"SIZE-SORTED PORTFOLIOS AND THE VIOLATION OF THE RANDOM WALK HYPOTHESIS:ADDITIONAL EMPIRICAL EVIDENCE AND IMPLICATION FOR TESTS OF ASSET PRICING MODELS"

by
Pierre H. HILLION*

N° 88 / 56

* Pierre H. HILLION, Assistant Professor of Finance, INSEAD
Fontainebleau, France

Director of Publication :

Charles WYPLOSZ, Associate Dean
for Research and Development

Printed at INSEAD,
Fontainebleau, France

Size–sorted portfolios and the violation of the random walk hypothesis: Additional empirical evidence and implication for tests of asset pricing models

Pierre H. Hillion

INSEAD

June 1988

I would like to thank my dissertation committee M. Brennan, T. Cameron, M. Grinblatt, Ker–Chau Li, R. Masulis, W. Torous, and especially my Committee Chair, S. Titman, for offering me helpful guidance and comments. All remaining errors are mine.

Abstract

The violation of the random walk hypothesis is documented in an increasing number of papers. This study provides additional evidence that risk–adjusted monthly returns do not follow a random walk. The goals of the paper are fourfold. It 1) documents the serial correlation displayed by the monthly returns as well as risk–adjusted returns on size–sorted portfolios, 2) determines its origins, 3) investigates its effects on the tests of asset pricing models and mean–variance efficiency, and 4) explains why the presence of serial correlation in the risk–adjusted monthly returns on small firm portfolios does not necessarily imply that the stock market is inefficient.

1 Introduction

The violation of the random walk hypothesis is documented in an increasing number of papers. Fama and French (1988) present evidence that stock prices have temporary components that are slowly eliminated. Long holding–period returns are negatively serially correlated. Their results indicate that 25% to 40% of the variation of longer horizon returns is predictable from past returns. Lo and MacKinlay (1987) reject the random walk hypothesis for weekly stock market returns. In contrast to the negative serial correlation found by Fama and French (1988) for longer horizon returns, they find significant positive serial correlation. The rejections are largely due to the behavior of small stocks. However, they conclude that there is weak evidence against the random walk hypothesis using monthly data. There seems to be an agreement that security monthly returns are not systematically related to their past returns at least in the short run horizon.

The empirical evidence of Jegadeesh (1987) is one exception. He finds that the monthly risk–adjusted returns of NYSE stocks are systematically related to their history of returns as far as thirty six months. This paper provides additional evidence that risk–adjusted monthly returns do not follow a random walk. The risk–adjusted monthly returns on small firm portfolios exhibit statistically significant positive serial correlation. This contrasts with Lo and MacKinlay's (1987) conclusion. The goals of the paper are fourfold. The paper 1) documents the serial correlation displayed by the monthly returns as well as risk–adjusted returns on size–sorted portfolios, 2) determines its origins, 3) investigates its effects on the tests of mean–variance (MV) efficiency, and 4) explains why the presence of se-

rial correlation in the risk–adjusted monthly returns on small firm portfolios is not necessarily inconsistent with the market efficiency hypothesis.

Consistent with previous studies, the first–order serial correlation estimate is found to be statistically insignificant for all the size–sorted portfolios. However, the twelve–order serial correlation is positive and statistically significant. Further, the twelve order serial correlation displayed by the monthly returns and risk–adjusted returns on the size–sorted portfolios are decreasing with firm size. The serial correlation is statistically significant for the smallest market value portfolios. Depending on the market index, the point estimates are in the .30 to .40 bracket. The twelve–order serial correlation is essentially related to the January seasonal.

The finding of serial correlation has important implications for the tests of MV efficiency performed on size–sorted portfolios. In particular, the use of a maximum likelihood estimator which accounts for the serial autocorrelation indicates that the twelve order serial correlation biases the standard error estimates of the abnormal excess returns, *i.e.*, the intercept of the simple index model. After controlling for the effects of autocorrelation, the OLS standard deviation estimates turn out to be systematically underestimated. The underestimation is serious enough to reverse inferences about the small firm effect.

The presence of twelve–order serial correlation in the risk–adjusted monthly returns on small firm portfolios is not necessarily inconsistent with the efficient market hypothesis. This presence is shown to be consistent with omitted risk factor and more specifically with the violation of the assumption of a stationary return generating process. When the non–stationarity induced by the January seasonal is controlled for, most of the twelve order serial correlation disappears.

The paper is organized as follows. Section 2 presents the tests of MV efficiency under the assumption of serially uncorrelated disturbances. Section 3 describes the data and the portfolio formation process. Section 4 documents the serial correlation displayed by the returns and the risk-adjusted returns on size-sorted portfolios. The econometric implications of serial correlation for the tests of MV efficiency are investigated in section 5. The origins of the twelve order serial correlation are examined in Section 6. Section 7 concludes the paper.

2 Tests of MV efficiency under the hypothesis of serially uncorrelated disturbances

The market model describes an asset return, $\tilde{R}_{it}$, as a linear function of the market return, $\tilde{R}_{mt}$,

$$\tilde{R}_{it} = \iota_T \alpha_i + \tilde{R}_{mt}\beta_i + \tilde{\epsilon}_{it}, \quad t = 1, \ldots, T, \tag{1}$$

where, ι_T is a $(T \times 1)$ vector of ones, $\tilde{R}_{it}$ is the percent return of security i in month t, $\tilde{R}_{mt}$ is the percent return of the market in month t, α_i is the market model intercept, β_i is the usual market beta or systematic risk of security i, and $\tilde{\epsilon}_i$ is an idiosyncratic disturbance normally distributed with mean zero and variance σ^2. It is generally assumed that $E(\epsilon\epsilon') = \sigma^2 I$, where ϵ is the $(T \times 1)$ vector of disturbances and I is the $(T \times T)$ identity matrix. The market model written in terms of risk premia becomes,

$$(\tilde{R}_{it} - R_{Ft}) = \iota_T \alpha_i^* + (\tilde{R}_{mt} - R_{Ft})\beta_i + \tilde{\epsilon}_{it}, \quad t = 1, \ldots, T, \tag{2}$$

where R_{Ft} is the ex-ante riskless rate in period t.

It is well-known that the Sharpe–Lintner (1964,1965) CAPM, which describes a relation between expected return and risk,

$$E(\tilde{R}_{it}) = R_{Ft} + E(\tilde{R}_{mt} - R_{Ft})\beta_i, \tag{3}$$

where E denotes the expected value operator, implies that α_i^* in equation (2) is equal to zero for each asset i. If there is a size anomaly relative to the Sharpe–Lintner (1964,1965) CAPM, α^* is related to firm size. Many studies such as those of Banz (1981) and Reinganum (1981) find a negative relation between α_i^* and size. Further, the empirical evidence of Keim (1983) and Roll (1983) indicates that the average risk–adjusted return to a portfolio of small firms is larger in January and much smaller for the rest of the year. About 67% of the annual return differential occurs in January and about 37% of the size effect occurs during the first trading days of January.

To simplify the notation, the "tildes" are dropped from equation (2), and it is rewritten as,

$$r_{it} = \iota_T \alpha_i + r_{mt}\beta_i + \epsilon_{it}, \quad t = 1, \ldots, T, \tag{4}$$

with $r_{it} = (\tilde{R}_{it} - R_{Ft})$ and $r_{mt} = (\tilde{R}_{mt} - R_{Ft})$. The intercept α_i in (4) is equal to α_i^* in (2), which differs from α_i in (1) by the constant $R_{Ft}(1-\beta_i)$. A test of the CAPM, or more specifically, as shown by Roll (1977), a test of the mean–variance (MV) efficiency of the proxy used as the market index, can be performed by examining the estimates and the statitical significance of the abnormal excess returns, α_i in (4), obtained for each individual portfolio i. Univariate t or F–tests can be used for that purpose. As Gibbons, Ross, and Shanken (1986) point out, it is difficult to draw a proper joint inference across a number of univariate t tests for the statistics may be highly dependent. A possible alternative to the univariate t or F test is

the multivariate F statistic which enables one to test the joint hypothesis that the abnormal excess returns of each portfolio are equal to zero.

Regardless of the statistical approach followed, *i.e.*, univariate or multivariate, the tests are generally performed on portfolios instead of individual securities to obtain efficient estimates. Firm size is used as an instrumental variable in an increasing number of papers. Power considerations justify the choice of firm size in those tests. Firm size seems to be an adequate instrumental variable to simultaneously maximize estimation efficiency and the power of the tests.

3 Description of the data and the portfolio formation process

Stock return data from January 1, 1963 through December 31, 1984 is extracted from the CRSP 1985 daily data files. These returns are then compounded on a daily basis to yield monthly stock returns. This yields a total of 264 monthly observations. The value–weighted (VW) and equally–weighted (EW) monthly indices of NYSE and AMEX listed stocks are used as market proxies.[1] To improve the precision of the regression estimates, asset returns are grouped into twenty portfolios, based on market value of equity at the end of the month. The portfolio formation procedure is as follows. On the preceding month, asset returns are ranked according to the market value of their equity. Those firms are then divided into 20 portfolios, each containing an equal number of stocks. Portfolio P_1 contains the smallest firms while P_{20} contains the largest companies. The returns for these portfolios are

[1] These two indices differ from the CRSP EW and VW monthly indices since the AMEX securities are not included in the construction of the CRSP (monthly) indices.

then collected for the current month. The firms are re-ranked every month and the process is repeated for twenty two years. The tests are performed on four subperiods and on the total period. For purposes of comparison, the subperiods are those of Brown, Kleidon, and Marsh (1983). They extend from 1) January 1963 to December 1968, 2) January 1969 to December 1973, 3) January 1974 to June 1979, and 4) July 1979 to December 1984. They approximately correspond to four subperiods of equal length.

It is of interest to compare the sample used here to the samples employed in past studies. They differ with respect to 1) the characteristics of the firms, 2) the rebalancing method, 3) the missing return requirement and 4) the indices used as market proxies. Both NYSE and AMEX firms are employed to construct the portfolios. The market value of small firm portfolios, especially P_1 is, therefore, smaller and the number of securities in each portfolio is larger than in past studies. Also, the portfolios are rebalanced every month instead of being calculated only once at the end of the previous subperiod. The market value of P_1 is likely to be smaller for that reason. Further, the survivorship bias is less likely to be a serious problem. Unlike past studies which require complete returns in each subperiod, firms are only removed in the months that display a missing return. The smallest firms are, therefore, not eliminated. Finally, the CRSP EW and VW monthly return indices are recomputed to include both the NYSE and AMEX firms.[2]

[2]The EW index is formed by first compounding the individual returns and then forming an equal weighted portfolio.

4 Tests for autocorrelated market model disturbances: Description of the tests and empirical evidence

4.1 Description of the statistical tests

Two tests are suggested to detect autocorrelated disturbances of order p in the stock return generating process.[3] The first test examines the sample autocorrelation,

$$\rho_p = \frac{\sum_{t=1}^{t=T-p} \hat{\epsilon}_t \hat{\epsilon}_{t+p}}{\sum_{t=1}^{t=T} \hat{\epsilon}_t^2} \tag{5}$$

where the $\hat{\epsilon}_t$ are the least squares residuals from the market model. It is well-known that this test assumes that each $\hat{\epsilon}_t$ converges in probability to the corresponding ϵ_t, the true unobservable residual. In small samples, the $\hat{\epsilon}_t$'s will be correlated even if the unobservable ϵ_t's are not. Also, Malinvaud (1970) shows that ρ_p exhibit substantial small sample bias. This test yields, therefore, useful information only if the sample is sufficiently large.

A Lagrange multiplier (LM) test can also be used. If the null hypothesis is $\rho_p = 0$, and the alternative hypothesis is the autoregressive $AR(p)$ process $\epsilon_t = \rho_p \epsilon_{t-p} + \eta_t$, with $\rho_p \neq 0$, then the LM test statistic $T\rho_p^2$ has a $\chi^2_{(1)}$ distribution asymptotically. For a more general alternative, such as $\epsilon_t = \sum_{j=1}^{j=p} \rho_j \epsilon_{t-j} + \eta_t$ in the $AR(p)$ case and the null hypothesis $\rho_1 = \rho_2 = \cdots = \rho_p = 0$, the LM test statistic $T\sum_{j=1}^{j=p} \rho_j^2$ has a $\chi^2_{(p)}$ distribution asympotically. This suggests, 1) estimating the autocorrelation of the residuals using (5) by letting p vary between, say, 1 and 12,

[3] It is assumed here that the disturbances follow an autoregressive process. More complicated processes, such as moving average (MA) and autoregressive and moving average (ARMA) processes are not examined.

and 2) computing the LM statistic for the orders of the autocorrelation function found significant.

The tests for autocorreled disturbances are applied to each of the 20 size–sorted portfolios. Potential relationships between firm size and the deviations from uncorrelated disturbances can, therefore, be tested. Also, these statistics are computed on a sample that successively includes and excludes the January observations. This enables us to control for possible effects of the January seasonal on the violation of the assumption of uncorrelated disturbances.

4.2 The empirical evidence

Table 1 displays the parameter estimates of the market model, as specified in regression equation (4), obtained in the four subperiods examined by Brown, Kleidon and Marsh (1983) and in the total period. The abnormal excess returns estimates, *i.e.*, the $\hat{\alpha}_i$'s, turn out to be insignifiant in two of the subperiods and in the total period using either the EW or the VW index. As in Brown, Kleidon, and Marsh (1983), small firms earn negative, though insignificant, abnormal excess returns in the second period. The small firm effect tends to be statistically more significant with the VW index than with the EW index. Also, with the VW index, large firms appear to earn significant negative abnormal excess returns in two of the four subperiods and also in the total period. The empirical evidence supports the existence of a "size effect". However, the abnormal excess returns are non–stationary, *i.e.*, vary over time, and are not always significant in the subperiods.

The autocorrelation function of the market model residuals obtained with both market indices is computed using equation (5). Though not reported here, the

autocorrelation functions indicate that $\hat{\rho}_{12}$ is the only serial correlation parameter estimate which displays a consistent pattern across portfolios and in the different time periods, being always positive and generally significant. The results obtained with the LM statistic, distributed as a $\chi^2_{(1)}$, testing the null hypothesis that $\hat{\rho}_{12} = 0$ against the hypothesis that $\epsilon_t = \rho_{12}\epsilon_{t-12} + \eta_t$, with $\hat{\rho}_{12} \neq 0$, are reported in table 2. The null hypothesis is rejected when the LM statistic is large. The probability values appear in parentheses. Three important results concerning the autocorrelation structure of the residuals of the market model appear in table 2. First, the LM test is sensitive to sample size. The hypothesis of uncorrelated disturbances is rejected for most portfolios in the total period but only for the smallest ones in the subperiods. Second, the rejection is sharper for portfolios of small firms than for portfolios of large firms in the total period. The p–values increase as the market value of the portfolios increases. Third, the January seasonality totally accounts for the autocorrelation observed at lag 12. Without the January observations, the null hypothesis is neither rejected in the subperiods, nor in the total period, with both indices.

It is surprising to find that the market model residuals mostly exhibit twelve order autocorrelation. One would expect the market model residuals to display first order autocorrelation as well. This would be consistent with Lo and MacKinlay's (1987) rejection of the random walk hypothesis for size–sorted portfolios. To better understand the lack of first order autocorrelation in the market model residuals, the first order ($\hat{\rho}_1$) and twelve order ($\hat{\rho}_{12}$) sample autocorrelation estimates of 1) the returns and 2) the market model residuals of the 20 size–sorted portfolios are reported in table 3. The January observations are included in panel A and excluded

in panel B. The standard deviation of the first and twelve order autocorrelation estimates, denoted $\hat{\sigma}_{\rho_1}$ and $\hat{\sigma}_{\rho_{12}}$, respectively, appear in the third column. Table 4 presents the sample autocorrelation estimates of the returns on the EW and VW market indices.

Panels A.1 and B.1 of table 3 indicate that the returns on the size-sorted portfolios do exhibit first order serial correlation. The estimate of first order autocorrelation, $\hat{\rho}_1$, is significant for P_1 and is decaying as the average market value of the portfolio increases. This result holds in the total period as well as in the subperiods. This is consistent with Lo and MacKinlay's (1987) finding. Though a sharp decrease in the estimate of first order autocorrelation is observed when a four week base interval replaces a one week interval, Lo and MacKinlay (1987) report a serial correlation of 23% for the smallest quintile portfolio. In panel A.1 of table 3, the estimate of first order autocorrelation is equal to .20 in the entire sample period. Like $\hat{\rho}_1$, $\hat{\rho}_{12}$ is positive. However, the twelve order autocorrelation estimates are comparatively larger and more significant than the first order autocorrelation estimates at least in the total period. For example, the estimate obtained for P_1, $\hat{\rho}_{12}$ is equal to .34 versus .20 for $\hat{\rho}_1$ in the total period. The twelve order autocorrelation estimates are also decaying as the average market value of the portfolios increases. When the January observations are removed from the sample, the first order autocorrelation estimates remain identical. The point estimates are even slightly larger as panel B.1 of table 3 indicates. The twelve order autocorrelation totally vanishes, however.

It is necessary to examine the sample first order and twelve order autocorrelation estimates obtained for the market indices to understand why, unlike the twelve

order autocorrelation, there are no traces of first order autocorrelation left in the market model residuals. The first order autocorrelation estimates displayed by the returns on the EW and VW market indices are similar to those obtained by Lo and MacKinlay (1987). They report an estimate of 15% for the returns on the EW index (with a base interval of four weeks), which is significant at the 5% level. In the total sample period, table 4 indicates that $\hat{\rho}_1$ is equal to .16. Also, Lo and MacKinlay (1987) find that the rejection of the random walk hypothesis is much weaker for the VW index. Consistent with their result, the first order autocorrelation estimate obtained for the VW index turns out to be equal to .05 in the total sample period. Also, table 4 shows that the twelve order autocorrelation estimate is marginally significant for the returns on the EW index but insignificant for the returns on the VW index.

It is important to notice, however, that the difference in the first order autocorrelation estimates obtained for the returns on P_1 and on the EW index is smaller than the difference in the twelve order autocorrelation estimates obtained for the returns on P_1 and on the EW index, namely $(\hat{\rho}_1(P_1) - \hat{\rho}_1(EW)) < (\hat{\rho}_{12}(P_1) - \hat{\rho}_{12}(EW))$. In the total period, the two diferences are (.201-.046)=.155 and (.343-.049)=.294, respectively. This is also true with the VW index. This difference is critical to understand why, unlike twelve order autocorrelation, the residuals obtained with the EW or VW index do not exhibit first order autocorrelation. This result holds in the total sample period but is less apparent in the subperiods.

5 Tests of MV efficiency in the presence of serially correlated disturbances

5.1 The econometric implications of serially correlated disturbances

The econometric implications of serially correlated disturbances are well-known. Autocorrelation implies that the OLS estimator is unbiased but inefficient. Also, the OLS variance estimator is biased and consequently the usual OLS test statistics are not valid. The direction of the bias is of great interest, especially for the problem at hand. Nicholls and Pagan (1977) find that, when the errors follow an $AR(1)$ process, that understatement of the variance is the more likely situation provided the autocorrelation is positive.[4]

An important issue is to determine if this result extends to higher order processes and in particular to the $AR(12)$ process found in the market model residuals. The unbiased standard deviation estimates of the market model intercept and slope coefficients are computed with a ML procedure that jointly estimates the $AR(12)$ autocorrelation coefficient and the regression parameters. For each portfolio, the market model is rewritten as,

$$\begin{cases} r_t &= r_{mt}\beta + \epsilon_t \\ \epsilon_t &= \rho\epsilon_{t-12} + \eta_t, \qquad t = 1, \ldots, T-13, \end{cases} \tag{6}$$

with $E(\eta_t) = 0$, $E(\eta_t^2) = \sigma_\eta^2$, $E(\eta_t\eta_s') = 0$ for $t \neq s$ and $|\rho| < 1$. Here, r_t is a $(T \times 1)$ vector containing the observations on the excess returns obtained for each portfolio i, r_{mt} is a $(T \times 2)$ matrix made of a vector of ones in the first column and a vector containing the observations on the excess market returns in the second

[4] This result has been established under very specific conditions. Nicholls and Pagan (1977) assume that there is only one explanatory variable.

column, and β is a (2×1) row vector with the intercept and slope coefficients of the market model in the first and second columns, respectively. The covariance matrix $E(\epsilon\epsilon') = \Phi$ is equal to $\sigma_\eta^2\Psi = \sigma_\eta^2(\Psi_o \otimes I)$ where Ψ_o is of dimension $(T/12 \times T/12)$ and has the same structure as the matrix obtained for an $AR(1)$ process, I is the identity matrix, and $\otimes$ designates the Kronecker product.[5] The structure of Ψ is such that $E(\epsilon_t\epsilon_{t-s}') = 0$ unless s is a multiple of 12 in which case it is equal to $\sigma_\epsilon^2\rho^{\frac{s}{12}}$ with $\sigma_\epsilon^2 = \frac{\sigma_\eta^2}{(1-\rho^2)}$.

The maximum likelihood estimates can be computed as follows. If y has a multivariate normal distribution, its density is,

$$f(y) = (2\pi)^{\frac{-T}{2}}(\sigma_\eta^2)^{-\frac{T}{2}}|\Psi|^{-\frac{1}{2}}\exp\left\{-\frac{(r - r_m\beta)'\Psi^{-1}(r - r_m\beta)}{2\sigma_\eta^2}\right\}, \tag{7}$$

and so for ML estimation, the log–likelihood is, apart from a constant,

$$L(\beta, \Psi) = -\frac{T}{2}\log\sigma_\eta^2 - \frac{1}{2}\log|\Psi| - \frac{(r - r_m\beta)'\Psi^{-1}(r - r_m\beta)}{2\sigma_\eta^2}. \tag{8}$$

Conditional on β and Ψ, the ML estimator for σ_η^2 is,

$$\sigma_\eta^2 = \frac{1}{T}(r - r_m\beta)'\Psi^{-1}(r - r_m\beta), \tag{9}$$

and substituting this into L and ignoring the constants gives the log–likelihood function,

$$L(\beta, \Psi) = -\frac{T}{2}\log\left[(r - r_m\beta)'\Psi^{-1}(r - r_m\beta)\right] - \frac{1}{2}\log|\Psi|, \tag{10}$$

[5] That is

$$\Psi_o = \frac{1}{1-\rho^2}\begin{bmatrix} 1 & \rho & \rho^2 & \cdots & \rho^{T-1} \\ \rho & 1 & \rho & \cdots & \rho^{T-2} \\ \rho^2 & \rho & 1 & \cdots & \rho^{T-3} \\ \vdots & \vdots & \vdots & \ddots & \vdots \\ \rho^{T-1} & \rho^{T-2} & \rho^{T-3} & \cdots & 1 \end{bmatrix}$$

The maximum likelihood estimates of β and the unknown elements in Ψ are those values that maximize (10). After some rearrangings, these values are shown to be equal to the values that maximize,

$$S_L(\beta, \Psi) = |\Psi|^{\frac{1}{T}}(r - r_m\beta)'\Psi^{-1}(r - r_m\beta). \tag{11}$$

Equation (11) can be rewritten as,

$$S_L(\beta, \Psi) = |\Psi|^{\frac{1}{T}}(r^* - r_m^*\beta)'(r^* - r_m^*\beta) \tag{12}$$

with $r^* = Pr$ and $r_m^* = Pr_m$ and $\Psi^{-1} = P'P$. The ML estimates are found by minimizing equation (12). The Generalized Least squares estimator is obtained by minimizing,

$$S(\beta, \Psi) = (r^* - r_m^*\beta)'(r^* - r_m^*\beta). \tag{13}$$

Therefore, as shown in Judge and *al.* (1985), the difference between the objective function for the ML estimates and that for the GLS estimator is that the former contains the tth root of the determinant of Ψ.[6]

Also, though the market model residuals do not exhibit any systematic autocorrelation at lag 1, the estimates obtained with the $AR(12)$ process can be compared to estimates obtained with an $AR(1,12)$ process defined as,

$$\begin{cases} r_t &= r_{mt}\beta + \epsilon_t \\ \epsilon_t &= \rho_1\epsilon_{t-1} + \rho_{12}\epsilon_{t-12} + \eta_t, \quad t = 1, \ldots, T-13. \end{cases} \tag{14}$$

The ML estimates employ a Gauss–Marquardt algorithm to minimize the sum of squares and maximize the log–likelihood function. Yule–Walker estimates are used as starting values. The relevant optimization is performed simultaneously for both the regression and $AR(12)$ or $AR(1,12)$ parameters.

[6] Hence, equation (12) can be rewritten as $S_L(\beta, \Psi) = |\Psi|^{\frac{1}{T}}S(\beta, \Psi)$.

5.2 The empirical evidence

Table 5 reports the results of the ML estimates of the market model computed under the assumption that the residuals follow an $AR(12)$ process. The regressions and the autocorrelation parameters are estimated jointly. The January observations are included in the estimation process in panel A and excluded in panel B.

The empirical evidence concerning the autocorrelation parameter at lag 12, ρ_{12}, is first examined. As displayed in panel A of table 5, $\hat{\rho}_{12}$ is systematically positive and is decreasing with firm size. In the total period, the point estimate obtained for P_1 is equal to .40 (.45) and to .25 (.15) for P_{20} with the EW (VW) index. For porfolios of small firms, the point estimates tend to be larger with the VW than with the EW index but the reverse is true for portfolios of large firms. Most of these results hold in the total period and also in the subperiods. The serial correlation parameter estimate, $\hat{\rho}_{12}$, is significant for portfolios of small firms, generally insignificant for portfolios of medium size firms, and marginally significant or insignificant depending on the market index for portfolios of large firms. Also, $\hat{\rho}_{12}$ is more significant in the total period than in the subperiods, and t–statistics greater than 6 are not uncommon. Panel B of table 5 supports the empirical finding of section 4 that the January seasonal is generating the twelve order serial correlation. Once the January observations are removed, a sharp drop in the point estimate of $\hat{\rho}_{12}$ is observed. For P_1, $\hat{\rho}_{12}$ decreases from .40 (.45) with the January returns to .02 (.03) with the EW (VW) index in the total period. The autocorrelation coefficient is also never significant in the subperiods after eliminating the January observations.

The autocorrelation coefficient, $\hat{\rho}_{12}$, is positive and significant. It is, therefore, important to verify if the existence of a significant twelve order serial correlation

parameter estimate implies, like an $AR(1)$ process, a downward bias in the OLS standard deviation estimates. First, the comparison of the OLS and ML point estimates of the market model parameters reveals a small difference between the two sets of estimates. This is especially true in the subperiods which suggests that sample size might partly be responsible for the discrepancy. However, even in the subperiods, no systematic pattern is observed. The ML estimates of the abnormal excess returns are slightly higher or lower than their OLS counterparts.

The most important result concerns the ML standard deviation estimates of the abnormal excess returns. They turn out to be systematically larger than their OLS counterparts, and are larger the higher the autocorrelation coefficient. In the total period, the t–statistic obtained for P_1 decreases from 3.44 (3.56) with OLS to 2.30 (2.07) with the ML estimator for the EW (VW) index. The probability levels decrease from 1% with OLS to 5% with the ML estimator. In the third subperiod, the ML estimates of the abnormal excess returns become insignificant. The only subperiod that still displays positive abnormal excess returns estimates is the first one. This, however, is not surprising since $\hat{\rho}_{12}$ is only marginally significant in that subperiod.[7] The evidence reported in table 5 is consistent with the result

[7] The market model regression parameters obtained with the EW and VW indices, respectively, are also estimated under the assumption that the errors follow an autoregressive process $AR(1,12)$. Though not reported, the results can be summarized as follows. The estimates of the parameter $\hat{\rho}_1$ are generally neither statistically significant in the subperiods nor in the total period. Portfolios of small firms do not exhibit autocorrelated returns at lag 1. However, unlike small firms, portfolios of large firms seem to display negative autocorrelated returns at lag 1. The coefficients are statistically significant in the third and in the total period with the EW index but are never significant with the VW index. The estimates of the parameter $\hat{\rho}_{12}$ are similar to those reported in table 5. Portfolios of small firms exhibit a large positive autocorrelation parameter estimate at lag 12. This parameter is decaying with firm size. It is finally interesting to note that the standard deviation estimates of the abnormal excess returns are generally larger with the $AR(1,12)$ than with the $AR(12)$ process. This may reflect an overfitting of the model. The lack of significance of the autoregressive parameter at lag 1 suggests that the simple $AR(12)$ is adequate to account for the autocorrelation in the market model residuals displayed by size–sorted portfolios.

reported in the econometric literature that the OLS standard deviation estimates are downward biased in presence of positive serial correlation.

The empirical findings of the two previous sections can be summarized as follows. The risk–adjusted returns on size–sorted portfolios display a systematic positive and significant autocorrelation estimate at lag 12. After controlling for the effects of serial correlation, the OLS standard deviation estimates turn out to be systematically underestimated. The underestimation is serious enough to reverse inferences about the small firm effect in most subperiods. As found in section 4, the main source of autocorrelation for the 20 size–sorted portfolios is the January seasonal. Before concluding that the presence of serial correlation in the risk adjusted returns is inconsistent with the market efficiency hypothesis, it is necessary to test whether the presence of twelve order serial correlation does not arise from a misspecified return generating process. This issue is addressed in the following section.

6 Origins of the serially correlated disturbances: The January seasonal and the misspecification of the return generating process

6.1 The January seasonal and the non–stationarity of the return generating process

The empirical evidence in the previous sections indicates that the residuals of the market model are serially correlated, and more so for portfolios of small firms than for portfolios of large firms. This econometric problem may arise from a misspecification of the market model and in particular from non–stationarities in the

return generating process. One obvious potential source of non-stationarity is the January seasonal. The January returns may be generated by factors not accounted for by the market model. As the empirical evidence of Keim (1983) and Roll (1983) indicates, the January seasonal is more pronounced for small than for large firms. This might explain why the market model disturbances obtained for small firm portfolios exhibit more serial correlation than those obtained for large firm portfolios. The goal of this section is to investigate whether the econometric problems of serially correlated disturbances vanish after controlling for non-stationarity.

Several recent papers have empirically examined the question of seasonality of market risk. Using twenty years of daily data, Tinic and Rogalski (1986) investigate the betas of NYSE and AMEX stocks. They find that the mean returns, betas, and residual variances of the size portfolios are not equal across months. Morgan and Morgan (1987) challenge the view that market risk is not constant throughout the year. By accounting for seasonal heteroskedasticity with the Autoregressive Conditional Heteroskedasticity (ARCH) model of Engle (1982), they conclude that 1) market risk does not rise in January and 2) all the apparent non-stationarity in the estimates of systematic risk is due to the heteroskedasticity of the market model residuals. This conflict is partly resolved in Hillion and Sirri (1987). They find that except for the smallest firms, it is heteroskedasticity and not non-stationarity that leads to an apparent seasonality in beta.

The dummy variable approach is a simple way to test whether the return generating process has time-varying parameters. In the subsequent analysis, the market model parameters are hypothesized to shift only in January, and remain constant

throughout the other 11 months.[8] In this case, the model tested for each portfolio i is,

$$r_{it} = \iota_T \alpha_i + D_{Jt}\alpha_{iJ} + r_{mt}\beta_i + r_{mt}D_{Jt}\beta_{iJ} + \epsilon_{it}, \qquad t = 1, \ldots, T, \tag{15}$$

where D_{Jt} takes on the value 1 in January and 0 otherwise. This regression can be also run without the shift in the January intercept. In that case, the model tested for each portfolio becomes,

$$r_{it} = \iota_T \alpha_i + r_{mt}\beta_i + r_{mt}D_{Jt}\beta_{iJ} + \epsilon_{it}, \qquad t = 1, \ldots, T. \tag{16}$$

The issue is to test if the serial correlation in the market model residuals disappears after controlling for the non–stationary in the risk estimates, and more generally in the return generating process, *i.e.*, after specifying a market model which let the intercept and slope coefficients vary in January. The LM test suggested in section 4 to detect serial correlation is applied to the residuals of the augmented market model, *i.e.*, the disturbance terms of equations (15) and (16).

Also, to investigate the potential relation between non–stationarity and the autocorrelation of the residuals, the parameters of the augmented market model are estimated under the additional assumption that the ϵ_{it}, with $i = 1, \ldots, K$, follow an autoregressive process of order p,

$$\epsilon_{it} = \rho_p \epsilon_{it-p} + \eta_{it}, \quad t = 1, \ldots, T - p - 1,$$

where η_{it}, with $i = 1, \ldots, K$, is asssumed to be normally and independently distributed with a mean of 0 and a variance σ_η^2. From the empirical evidence in section 4, the order of the autoregressive process, *i.e.*, p, is set equal to 12. The ML

[8] See Hillion and Sirri (1987) for alternative models.

procedure described in the previous section is used to jointly estimate the regression parameters and the $AR(12)$ serial correlation coefficient, *i.e.*, ρ_{12}.

6.2 The empirical evidence

6.2.1 Estimates of the augmented market model under the assumption of no serial correlation

Estimation of the parameters of the augmented market model is given in table 6. The table results from running OLS regressions like (15) and (16), without controlling for serially correlated disturbances. In panel A, both the intercept and the slope coefficients are allowed to shift in January, as in equation (15), while in panel B, only the slope coefficient is allowed to vary, as in equation (16). Panel A shows that the beta of small firm rises significantly in January with the EW index while the January beta of all firms rise with the VW index. The beta for P_1 rises from 1.21 to 1.61 in January with the EW index. It is particularly striking that the VW P_1 beta almost doubles, shifting from 1.13 to 2.13 in January. Since monthly data are used, this effect is unlikely to be due to thin trading. Panel B shows that the shift in beta is larger than in panel A when the intercept is not allowed to vary. A possible reason for this is that the mean returns in January are higher than the mean returns in other months. With the intercept constrained to be equal, the beta may pick up the shift in the mean, as well as the covariance effects.

Given the conflicting empirical evidence of Tinic and Rogalski (1986) and Morgan and Morgan (1987), table 6 also presents three different heteroskedastic– consistent standard deviation estimates of the regression parameters. Corrected standard errors are presented for the White (1980) covariance matrix (HC1), the MacKinnon

and White (1985) jacknife estimator (HC2), and the weighted jacknife of Hinkley (1977) (HC1).[9] No difference for the beta coefficient, *i.e.*, β_i, is apparent between the three corrected standard error estimates, although for most portfolios, the OLS covariance matrix underestimates the standard errors relative to the other three.

The results obtained for the standard error estimates of the January beta dummy are more interesting. For the EW index, the standard errors of P_1 increases by 40%, and therefore the t–statistic drops from 3.63 to 2.75. The results for other portfolios are similar, though less dramatic. The results obtained with the VW index are even more pronounced. The standard errors for the MacKinnon and White (1985) estimate are 2 times higher than the OLS estimate. The shift in the P_1 January beta from 1.13 to 2.13 becomes only marginally significant.[10] The above results lend some support to the contention of Morgan and Morgan (1987) that except for the smallest firms, it is heteroskedasticity and not non–stationarity that leads to an apparent seasonality in beta.

From table 6, it also appears that the estimates of the non–January abnormal excess returns are never significant for any of the 20 portfolios, including P_1, and for the two market indices used. The estimates of the January abnormal excess returns

[9]The three heteroskedastic–consistent covariance matrix estimators are asymptotically equivalent. Their small sample properties differ, however. MacKinnon and White (1985) show that White (1980) estimator only gives correct results asymptotically, and is biased in finite samples. MacKinnon and White (1985) show that Hinkley's (1977) weighted jacknife estimator differs from White's (1980) covariance matrix by a degree of freedom correction similar to the one conventionally used to obtain unbiased estimates of δ^2, the residual variance. Unfortunately, Hinkley's (1977) covariance matrix estimator is also biased in the case of unbalanced data, though the degree of the bias is less than previous estimators. MacKinnon and White (1985) develop an alternative jacknife covariance matrix estimator which according to a Monte Carlo study appears to the least biased in small samples.

[10]This sharp increase in the standard error can be explained as follows. Since the variance of the residuals are higher in January, a January dummy will be positively correlated with the residuals squared. Thus, OLS standard errors are biased down as the bias depends on the correlation between the variance of the individual residuals and the square of any column of the matrix of explanatory variables. See Hillion (1988) chapter 3 and Hillion and Sirri (1987).

are, however, high and extremely significant, positive for small firms but negative for large firms. As displayed in tables 5.1 and 5.2, the January abnormal excess returns estimates remain significant with the three heteroskedastic–consistent covariance matrix estimators, though a sharp drop in the significance of α_{iJ} is observed for certain portfolios with the VW index.

6.2.2 Estimates of the augmented market model under the assumption of twelve order serial correlation

Table 7 displays the results of the LM statistic testing for the presence of serially correlated disturbances. Also, table 7 presents the parameters of the augmented market model, as specified in equations (15) and (16), estimated under the assumption that the residuals follow an autoregressive process of order 12, *i.e.*, an $AR(12)$ process. The estimate of the autoregressive process, $\hat{\rho}_{12}$ and the LM statistic distributed as a χ^2 with 1 degree of freedom, yield identical conclusions about the impact of non–stationarity on the autocorrelation of the residuals. The relation between non–stationarity and the autocorrelation of the residuals depends, on the version of the augmented market model that is being tested. When only the intercept is assumed to be stationary, as specified in regression equation (16), panel B of table 7 displays a relatively modest decrease in the estimate of the autocorrelation parameter $\hat{\rho}_{12}$, relative to its market model counterpart. The LM statistic confirms this finding. The residuals remain serially correlated at lag 12 after controlling for non–stationary risk estimates.

However, when both the intercept and the slope coefficients are assumed to be non–stationary, as specified in regression equation (15), panel A of table 7 displays

a sharp decrease in the point estimate of $\hat{\rho}_{12}$, relative to its market model counterpart. For example, with the EW index, the estimate of $\hat{\rho}_{12}$ obtained for P_1 decreases from .36 in the simple version of the market model to .14 in the dummy variable augmented version of the market model as specified in (15). The estimates are .44 and .16 with the VW index, respectively. The autoregressive process becomes only marginally significant. This result is also confirmed by the LM statistic which does not reject the hypothesis of uncorrelated residuals at lag 12 at the 1% level for P_1. However, the hypothesis is still rejected at the 5% level for P_1, P_2 and P_{20} with the EW index and for P_2 through P_{18} with the VW index. These results indicate that the serially correlated disturbances originate mostly from a misspecificied return generating process. The serially correlated disturbances almost vanish after controlling for the non–stationarities induced by the January seasonal.

The comparison of the parameters of the augmented market model, as specified in (15) and in (16), estimated with and without the assumption that the residuals follow an autoregressive process of order 12 is also very instructive. Tables 6 and 7 reveal that the point estimates and the significance of the slope and the slope dummy are almost identical. It is particularly interesting to notice that in panels A and B of table 7 that the estimate of the slope dummy obtained for P_1 is not driven to insignificance when the autoregressive parameter $\hat{\rho}_{12}$ is jointly estimated with the parameters of the augmented market model. This result holds regardless 1) of the market index, and 2) of the specification of the market model, namely with a stationary or non–stationary intercept. However, the twelve order serial correlation parameter estimate, *i.e.*, $\hat{\rho}_{12}$, is driven to insignificance after controlling

for risk and intercept non–stationarity.[11] These results hold only for P_1. For the other portfolios, the beta dummy is insignificant but the autocorrelation parameter estimate remains marginally significant. This is more consistent with Morgan and Morgan's (1987) finding.

7 Conclusion

Monthly stock returns or risk–adjusted returns are believed to be serially uncorrelated. This paper shows that the risk–adjusted monthly returns on size–sorted portfolios exhibit no or little first order serial correlation but displays positive and significant twelve order serial correlation. Except for the paper of Jegadeesh (1987), little research has been devoted to test the presence of high order serial correlation in risk–adjusted returns and to assess its impact on the empirical tests of MV efficiency.

This paper finds that twelve order serial correlation mostly affects the risk–adjusted returns on small firms' securities. Depending on the market index, the point estimates of the twelve order serial correlation parameter obtained for the portolio of smallest firms are in the .30 to .40 bracket. Most, if not all, of the twelve order serial correlation originates from the January seasonal. Is the presence of serial correlation in the risk adjusted returns on small firm portfolios inconsistent with the market efficiency hypothesis? The answer is negative. The twelve order serial correlation is shown to arise from a misspecification of the return generating process. When the return generating process is let to have time varying parameters,

[11] Therefore, unlike what happens during the estimation of the ARCH process of Morgan and Morgan (1987), the slope dummy variable taking on the value of 1 in January is not driven to insignificance when the autocorrelation of order 12 is introduced.

i.e., when the intercept and slope coefficients are let to vary in January, most of the twelve order serial correlation disappears. This result is obtained with a ML procedure which jointly estimates the parameters of the return generating process and the twelve order serial correlation parameter.

The presence of twelve order serial correlation in risk–adjusted returns has important implications for tests of MV efficiency. Autocorrelation implies that the OLS variance estimator is biased and consequently the usual OLS test statistics are not valid. This result is particularly important to assess the true statistical significance of the intercept of the market model, *i.e.*, to test the statistical significance of the abnormal excess returns estimates. After controlling for the effects of serial correlation, the OLS standard deviation of the abnormal excess returns estimates are found to be systematically underestimated. The underestimation is serious enough to reverse inferences about the small firm effect in most subperiods. Though most of the tests performed in this paper are based on univariate statistics, serial correlation is also likely to affect multivariate tests of MV efficiency since they are also based on the assumption that the market model disturbances are serially uncorrelated. This point needs, however, to be empirically tested.

The presence of twelve order serial correlation in the risk adjusted returns on small firm portfolios prevents from drawing valid inferences from univariate or multivariate tests of MV efficiency based on size–sorted portfolios. Fortunately, the twelve order serial correlation disappears when the return generating process is correctly specified. Therefore, valid inferences about the MV efficiency of a market index can be drawn when an intercept dummy and a slope dummy variable for the January observations are added to the market model. This is one alter-

native. A second alternative is to split the data into January and non–January observations and to perform tests of conditional MV efficiency as opposed to tests of unconditional MV efficiency. Which approach is preferable, *i.e.*, should the return generating process be modified to have time–varying parameters or should the observations be split according to a calendar criterion is an unanswered question which deserves to be carefully investigated.

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TABLE 1

The parameters of the market model,

$$r_{it} = \iota_T \alpha_i + r_{mt}\beta_i + \epsilon_{it}, \quad t = 1, \ldots T.$$

The OLS solution.

INDEX		EW INDEX				VW INDEX			
		α ($\times 10^3$)	$t(\alpha)$	β	$t(\beta)$	α ($\times 10^3$)	$t(\alpha)$	β	$t(\beta)$
Sub 1	P_1	13.31	2.52	1.42	12.98	29.45	3.66	1.26	4.88
	P_2	4.95	1.43	1.32	18.53	18.92	2.98	1.31	6.47
	P_{10}	-.12	-.11	1.06	46.86	8.76	2.74	1.34	13.12
	P_{19}	-2.30	-1.05	.58	12.71	.51	.52	.99	34.50
	P_{20}	-3.47	-1.44	.51	10.24	-1.66	-2.65	.96	47.75
Sub 2	P_1	2.98	.49	1.27	14.46	-7.31	-.77	1.47	6.91
	P_2	-.99	-.21	1.23	18.31	-10.83	-1.31	1.47	7.96
	P_{10}	-.43	-3.96	1.07	67.71	-11.98	-3.16	1.50	17.66
	P_{19}	3.05	1.12	.66	16.63	-1.18	-.80	1.08	32.57
	P_{20}	5.30	1.85	.53	12.73	2.11	2.22	.93	43.51
Sub 3	P_1	16.26	2.21	1.62	16.60	37.76	2.75	1.49	5.49
	P_2	5.44	1.06	1.42	20.95	24.28	2.17	1.36	6.14
	P_{10}	.53	.39	1.00	56.18	13.32	2.50	1.22	11.55
	P_{19}	-6.44	-1.91	.62	13.79	.96	.76	1.03	40.76
	P_{20}	-8.56	-2.00	.50	8.89	-2.75	-2.59	.98	46.41
Sub 4	P_1	6.24	1.09	1.14	10.97	11.07	1.44	1.00	5.97
	P_2	4.06	.12	1.12	18.48	4.86	.85	1.05	8.50
	P_{10}	-2.35	-1.74	1.05	42.86	1.13	.40	1.17	19.03
	P_{19}	-1.90	-.82	.79	18.85	2.53	.20	1.00	37.38
	P_{20}	-2.89	-1.08	.74	14.99	-1.12	-1.44	.98	57.76
Sub 5	P_1	10.59	3.44	1.41	28.58	18.02	3.56	1.34	11.51
	P_2	2.96	1.43	1.30	39.15	9.62	2.35	1.31	13.96
	P_{10}	-1.59	-2.59	1.04	105.79	3.03	1.49	1.31	27.91
	P_{19}	-2.25	-1.67	.65	30.12	.02	.04	1.03	72.76
	P_{20}	-2.52	-1.59	.54	21.38	-.89	-2.02	.96	94.61

Remarks:
Subperiod 1: January 1963–December 1968 (T=72).
Subperiod 2: January 1969–December 1973 (T=60).
Subperiod 3: January 1974–June 1979 (T=66).
Subperiod 4: July 1979–December 1984 (T=66).
Total period: January 1963–December 1984 (T=264).

TABLE 2

LM statistic testing for autocorrelation at lag 12.
p–values appear in parentheses.

Panel A January observations included.

		P_1	P_2	P_3	P_5	P_{10}	P_{15}	P_{18}	P_{19}	P_{20}
EW Index	Sub 1	1.92	.00	.29	.03	1.17	.55	.05	.64	.57
		(>.15)	(>.15)	(>.15)	(>.15)	(>.15)	(>.15)	(>.15)	(>.15)	(>.15)
	Sub 2	12.98	17.32	6.51	2.53	3.92	6.53	2.48	4.17	2.37
		(<.01)	(<.01)	(<.01)	(.111)	(.0475)	(.106)	(.115)	(.0410)	(.124)
	Sub 3	7.04	4.07	5.77	.03	.02	1.84	4.08	4.85	2.41
		(<.01)	(.043)	(.016)	(>.15)	(>.15)	(>.15)	(.04)	(.027)	(.106)
	Sub 4	3.72	11.73	2.84	.90	.02	2.28	4.39	5.00	5.60
		(.054)	(<.01)	(.092)	(.15)	(>.15)	(.13)	(.036)	(.025)	(.018)
	Sub 5	33.62	36.54	18.21	2.03	.00	10.77	13.27	14.23	15.96
		(<.01)	(<.01)	(<01)	(>.15)	(>.15)	(<.01)	(<.01)	(<.01)	(<.01)
VW Index	Sub 1	4.99	3.51	2.84	3.94	2.57	3.07	.00	.10	.34
		(.025)	(.11)	(.092)	(.047)	(.102)	(.079)	(>.15)	(>.15)	(>.15)
	Sub 2	15.73	15.75	9.36	7.83	3.39	.33	1.05	.49	2.36
		(<.01)	(<.01)	(<.01)	(<.01)	(.065)	(>.15)	(>.15)	(>.15)	(>.15)
	Sub 3	12.66	12.02	12.03	7.03	5.09	5.12	2.07	.04	1.06
		(<.01)	(<.01)	(<.01)	(<.01)	(<.01)	(.023)	(>.15)	(>.15)	(>.15)
	Sub 4	2.53	4.79	1.98	1.33	.83	.00	.01	.25	.18
		(.106)	(.028)	(>.15)	(>.15)	(>.15)	(>.15)	(>.15)	(>.15)	(>.15)
	Sub 5	51.41	52.83	44.01	30.46	28.75	19.29	3.95	.00	5.85
		(<.01)	(<.01)	(<.01)	(<.01)	(<.01)	(<.01)	(.047)	(>.15)	(>.15)

Panel B: January observations excluded.

		P_1	P_2	P_3	P_5	P_{10}	P_{15}	P_{18}	P_{19}	P_{20}
EW Index	Sub 1	.023	.53	.01	1.57	.09	.04	.58	.60	.52
		(>.15)	(>.15)	(>.15)	(>.15)	(>.15)	(>.15)	(>.15)	(>.15)	(>.15)
	Sub 2	1.11	2.00	.38	.11	.39	.22	.29	1.93	.02
		(>.15)	(>.15)	(>.15)	(>.15)	(>.15)	(>.15)	(>.15)	(>.15)	(>.15)
	Sub 3	2.00	3.71	2.59	1.96	9.86	.01	6.09	.77	.67
		(>.15)	(.053)	(.11)	(>.15)	(<.01)	(>.15)	(.013)	(>.15)	(>.15)
	Sub 4	.00	.62	.00	.00	.03	.99	.26	.82	9.56
		(>.15)	(>.15)	(>.15)	(>.15)	(>.15)	(>.15)	(>.15)	(>.15)	(<.01)
	Sub 5	.20	.53	.02	.00	.015	1.86	.53	.70	.46
		(>.15)	(>.15)	(>.15)	(>.15)	(>.15)	(>.15)	(>.15)	(>.15)	(>.15)
VW Index	Sub 1	.06	1.27	.65	1.91	.03	.49	.09	.85	.11
		(>.15)	(>.15)	(>.15)	(>.15)	(>.15)	(>.15)	(>.15)	(>.15)	(>.15)
	Sub 2	.73	3.54	.13	.21	.00	2.44	.02	.93	.12
		(>.15)	(.06)	(>.15)	(>.15)	(>.15)	(.12)	(>.15)	(>.15)	(>.15)
	Sub 3	.04	2.11	.19	.073	3.04	.01	.24	1.93	2.03
		(>.15)	(>.15)	(>.15)	(>.15)	(.08)	(>.15)	(>.15)	(>.15)	(>.15)
	Sub 4	.05	.94	.034	.082	1.86	.51	.17	.51	1.02
		(>.15)	(>.15)	(>.15)	(>.15)	(>.15)	(>.15)	(>.15)	(>.15)	(>.15)
	Sub 5	.75	.37	1.06	.07	.45	.19	.51	.09	.08
		(>.15)	(>.15)	(>.15)	(>.15)	(>.15)	(>.15)	(>.15)	(>.15)	(>.15)

Remarks:
Sub 1: January 1963–December 1968 (T=72).
Sub 2: January 1969–December 1973 (T=60).
Sub 3: January 1974–June 1979 (T=66).
Sub 4: July 1979–December 1984 (T=66).
Sub 5: Total period January 1963–December 1984 (T=264).

TABLE 3

First and twelve order sample autocorrelation estimates of the 20 size-sorted portfolio returns and market model residuals.

<u>Panel A</u>: January observations included.

<u>Panel A.1</u>: Portfolio returns.

Sub	$\hat{\rho}(\cdot)$	$\hat{\sigma}_{\hat{\rho}(\cdot)}$	P_1	P_2	P_3	P_5	P_{10}	P_{15}	P_{18}	P_{19}	P_{20}
1	$\hat{\rho}_1$	.117	.228	.201	.244	.188	.099	.064	.048	-.010	-.050
1	$\hat{\rho}_{12}$	.181	.248	.179	.156	.180	-.005	-.070	-.048	-.102	-.053
2	$\hat{\rho}_1$	.129	.260	.254	.255	.248	.227	.218	.085	.185	.007
2	$\hat{\rho}_{12}$	.153	.226	.180	.126	.059	-.048	-.051	-.162	-.156	-.163
3	$\hat{\rho}_1$	.123	.101	.050	.086	-.006	.032	.077	.018	.054	-.028
3	$\hat{\rho}_{12}$	.137	.445	.444	.401	.349	.315	.252	.230	.196	.147
4	$\hat{\rho}_1$	.123	.338	.186	.264	.208	.136	.105	.062	.046	-.015
4	$\hat{\rho}_{12}$	.149	-.119	-.150	-.162	-.162	-.211	-.157	-.129	-.066	.022
5	$\hat{\rho}_1$	.062	.201	.153	.203	.144	.151	.138	.065	.077	-.014
5	$\hat{\rho}_{12}$	.065	.343	.317	.270	.219	.124	.051	.040	.016	.042

<u>Panel A.2</u>: Market model residuals obtained with the EW index.

Sub	$\hat{\rho}(\cdot)$	$\hat{\sigma}_{\hat{\rho}(\cdot)}$	P_1	P_2	P_3	P_5	P_{10}	P_{15}	P_{18}	P_{19}	P_{20}
1	$\hat{\rho}_1$	.117	.073	.018	.046	.032	-.210	.097	.123	.109	.077
1	$\hat{\rho}_{12}$	.129	.163	-.010	-.063	.020	.127	.088	-.027	-.089	.095
2	$\hat{\rho}_1$	.129	-.027	-.100	-.088	.015	.071	-.057	-.094	-.137	-.185
2	$\hat{\rho}_{12}$	.157	.465	.536	.337	.206	-.256	.330	.203	.264	.199
3	$\hat{\rho}_1$	.123	.074	-.038	.084	-.050	-.196	.004	-.042	-.044	-.200
3	$\hat{\rho}_{12}$	.152	.327	.249	.296	.021	-.016	.143	.249	.271	.199
4	$\hat{\rho}_1$	.123	.208	.031	.050	-.015	-.319	-.037	.013	-.109	-.103
4	$\hat{\rho}_{12}$	.140	.237	.422	.207	.117	-.019	.186	.258	.275	.291
5	$\hat{\rho}_1$	.062	.060	-.057	.002	-.017	-.185	-.003	-.032	-.055	-.130
5	$\hat{\rho}_{12}$	.067	.357	.372	.262	.088	-.002	.202	.224	.232	.246

<u>Panel A.3</u>: Market model residuals obtained with the VW index.

Sub	$\hat{\rho}(\cdot)$	$\hat{\sigma}_{\hat{\rho}(\cdot)}$	P_1	P_2	P_3	P_5	P_{10}	P_{15}	P_{18}	P_{19}	P_{20}
1	$\hat{\rho}_1$	.117	.240	.248	.125	.193	.273	.122	.056	.239	-.097
1	$\hat{\rho}_{12}$	.184	.263	.187	.199	.234	.193	.207	.009	-.037	-.069
2	$\hat{\rho}_1$	.129	.023	-.024	-.013	-.001	-.035	.110	.238	.007	-.038
2	$\hat{\rho}_{12}$	.158	.512	.518	.395	.261	.238	.118	-.132	-.090	.199
3	$\hat{\rho}_1$	.123	-.014	-.092	-.054	-.196	-.224	-.256	-.343	-.107	-.043
3	$\hat{\rho}_{12}$	.153	.427	.427	.427	.326	.350	.278	.177	-.025	.127
4	$\hat{\rho}_1$	.123	.241	.069	.134	.069	-.037	.006	-.010	-.202	.022
4	$\hat{\rho}_{12}$	.142	.198	.269	.173	.142	.116	.007	-.012	.062	.053
5	$\hat{\rho}_1$	.062	.078	.002	.057	-.020	-.021	-.035	-.056	-.019	.019
5	$\hat{\rho}_{12}$	.067	.441	.447	.408	.339	.330	.270	.127	.002	.149

Sub 1: January 1963–December 1968 (T=72).
Sub 2: January 1969–December 1973 (T=60).
Sub 3: January 1974–June 1979 (T=66).
Sub 4: July 1979–December 1984 (T=66).
Sub 5: Total period January 1963–December 1984 (T=264).

TABLE 3 (Cont.)
First and twelve order sample autocorrelation estimates of the 20 size-sorted portfolio returns and market model residuals.

Panel B: January observations excluded.

Panel B.1: Portfolio returns.

Sub	$\hat{\rho}_{(\cdot)}$	$\hat{\sigma}_{\rho(\cdot)}$	P_1	P_2	P_3	P_5	P_{10}	P_{15}	P_{18}	P_{19}	P_{20}
1	$\hat{\rho}_1$	.123	.333	.234	.212	.180	.074	.028	.024	-.030	-.039
1	$\hat{\rho}_{12}$	.146	-.018	-.103	-.083	.007	-.052	-.088	-.057	-.033	-.039
2	$\hat{\rho}_1$	.137	.135	.140	.118	.121	.125	.144	.044	.102	.016
2	$\hat{\rho}_{12}$	.161	-.110	-.097	-.070	-.037	-.058	-.040	-.001	-.010	-.018
3	$\hat{\rho}_1$	.129	.113	.108	.163	.070	.099	.107	.035	.057	-.032
3	$\hat{\rho}_{12}$	.146	-.133	-.141	-.107	-.077	-.061	-.044	-.020	-.003	.018
4	$\hat{\rho}_1$	.147	.407	.236	.303	.215	.102	.080	.040	.024	-.060
4	$\hat{\rho}_{12}$	.162	-.094	-.069	-.179	-.172	-.170	-.205	-.190	-.205	-.225
5	$\hat{\rho}_1$	.064	.254	.209	.232	.161	.146	.123	.033	.066	-.019
5	$\hat{\rho}_{12}$	.072	-.070	-.066	-.046	-.044	-.039	-.035	-.006	-.023	-.029

Panel B.2: Market model residuals obtained with the EW index.

Sub	$\hat{\rho}_{(\cdot)}$	$\hat{\sigma}_{\rho(\cdot)}$	P_1	P_2	P_3	P_5	P_{10}	P_{15}	P_{18}	P_{19}	P_{20}
1	$\hat{\rho}_1$	.123	.188	.047	.058	-.006	-.025	.099	.061	.074	.048
1	$\hat{\rho}_{12}$	.136	-.024	-.102	.084	.096	.018	-.061	.080	-.018	.032
2	$\hat{\rho}_1$	.135	-.053	-.025	-.070	-.001	.039	-.029	-.039	-.040	-.051
2	$\hat{\rho}_{12}$	.179	-.096	-.189	-.036	.051	.007	-.155	-.015	.108	-.039
3	$\hat{\rho}_1$	.129	.069	-.010	.080	.027	-.124	-.048	-.074	.024	-.149
3	$\hat{\rho}_{12}$	.164	.018	-.146	-.045	.025	.195	.011	.209	-.113	-.111
4	$\hat{\rho}_1$	.128	.244	.048	.053	-.007	-.307	-.152	.089	-.133	-.152
4	$\hat{\rho}_{12}$	.142	-.030	.124	.024	.037	-.175	.091	-.053	-.091	-.129
5	$\hat{\rho}_1$	.064	.126	.050	.054	.021	-.077	.009	.019	.039	-.048
5	$\hat{\rho}_{12}$	.069	-.018	-.039	.066	.064	.045	.023	.046	-.019	-.018

Panel B.3: Market model residuals obtained with the VW index.

Sub	$\hat{\rho}_{(\cdot)}$	$\hat{\sigma}_{\rho(\cdot)}$	P_1	P_2	P_3	P_5	P_{10}	P_{15}	P_{18}	P_{19}	P_{20}
1	$\hat{\rho}_1$	.123	.329	.245	.268	.151	.180	.029	-.042	.174	-.043
1	$\hat{\rho}_{12}$	.144	.015	-.065	.012	.106	.098	-.022	.084	.090	-.074
2	$\hat{\rho}_1$	.135	.001	.017	-.008	-.015	.020	.121	.163	-.113	-.057
2	$\hat{\rho}_{12}$	.188	-.118	-.143	-.065	-.038	-.068	.038	.036	.099	.015
3	$\hat{\rho}_1$	.129	.031	-.007	.044	-.071	-.118	-.119	-.193	-.028	.00
3	$\hat{\rho}_{12}$	.153	-.151	-.202	-.165	-.131	-.234	-.196	-.008	-.060	-.056
4	$\hat{\rho}_1$	.128	.286	.089	.144	.051	-.172	-.013	-.083	-.291	-.020
4	$\hat{\rho}_{12}$	.143	-.005	.101	.00	-.003	.021	-.128	.066	.116	-.396
5	$\hat{\rho}_1$	.064	.172	.114	.144	.052	.053	.008	-.068	-.079	.077
5	$\hat{\rho}_{12}$	.071	-.028	-.047	.010	.00	-.008	-.088	.047	.054	-.041

Sub 1: January 1963–December 1968 (T=72).
Sub 2: January 1969–December 1973 (T=60).
Sub 3: January 1974–June 1979 (T=66).
Sub 4: July 1979–December 1984 (T=66).
Sub 5: Total period January 1963–December 1984 (T=264).

TABLE 4

First and twelve order sample autocorrelations of the equally–weighted and value–weighted market market indices.

Panel A: January observations included.

	EW INDEX				VW INDEX			
Subperiod	$\hat{\rho}_1$	$\hat{\sigma}_{\rho_1}$	$\hat{\rho}_{12}$	$\hat{\sigma}_{\rho_{12}}$	$\hat{\rho}_1$	$\hat{\sigma}_{\rho_1}$	$\hat{\rho}_{12}$	$\hat{\sigma}_{\rho_{12}}$
Sub 1	.130	.118	.046	.131	-.027	.118	-.034	.133
Sub 2	.239	.129	-.028	.151	.121	.129	-.124	.143
Sub 3	.054	.123	.334	.136	.00	.123	.180	.137
Sub 4	.181	.123	-.199	.136	.038	.123	-.050	.134
Sub 5	.163	.062	.153	.064	.046	.062	.049	.064

Panel B: January observations excluded.

	EW INDEX				VW INDEX			
Subperiod	$\hat{\rho}_1$	$\hat{\sigma}_{\rho_1}$	$\hat{\rho}_{12}$	$\hat{\sigma}_{\rho_{12}}$	$\hat{\rho}_1$	$\hat{\sigma}_{\rho_1}$	$\hat{\rho}_{12}$	$\hat{\sigma}_{\rho_{12}}$
Sub 1	.114	.123	-.061	.135	-.037	.123	-.069	.137
Sub 2	.131	.135	-.054	.153	.093	.135	-.011	.147
Sub 3	.116	.129	-.062	.145	-.021	.129	-.002	.143
Sub 4	.167	.128	-.194	.142	.00	.128	-.217	.142
Sub 5	.164	.066	-.043	.067	.027	.064	-.038	.066

Remarks:

Sub 1: January 1963–December 1968 (T=72).
Sub 2: January 1969–December 1973 (T=60).
Sub 3: January 1974–June 1979 (T=66).
Sub 4: July 1979–December 1984 (T=66).
Sub 5: Total period January 1963–December 1984 (T=264).

TABLE 5

Maximum likelihood estimates of the market model parameters under the assumption that the residuals follow an $AR(12)$ process,

$$\begin{cases} r_{it} = \iota_i \alpha_i + r_{mt}\beta_i + \epsilon_{it} \\ \epsilon_{it} = \rho_i \epsilon_{it-12} + \eta_{it}, \qquad t = 1, \ldots, 13. \end{cases}$$

Panel A: January observations included and EW index.

INDEX		EW INDEX								
Portfolio		P_1	P_2	P_3	P_5	P_{10}	P_{15}	P_{18}	P_{19}	P_{20}
Sub 1	$\hat{\alpha}_i$	.01487	.00488	.00406	-.00036	-.00026	-.00246	-.00102	-.00353	-.00224
	$(\sigma(\hat{\alpha}_i))$	(.00647)	(.00346)	(.00259)	(.00204)	(.00119)	(.00180)	(.00197)	(.00258)	(.00206)
	$\hat{\beta}_i$	1.34	1.33	1.28	1.21	1.06	.83	.63	.52	.57
	$(\sigma(\hat{\beta}_i))$	(.108)	(.074)	(.059)	(.041)	(.022)	(.035)	(.042)	(.050)	(.045)
	$\hat{\rho}$	.255	-.0135	-.109	.048	.127	.088	-.027	.095	-.089
	$(\sigma(\hat{\rho}))$	(.138)	(.144)	(.151)	(.168)	(.119)	(.120)	(.120)	(.120)	(.120)
Sub 2	$\hat{\alpha}_i$	-.00006	-.00372	-.00527	-.00356	-.00438	.00297	.00466	.00339	.00568
	$(\sigma(\hat{\alpha}_i))$	(.00914)	(.00748)	(.00504)	(.00266)	(.00087)	(.00280)	(.00320)	(.00331)	(.00330)
	$\hat{\beta}_i$	1.18	1.15	1.18	1.15	1.06	.90	.77	.70	.56
	$(\sigma(\hat{\beta}_i))$	(.067)	(.045)	(.047)	(.030)	(.015)	(.024)	(.038)	(.037)	(.041)
	$\hat{\rho}$	.553	.645	.405	.262	.255	.330	.203	.263	.198
	$(\sigma(\hat{\rho}))$	(.119)	(.105)	(.130)	(.144)	(.128)	(.125)	(.130)	(.128)	(.130)
Sub 3	$\hat{\alpha}_i$	.01949	.00776	.00115	.00186	.00054	-.00298	-.00473	-.00684	-.00895
	$(\sigma(\hat{\alpha}_i))$	(.01078)	(.00749)	(.00440)	(.00253)	(.00134)	(.00221)	(.00374)	(.00407)	(.00494)
	$\hat{\beta}_i$	1.51	1.33	1.24	1.23	1.00	.83	.73	.64	.53
	$(\sigma(\hat{\beta}_i))$	(.099)	(.072)	(.042)	(.033)	(.018)	(.027)	(.043)	(.045)	(.058)
	$\hat{\rho}$	.476	.452	.435	.028	-.018	.142	.248	.271	.199
	$(\sigma(\hat{\rho}))$	(.119)	(.120)	(.118)	(.136)	(.126)	(.125)	(.122)	(.121)	(.123)
Sub 4	$\hat{\alpha}_i$	.00435	-.00096	-.00067	-.00248	-.00235	-.00016	-.00054	-.00109	-.00209
	$(\sigma(\hat{\alpha}_i))$	(.00780)	(.00485)	(.00340)	(.00227)	(.00134)	(.00192)	(.00242)	(.00288)	(.00336)
	$\hat{\beta}_i$	1.17	1.14	1.06	1.17	1.05	.91	.84	.78	.71
	$(\sigma(\hat{\beta}_i))$	(.094)	(.046)	(.047)	(.036)	(.024)	(.028)	(.038)	(.038)	(.043)
	$\hat{\rho}$	.305	.486	.244	.143	-.019	.186	.258	.275	.291
	$(\sigma(\hat{\rho}))$	(.129)	(.121)	(.135)	(.143)	(.126)	(.124)	(.122)	(.121)	(.121)
Sub 5	$\hat{\alpha}_i$	.01051	.00303	.00099	-.00039	-.00159	-.00107	-.00108	-.00227	-.00256
	$(\sigma(\hat{\alpha}_i))$	(.00455)	(.00321)	(.00206)	(.00117)	(.00618)	(.00107)	(.00153)	(.00167)	(.00199)
	$\hat{\beta}_i$	1.31	1.22	1.17	1.18	1.04	.88	.75	.68	.57
	$(\sigma(\hat{\beta}_i))$	(.045)	(.030)	(.024)	(.017)	(.010)	(.014)	(.020)	(.021)	(.025)
	$\hat{\rho}$	.397	.433	.303	.101	-.002	.202	.224	.232	.246
	$(\sigma(\hat{\rho}))$	(.058)	(.057)	(.061)	(.064)	(.062)	(.061)	(.060)	(.060)	(.060)

Remarks:

Subperiod 1: January 1963–December 1968 (T=72).
Subperiod 2: January 1969–December 1973 (T=60).
Subperiod 3: January 1974–June 1979 (T=66).
Subperiod 4: July 1979–December 1984 (T=66).
Subperiod 5: January 1963–December 1984 (T=264).

TABLE 5 (Cont.)

Maximum likelihood estimates of the market model parameters under the assumption that the residuals follow an $AR(12)$ process,

$$\begin{cases} r_{it} = \iota_i \alpha_i + r_{mt}\beta_i + \epsilon_{it} \\ \epsilon_{it} = \rho_i \epsilon_{it-12} + \eta_{it}, \qquad t = 1, \ldots, 13. \end{cases}$$

<u>Panel A</u>: January observations included and VW index.

INDEX		VW INDEX								
Portfolio		P_1	P_2	P_3	P_5	P_{10}	P_{15}	P_{18}	P_{19}	P_{20}
Sub 1	$\hat{\alpha}_i$	.02952	.01889	.01688	.01091	.00835	.00294	.00251	.00051	-.00166
	$(\sigma(\hat{\alpha}_i))$	(.01066)	(.00787)	(.00702)	(.00647)	(.00371)	(.00212)	(.00115)	(.00097)	(.00059)
	$\hat{\beta}_i$	1.19	1.25	1.27	1.32	1.35	1.21	1.02	.99	.96
	$(\sigma(\hat{\beta}_i))$	(.234)	(.194)	(.172)	(.144)	(.099)	(.056)	(.036)	(.031)	(.020)
	$\hat{\rho}$	.343	.260	.266	.335	.192	.206	.009	-.037	-.069
	$(\sigma(\hat{\rho}))$	(.131)	(.136)	(.136)	(.137)	(.138)	(.118)	(.118)	(.120)	(.120)
Sub 2	$\hat{\alpha}_i$	-.01040	-.01338	-.01481	-.01302	-.01217	-.00342	-.00076	-.00119	.00234
	$(\sigma(\hat{\alpha}_i))$	(.01398)	(.01232)	(.00979)	(.00753)	(.00456)	(.00266)	(.00157)	(.00138)	(.00111)
	$\hat{\beta}_i$	1.45	1.45	1.41	1.49	1.49	1.31	1.17	1.07	.93
	$(\sigma(\hat{\beta}_i))$	(.154)	(.132)	(.133)	(.110)	(.079)	(.053)	(.039)	(.038)	(.020)
	$\hat{\rho}$	.533	.550	411	.366	.257	.118	-.132	-.090	.198
	$(\sigma(\hat{\rho}))$	(.117)	(.116)	(.128)	(.130)	(.128)	(.131)	(.131)	(.132)	(.130)
Sub 3	$\hat{\alpha}_i$	.03524	.02131	.01372	.01615	.01211	.00688	.00416	.00096	-.00268
	$(\sigma(\hat{\alpha}_i))$	(.02115)	(.01799)	(.01400)	(.01139)	(.00682)	(.00384)	(.00228)	(.00125)	(.00118)
	$\hat{\beta}_i$	1.29	1.17	1.23	1.23	1.17	1.11	1.09	1.02	.98
	$(\sigma(\hat{\beta}_i))$	(.240)	(.192)	(.153)	(.152)	(.099)	(.061)	(.039)	(.025)	(.021)
	$\hat{\rho}$	.536	.572	.559	.432	.350	.278	.177	-.025	.127
	$(\sigma(\hat{\rho}))$	(.107)	(.103)	(.108)	(.119)	(.118)	(.121)	(.124)	(.126)	(.125)
Sub 4	$\hat{\alpha}_i$	.00797	.00235	.00261	.00092	.00075	.00251	.00134	.00283	-.00109
	$(\sigma(\hat{\alpha}_i))$	(.00984)	(.00763)	(.00611)	(.00522)	(.00312)	(.00205)	(.00157)	(.00216)	(.00081)
	$\hat{\beta}_i$	1.16	1.19	1.10	1.18	1.20	1.06	1.02	1.00	.97
	$(\sigma(\hat{\beta}_i))$	(.165)	(.114)	(.108)	(.097)	(.061.)	(.044)	(.034)	(.027)	(.017)
	$\hat{\rho}$	.287	.354	.229	.178	.115	.007	-.012	.061	.053
	$(\sigma(\hat{\rho}))$	(.136)	(.129)	(.137)	(.134)	(.125)	(.126)	(.126)	(.126)	(.126)
Sub 5	$\hat{\alpha}_i$	.01622	.00826	.00592	.00407	.00260	.00222	.00161	.00002	-.00089
	$(\sigma(\hat{\alpha}_i))$	(.00782)	(.00637)	(.00514)	(.00419)	(.00270)	(.00158)	(.00092)	(.00060)	(.00051)
	$\hat{\beta}_i$	1.29	1.27	1.26	1.31	1.30	1.17	1.09	1.03	.96
	$(\sigma(\hat{\beta}_i))$	(.097)	(.078)	(.068)	(.063)	(.043)	(.026)	(.019)	(.014)	(.010)
	$\hat{\rho}$	.442	.450	.408	.341	.330	.270	.122	.002	.149
	$(\sigma(\hat{\rho}))$	(.056)	(.055)	(.057)	(.059)	(.058)	(.060)	(.061)	(.062)	(.061)

Remarks:

Subperiod 1: January 1963–December 1968 (T=72).
Subperiod 2: January 1969–December 1973 (T=60).
Subperiod 3: January 1974–June 1979 (T=66).
Subperiod 4: July 1979–December 1984 (T=66).
Subperiod 5: January 1963–December 1984 (T=264).

TABLE 5 (Cont.)

Maximum likelihood estimates of the market model parameters under the assumption that the residuals follow an $AR(12)$ process,

$$\begin{cases} r_{it} &= s_i\alpha_i + r_{mt}\beta_i + \epsilon_{it} \\ \epsilon_{it} &= \rho_i\epsilon_{it-12} + \eta_{it}, \qquad t = 1,\ldots,13. \end{cases}$$

Panel B: –January observations excluded – EW index.

INDEX		EW INDEX								
Portfolio		P_1	P_2	P_3	P_5	P_{10}	P_{15}	P_{18}	P_{19}	P_{20}
Sub 1	$\hat{\alpha}_i$	.01015	.00329	.00236	-.00093	.00061	-.00121	-.00029	-.00155	-.00259
	$(\sigma(\hat{\alpha}_i))$	(.00477)	(.00296)	(.00267)	(.00207)	(.00107)	(.00150)	(.00202)	(.00212)	(.00240)
	$\hat{\beta}_i$	1.26	1.26	1.22	1.15	1.07	.85	.66	.60	.53
	$(\sigma(\hat{\beta}_i))$	(.114)	(.076)	(.057)	(.041)	(.025)	(.037)	(.044)	(.050)	(.054)
	$\hat{\rho}$	-.029	-.139	.099	.170	.021	-.086	.039	-.023	.040
	$(\sigma(\hat{\rho}))$	(.145)	(.138)	(.138)	(.164)	(.142)	(.150)	(.144)	(.139)	(.137)
Sub 2	$\hat{\alpha}_i$	-.00808	-.00953	-.01039	-.00591	-.00351	.00529	.00821	.00684	.00913
	$(\sigma(\hat{\alpha}_i))$	(.00418)	(.00284)	(.00279)	(.00228)	(.00118)	(.00142)	(.00242)	(.00271)	(.00244)
	$\hat{\beta}_i$	1.14	1.14	1.09	1.14	1.07	.92	.78	.69	.57
	$(\sigma(\hat{\beta}_i))$	(.069)	(.051)	(.048)	(.032)	(.017)	(.025)	(.037)	(.036)	(.037)
	$\hat{\rho}$	-.115	-.253	-.049	.062	.008	-.211	-.0199	.134	-.048
	$(\sigma(\hat{\rho}))$	(.161)	(.160)	(.162)	(.161)	(.156)	(.164)	(.175)	(.161)	(.159)
Sub 3	$\hat{\alpha}_i$	.00984	.00061	-.00162	.00056	.00001	-.00098	-.00174	-.00366	-.00519
	$(\sigma(\hat{\alpha}_i))$	(.00679)	(.00894)	(.00277)	(.00248)	(.00167)	(.00187)	(.00277)	(.00268)	(.00339)
	$\hat{\beta}_i$	1.33	1.21	1.20	1.16	1.02	..87	.77	.74	.66
	$(\sigma(\hat{\beta}_i))$	(.118)	(.080)	(.051)	(.042)	(.023)	(.033)	(.051)	(.054)	(.069)
	$\hat{\rho}$	.026	-.187	-.056	.035	.225	.016	-.068	-.179	-.184
	$(\sigma(\hat{\rho}))$	(.151)	(.140)	(.138)	(.147)	(.138)	(.139)	(.138)	(.137)	(.137)
Sub 4	$\hat{\alpha}_i$	.00018	-.00340	-.00269	-.00416	-.00157	.00119	.00033	-.00017	-.00131
	$(\sigma(\hat{\alpha}_i))$	(.00500)	(.00344)	(.00272)	(.00202)	(.00115)	(.00173)	(.00191)	(.00210)	(.00246)
	$\hat{\beta}_i$	1.07	1.08	1.05	1.10	1.07	.93	.86	.81	.73
	$(\sigma(\hat{\beta}_i))$	(.093)	(.054)	(.048)	(.035)	(.024)	(.028)	(.036)	(.041)	(.049)
	$\hat{\rho}$	-.029	.124	.024	.037	-.174	.091	-.053	-.091	-.129
	$(\sigma(\hat{\rho}))$	(.131)	(.130)	(.131)	(.131)	(.129)	(.131)	(.131)	(.131)	(.130)
Sub 5	$\hat{\alpha}_i$	.00363	-.00162	-.00248	-.00254	-.00107	.00074	.00121	.00010	-.00005
	$(\sigma(\hat{\alpha}_i))$	(.00262)	(.00174)	(.00145)	(.00108)	(.00064)	(.00086)	(.00122)	(.00124)	(.00146)
	$\hat{\beta}_i$	1.21	1.19	1.15	11.15	1.06	.89	.77	.71	.61
	$(\sigma(\hat{\beta}_i))$	(.048)	(.032)	(.024)	(.018)	(.011)	(.015)	(.021)	(.022)	(.027)
	$\hat{\rho}$	-.017	-.039	.066	.068	.044	.028	.046	-.019	-.018
	$(\sigma(\hat{\rho}))$	(.065)	(.065)	(.065)	(.065)	(.065)	(.065)	(.065)	(.065)	(.065)

Remarks:

Subperiod 1: January 1963–December 1968 (T=72).
Subperiod 2: January 1969–December 1973 (T=60).
Subperiod 3: January 1974–June 1979 (T=66).
Subperiod 4: July 1979–December 1984 (T=66).
Subperiod 5: January 1963–December 1984 (T=264).

TABLE 5 (Cont.)

Maximum likelihood estimates of the market model parameters under the assumption that the residuals follow an $AR(12)$ process,

$$\begin{cases} r_{it} = \iota_i \alpha_i + r_{mt}\beta_i + \epsilon_{it} \\ \epsilon_{it} = \rho_i \epsilon_{it-12} + \eta_{it}, \qquad t = 1, \ldots, 13. \end{cases}$$

<u>Panel B</u>: January observations excluded and VW index.

INDEX		VW INDEX								
Portfolio		P_1	P_2	P_3	P_5	P_{10}	P_{15}	P_{18}	P_{19}	P_{20}
Sub 1	$\hat{\alpha}_i$	.02054	.01287	.01135	.00688	.00690	.00300	.00228	.00052	-.00140
	$(\sigma(\hat{\alpha}_i))$	(.00729)	(.00552)	(.00520)	(.00511)	(.00349)	(.00184)	(.00129)	(.00111)	(.00565)
	$\hat{\beta}_i$	1.05	1.21	1.22	1.23	1.29	1.17	1.02	.98	.96
	$(\sigma(\hat{\beta}_i))$	(.244)	(.200)	(.173)	(.148)	(.105)	(.064)	(.040)	(.034)	(.021)
	$\hat{\rho}$	.018	-.089	.014	.154	.123	-.0256	.094	.095	-.038
	$(\sigma(\hat{\rho}))$	(.146)	(.147)	(.144)	(.154)	(.141)	(.136)	(.134)	(.134)	(.136)
Sub 2	$\hat{\alpha}_i$	-.02167	-.02301	-.02234	-.01944	-.01631	-.00510	-.00061	-.00083	.00288
	$(\sigma(\hat{\alpha}_i))$	(.00633)	(.00536)	(.00522)	(.00484)	(.00316)	(.00246)	(.00196)	(.00166)	(.00091)
	$\hat{\beta}_i$	1.42	1.44	1.41	1.51	1.50	1.38	1.18	1.07	.92
	$(\sigma(\hat{\beta}_i))$	(.160)	(.139)	(.125)	(.114)	(.076)	(.052)	(.042)	(.032)	(.020)
	$\hat{\rho}$	-.142	-.191	-.083	-.044	-.084	.064	.073	.137	.017
	$(\sigma(\hat{\rho}))$	(.159)	(.165)	(.163)	(.158)	(.167)	(.195)	(.232)	(.187)	(.161)
Sub 3	$\hat{\alpha}_i$	.01619	.00717	.00487	.00465	.00628	.00224	.00393	.00034	-.00149
	$(\sigma(\hat{\alpha}_i))$	(.00878)	(.00681)	(.00579)	(.00545)	(.00315)	(.00164)	(.00195)	(.00121)	(.00092)
	$\hat{\beta}_i$	.95	.96	1.05	1.07	1.06	1.02	1.03	1.00	1.00
	$(\sigma(\hat{\beta}_i))$	(.216)	(.174)	(.144)	(.134)	(.085)	(.055)	(.052)	(.028)	(.021)
	$\hat{\rho}$	-.182	-.248	-.208	-.180	-.238	-.0129	-.319	-.113	-.106
	$(\sigma(\hat{\rho}))$	(.146)	(.141)	(.140)	(.150)	(.137)	(.146)	(.141)	(.138)	(.163)
Sub 4	$\hat{\alpha}_i$	.00256	-.00131	-.00053	-.00209	.00	.00228	.00134	.00068	-.00091
	$(\sigma(\hat{\alpha}_i))$	(.00707)	(.00601)	(.00496)	(.00443)	(.00302)	(.00187)	(.00178)	(.00144)	(.00056)
	$\hat{\beta}_i$	.97	1.05	1.05	1.13	1.18	1.08	1.03	1.00	.97
	$(\sigma(\hat{\beta}_i))$	(.152)	(.116)	(.106)	(.095)	(.063)	(.045)	(.036)	(.027)	(.016)
	$\hat{\rho}$	-.005	.100	-.000	-.003	.021	-.127	.066	.116	-.395
	$(\sigma(\hat{\rho}))$	(.131)	(.131)	(.131)	(.131)	(.131)	(.130)	(.131)	(.130)	(.120)
Sub 5	$\hat{\alpha}_i$	.00470	-.00067	-.00165	-.00182	-.00066	.00093	.00117	.00004	-.00026
	$(\sigma(\hat{\alpha}_i))$	(.00397)	(.00324)	(.00295)	(.00263)	(.00183)	(.00100)	(.00084)	(.00066)	(.0040)
	$\hat{\beta}_i$	1.13	1.17	1.19	1.24	1.25	1.14	1.07	1.02	.96
	$(\sigma(\hat{\beta}_i))$	(.097)	(.080)	(.069)	(.062)	(.044)	(.027)	(.019)	(.015)	(.010)
	$\hat{\rho}$	-.028	-.047	.010	.001	-.007	-.088	.047	.054	-.040
	$(\sigma(\hat{\rho}))$	(.065)	(.065)	(.065)	(.065)	(.065)	(.065)	(.065)	(.065)	(.065)

Remarks:
Subperiod 1: January 1963–December 1968 (T=72).
Subperiod 2: January 1969–December 1973 (T=60).
Subperiod 3: January 1974–June 1979 (T=66).
Subperiod 4: July 1979–December 1984 (T=66).
Subperiod 5: January 1963–December 1984 (T=264).

TABLE 6

Point estimates and heteroskedastic consistent standard error estimates for the intercept of the augmented market model.

Panel A: Intercept dummy included.

$$r_{it} = \iota_T \alpha_i + D_{Jt}\alpha_{iJ} + r_{mt}\beta_i + r_{mt}D_{Jt}\beta_{iJ} + \epsilon_{it}, \quad t = 1, \ldots, T.$$

		P_1	P_2	P_3	P_5	P_{10}	P_{15}	P_{18}	P_{19}	P_{20}
EW	α_i	.00361	-.00165	-.00240	-.00249	-.00107	.00073	.00118	.00012	-.00004
	σ_i(LS)	.00265	.00184	.00139	.00103	.00063	.00082	.00118	.00127	.00153
	σ_i(HC)	.00256	.00178	.00133	.00100	.00061	.00081	.00114	.00123	.00145
	σ_i(HC2)	.00258	.00177	.00134	.00101	.00061	.00082	.00115	.00124	.00146
	α_{iJ}	.07646	.05537	.04207	.01810	-.00532	-.02182	-.02901	-.02997	-.02835
	σ_{iJ}(LS)	.01126	.00780	.00591	.00436	.00265	.00350	.00502	.00539	.00649
	σ_{iJ}(HC)	.00952	.00756	.00582	.00378	.00251	.00267	.00464	.00551	.00753
	σ_{iJ}(HC2)	.01065	.00832	.00622	.00432	.00281	.00284	.00501	.00610	.00881
VW	α_i	.00464	-.00076	-.00164	-.00182	-.00066	.00090	.00118	.00004	-.00026
	σ_i(LS)	.00428	.00357	.00305	.00275	.00195	.00122	.00084	.00064	.00044
	σ_i(HC)	.00401	.00333	.00287	.00259	.00182	.00114	.00080	.00063	.00042
	σ_i(HC2)	.00404	.00334	.00290	.00262	.00184	.00115	.00081	.00063	.00043
	α_{iJ}	.15470	.12140	.10023	.07690	.04320	.01770	.00450	-.00050	-.00761
	σ_{iJ}(LS)	.01517	.01265	.01052	.00974	.00692	.00433	.00297	.00227	.00156
	σ_{iJ}(HC)	.01933	.01652	.01373	.01215	.00935	.00582	.00322	.00274	.00188
	σ_{iJ}(HC2)	.02093	.01790	.01463	.01327	.01010	.00643	.00364	.00298	.00206

Panel B: Intercept dummy excluded.

$$r_{it} = \iota_T \alpha_i + r_{mt}\beta_i + r_{mt}D_{Jt}\beta_{iJ} + \epsilon_{it}, \quad t = 1, \ldots, T.$$

		P_1	P_2	P_3	P_5	P_{10}	P_{15}	P_{18}	P_{19}	P_{20}
EW	α_i	.00786	.00163	-.00006	-.00149	-.00187	-.00048	-.00043	-.00144	-.00161
	σ(LS)	.00280	.00195	.00148	.00108	.00061	.00086	.00122	.00129	.00154
	σ_i(HC)	.00273	.00192	.00145	.00101	.00059	.00084	.00118	.00128	.00151
	σ_i(HC2)	.00275	.00193	.00146	.00102	.00060	.00085	.00119	.00129	.00152
VW	α_i	.01696	.00891	.00634	.00430	.00278	.00231	.00158	.00000	-.00087
	σ_i(LS)	.00485	.00397	.00337	.00293	.00200	.00121	.00081	.00061	.00044
	σ_i(HC)	.00458	.00400	.00339	.00293	.00199	.00119	.00078	.00062	.00044
	σ_i(HC2)	.00499	.00410	.00347	.00300	.00204	.00122	.00080	.00062	.00045

Remarks: $\sigma_i(\cdot)$ is the standard deviation estimate obtained with:
"$(\cdot)$=LS " the least squares covariance matrix.
"$(\cdot)$=HC " White (1980) covariance matrix.
"$(\cdot)$=HC2" MacKinnon and White (1985) jacknife covariance matrix.

TABLE 7

Parameters of the augmented market model under the assumption that the residuals follow an autoregressive process $AR(12)$.

Panel A: Intercept dummy included.

$$\begin{cases} R_{it} &= \iota_T \alpha_i + \alpha_{iJ} + R_{mt}\beta_i + R_{mt}D_J\beta_{iJ} + \epsilon_{it}, \\ \epsilon_{it} &= \rho_{12}\epsilon_{it-12} + \eta_{it}, \quad t = 1, \ldots, (T-12). \end{cases}$$

Index		P_1	P_2	P_3	P_5	P_{10}	P_{15}	P_{18}	P_{19}	P_{20}
EW	α_i	.00540	-.00188	-.00248	-.00249	-.00108	.00074	.00123	.00019	.00004
		(.00302)	(.00222)	(.0015)	(.00108)	(.00060)	(.00086)	(.0018)	(.00142)	(.00170)
	α_{iJ}	.07817	.05778	.0425	.0181	-.00526	-.02199	-.0293	-.02867	-.0292
		(.01288)	(.00894)	(.00630)	(.0044)	(.00258)	(.00362)	(.0054)	(.0058)	(.0072)
	β_i	1.21	1.18	1.14	1.15	1.06	.89	.77	.71	.62
		(.047)	(.032)	(.025)	(.018)	(.011)	(.015)	(.021)	(.023)	(.027)
	β_{iJ}	.369	.132	.052	.077	-.039	-.042	-.109	-.114	-.149
		(.111)	(.0764)	(.0582)	(.0424)	(.0255)	(.0344)	(.0495)	(.0532)	(.0640)
	ρ_{12}	.185	.198	.061	.002	.089	.042	.109	.114	.149
		(.0616)	(.0609)	(.0619)	(.0621)	(.0621)	(.0620)	(.0617)	(.0617)	(.0614)
	TR^2	4.85	10.37	1.74	.00	.40	.47	3.11	3.46	5.89
	(p-val)	(.0277)	(.0013)	(.186)	(.97)	(.53)	(.49)	(.078)	(.063)	(.015)
VW	α_i	.00419	-.00128	-.00193	-.00205	-.00087	.00077	.00116	.0004	-.00026
		(.0049)	(.0043)	(.0036)	(.0031)	(.0023)	(.0015)	(.0008)	(.0006)	(.0005)
	α_{iJ}	.1557	.1225	.1007	.0772	.0485	.0178	.0046	-.0005	-.0076
		(.0178)	(.015)	(.0127)	(.0111)	(.0085)	(.0051)	(.0032)	(.0022)	(.0016)
	β_i	1.15	1.19	1.20	1.24	1.27	1.15	1.07	1.02	.96
		(.100)	(.082)	(.071)	(.064)	(.045)	(.028)	(.020)	(.015)	(.011)
	β_{iJ}	.847	.447	.343	.412	.149	.116	.185	.036	.00
		(.275)	(.226)	(.195)	(.177)	(.124)	(.078)	(.055)	(.042)	(.029)
	ρ_{12}	.155	.207	.171	.139	.194	.187	.069	.006	.063
		(.061)	(.061)	(.061)	(.062)	(.061)	(.061)	(.062)	(.062)	(.062)
	TR^2	6.34	11.34	7.73	5.17	9.96	9.27	1.26	.008	1.05
	(p-val)	(.012)	(.0007)	(.0054)	(.0229)	(.0016)	(.0023)	(.026)	(.93)	(.31)

Panel B: Intercept dummy excluded.

$$\begin{cases} R_{it} &= \iota_T \alpha_i + R_{mt}\beta_i + R_{mt}D_J\beta_{iJ} + \epsilon_{it}, \\ \epsilon_{it} &= \rho_{12}\epsilon_{it-12} + \eta_{it}, \quad t = 1, \ldots, (T-12). \end{cases}$$

Index		P_1	P_2	P_3	P_5	P_{10}	P_{15}	P_{18}	P_{19}	P_{20}
EW	α_i	.00817	.00182	.00018	-.00148	-.00136	-.00019	-.0005	-.00152	-.00174
		(.0033)	(.0024)	(.0017)	(.0011)	(.00059)	(.0008)	(.0014)	(.0015)	(.0018)
	β_i	1.21	1.18	1.14	1.15	1.06	.88	.78	.72	.62
		(.050)	(.034)	(.027)	(.019)	(.011)	(.018)	(.022)	(.023)	(.028)
	β_{iJ}	.668	.386	.226	.168	-.066	-.209	-.151	-.206	-.223
		(.1055)	(.0733)	(.0556)	(.0378)	(.0218)	(.0357)	(.046)	(.0487)	(.0581)
	ρ_{12}	.199	.249	.165	.045	.031	.105	.168	.152	.175
		(.0607)	(.0600)	(.0611)	(.0619)	(.0619)	(.0616)	(.0611)	(.0613)	(.0137)
	TR^2	10.49	16.43	7.21	.54	.26	2.92	7.44	6.07	8.05
	(p-val)	(.0012)	(.00)	(.0072)	(.461)	(.609)	(.087)	(.0064)	(.0137)	(.0045)
VW	α_i	.01613	.00825	.00494	.00391	.00255	.00217	.00152	.00	-.00088
		(.0068)	(.0057)	(.0031)	(.0028)	(.0026)	(.0015)	(.00087)	(.0006)	(.0005)
	β_i	1.18	1.21	1.24	1.24	1.28	1.15	1.07	1.02	.97
		(.106)	(.086)	(.076)	(.067)	(.046)	(.029)	(.020)	(.015)	(.011)
	β_{iJ}	.816	.514	.363	.506	.205	.148	.149	.035	-.017
		(.289)	(.234)	(.203)	(.184)	(.126)	(.078)	(.054)	(.041)	(.029)
	ρ_{12}	.359	.379	.346	.282	.281	.228	.074	.005	.184
		(.058)	(.057)	(.058)	(.059)	(.059)	(.060)	(.061)	(.062)	(.061)
	TR^2	34.18	37.98	31.59	21.06	20.93	18.80	1.44	.007	4.71
	(p-val)	(.50E-8)	(.71E-10)	(.19E-7)	(.44E-5)	(.47E-5)	(.0002)	(.2301)	(.931)	(.0299)

Remark: The TR^2 statistics is distributed as a χ^2 distribution with 1 degree of freedom.

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