficient estimate, is for the test of whether January's average rate of return is significantly different from zero. The coefficients $\hat{a}_{p2},\hat{a}_{p3},\ldots,\hat{a}_{p12}$ are estimates of the differences in average rates of return in February$(=D_{2t})$, March $(=D_{3t}),\ldots$, December$(=D_{12t})$, from that in January. The t-statistics under $\hat{a}_{p2},\hat{a}_{p3},\ldots,\hat{a}_{p12}$ test whether there is significant variation in the February, March, …, December average rates of return relative to that of January.

[d]The standard error of these autocorrelation coefficients is approximately 0.06.

[e]The F-statistic tests the hypothesis that a_{p2} through a_{p12} are jointly equal to zero for each $p=1,\ldots,10$: $F_{11,268}(95\%)=1.79$; $F_{11,268}(99\%)=2.25$.

[f]The Prob$>F$-statistic is the p-value for the F-test, i.e., it is the area to the right of the F-statistic on the $F(11,268)$ distribution. A sampling theorist would reject the null hypothesis that a_{p2} through a_{p12} jointly equal zero at the $\alpha\%$ level if (Prob$>F)<\alpha$.

null hypothesis of no difference across months cannot be rejected for portfolio 1 at conventional significance levels. However, the null hypothesis is rejected at the 5% level for portfolios 2 through 6 and portfolio 8, and barely accepted at this level for portfolios 7, 9 and 10. For most portfolios, two strong seasonals emerge. December–January and July–August appear to earn consistently higher returns than do other months, with the largest returns in January and July. For example, for portfolio 2, the average December–January and July–August returns are 4.91% and 5.01%, respectively, while the average for other months is 0.87%. (Interestingly, fig. 3 shows that the largest average returns in the U.S. data also occur in January and July.) The January and the July–August seasonals reported by Praetz (1973) for a market index are confirmed across size-ranked portfolios, but we find little evidence of his reported December trough and February peak.

Examination of the returns before and after December 1973 yields subperiod results whose interpretation is qualitatively equivalent to that for the overall period.[12] The primary difference is that the sample from January 1974 to June 1981 contains more small firms, and average returns for the smaller firm portfolios during this period are much larger than returns from March 1958 to December 1973. For example, portfolio 1 returns are, on average, 9.84% per month in contrast to an average of 5.29% during the earlier subperiod. For the largest firm portfolio 10, the average first and second subperiod returns differ by only 0.94%.

In summary, analysis of the raw returns indicates a strong 'size effect', and the average small firm premium appears to be roughly equal across all months. Second, there are pronounced December–January and July–August seasonals in average portfolio returns. The tax-loss selling hypothesis predicts a July seasonal based on Australian tax law, and hence we conclude that the Australian evidence is inconsistent with at least the usual form of the tax-loss hypothesis. Even if one invoked an integrated capital markets hypothesis to explain the January seasonal as a reflection of U.S. tax-loss selling, neither story explains the December and August results.

5.2. Excess returns relative to the CAPM

In the previous section, we examined the seasonality of raw returns on Australian stocks. In this section, we discuss the seasonality of abnormal returns relative to the one-period Sharpe–Lintner CAPM. In place of (1), the model is

$$R_{pt}-R_{Ft}=a'_{p1}+\beta_p(R_{Mt}-R_{Ft})+\sum_{i=2}^{12}a'_{pi}D_{it}+\varepsilon'_{pt}, \quad p=1,\ldots,10, \tag{2}$$

[12]However, the second subperiod has only eight observations per month, and shows greater sample variation than does the earlier period.

where the additional variables are R_{Ft}, the riskless rate of return, and R_{Mt}, the rate of return on the value-weighted index. The coefficients a'_{p1}, $a'_{p2}, \ldots, a'_{p12}$ now detect any seasonality of average abnormal returns (relative to the Sharpe–Lintner model). If the abnormal returns display a seasonal pattern and markets are efficient, there is some systematic seasonal risk not completely accounted for by the CAPM, or perhaps the beta measure of risk itself is not seasonally invariant.

In table 3 and fig. 2 we present estimates of (2) for the overall period March 1958 to June 1981. There seems to be little serial correlation left in the residuals, consistent with the removal of the autocorrelated market returns reported in table 1. Statistically, the lack of residual autocorrelation means that the test statistics are more reliable in table 3 than in table 2.

Table 3 and fig. 2 show that a 'size effect' similar to that in the raw returns is also found in excess returns relative to the Sharpe–Lintner model. For portfolios 1 and 2, average monthly excess returns (across all months) are 5.66% and 1.27%, respectively. Again, the premium of about 4% per month for the smallest firms seems fairly even across months (see fig. 2).

Although average excess returns in January and July are larger than most other months in table 3, the effect is not as pronounced as for the raw returns in table 2. For example, the F-statistics in table 3 indicate that we

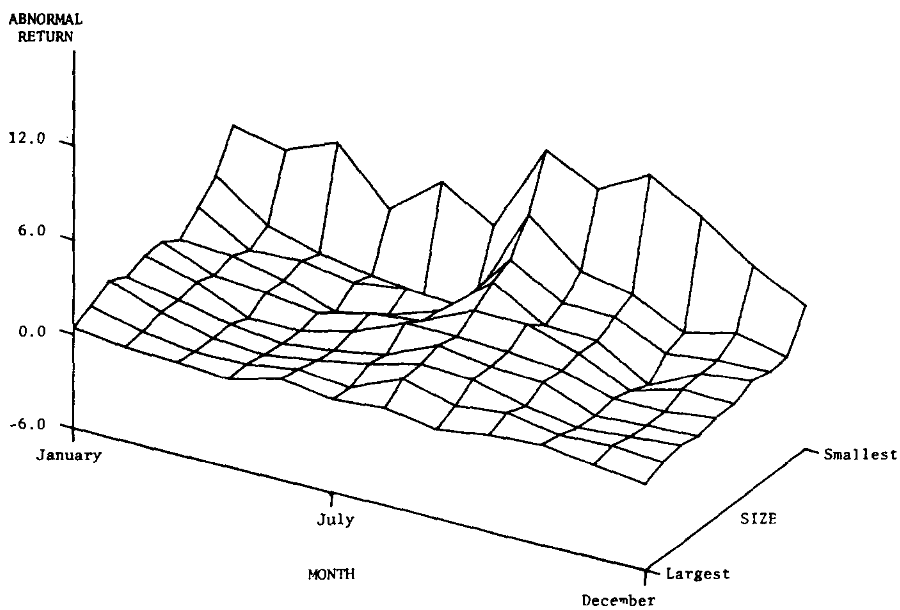

Fig. 2. Plot of the average monthly abnormal rates of return (in percent) for ten market value portfolios constructed from Australian industrial and mining stocks for each month over the period March 1958 to June 1981.

Table 3

Estimates of variation in month-by-month excess monthly (%) rates of return (t-statistics in parentheses), relative to the Sharpe–Lintner capital asset pricing model, on ten portfolios constructed by equally weighting monthly rates of return on all Australian industrial, mining and oil stocks[a] in each decile of size (measured by market value of equity),[b] over the period March 1958 to June 1981.

$$(R_{pt}-R_{Ft})=a'_{p1}+\beta_p(R_{Mt}-R_{Ft})+\sum_{i=2}^{12}a'_{pi}D_{it}+\varepsilon_{pt},\qquad p=1,\dots,10,\quad t=1,\dots,T^{c}$$

Portfolio	Jan.	$\hat{\beta}_p{}^{d}$	Feb.	March	April	May	June	July	Aug.	Sept.	Oct.	Nov.	Dec.
Smallest	5.54	1.07	−0.74	0.54	−2.87	−0.31	−2.27	3.37	1.74	3.47	1.65	−0.70	−2.40
	(2.53)	(0.53)	(−0.24)	(0.18)	(−0.95)	(−0.10)	(−0.75)	(1.10)	(0.57)	(1.12)	(0.54)	(−0.23)	(−0.78)
2	3.24	0.86	−2.40	−3.34	−3.97	−4.43	−3.64	2.40	−0.37	−1.02	−2.63	−1.81	−2.48
	(3.20)	(−2.42)	(−1.68)	(−2.37)	(−2.83)	(−3.16)	(−2.59)	(1.70)	(−0.26)	(−0.71)	(−1.86)	(−1.27)	(−1.75)
3	2.05	0.65	−1.89	−1.79	−2.34	−1.91	−2.15	1.59	−0.33	−0.43	−1.87	−1.47	−1.37
	(2.49)	(−7.03)	(−1.62)	(−1.55)	(−2.05)	(−1.67)	(−1.88)	(1.37)	(−0.29)	(−0.37)	(−1.62)	(−1.26)	(−1.19)
4	0.86	0.66	−0.22	−0.99	−2.19	−1.52	−0.42	2.19	0.25	0.13	−1.83	−0.29	0.07
	(1.33)	(−8.87)	(−0.24)	(−1.11)	(−2.46)	(−1.70)	(−0.47)	(2.43)	(0.28)	(0.15)	(−2.03)	(−0.32)	(0.07)
5	1.57	0.61	−1.31	−1.61	−2.08	−1.28	−0.94	0.54	0.51	−0.52	−2.08	−1.18	−1.24
	(2.85)	(−11.69)	(−1.69)	(−2.09)	(−2.71)	(−1.67)	(−1.23)	(0.70)	(0.66)	(−0.67)	(−2.69)	(−1.52)	(−1.61)
6	1.60	0.61	−0.93	−2.40	−2.39	−1.60	−0.27	−0.30	−0.45	−0.78	−1.78	−1.75	−1.35
	(2.99)	(−12.00)	(−1.23)	(−3.22)	(−3.22)	(−2.15)	(−0.36)	(−0.41)	(−0.60)	(−1.02)	(−2.37)	(−2.33)	(−1.80)
7	1.06	0.66	−0.62	−1.35	−1.75	−1.40	−0.87	0.58	0.05	−0.14	−1.46	−1.24	−1.07
	(2.16)	(−11.50)	(−0.89)	(−1.98)	(−2.58)	(−2.06)	(−1.27)	(0.84)	(0.08)	(−0.19)	(−2.13)	(−1.80)	(−1.56)
8	1.70	0.67	−1.22	−2.01	−2.27	−1.80	−1.23	−0.40	−0.58	−1.46	−2.03	−1.73	−1.45
	(3.66)	−11.75)	(−1.87)	(−3.11)	(−3.52)	(−2.79)	(−1.90)	(−0.61)	(−0.89)	(−2.22)	(−3.12)	(−2.65)	(−2.24)
9	0.89	0.76	−0.74	−1.23	−1.18	−0.68	−1.04	0.31	−0.58	−0.19	−0.78	−0.64	−0.87
	(2.31)	(−10.23)	(−1.36)	(−2.28)	(−2.19)	(−1.27)	(−1.94)	(0.58)	(−1.06)	(0.35)	(−1.45)	(−1.18)	(−1.61)
Largest	0.39	0.95	−0.41	−0.50	−0.78	0.07	−0.48	−0.23	−0.76	−0.30	−0.05	−0.52	−0.91
	(1.50)	(−3.26)	(−1.14)	(−1.38)	(−2.18)	(0.20)	(−1.35)	(−0.64)	(−2.11)	(−0.82)	(−0.14)	(−1.44)	(−2.52)

Portfolio	Autocorrelations of $\hat{\varepsilon}_{pt}$[e] $\hat{\rho}_1$	$\hat{\rho}_6$	$\hat{\rho}_9$	$\hat{\rho}_{12}$	$\hat{\rho}_{13}$	$\hat{\rho}_{14}$	DW	F-stat.[f]	Prob $> F$[g]
Smallest	0.10	0.02	0.08	−0.04	0.02	−0.08	1.80	0.97	0.477
2	0.01	0.05	0.06	−0.06	0.11	−0.05	1.98	3.89	0.000
3	−0.02	−0.05	−0.00	−0.05	−0.02	−0.11	2.04	2.02	0.027
4	−0.02	−0.00	0.13	−0.07	0.03	−0.10	2.05	3.35	0.000
5	0.03	0.02	0.17	−0.09	−0.03	−0.16	1.93	2.71	0.002
6	0.07	0.03	0.15	−0.04	0.01	−0.12	1.85	2.52	0.005
7	0.01	0.12	0.12	−0.15	0.08	−0.15	1.97	2.36	0.009
8	0.16	−0.01	0.11	−0.12	−0.03	−0.16	1.68	2.38	0.008
9	0.12	0.11	0.05	−0.06	−0.06	−0.11	1.75	1.52	0.123
Largest	−0.00	0.00	0.05	−0.05	−0.06	−0.02	2.00	1.58	0.104

[a]The industrial stocks comprise all those listed on an Australian stock exchange, but prior to December 1973, include only those with a par value of at least $A1 million. The mining and oil stocks comprise all those which are listed and have their main operations in Australia.

[b]Size rankings are up-dated monthly, and are based on market values of equity two months earlier, i.e., in month t, portfolios are formed using equity values at the end of month $t-2$, $t=1,\dots,T$.

[c]For each portfolio p, the intercept coefficient $\hat{a}'_{p1}$ is an estimate of that portfolio's average excess rate of return in January. The t-statistic for $\hat{a}'_{p1}$, in parentheses under that coefficient's estimate, is for the test of whether January's average excess rate of return is significantly different from zero. The coefficients $\hat{a}'_{p2},\hat{a}'_{p3},\dots,\hat{a}'_{p12}$ are estimates of the differences in average excess rates of return in February$(=D_{2t})$, March$(=D_{3t})$, …, December$(=D_{12t})$, from that in January. The t-statistics under $\hat{a}'_{p2},\hat{a}'_{p3},\dots,\hat{a}'_{p12}$ test whether there is significant variation in the February, March, …, December average excess rates of return relative to that of January.

[d]The t-statistics for beta are for the null hypothesis $\beta_p=1$.

[e]The standard error of these autocorrelation coefficients is approximately 0.06.

[f]The F-statistic tests the hypothesis that a'_{p2} through a'_{p12} are jointly equal to zero for each $p=1,\dots,10$: $F_{11,268}(95\%)=1.79$; $F_{11,268}(99\%)=2.25$.

[g]The Prob $>F$-statistic is the p-value for the F-test, i.e., it is the area to the right of the F-statistic on the $F(11,268)$ distribution. A sampling theorist would reject the null hypothesis that a'_{p2} through a'_{p12} jointly equal zero at the $\alpha\%$ level if $(\text{Prob}>F)<\alpha$.

cannot reject at conventional significance levels the hypothesis of equal excess returns across months for portfolios 9 and 10. As for the raw returns, despite relatively large absolute differences in the point estimates of average excess returns across months for portfolio 1, the sampling variation is sufficiently high that the null hypothesis of equal excess returns for all months is accepted for the smallest firms. For portfolios 2–8, the F-statistics indicate rejection of the null hypothesis, with the largest average excess returns in January and July.[13] The major implication of table 3 and fig. 2 is that although the levels of average excess returns are generally highest around the months of January and July, the size-related premium for small firms is fairly constant across months (this is readily apparent in fig. 2). This

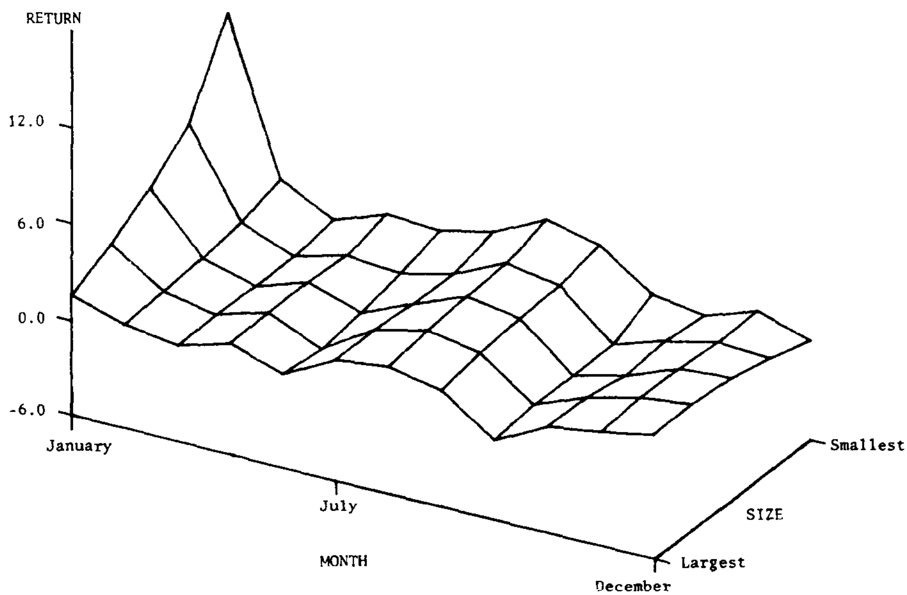

Fig. 3. Plot of the average monthly rates of return (in percent) for five market value portfolios constructed by Keim (1982) from firms on the NYSE for each month over the period January 1931 to December 1978.

[13]Although the seasonals for January, July and August show up in excess returns as in raw returns, the excess returns are influenced by the seasonal pattern for the market index, and one effect is that in general the December seasonal disappears. The raw returns in table 2 show that the market index has higher average returns in December than in any other month (December has the only positive dummy coefficient, 0.31%). Even for the (equally-weighted) largest firm portfolio 10, the December dummy coefficient is negative. Portfolio returns are equally-weighted, so large December returns for the largest firms in portfolio 10 can result in a negative dummy coefficient for portfolio 10 and a positive coefficient for the value-weighted market index. [For an equally-weighted index (of continuously compounded returns), the December dummy coefficient is negative.] The removal of the large December returns for the market index leaves no residual December seasonal in excess returns. Similar effects can be seen in, say, January and July. The index shows a negative July dummy coefficient in table 2, and so a positive coefficient for July relative to January in the small firm portfolios is accentuated in the excess returns in table 3.

finding is in contrast to the strong January seasonal in the size-related premium in the U.S. data in fig. 3.

The subperiods show basically similar results to the overall period, although again the size effect is more pronounced in the second period January 1974 to June 1981 in which there are more small firms.

We have also fitted a variation of (2) which examines the January and July abnormal returns separately,

$$R_{pt}-R_{Ft}=\alpha_p+\beta_p(R_{Mt}-R_{Ft})+\gamma_{p,1}D_t^{\mathrm{Jan}}+\gamma_{p,2}D_t^{\mathrm{Jul}}+e_{pt}, \qquad p=1,\ldots,10, \quad t=1,\ldots,T, \tag{3}$$

Table 4

Estimates of average excess monthly (%) rate of return (t-statistics in parentheses) for the months of January and July and for all other calendar months combined, relative to the Sharpe–Lintner capital asset pricing model, on ten portfolios constructed by equally weighting monthly rates of return on all Australian industrial, mining and oil stocks[a] in each decile of size (measured by market value of equity),[b] over the period March 1958 to June 1981.

Part A: Estimation with OLS betas

$$R_{pt}-R_{Ft}=\alpha_p+\beta_p(R_{Mt}-R_{Ft})+\gamma_{p,1}D_t^{\mathrm{Jan}}+\gamma_{p,2}D_t^{\mathrm{Jul}}+e_{pt}, \qquad p=1,\ldots,10, \quad t=1,\ldots,T.$$

Portfolio	$\hat{\alpha}_p$	$\hat{\gamma}_{p,1}$	$\hat{\gamma}_{p,2}$	$\hat{\beta}_p$[c]
Smallest	5.37	0.29	3.57	1.02
	(7.85)	(0.13)	(1.57)	(7.95)
2	0.62	2.67	5.04	0.84
	(1.94)	(2.49)	(4.74)	(13.93)
3	0.49	1.59	3.16	0.64
	(1.90)	(1.85)	(3.70)	(13.21)
4	0.14	0.74	2.92	0.65
	(0.69)	(1.08)	(4.30)	(16.94)
5	0.41	1.20	1.71	0.60
	(2.34)	(2.04)	(2.95)	(18.03)
6	0.21	1.40	1.09	0.60
	(1.24)	(2.46)	(1.93)	(18.75)
7	0.06	1.03	1.59	0.65
	(0.39)	(1.98)	(3.09)	(22.29)
8	0.11	1.60	1.20	0.66
	(0.75)	(3.28)	(2.47)	(24.18)
9	0.10	0.81	1.11	0.75
	(0.82)	(2.01)	(2.79)	(33.10)
Largest	−0.07	0.47	0.23	0.95
	(−0.88)	(1.71)	(0.85)	(61.76)

Table 4 (continued)

Part B: Estimation with Dimson betas

$$R_{pt}-R_{Ft}=\alpha'_p+\sum_{k=-1}^{+1}\beta'_{p,k}(R_{m,t+k}-R_{F,t+k})+\gamma'_{p,1}D_t^{\mathrm{Jan}}+\gamma'_{p,2}D_t^{\mathrm{Jul}}+U_{pt},$$

$$p=1,\ldots,10,\quad t=1,\ldots,T.$$

Portfolio	$\hat{\alpha}'_p$	$\hat{\gamma}'_{p,1}$	$\hat{\gamma}'_{p,2}$	$\hat{\beta}'_{p,-1}$	$\hat{\beta}'_{p,0}$	$\hat{\beta}'_{p,1}$
Smallest	5.21	−0.72	3.51	0.47	0.95	0.02
	(7.69)	(−0.32)	(1.57)	(3.64)	(7.34)	(0.18)
2	0.56	2.03	4.96	0.26	0.81	−0.03
	(1.79)	(1.90)	(4.81)	(4.39)	(13.44)	(−0.49)
3	0.42	1.13	3.13	0.21	0.61	0.01
	(1.66)	(1.34)	(3.77)	(4.42)	(12.60)	(0.19)
4	0.04	0.27	2.92	0.24	0.61	0.04
	(0.20)	(0.42)	(4.60)	(6.52)	(16.57)	(1.09)
5	0.34	0.70	1.68	0.22	0.56	0.00
	(2.08)	(1.28)	(3.13)	(7.23)	(18.04)	(0.07)
6	0.15	1.04	1.07	0.17	0.58	0.01
	(0.93)	(1.89)	(1.97)	(5.33)	(18.31)	(0.25)
7	−0.01	0.58	1.55	0.20	0.62	0.00
	(−0.04)	(1.22)	(3.30)	(7.39)	(22.61)	(0.16)
8	0.05	1.26	1.18	0.16	0.64	0.02
	(0.34)	(2.72)	(2.60)	(6.19)	(24.08)	(0.62)
9	0.07	0.57	1.09	0.10	0.74	−0.01
	(0.62)	(1.44)	(2.81)	(4.50)	(32.71)	(−0.35)
Largest	−0.07	0.39	0.21	0.02	0.94	−0.02
	(−0.81)	(1.41)	(0.78)	(1.45)	(60.26)	(−1.04)

[a]The industrial stocks comprise all those listed on an Australian stock exchange, but prior to December 1973, include only those with a par value of at least $A1 million. The mining and oil stocks comprise all those which are listed and have their main operations in Australia.

[b]Size rankings are up-dated monthly, and are based on market values of equity two months earlier, i.e., in month t, portfolios are formed using equity values at the end of month $t-2$, $t=1,\ldots,T$.

[c]To maintain comparability with the Dimson beta estimates, the t-statistics for the OLS betas in this table are for the null hypothesis $\beta_p=0$.

where D_t^{Jan} and D_t^{Jul} are defined as seasonal dummies for January and July. In addition to model (3), which we estimated with OLS, we estimated the following time series model employing Dimson's (1979) estimator for beta:

$$R_{pt}-R_{Ft}=\alpha'_p+\sum_{k=-1}^{+1}\beta'_{p,k}(R_{m,t+k}-R_{F,t+k})+\gamma'_{p,1}D_t^{\mathrm{Jan}}+\gamma'_{p,2}D_t^{\mathrm{Jul}}+U_{pt},\qquad p=1,\ldots,10,\quad t=1,\ldots,T. \tag{4}$$

The results are given in table 4. As in table 3, the removal of market effects by model (3) accentuates the July seasonal relative to January, and this is more pronounced using Dimson's estimator in model (4).[14] However, even the use of the Dimson estimator and January and July dummies does not remove the 'size effect' — the average abnormal monthly returns across other months for portfolio 1 still exceed abnormal returns for portfolio 2 by over 4%.

6. Summary and conclusion

Evidence from U.S. stock returns suggests that a large proportion of the 'size effect' consists of a premium for small firms in January. Australian returns show an average premium of at least 4% per month for the smallest-firm decile (portfolio 1) relative to any other decile, and this premium appears to be fairly constant across months in contrast to U.S. data.

While recent research has attributed the U.S. January premium in large part to tax-loss selling, we have argued that there are persuasive a priori reasons for questioning the importance of tax-loss selling as an explanation. In Australia where the tax year-end is June 30, we should expect a perhaps even stronger 'July effect' under that hypothesis because, for market participants classified as 'traders', there is no $3,000 limit for the loss deduction. What we find is more complex than the simple tax-loss selling hypothesis predicts. Pronounced seasonals (in raw returns) occur in December–January and July–August, with the largest (and roughly equal) effects in January and July.

It is conceivable that a tax-induced January effect in the U.S. could show up in Australian data due to arbitrage across capital markets. But this response is, at best, incomplete. The tax-loss selling hypothesis itself relies on an *absence* of arbitrage from those not forced to sell a particular security for tax purposes. It seems difficult to reconcile an integrated capital market, that functions so well that a U.S. tax-induced January seasonal shows up in even penny stocks in Australia, with simultaneous mispricing of securities to create the original U.S. January seasonal.

Moreover, if the January and July premiums are induced by U.S. and Australian tax law, what is the cause of the December and August premiums? For August, one might argue that non-trading causes the July effect to show up in August returns, or that the tax-effect is somewhat continuous and *should* appear in months subsequent to July. But if so, why is there no apparent February seasonal in both Australia and the U.S., and what explains the December results?

[14]For all portfolios other than portfolio 10, the coefficient on the lagged market index appears significant. Preliminary work showed that lags greater than 1, and leads, were not significantly different from zero.

On balance, we regard the Australian evidence as difficult to reconcile with the tax-loss selling hypothesis. Although the original hypothesis is at least consistent with the U.S. January premium, the story seems to be much more complicated if it is to be reconciled with the Australian data. It seems more promising to investigate possible equilibrium causes of the U.S. January seasonal, perhaps by comparing the relative timing of events other than tax years in Australia and the U.S.

This conclusion is further supported by Keim's (1982) finding that the January effect is a significant phenomenon in every year (except two) from 1931 to 1979, even during periods (e.g., prior to World War II) when personal tax rates were relatively low and when the benefit of the capital loss offset was lower. The magnitude of the January effect does not seem sensitive to variations in the tax rate. In one sentence, tax-loss selling still leaves us at a loss for an explanation of the January effect.

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