An Analysis of Merger-Induced Nonstationarity in Market Model Parameters of Acquiring Firms

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Abstract

This paper formulates a modification of the market model methodology that allows direct testing for the impact of mergers on market model parameters, while simultaneously estimating acquiring firm abnormal stockholder returns. A large sample of merger events is stratified in three ways: conglomerate versus nonconglomerate, stock versus cash transactions, and by merger activity level. Results of prior studies are examined for sensitivity to a failure to appropriately incorporate the impact of merger on market model parameters. Results indicate that negative drift in postannouncement returns for acquiring firms is partly the result of parameter shifts.

Introduction

Many papers have addressed the impact of merger on acquiring firm stockholders using variations of the market model or capital asset pricing model [1, 2, 5, 6, 7, 8, 9, 13, 16, 18, 19, 21, 25, and 26]. Researchers have identified a number of issues in connection with the application of this methodology, all dealing with whether and how merger events impact the return-generating processes of acquiring firms.

One problem that researchers generally have recognized is that a systematic relationship between the merger event and measured market model parameters could lead to biases in estimates of merger-related abnormal performance.[1] Concerns have arisen that the assumptions of parameter stationarity underlying application of the market model may be invalidated by the events under study. Numerous methodological conventions have been used for the estimation of market model parameters. Among these are: estimation of parameters over an interval that excludes a reasonable period surrounding the event under study; distinction between pre- and postmerger intervals; and use of a moving window

[1] Mandelker [19] notes about an 8 percent decline in beta for 241 mergers; Joehnk and Nielsen [11] report an insignificant decline for conglomerates and an insignificant increase for nonconglomerates; Langetieg, Haugen, and Wichern [14] report increases in beta.

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interval for parameter estimation, where the interval excludes the period for which the abnormal return estimate is desired. Although these conventions may mitigate the consequences of parameter nonstationarity, they will not eliminate the problem when the firm in question participates in similar events during the postmerger parameter estimation period.[2]

A second concern expressed by researchers is that the market model methodology should yield estimates of abnormal performance that reflect only the merger event under study, i.e., that the return series not be contaminated by the presence of other significant corporate events. Thus, a sampling convention employed in a number of early studies was the exclusion of mergers for which the acquiring firm engaged in another significant acquisition for some specified period (e.g., three years) before or after the event of interest. A problem with this sampling convention is that the resulting samples of relatively inactive acquiring firms may not be representative of the population of acquiring firms [13, 14]. In recognition of this sampling problem, some recent studies have focused exclusively on firms that were active acquirers and have developed and tested hypotheses unique to this class of acquiring firms [2, 18, 21]. This group of studies, however, made no systematic attempt to come to grips with the potential event-induced nonstationarities in the return-generating process underlying measured abnormal returns.

A third issue regarding the application of market model methodology to merger events is whether the means of payment (financing) of an acquisition significantly affects market model parameters of the acquiring firms, hence measured abnormal returns at announcement dates as well as subsequent dates. Again, studies investigating the relationship of means of payment to merger-related returns have not addressed this issue fully [3, 8, 25, and 26].

A final question raised in recent research concerns apparently abnormal postmerger returns to acquiring firms that have been documented by a number of researchers. Malatesta [16] suggests failure to account for shifts in risk parameters around the time of the merger as one explanation for these findings. Other possible explanations mentioned by Malatesta are an inefficient securities market and sample specific biases, e.g., choice of a sampling period that was characterized by unanticipated increases in the frequency of antitrust litigation.

This paper utilizes an approach that allows direct testing for the systematic impact of merger events on market model parameters, while simultaneously

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[2]For example, Dodd and Ruback [5] use intercept and slope dummies to determine whether individual firms experienced parameter shifts consequent to tender offers. They decided to use both before and after parameter estimates to assess abnormal performance. They could not test for parameter shifts for many firms in their sample due to data limitations, and their methodology did not consider possible confounding of the effects of multiple events.

estimating abnormal stockholder returns.[3] These issues are examined in three sampling periods and in the context of three means of differentiating merger events: the conglomerate vs. nonconglomerate type classification, the medium of exchange, and the merger activity level of the acquiring firm. The results are compared to more conventional market model methodology,[4] with emphasis on postmerger abnormal returns.

Evidence is found that mergers affect market model intercept and slope parameters, but that the direction and magnitude of the impact is sensitive to the sampling period. The magnitude of these effects does not appear to influence substantially conclusions relating to overall merger announcement-related gains to acquiring firm stockholders. Nevertheless, evidence is provided that failure of conventional methodology to accommodate merger influences on market model parameters may partially account for findings in some studies of postmerger abnormal returns.

This paper develops a general model of the impact of merger events on the return-generating process of acquiring firms. A significant advantage of this approach is that it uses the historical record of merger activity and returns for each acquiring firm in testing for the impact of merger, thereby accommodating the fact that many acquiring firms engage in merger relatively frequently.

A Modified Market Model of Merger Effects

Suppose the returns on security j in period t are generated by the process

$$(1) \quad \tilde{r}_{jt} = \alpha_{jt} + \beta_{jt}\tilde{r}_{mt} + \tilde{\varepsilon}_{jt},$$

where:

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[3]Recently, numerous researchers have shown how multiple regression techniques may be applied in event studies. Unlike the present study, these techniques involve variations of the seemingly unrelated regression (SUR) model. See Thompson [24] and Malatesta [17] for comparisons of these approaches with the more traditional residuals technique and Schipper and Thompson [22] for an example of an application to event studies.

[4]This paper uses the phrase *conventional market model methodology* throughout in reference to approaches that utilize regression or forecast errors generated from any variation of the market model or CAPM as measures of event-related abnormal returns. Examples of the use of conventional methodology are [1, 2, 5, 6, 7, 9, 11, 13, 16, 19, and 21]. To the extent that other methodologies (e.g., the mean-adjusted returns approach) rely on benchmarks established without due consideration of the impact of the event on the benchmark, many of the concerns raised above remain valid.

$\tilde{r}_{jt}, \tilde{r}_{mt}$ = excess returns on security and a market index portfolio, respectively;

α_{jt} and β_{jt} = firm and period specific parameters; and

$\tilde{\varepsilon}_{jt}$ = a random noise term;

$Cov(\tilde{\varepsilon}_{jt}, \tilde{r}_{mt})$ = 0

Let t = 1, . . . , H index the periods over which returns are measured for estimation of market model parameters according to

(2) $\tilde{r}_{jt} = \hat{\alpha}_j + \hat{\beta}_j \tilde{r}_{mt} + \tilde{e}_{jt}$,

where:

$\hat{\alpha}_j, \hat{\beta}_j$ = the market model intercept and slope coefficient estimates (by OLS) for security j, and

$\tilde{e}_{jt}$ = the market model residual for security j in period t.

Suppose for each security j, a sequence of similar events, e.g., mergers, occurs over the estimation period, t = [1, . . . , H]. Let D_{jt} be a dummy variable such that D_{jt} = 1 if firm j experiences such an event in period t and Djt = 0 if firm j does not experience such an event in period t. Suppose further that the similar events affect returns on security j and the stochastic process generating returns on security j in the following ways:

(i) Security j realizes an abnormal return of amount z in the month in which the event occurs, so that $\tilde{\varepsilon}_{jt} = \tilde{e}_{jt} + zD_{jt}$, where $\tilde{e}_{jt}$ is a random noise term;

(ii) The merger results in a shift in the market model intercept parameter, i.e., $\alpha_{jt} = \alpha_{j,t-1} + \Delta\alpha D_{jt}$; and

(iii) The merger results in a change in the firm's systematic risk; i.e., $\beta_{jt} = \beta_{j,t-1} + \Delta\beta D_{jt}$.

Effects (i) and (iii) posited above are recognized in the literature and have simple economic interpretations that are consistent with the capital asset pricing model. The $\Delta\alpha$ effect, however, does not have an easy interpretation within a single factor CAPM context. According to the CAPM, α_{jt} should be zero for all t. Nevertheless, empirical studies regularly document nonzero intercepts. Negative market model intercepts, as well as decreases in the intercepts, generally are interpreted as adverse to the welfare of stockholders, because, *ceteris paribus*, they imply a lower level of return per unit of risk.[5]

The effects of the events on the coefficients α_{jt} and β_{jt} are cumulative, given that $\Delta\alpha$ and $\Delta\beta$ are constants, and the true coefficients of equation (1) for the period t = [1, . . . , H] may be defined recursively as follows:

$$(3)\quad \alpha_{jt} = \alpha_{jo} + \Delta\alpha \sum_{s=1}^{t} D_{js} = \alpha_{jo} + \Delta\alpha C_{jt}$$

and

$$(4)\quad \beta_{jt} = \beta_{jo} + \Delta\beta \sum_{s=1}^{t} D_{js} = \beta_{jo} + \Delta\beta C_{jt},$$

where:

C_{jt} = the cumulative number of events for firm j through period t.[6]

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[5]Security returns may be generated by a multifactor CAPM or by a more general APT process. For example, ignoring merger event effects, suppose that true returns are given by

$$\tilde{r}_{jt} = \alpha_j + \beta_{1j}\tilde{r}_{1t} + \beta_{2j}\tilde{r}_{2t} + \tilde{\varepsilon}_{jt},$$

where r_{1t} and r_{2t} are orthogonal. If the parameters α_j and β_{1j} are estimated from the misspecified model given by equation (2), then it easily may be shown that

$$E(\hat{\alpha}_j) = \alpha_j + \beta_{2j}E(r_2).$$

It follows that a change in the intercept term in equation (2) may be decomposed into three elements:

$$E(\Delta\hat{\alpha}_j) = \Delta\alpha_j + (\Delta\beta_{2j})E(r_2) + \beta_{2j}(\Delta E(r_2)).$$

Thus, finding an intercept effect may be explained by either changing volatility with respect to an omitted factor ($\Delta\beta_{2j}$), a change in the level of an omitted factor ($\Delta E(r_2)$), or some combination of the two, as opposed to finding a pure intercept effect ($\Delta\alpha_j$).

[6]The effects of an event on market model parameters are modeled as permanent effects as presented in conditions (ii) and (iii) and equations (3) and (4). It is a simple matter to redefine them as transitory, e.g., assume that the impact on α lasts for only 12 months, after which α reverts to its original level. Such a redefinition of the event effects implies a straightforward redefinition of C_{jt}. Some empirical findings to be reported later reflect on the choice between permanent or transitory effects. Additionally, the beta effect was modelled as proportional to the magnitude of the firm's premerger beta coefficient, the rationale being that firms with very high betas may be more likely to engage in mergers that lowered beta. This specification did not appear to represent an improvement over the one presented in the text.

Given effect (i) of the impact of an event on security return, the return-generating process (1) may be rewritten as:

(5) $\tilde{r}_{jt} = [\alpha_{jo} + \Delta\alpha C_{jt}] + [\beta_{jo} + \Delta\beta C_{jt}]\tilde{r}_{mt} + zD_{jt} + \tilde{e}_{jt}$,

where:

$Cov(\tilde{e}_{jt}, \tilde{e}_{it}) = 0$.[7]

If the true return-generating process is given by equation (5) but is estimated by using ordinary least squares in equation (2), then under plausible conditions, expected values of the least squares estimators for the parameters of equation (2) are given by (Appendix):

(6) $E[\hat{\beta}_j] = \beta_{jo} + \Delta\beta\bar{C}_j$ and

(7) $E[\hat{\alpha}_j] = \alpha_{jo} + \Delta\alpha\bar{C}_j + z\,\bar{D}_j$,

where:

$\bar{C}_j$ = the average value of C_{jt} for firm j over the estimation period; and
$\bar{D}_j$ = the average value of D_{jt}.

Using equations (6) and (7) to substitute for α_{jo} and β_{jo} in equation (5) gives the following expression for return in period t:

$$(8)\quad \tilde{r}_{jt} = (E[\hat{\alpha}_j] + E[\hat{\beta}_j]\tilde{r}_{mt}) + \Delta\alpha(C_{jt} - \bar{C}_j) + \Delta\beta(C_{jt} - \bar{C}_j)\,\tilde{r}_{mt} + z(D_{jt} - \bar{D}_j) + \tilde{e}_{jt}$$

[7]It may be shown that this formulation of merger effects may be restated as a special case of what elsewhere has been referred to as *intervention analysis*. See Box and Tiao [4] for a general statement of the methodology, and Larker *et al.* [15] for a general discussion of its application to event studies. This paper's formulation makes a number of assumptions regarding the timing of merger effects (e.g., the market model intercept and slope coefficient effects are permanent, become effective at the merger announcement date, and are of the same magnitude for all mergers within a given category) that make the approach tractable for large sample analysis. Such assumptions are necessary to any application of intervention analysis.

The advantage of equation (8) relative to equation (5) for empirical purposes is that the unobservable initial values of market model parameters, α_{jo} and β_{jo}, are replaced with expected values of the unconditional market model parameters that can be estimated from available data.[8] This equation represents the basis for subsequent testing for the impact of mergers on the risk and return of acquiring firms.

Methodological Issues and Data

Equation (8) may be viewed as an alternative to the more conventional event study/abnormal returns analysis for estimating the impact of merger on the returns of acquiring firms (hereafter, equation (8) and its attendant assumptions is called the *modified market model*). Inasmuch as its derivation explicitly incorporates estimation of the impact of the event of interest on market model parameters, it has an advantage over the conventional market model approach. The periods over which the model is estimated may be defined in terms of event time (and therefore vary from firm to firm), or the same calendar time interval may be chosen for all firms. In either case, equation (8) may be estimated for each acquiring firm independently or cross-section and time series data may be pooled.

For ease in exposition of specific hypotheses to be tested, equation (8) is restated in the following form:

$$(8')\quad \tilde{r}_{jt} = a\,E[\hat{\alpha}_j] + b\,(E[\hat{\beta}_j]\tilde{r}_{mt}) + \Delta\alpha(C_{jt} - \bar{C}_j)$$

$$+ \Delta\beta[(C_{jt} - \bar{C}_j)\,\tilde{r}_{mt}] + z(D_{jt} - \bar{D}_j) + \tilde{e}_{jt}$$

A two stage approach is used. The first stage involves estimation of unconditional market model intercept and slope parameters for each acquiring firm over the calendar periods chosen. The second stage utilizes these as estimates of $E[\hat{\alpha}_j]$ and $E[\hat{\beta}_j]$ with market returns data and merger activity data for each firm in the sample to estimate the parameters of equation (8').

The results of estimation of equation (8') provide direct estimates of the effects of merger events on market model parameters ($\Delta\alpha$, $\Delta\beta$), as well as abnormal returns to merger announcements (z) and provide insights on the validity of the often used market model methodology for establishing merger

[8]The conditional market model approach involves first estimating market model parameters from nonevent period data, then using those parameters to estimate abnormal returns within the event period. The unconditional market model approach estimates market model parameters from both event and nonevent periods.

benefits to stockholders of acquiring firms. If the effect coefficients $\Delta\alpha$ and $\Delta\beta$ are nonzero, then conventional market model estimates of abnormal returns will be biased (see equation (A8) in the Appendix). The null hypotheses regarding the impact of mergers on the returns of acquiring firms may be stated as:

(H1) $\Delta\alpha = 0$
(H2) $\Delta\beta = 0$
(H3) $z = 0$.

The assumptions under which the model is developed imply that the intercept term is zero, that the parameters a and b are unity, and that the other slope parameters are constant over time and firms. Thus, it appears reasonable to estimate equation (8') from a pooling of cross-section and time series data. Because prior research indicates that stock returns are likely to be heteroskedastic across firms and over time, White's [27] test for heteroskedasticity and appropriately modified procedures for tests of hypotheses concerning the parameters estimated are employed.[9]

It is arguable that the impact of merger may not be uniform across all mergers. It easily is shown that equation (8) generalizes to accommodate multiple merger types wherein each type (k) may be characterized by distinct merger effect coefficients ($\Delta\alpha_k$, $\Delta\beta_k$, and z_k):

$$(9)\quad \tilde{r}_{jt} = (E[\hat{\alpha}_j] + E[\hat{\beta}_j]\tilde{r}_{mt}) + \sum_k \Delta\alpha_k (C_{kjt} - \bar{C}_{kj})$$

$$+ \sum_k \Delta\beta_k (C_{kjt} - \bar{C}_{kj})\,\tilde{r}_{mt} + \sum_k z_k(D_{kjt} - \bar{D}_{kj}) + \tilde{e}_{jt}$$

The empirical issue, however, is specifying *a priori* the merger types and assigning individual mergers to a type category. Three such classification schemes are tested. In the first, mergers are classified as pure conglomerate or nonconglomerate by the Federal Trade Commission (FTC). The second means of differentiating among mergers is on the basis of the means of payment to the acquired firms' stockholders, whether stock or cash. Finally, mergers were differentiated based on the level of acquisition activity by the acquiring firm.

[9] See Judge *et al.* [12, Chapter 8] for a discussion of the econometric alternatives when using pooled data. Although the model could be generalized somewhat by allowing parameters to vary across firms and using the seemingly unrelated regression technique, such an approach would raise other problems when the number of firms is large relative to the number of time periods, as is the case herein (cf. Thompson [24, p. 165] and Schipper and Thompson [21, p. 201].

It is also arguable that the impact of mergers may not be uniform over time due to changing economic or regulatory environments. Therefore, samples were analyzed from three separate five year periods, 1961-1965, 1965-1970, and 1971-1975, as well as the 15 year period 1961-1975.[10, 11]

Merger announcement dates are utilized as the relevant event dates, as prior research indicates that announcement dates are superior to consummation dates for revealing market reaction to mergers.

The merger history of all acquiring firms listed in the Federal Trade Commission large merger series was compiled for the period 1961-1975. For each of the periods 1961-1965, 1966-1970, 1971-1975, and 1961-1975, all acquiring firms meeting the following conditions were selected: CRSP monthly return data were available for the period, and announcement dates were available from the *Wall Street Journal Index* for all large mergers by the firm during the period. Table 1 summarizes the number of acquiring firms for each period, the total number of mergers by those firms during the period, and a breakdown of the total by FTC type category. Table 1 also indicates the distribution of means of payment in the two five year periods and the ten year period for which the data were available from the FTC merger series.

For each acquiring firm and each period, market model parameters and residuals were estimated according to equation (2). The sample average market model results for each of the four samples are presented in Table 2.

Results of Merger Effects Estimation: Nonstratified Samples

Equation (8'), where all mergers are considered homogeneous in terms of their impact on returns, is estimated first. The following section considers stratification of mergers by FTC type, means of payment, and merger activity level.

Table 3 presents the result of ordinary least squares estimation of the parameters of equation (8'), assuming that all events are homogeneous in terms

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[10]Although the model developed in the previous section is equally applicable conceptually to either daily or monthly return data, monthly data were chosen to avoid confounding the longer term structural impact of merger with more temporary trading phenomena that might be present in proximity to merger announcements.

[11]The period 1968-1970 often is selected in analyses of merger events for the impact of several regulatory events that may impact the profitability or form of takeovers: the 1968 Williams amendment and its 1970 extension, Accounting Principals Board's Opinions 16 and 17, and the Tax Reform Act of 1969. See Schipper and Thompson [22] for an analysis of the impact of these events on acquiring firm shareholder returns.

of their impact on returns. Application of White's [27] test for heteroskedasticity resulted in a rejection of the hypothesis of homoskedastic errors. Therefore, the t-statistics presented in Table 3 reflect the use of a heteroskedasticity-consistent covariance matrix estimator in drawing inferences regarding the regression coefficients (see White [27]).

In two of the three five year samples investigated, there were statistically significant merger announcement month effects of 1.5 percent and 2.0 percent. The results indicate a significant positive impact on the market model intercept in the 1961-1965 sample and a significant negative impact in the 1966-1970 sample. The impact of merger on the beta coefficient was negative and highly significant in 1971-1975, with an estimated decline of .13. There was a decline of .05 in 1966-1970 that was not quite significant at 5 percent.

The 15 year sample results indicate a significantly positive announcement effect of 1.3 percent, a small, but significantly negative impact on the market model intercept coefficient, and no significant impact on acquiring firm betas.[12] Although it is tempting to conclude from the positive announcement effects noted in Table 3 that acquiring firms benefit from mergers, such a conclusion is not warranted when the intercept effects are taken into account. These effects represent a more or less permanent impact on monthly returns of the acquiring firm.

Comparison of the announcement effect results with prior research is facilitated by referring to the Jensen and Ruback [10] summary of one month announcement effects from five merger studies published between 1980 and 1983 (see [10, p. 12]): ". . . all five of the one month announcement effects . . . are

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[12]An alternative approach to the problem of heteroskedasticity in the context of this study involves use of weighted regression. If the residuals from equation (8') for each firm are not identically distributed, then the heteroskedasticity encountered in the pooling approach may be dealt with by first estimating the residual standard deviation for each firm, then weighting the observations for each firm in the pooled regression by the reciprocals of the respective estimated standard deviations. This is analogous to the general practice of computing standardized residuals or prediction errors in the more conventional market model methodology. Standardized residuals are presumed to be identically and independently distributed, so that the assumptions underlying standard significance tests on average residuals (prediction errors) are met. This research utilizes the standard error of estimate from the unconditional market model regression for each firm over the time period in question to obtain estimates of the error variances. The results of the weighted regressions are essentially the same as with the ordinary least squares results in Table 3. Estimated magnitudes of the various effects are similar. Some notable differences in significance levels are the lack of statistical significance of the intercept parameter effects in the 1961-1965 and 1961-1975 samples and the substantial increase in significance of the slope parameter effects in the 1966-1970 sample.

positive, but only the estimate of 3.48 percent by Asquith, Bruner, and Mullins (1983) is significantly different from zero. The weighted averages are 1.37 percent for the one month announcement effect . . ."

To facilitate comparison of announcement effects between the modified market model methodology and the conventional approach, two variations of the conventional approach were applied to the 15 year sample. The variations of the conventional approach follow closely methodology recently employed by Malatesta [16]. One variation uses preannouncement market model parameters to estimate announcement period prediction errors; the other, postannouncement parameters. In the former instance, 60 months of returns from event month -72 to -13 were utilized to estimate preannouncement market model parameters for each merger event. Prediction errors then were calculated for event months -12 to +12. This results in a sample of 269 merger events occurring during the period 1967 through 1974. Similarly, to replicate the methodology that uses postannouncement parameters, 60 months of returns from event months +13 to +72 were used to estimate postannouncement market model parameters for each merger event; prediction errors then were calculated for event months -12 to +12 based on these parameters. This approach results in a sample of 307 merger events during the period 1962 through 1969.[13] The average announcement month prediction errors for the two variations of conventional methodology were 1.6 percent and 1.2 percent respectively, both significant at 1 percent.[14]

The evidence from nonstratified merger samples on hypotheses (H1) through (H3) is summarized below.

(H1) Mergers impact market model intercepts for acquiring firms, with evidence of statistically significant effects noted in three of the four samples (two negative, one positive).

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[13]It is noteworthy that the same data base of 15 years of monthly returns produces a considerably larger sample of merger events for analysis when using the modified market model methodology, as the pre- or postannouncement parameter estimation period merger events are lost when using the conventional approach. In this case, 473 merger events were included in the sample using the modified approach versus 269 or 307 for the two variations of the conventional approach. Furthermore, if the conventional approach employs screening on the basis of multiple merger events by the same firm, which is not necessary when employing the modified approach, then the samples would be reduced even further. To the authors' knowledge, the 473 merger events in the 15 year sample period represent the most comprehensive sample ever used in this type of study.

[14]Significance levels were established using appropriately standardized prediction errors (Malatesta [16, p. 166]).

(H2) Mergers appear to lower beta coefficients in two of the three five year samples, although the effect was only of marginal statistical significance in one of the samples.

(H3) Mergers appear to have a small, but statistically significant positive announcement effect on acquiring firm stock returns.

Furthermore, on the basis of comparisons between other merger studies as summarized in Jensen and Ruback and application of conventional market model methodology to the data base, the magnitude of estimated announcement period returns to acquiring firms appears not to be influenced significantly by merger-induced nonstationarities in market model parameters.

Results of Merger Effects Estimation: Mergers Stratified by FTC Type, Means of Payment, and Merger Activity Level

Conglomerates vs. Nonconglomerate Mergers

As indicated earlier, there is no reason *a priori* to assume that the merger effects posited are homogeneous across merger groups. Merger events were stratified according to whether they were pure conglomerate or one of the other four Federal Trade Commission types (horizontal, vertical, product extension, and market extension). This division seems logical in light of the objective of discerning differential market model parameter effects (especially beta effects) because, by definition, merging entities are unrelated in a pure conglomerate merger and not so in the other categories. Results are presented in Table 4. Using White's heteroskedastic consistent covariance matrix estimator, this research tested for equality of the parameter estimates between the two categories. The only significant differences found were in the 1961-1965 and 1971-1975 sample periods, in which the estimated impacts on the market model intercept coefficients were positive for nonconglomerate acquisitions, but negative for conglomerates. This differential in intercepts implies that a significant differential merger-related return in favor of nonconglomerate mergers was realized over an extended period subsequent to the announcement month in two of the three five year periods. None of the differences in estimated systematic risk effects or announcement effects was statistically different. Because the differential intercept effect did not appear in the 1961-1975 period, it appears that the overall evidence does not unambiguously support a distinction between conglomerate and nonconglomerate acquisitions based on their impact on acquiring firm stockholder returns.

The Medium of Exchange

The literature suggests that there are two important aspects of the means of payment in merger transactions that potentially could impact merger performance as well as the risk of the merging firms. One of these aspects is the leverage effect associated with the form of financing. If cash for stock transactions tends to be associated with significant increases in leverage in the acquiring firm, then market model beta estimates should increase accordingly (see [23] for a discussion and some evidence on means of payment and leverage changes in acquiring firms). Some recent research emphasizes signalling implications of the alternative payment means, based primarily on the theory espoused by Myers and Majluf [20] that investors feel that firms tend to issue stock when it is overvalued. Thus, the use of cash as the medium of exchange represents a positive signal (relative to stock) regarding the value of the acquiring firm. Recent investigations into the merger-related gains to acquiring firms have supported this positive information effect, finding that returns to acquiring firms are significantly lower in stock-financed transactions [3, 8, and 25]. Additionally, Franks *et al.* [8], by separately estimating pre- and postmerger market model parameters, document declines in market model intercepts for equity-financed acquisitions. They find no evidence that the hypothesized signal associated with the means of payment may resolve uncertainty in the sense of affecting the acquiring firm's beta coefficient, i.e., that the signal is associated with a change in volatility of the firm's returns relative to the market.

As with the merger type comparisons above, merger events were stratified according to stock or cash as the medium of exchange, and equation (9) was estimated for the periods for which FTC data on the medium of exchange were available, 1966-1970, 1971-1975, and 1966-1975 (see Table 1). The results are presented in Table 5. Market model intercept effects were about the same for stock and cash transactions in 1966-1970, but lower for stock transactions in the other two samples (1971-1975 and 1966-1975), though the differences were never statistically significant. Somewhat surprisingly, cash acquisitions tended to lower acquiring firm betas relative to stock transactions, and the difference was significant at 1 percent in the 1971-1975 sample. Thus, it appears that information effects related to the resolution of uncertainty are overwhelming any leverage effects in terms of the impact of mergers on acquiring firm systematic risk. Finally, after allowing for market model parameter effects, the results in Table 5 indicate significantly positive announcement (z) effects for stock transactions in all three samples and smaller statistically insignificant announcement effects for cash acquisitions. The differences between stock and cash announcement effects are in favor of stock transactions, but are never statistically significant.

For comparison with conventional methodology, the conventional approach using both preannouncement and postannouncement market model parameters was applied to the stratified samples. Using preannouncement parameter estimates for 174 stock and 34 cash transactions occurring during 1967-1974, announcement month returns of 1.5 percent for stock transactions and 2.4 percent for cash were found. When postannouncement parameters were used for 131 stock and 22 cash transactions occurring during 1962-1969, announcement month returns of 1.6 percent for stock transactions and 0.9 percent for cash were found. The modified market model approach reaches similar conclusions regarding announcement month effects as the conventional methodology with postannouncement market model parameters. What the conventional approach fails to provide are estimates of the effects of the merger event on the parameters themselves. The results of the modified market model approach suggest that over an extended period, the relatively unfavorable market model intercept ($\Delta\alpha$) effect for stock transactions would offset the difference in announcement month effects (z) between stock and cash transactions.[15]

Acquiring Firm Merger Activity Level

A number of recent studies have differentiated among acquiring firms on the basis of information embodied in merger announcements by such firms (see [2, 18, and 21]). Schipper and Thompson [21] find evidence to support the argument that the positive net present value benefits of acquisition programs may be anticipated due to explicit announcement of merger programs by some acquiring firms and, therefore, that specific merger event announcement effects might understate the value of acquisitions to such firms.

Malatesta and Thompson [18] formalize the argument that merger events are partially anticipated and provide evidence that announcement effects understate the economic impact of acquisitions for active acquirers. Using acquisitions made by 30 firms identified in the Schipper and Thompson study, they estimate announcement effects averaging 2.9 percent and total economic impact averaging 3.5 percent [18, p. 248], yielding a difference of about 0.6 percent. If merger events are unanticipated for relatively inactive acquirers, it would be expected that

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[15]For example, for the 1966-1975 period, the differential α effect is 0.2 percent (per month) in favor of cash transactions. This would overwhelm the announcement month (z) effect differential of 1.3 percent within six months after the event. Franks *et al.* [8] also report finding negative intercept effects in comparing pre- and postannouncement date parameter estimates for equity transactions. Conclusions drawn from market model intercept effects must be qualified in light of the discussion in note 5 above concerning interpretation of such effects.

the announcement (z) effect in the model would differ between active and inactive acquiring firms.

In order to test for this differential announcement effect, acquiring firms in the sample were identified as active acquirers if they made three or more announcements within a five year period or four or more announcements in the 15 year sample period. Table 6 gives the frequency distribution of total merger announcements by each acquiring firm for each sample period. By the criterion stated above, there are 10, 27, 4, and 29 active acquirers, respectively, in the four samples.[16]

Slope dummy variables with a value of one are used to identify active acquirers and with a value of zero to identify inactive acquirers. The slope coefficients estimated for the interaction terms represent the estimated difference between the value of the merger effect parameter for active acquirers and those for inactive acquirers. Results are shown in Table 7. As predicted by the partial anticipation hypothesis, the announcement impact for active acquirers is lower than that of inactive acquirers in every sample period and significantly lower in the 1961-1965 period and the 1961-1975 period. The weighted average of the three five year differences is -1.70 percent; the difference for the 15 year sample is -2.22 percent.

Neither of the other merger effect parameters ($\Delta\alpha$, $\Delta\beta$) are significantly different at 5 percent between the active and inactive acquiring firms, nor are their algebraic signs consistent across the four sampling periods.

An Analysis of Pre- and Postannouncement Abnormal Returns

As indicated in the introduction, concern has recently been expressed over the large negative abnormal returns to acquiring firm shareholders in the periods following acquisition.[17] One explanation is that the securities market is inefficient with regard to acquisition-related information. An alternative hypothesis concerning these negative returns is that they result from merger-induced nonstationarity in market model parameters. To place the estimates of merger-related parameter nonstationarity in perspective, it is noted that of the

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[16] The current sample differs from those of Schipper and Thompson [21] and Malatesta and Thompson [18] in a potentially important respect. All of the mergers used in this study originated from the FTC large merger series. Therefore, many smaller acquisitions by these acquiring firms are omitted.

[17] In some studies, the period is measured relative to merger announcement in the financial press; in others, the merger consummation or board approval dates are used. Jensen and Ruback [10] provide a concise summary of a number of studies that analyze postacquisition performance.

five studies summarized by Jensen and Ruback, all but one find negative abnormal returns over the year following the event. The negative outcomes range from -13.7 percent to -1.32 percent. For the four samples reported in Table 3 above, the estimates of the permanent impact of merger on market model intercepts translate into twelve month postannouncement abnormal returns ranging from -12.0 percent to +3.6 percent.[18]

Table 8 presents twelve month pre- and postannouncement average cumulative residuals from the modified market model for the four samples described earlier, plus other samples to be described below.[19] As with a number of other studies, positive cumulative abnormal returns in the pre-announcement period were found for all four of the previously described samples, with statistically significant values for the 1971-1975 and 1961-1975 periods (first column, rows 1-4).[20] These results may reflect the fact that firms that have

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[18]In an alternative specification, the market model intercept effect, $\Delta\alpha$, was modeled as transitory, with a 12 month life. This is accomplished by redefining the cumulative merger activity variable relating to $\Delta\alpha$ as a moving 12 month cumulant. The signs of the resulting estimates were the same as those shown in Table 3, but the magnitudes ranged from -0.8 percent to 0.0 percent. Only the coefficient for the 1971-1975 sample was statistically significant. Other parameter estimates were not affected in any noteworthy respect by the revision of the model, and overall F-statistics were insignificantly lower for the transitory effects model. Experimentation with a transitory beta effect also did not appear to improve the performance of the model. Subsequent estimates are made with the permanent effects version of the model.

[19]In the parlance of the typical event-relative time framework, the cumulative residuals shown are for the months -12 to -1 and +1 to +12 relative to the merger announcement month.

[20]Letting se_j represent the cumulative error for firm j from equation (8'), then the mean preannouncement period abnormal return is

$$\overline{se} = (1/N)\sum_{j=1}^{N} se_j = (1/N)\sum_{j=1}^{N}\sum_{t=-12}^{-1} e_{jt},$$

and the standard deviation of se_j is

$$S(se_j) = [\frac{1}{N-1}\sum_{j=1}^{N}(se_j - \overline{se})^2]^{.5}$$

t-statistics are computed as

$$t(se) = \frac{\overline{se}}{S(se_j)\,/\,\sqrt{N}} .$$

experienced good fortune or superior management are more likely to become acquirers, or they may reflect prior capitalization of anticipated merger-related gains.[21]

Table 8 presents the postannouncement period average cumulative residuals. These are negative and significant in the 1961-1965 and 1971-1975 samples, positive and significant in the 1966-1970 sample, and not significantly different from zero in the 15 year sample period. It appears on the basis of these findings that any postannouncement abnormal returns observed are sample specific. Given that these postannouncement returns were generated from a model that specifically allowed for systematic parameter nonstationarity, it seems reasonable to conclude that the findings of significant postannouncement abnormal returns in the five year samples were not a result of the form of parameter nonstationarity that is modelled herein.

The question still remains whether pre- or postannouncement returns found in prior studies may have been at least partially an artifact of the methodology used, which either ignored parameter nonstationarity by using preannouncement market model parameters for estimating postannouncement abnormal returns or attempted to deal with it by using market model parameters estimated over a postannouncement period.[22] To address this question, pre- and postannouncement abnormal returns were reestimated for the 15 year sample using conventional methods. Details of the methodology are described above.[23]

An analogous procedure is followed for the postannouncement period abnormal returns.

[21]The association was investigated between the abnormal return over the preannouncement period and the activity level of the acquiring firm, but no evidence of a strong systematic relationship was found. This may be interpreted as a weak refutation of the argument that abnormal returns during the 12 months immediately preceding announcement are the result of anticipation of benefits from yet unannounced merger events by merger active firms.

[22]If each acquiring firm participates in only one merger, then the conventional market model approach using postannouncement market model parameters would give unbiased estimates of the parameters α_{jt} and β_{jt} of equation (1) that are relevant to the postannouncement period. As the acquiring firms in the sample have, on average, 1.99 acquisitions each over the 15 year period, however, it is likely that many of the firms will have postannouncement market model parameters that confound the effects of two or more merger events. The modified market model methodology readily accommodates this phenomenon.

[23]The cumulative prediction errors from the conventional methodology first were standardized by dividing the error for each period by the associated standard deviation (unique in each period due to the index portfolio component). These standardized errors were summed over the appropriate 12 month interval for each

To ensure that the same sets of merger events are used in the comparison between methodologies, the pre- and postannouncement abnormal returns for the same merger events are presented using the modified market model methodology developed in this paper.

The results are presented in Table 8. Row 5 presents pre- and postannouncement period cumulative prediction errors for the conventional methodology employing preannouncement market model parameters; row 6, the corresponding results from the modified market model. Both models indicate statistically significant abnormal returns during the 12 month preannouncement period. Similar results for the preannouncement period are obtained when comparing the postannouncement market model parameter method (rows 7 and 8) with the approach developed in this paper.

Comparison of postannouncement period abnormal returns reveals noteworthy differences. The conventional approach using preannouncement model parameters results in a statistically significant -5.5 percent 12 month postannouncement period abnormal return (row 5). The modified market model approach yields a smaller, statistically insignificant -2.4 percent return for the same set of merger events (row 6). The postannouncement parameters form of the conventional model gives a positive (not quite significant) return of 3.3 percent, whereas the modified model yields a much smaller and statistically insignificant 0.7 percent return (rows 7 and 8). At least part of the observed abnormal postannouncement period returns may be attributed to parameter nonstationarity that is not dealt with adequately in the conventional application of market model methodology, even when using postannouncement market model parameters.

Returning momentarily to the distinction between mergers based on the means of payment, Franks *et al.* [8], utilizing preannouncement parameters, found that acquirers utilizing equity experienced postannouncement losses over 24 months following the event, whereas cash bidders did not experience significant postannouncement abnormal returns. With postannouncement parameters, they found no statistically significant postannouncement period returns for either class of mergers. Table 9 presents 12 month pre- and postannouncement period abnormal returns using the modified market model approach for stock and cash transactions, as well as for the conventional preannouncement parameters methodology applied to the data for the period 1967-1974. Table 9 indicates positive and mostly significant preannouncement period returns, but no significant postannouncement period returns. Row 4 of the table replicates the findings of Franks *et al.* using conventional

merger. t-statistics were computed by dividing the sample mean cumulative error by the estimated standard error of the mean for the N-firm sample.

methodology, with significantly negative postannouncement period returns for acquiring firms using stock as the exchange medium. Row 5 represents the results when the modified methodology is applied to the same period.[24] The negative postannouncement abnormal returns to equity transactions are insignificantly different from zero when the modified methodology is utilized. These results further support the conclusion that abnormal postannouncement period returns may be attributed to parameter nonstationarity that is not dealt with adequately in the conventional application of market model methodology.

Summary and Conclusions

This paper proposes and tests an alternative approach to the estimation of the effects of merger on returns to acquiring firm stockholders. The modified market model methodology developed herein has a significant advantage over the conventional market model approach in that it simultaneously evaluates the impact of mergers on market model parameters and residuals, makes use of the merger experience of acquiring firms over the full period for which returns data are collected, and does not require elimination of merger-active firms from the sample. It eliminates the need for separate estimation of market model parameters for pre- and postannouncement periods and the potential confusion over which set of parameters is relevant. It is amenable to a number of modifications, such as allowing distinctions for various classes of events or modelling effects as transitory rather than permanent.

The empirical findings indicate that mergers have a significant impact on the level and volatility of returns to acquiring firm stockholders. Specifically, strong statistical evidence is found of positive one month announcement effects of about 1.3 percent.

Findings relative to the impact of merger events on acquiring firm market model parameters indicate statistically significant effects on both intercept and slope, but these effects vary in magnitude and statistical significance over sampling periods.

Attempts to improve model specification by differentiating merger types met with mixed results. Over three sampling periods analyzed, no consistent distinction between pure conglomerate and nonconglomerate merger events was found. Evidence is found, however, to suggest that mergers involving cash as

[24] Note that with the modified methodology, many mergers in this period are lost because the means of payment information is not available for all mergers by a given firm during the sample period. The nonavailability of means of payment data for earlier years prevented a meaningful comparison between conventional methodology using postannouncement parameters and modified market model methodology.

the means of payment reduces stock returns volatility, while mergers using stock do not. This result may be explained in terms of the information content of the medium of exchange used.

Evidence is found that abnormal returns associated with announcements are smaller for active acquirers. This is consistent with the hypothesis that merger events by active acquirers are partially anticipated by the market. No significant differences in either market model intercept or slope effects were found among active and inactive acquirers.

Comparison of postannouncement period abnormal returns from the modified market model methodology to those found with more conventional methodology on the same samples of merger events indicates that post-announcement abnormal returns are smaller in magnitude when the modified model is applied to the data. This is especially true when transactions are stratified on the basis of stock vs. cash as the means of payment, in which case there were no instances of statistically significant postannouncement abnormal returns. Thus, the conjecture by Jensen and Ruback [10] that observed negative drift in postannouncement returns may be caused by market model parameter nonstationarity is supported. Of course, the possibility remains that observed nonstationarities result to some extent from failure of the single factor (or zero beta) market model to adequately describe security returns, in which case postacquisition abnormal returns may be a manifestation of specification error due to omitted factors.

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Table 1
Sample Description

Panel A -- Merger Distribution by FTC Type Within Samples

Sample Period	Number of Acquiring Firms	FTC Type* 1	2	3	4	5	Total Number of Mergers
1961-1965	124	17	28	93	10	28	176
1966-1970	185	16	23	167	5	72	283
1971-1975	113	31	19	43	4	53	150
1961-1975+	238	44	59	234	13	123	473

Panel B -- Merger Distribution by Means of Payment

Sample Period	Total Number of Mergers	Means of Payment*: Stock	Cash	Other or Unknown
1966-1970	283	181	28	74
1971-1975	150	95	34	21
1966-1975+	333	210	48	75

*Source: Federal Trade Commission, *1975 Statistical Report on Mergers and Acquisitions*. The FTC classification scheme is: 1 = horizontal; 2 = vertical, 3 = conglomerate market extension; 4 = conglomerate product extension; and 5 = other conglomerate

+ The sampling procedure used implies that the 15 year (1961-1975) and ten year (1966-1975) sample are not the union of the respective five year samples, as additional data requirements are imposed

Table 2
Sample Average Market Model Results*

Sample Period	Number of Firms	$\bar{\hat{\alpha}}$	$\bar{\hat{\beta}}$	$\bar{R}^2$
1961-1965	124	-.0003	1.186	.3397
1966-1970	185	.0039	1.266	.3448
1971-1975	113	.0055	1.154	.3482
1961-1975	238	.0025	1.165	.3260

The excess returns form of the market model with the CRSP value-weighted index was used. Monthly risk free rate data were obtained from the CRSP government bond file

$\bar{\hat{\alpha}}$, $\bar{\hat{\beta}}$, $\bar{R}^2$ are sample mean intercept, slope, and coefficient of determination statistics respectively, from the market model regressions

Table 3
Estimates of Merger Effects for Nonstratified Sample

	1961-1965 60 Months	1966-1970 60 Months	1971-1975 60 Months	1961-1975 180 Months
Summary Statistics:				
R-squared	0.353	0.360	0.344	0.311
F-statistic	810.553**	1,246.072**	709.592**	3,865.800**
m.s. Error	0.0037	0.0060	0.0074	0.0061
Number of Observations	7,440	11,100	6,780	42,840
Coefficient:				
a	1.000	0.996	0.992	0.999
	(0.002)	(-0.049)	(-0.067)	(-0.019)
b	1.002	0.999	1.018	1.000
	(0.139)	(-0.076)	(0.766)	(0.002)
$\Delta\alpha$	0.003**	-0.010**	-0.002	-0.001*
	(2.683)	(-9.390)	(-1.042)	(-2.075)
$\Delta\beta$	0.013	-0.050	-0.130**	-0.006
	(0.402)	(-1.912)	(-3.099)	(-0.553)
z	-0.002	0.015**	0.020**	0.013**
	(-0.385)	(2.946)	(2.787)	(3.555)

Notes: T-values computed from White's [27] heteroskedastic-consistent covariance estimator are reported in parentheses beneath the coefficient estimates. The coefficients a and b were tested for their difference from unity. One asterisk indicates significance of the coefficient at the 5 percent level; two, at the 1 percent level

Table 4
Estimates of Merger Effects for Mergers Stratified by FTC Type

	1961-1965 60 Months	1966-1970 60 Months	1971-1975 60 Months	1961-1975 180 Months
Summary Statistics:				
R-squared	0.354	0.360	0.345	0.311
F-statistic	508.273**	778.669**	444.897**	2,416.243**
m.s. Error	0.0037	0.0060	0.0074	0.0061
Number of Observations	7,440	11,100	6,780	42,840
Coefficient:				
a	1.001	0.995	0.993	0.999
	(0.008)	(-0.052)	(-0.057)	(-0.018)
b	1.002	0.999	1.019	1.000
	(0.101)	(-0.081)	(0.781)	(0.010)
$\Delta\alpha_{nc}$	0.004**	-0.011**	0.001	-0.001
	(3.456)	(-11.283)	(0.625)	(-1.546)
$\Delta\alpha_{c}$	-0.006	-0.010**	-0.009*	-0.001
	(-1.598)	(-3.555)	(-2.418)	(-0.915)
$\Delta\beta_{nc}$	0.033	-0.060	-0.092	-0.002
	(0.926)	(-1.922)	(-1.827)	(-0.146)
$\Delta\beta_{c}$	-0.128	-0.020	-0.214*	-0.018
	(-1.214)	(-0.331)	(-2.450)	(-0.628)
z_{nc}	-0.001	0.014*	0.013	0.011**
	(-0.260)	(2.467)	(1.487)	(2.605)
z_{c}	-0.003	0.016	0.033**	0.020*
	(-0.303)	(1.602)	(2.767)	(2.514)

Notes: The nc and c subscripts denote nonconglomerate and conglomerate merger effect coefficients, respectively. T-values computed from White's [27] heteroskedastic-consistent covariance estimator are reported in parentheses beneath the coefficient estimates. The coefficients a and b were tested for their difference from unity. One asterisk indicates significance of the coefficient at the 5 percent level; two, at the 1 percent level

Table 5
Estimates of Merger Effects for Mergers Stratified by Medium of Exchange

	1966-1970 60 Months	1971-1975 60 Months	1966-1975 120 Months
Summary Statistics:			
R-squared	0.349	0.343	0.335
F-statistic	483.682**	368.385**	961.837**
m.s. error	0.0058	0.0072	0.0065
Number of Observations	7,200	5,640	15,240
Coefficient:			
a	0.994	1.001	1.047
	(-0.053)	(0.008)	(0.342)
b	1.003	1.014	1.005
	(0.142)	(0.517)	(0.339)
$\Delta\alpha_{stock}$	-0.010**	-0.003	-0.005**
	(-6.770)	(-1.306)	(-4.587)
$\Delta\alpha_{cash}$	-0.010	-0.002	-0.002
	(-1.821)	(-0.491)	(-0.716)
$\Delta\beta_{stock}$	-0.081*	-0.035	0.000
	(-2.195)	(-0.614)	(-0.013)
$\Delta\beta_{cash}$	-0.157	-0.330**	-0.105
	(-1.252)	(-3.712)	(-1.611)
z_{stock}	0.018**	0.022*	0.019**
	(2.577)	(2.571)	(3.174)
z_{cash}	0.016	0.000	0.006
	(0.900)	(0.023)	(0.465)

Notes: The stock and cash subscripts denote stock and cash merger effect coefficients, respectively. T-values computed from White's [27] heteroskedastic-consistent covariance estimator are reported in parentheses beneath the coefficient estimates. The coefficients a and b were tested for their difference from unity. One asterisk indicates significance of the coefficient at the 5 percent level; two, at the 1 percent level

Table 6
Frequency Distributions of the Number of Merger Announcements by Acquiring Firms

	1961-1965		1966-1970		1971-1975		1961-1975	
Number of Mergers	Number of Firms	%	Number of Firms	%	Number of Firms	%	Number of Firms	%
1	90	72.6	129	69.7	86	76.1	132	55.5
2	24	19.4	29	15.7	23	20.4	48	20.2
3	5	4.0	17	9.2	1	0.9	29	12.2
4	3	2.4	8	4.3	1	0.9	14	5.9
5	1	0.8	1	0.5	1	0.9	4	1.7
6	1	0.8	1	0.5	1	0.9	4	1.7
7							3	1.3
8							3	1.3
9							0	0
10							1	0.4

Table 7
Test for Differences Between Active and Inactive Acquirers

	1961-1965 60 Months	1966-1970 60 Months	1971-1975 60 Months	1961-1975 180 Months
Summary Statistics:				
R-squared	0.354	0.360	0.344	0.311
F-statistic	508.386**	779.732**	443.938**	2,417.949**
m.s. Error	0.0037	0.0060	0.0074	0.0061
Number of Observations	7,440	11,100	6,780	42,840
Coefficient:				
a	1.000	0.995	0.991	0.998
	(0.001)	(-0.061)	(-0.074)	(-0.031)
b	1.003	1.000	1.016	1.001
	(0.196)	(0.006)	(0.669)	(0.081)
$\Delta\alpha$	0.004*	-0.012**	-0.003	-0.001*
	(2.293)	(-7.746)	(-1.598)	(-2.126)
$\Delta(\Delta\alpha)_a$	-0.001	0.004	0.007	0.001
	(-0.574)	(1.762)	(1.695)	(1.145)
$\Delta\beta$	0.024	-0.093*	-0.105*	-0.018
	(0.521)	(-2.330)	(-2.206)	(-1.113)
$\Delta(\Delta\beta)_a$	-0.019	0.079	-0.111	0.022
	(-0.304)	(1.554)	(-1.113)	(0.964)
z	0.006	0.018**	0.022**	0.020**
	(0.992)	(2.917)	(2.850)	(4.491)
$\Delta(z)_a$	-0.034**	-0.007	-0.016	-0.022**
	(-3.472)	(-0.617)	(-0.883)	(-3.007)

The a subscript denotes the differential effect of an active acquirer, relative to the effect of an inactive acquirer, which is represented by the corresponding nonsubscripted coefficient. T-values computed from White's [27] heteroskedastic-consistent covariance estimator are reported in parentheses beneath the coefficient estimates. The coefficients a and b were tested for their difference from unity. One asterisk indicates significance of the coefficient at the 5 percent level; two, at the 1 percent level

Table 8
Pre- and Postannouncement Period Cumulative Abnormal Returns

Row	Period:	Preannouncement	Postannouncement
1	1961-1965 Modified Model	.014 (147, 1.13)	-.045** (134, -3.32)
2	1966-1970 Modified Model	.016 (242, 1.03)	.036* (246, 2.46)
3	1971-1975 Modified Model	.065** (127, 2.90)	-.074** (109, -2.73)
4	1961-1975 Modified Model	.077** (442, 6.32)	-.012 (439, -1.00)
5	1967-1974 Conventional Model Pre-event Parameters	.089** (269, 5.35)	-.055* (269, -2.16)
6	1967-1974 Modified Model	.102** (269, 6.33)	-.024 (269, -1.52)
7	1962-1969 Conventional Model Postevent Parameters	.114** (307, 6.12)	.033 (307, 1.85)
8	1962-1969 Modified Model	.074** (307, 5.05)	.007 (307, 0.49)

The number of merger events and T-values are reported in parentheses beneath the coefficient estimates. One asterisk indicates significance at the 5 percent level; two, at the 1 percent level

Table 9
Pre- and Post-Announcement Period Cumulative Abnormal Returns by Medium of Exchange

Row	Period:	Preannouncement		Postannouncement	
		Stock	Cash	Stock	Cash
1	1966-1970 Modified Model	0.04* (126, 1.94)	0.059 (19, 1.24)	0.031 (124, 1.57)	0.047 (16, 0.92)
2	1971-1975 Modified Model	0.057** (79, 2.06)	0.086 (29, 1.94)	-0.050 (65, -1.57)	-0.121 (21, -1.66)
3	1966-1975 Modified Model	0.081** (146, 3.79)	0.144** (37, 3.03)	-0.031 (144, -1.50)	0.013 (31, 0.28)
4	1967-1974 Conventional Model Pre-event Parameters	0.097** (174, 4.94)	0.190** (34, 3.23)	-0.052* (174, -1.95)	0.042 (34, 1.00)
5	1967-1974 Modified Model	0.136** (85, 4.99)	0.134* (17, 2.07)	-0.010 (99, -0.37)	-0.015 (21, -0.25)

The number of merger events and T-values are reported in parentheses beneath the coefficient estimates. One asterisk indicates significance at the 5 percent level; two, at the 1 percent level

Appendix A
Derivation of Merger Effects on Measured Market Model Parameters and Residuals

Consider $\hat{\beta}_j$ (hats above variables will be used to indicate sample variances and covariances; bars, sample means). By definition,

$$\hat{\beta}_j = \hat{Cov}\,(r_{jt}, r_{mt}) / \hat{\sigma}^2_m.$$

The numerator may be rewritten, using equation (5), as:

$$\hat{Cov}\,(r_{jt}, r_{mt}) = \hat{Cov}\,(\alpha_{jt} + \beta_{jt}r_{mt} + e_{jt} + zD_{jt}, r_{mt}),$$

where $\alpha_{jt} = \alpha_{jo} + \Delta\alpha C_{jt}$, and $\beta_{jt} = \beta_{jo} + \Delta\beta C_{jt}$.

After some manipulation employing the identity $\hat{Cov}\,(x,y) = \overline{xy} - \bar{x}\ \bar{y}$:

$$\hat{Cov}\,(r_{jt}, r_{mt}) = \Delta\alpha\hat{Cov}\,(C_{jt}, r_{mt}) + \beta jo\, \hat{\sigma}^2_m + \Delta\beta\ [\bar{C}_j\, \hat{\sigma}^2_m + \hat{Cov}\,(C_{jt}, r^2_{mt}) - \bar{r}_m\, \hat{Cov}\,(C_{jt}, r_{mt})] + z\hat{Cov}(D_{jt}, r_{mt}) + \hat{Cov}(e_{jt}, r_{mt})$$

where:

$$\bar{C}_j = \sum_{t=1}^{H} C_{jt}/H \text{ is the longitudinal mean of } C_{jt} \text{ for firm j.}$$

Thus, the estimate, $\hat{\beta}_j$, is given by

(A1) $$\hat{\beta}_j = \beta jo + \Delta\beta\ \bar{C}_j + \varphi_j/\hat{\sigma}^2_m,$$

where

(A2) $$\varphi_j = (\Delta\alpha - \Delta\beta\bar{r}_m)\hat{Cov}(C_{jt}, r_{mt}) + \Delta\beta\hat{Cov}(C_{jt}, r^2_{mt}) + z\ Cov(D_{jt}, r_{mt}) + Cov(e_{jt}, r_{mt}).$$

Now consider $\hat{\alpha}_j = \bar{r}_j - \hat{\beta}_j\ \bar{r}_m$. First, note that

(A3) $$\bar{r}_j = \bar{\alpha}_j + \overline{\beta_{jt}r_{mt}} + z\bar{D}_j + \bar{e}_j.$$

Now $(\overline{\beta_{jt}r_{mt}}) = \hat{Cov}(\beta_{jt}, r_{mt}) + \bar{\beta}_j\ \bar{r}_m = \Delta\beta\hat{Cov}(C_{jt}, r_{mt}) + \bar{\beta}_j\ \bar{r}_m$.

Combining (A1) and (A3) with the definition of $\hat{\alpha}_j$ and simplifying yields:

(A4) $\hat{\alpha}_j = \alpha_{jo} + \Delta\alpha\bar{C}_j + z\bar{D}_j - \varphi'_j$, where

$$\varphi'_j = -\Delta\beta \hat{Cov}(C_{jt}, r_{mt}) + (\varphi_j \bar{r}_m)/\sigma_j^2.$$

Both φ and φ' contain terms that depend on the association between merger activity dummy variable, D_{jt}, and the market index return because

(A5) $$Cov(C_{jt}, r_{mt}) = \sum_{s \leq t} Cov(D_{js}, r_{ms}).$$

Given that the average number of mergers per firm in the sample employed in this study is around two, these sample covariances are small. Under the assumption that these covariances are zero and $\hat{Cov}(e_{jt}, r_{mt}) = 0$, the expressions for the expected values of the market model parameter estimates may be written as:

(A6) $E[\hat{\beta}_j] = \beta_{jo} + \Delta\beta\bar{C}_j$, and

(A7) $E[\hat{\alpha}_j] = \alpha_{jo} + \Delta\alpha\bar{C}_j + z\bar{D}_j$.

Thus, under the assumption that merger events are uncorrelated with the market index return, the following expression for the expected difference between measured market model residuals and true residuals may be derived:

(A8) $E[r_{jt} - (\hat{\alpha}_j + \hat{\beta}_j r_{mt})] =$

$$\Delta\alpha(C_{jt}-\bar{C}_j) + \Delta\beta(C_{jt} - \bar{C}_j)\, E[r_{mt}] + z(D_{jt} - \bar{D}_j).$$

To illustrate the significance of this expression, consider the more or less standard event study, where an acquiring firm experiences a single event in period T. In the event period, both $(C_{jt}-\bar{C}_j)$ and $(D_{jt} - \bar{D}_j)$ are positive, as C_{jt} and D_{jt} are above their average values.

In nonevent periods prior to T, because $C_{jt} = D_{jt} = 0$,

$$(C_{jt}-\bar{C}_j) = -\bar{C}_j < 0, \text{ and } (D_{jt} - \bar{D}_j) = -\bar{D}_j < 0,$$

whereas in nonevent periods subsequent to T,

$$(D_{jt} - \bar{D}_j) < 0, \text{ but } (C_{jt}-\bar{C}_j) > 0.$$

Thus, positive (negative) merger intercept ($\Delta\alpha$) or beta ($\Delta\beta$) effects would result, *ceteris paribus*, in negative (positive) expected residuals in the pre-event periods and positive (negative) expected residuals in the event and postevent periods. A positive (negative) abnormal return (z) effect would result in negative (positive) expected residuals in both pre-event and postevent periods, but a positive expected residual in the event period. Obviously, conventional market model residuals either can under- or overstate true residuals, depending on the signs and magnitudes of the merger effects that are present.