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Mergers and the Market Concentration Doctrine: Evidence from the Capital Market

The market concentration doctrine predicts that a horizontal merger is more likely to have collusive, anticompetitive effects the greater the merger-induced change in industry concentration. Since a collusive, anticompetitive merger generates an increase in the industry's quality-adjusted product price (or a decrease in factor prices), it also follows from the doctrine that the merger-induced expected benefits to the product market *rivals* of the merging firms should be an increasing function of the concentration change. The empirical results of this paper, which are based on the industry wealth effect of a large sample of horizontal mergers, including cases found in violation of antimonopoly laws, fail to support this prediction. This conclusion is robust with respect to assumptions concerning the probability that a proposed merger will be prevented by the law enforcement agencies, and it continues to hold after transforming the industry wealth effect into a hypothetical, constant expected change in the industry's product price. The results imply that the levels of concentration and market shares found in the Department of Justice's merger guidelines are unlikely to identify truly anticompetitive mergers.

I. Introduction

The last 3 decades of vigorous enforcement of U.S. antimerger laws is based on the market concentration doctrine, one of the oldest and most controversial propositions in industrial economics. This doctrine, which is an implication of oligopoly models in the tradition of Cournot ([1838] 1927) and Nash (1950), holds that the level of industry concentration is a reliable index of the industry's market power. The empirical implication is that a relatively high level of industry concentration, which in the presence of entry barriers is believed to facilitate intraindustry collusion or dominant-firm pricing, should be associated with relatively large industrywide monopoly rents.[1] Following Bain (1951), numerous

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1. A closely related but somewhat less general prediction is that high levels of concentration should be associated with relatively high, supracompetitive *product prices*. Although

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studies have attempted to test this proposition by estimating the cross-sectional correlation between accounting measures of industry profits and the level of concentration.[2] However, although this correlation is indeed typically found to be positive, the same studies have generally failed to discriminate between the market concentration doctrine and alternative propositions, including the hypothesis that the positive correlation is simply driven by cross-industry differences in risk or average costs of production.[3] Furthermore, although mergers between producers of close substitutes represent an important cause of rising industry concentration, systematic tests of the market concentration doctrine based directly on the history and wealth efects of merger activity are virtually nonexistent.[4] Thus, the rather extreme view of the empirical validity of the market concentration doctrine reflected in antimerger policy is based on a relatively weak set of empirical evidence, which is at best only indirectly relevant for assessing the anticompetitive impact of horizontal mergers.[5]

This paper uses capital market data to test the antitrust thesis that a horizontal corporate combination is more likely to have anticompetitive effects the higher the premerger level of, and merger-induced change in, industry concentration. In an efficient capital market,[6] a merger-induced change in expected future product and factor prices translates into merger-induced abnormal stock returns to firms competing in the same industry as the bidder and target firms. In particular, the announcement of a horizontal merger that is expected to increase the probability of successful collusion among rival producers, or that eliminates one of the competitors of the price leader, will cause the market to impound the resulting expected increase in the industry's monopoly rents in the market values of the merging firms' close competitors. Eckbo (1983), Stillman (1983), and Eckbo and Wier (1985) reject this prediction based on the merger-induced abnormal performance of rivals of horizontal mergers challenged with violating Section 7 of the Clayton Act, the major federal antitrust law regulating merg-

evidence of supracompetitive pricing is sufficient to conclude that market power is present, it is clearly not necessary: monopoly rents can also be generated by means of collusion on nonprice variables, by monopsonizing inputs, or by sophisticated price discrimination schemes that need not be evident through the observed product price.

2. For a survey of the literature on the "structure-conduct-performance" paradigm, see, e.g., Scherer (1970) and Weiss (1974).

3. Brozen (1970), Demsetz (1973), Peltzman (1977), and Carter (1978) present evidence that supports the theory that industry concentration is predominantly a result of the expansion of relatively cost-efficient producers. The issue of cross-industry variation in risk is not explicitly addressed in the literature.

4. See, e.g., the survey of the literature by Brozen (1982).

5. See the merger guidelines issued by the Department of Justice in 1968 and revised in 1982 for a policy description.

6. See Muth (1961) and Fama (1970) for an elaboration of the efficient markets/rational expectations hypothesis under which security prices "fully reflect" available information.

ers.[7] These studies do not, however, attempt to exploit information on firm- and industry-specific characteristics that can shed light on the market concentration doctrine itself.

A horizontal merger produces a measurable change in the industry's level of concentration, as well as a change in the risk-adjusted present value of industry rents that is *directly associated* with the concentration change. Under the market concentration doctrine, this change in industry rents should, ceteris paribus, be positively correlated with the change in concentration. This value-based test in the changes of the two variables allows more specific inferences than can be drawn from the traditional cross-sectional correlation between the levels of (accounting) profits and concentration. The test results are also directly relevant for the issue concerning whether the levels of concentration and market shares given in the Department of Justice's merger guidelines are likely to identify truly anticompetitive mergers.

The rest of the paper is organized as follows. In Section II, I outline the nature of the cross-sectional regressions and how the framework relates to the more traditional "structuralist" approach to testing the market concentration doctrine. The basic dependent variable is the industry wealth effect of a successful merger. However, I also focus on a particular transformation of this variable, which, in a "constant-growth" valuation framework, can be interpreted as the expected product price change caused by a hypothetically anticompetitive merger. As discussed in Section III, the results are based on a sample of 196 large horizontal mergers, of which 80 were challenged by the government under Section 7 of the Clayton Act, as well as on a "control" group of 70 nonhorizontal (vertical or conglomerate) mergers. The estimation procedure is outlined in Section IV. I measure the industry wealth effect relative to the day of the first public announcement of the merger *proposal,* adjusting the abnormal returns by a maximum likelihood estimate of the probability the proposed merger will be prevented under antitrust laws. The regression results are discussed in Section V, and Section VI concludes the paper. The results of this paper provide no support for the market concentration doctrine.

II. Concentration and the Industry Wealth Effect of Anticompetitive Mergers

A. The Regression Model

Consider the public announcement of a horizontal merger that, because of the new (and higher-valued) production/investment policy to

7. Section 7 makes illegal any merger where the effect "may be substantially to lessen competition or to tend to create a monopoly." Since the Celler-Kefauver Amendment of 1950, the Department of Justice and the Federal Trade Commission (FTC) have challenged more than 500 mergers under this statue.

be implemented by the merged firm, is expected to change equilibrium product and factor prices in the merging firms' industry. Let dv denote the total change in the equity value of the representative rival of the merged firm implied by this change in industry equilibrium. If the merger announcement drives the market's estimate of the probability that the merger will be successfully consummated from zero to one, shareholders of this rival firm will earn announcement-induced abnormal returns of $AR \equiv dv/v$. Assume that (1) the merger is anticompetitive (e.g., by facilitating a collusive agreement in the industry or by eliminating competition from one of the rivals of a price leader); (2) the merger does not signal opportunities for efficiency gains available to the rival firm; and (3) the rival firm has no risky bonds in its capital structure. In this case, dv is the present value of the expected increase in monopoly rents following the merger-induced increase in the (quality-adjusted) product price or the decrease in prices of specialized factors of production. Under the market concentration doctrine, this increase in monopoly rents is a positive function of the increase in concentration caused by the horizontal merger.[8]

We examine this prediction by estimating the coefficients of a cross-sectional regression model of the following form:

$$AR_j = \alpha_0 + \alpha_1 CR_j + \alpha_2 dCR_j + e_j, \tag{1}$$

where CR_j represents the premerger level of concentration in the industry of the jth horizontal merger, dCR_j is the resulting change in concentration, and e_j is a mean zero random error term possibly containing the influence on AR_j of other relevant factors, which are assumed to be

8. Suppose, e.g., that the bidder firm is part of a coalition of producers in the industry who jointly determine the industry's price by unilaterally reducing their output, letting the rival "fringe" firms taking full advantage of the output restriction (i.e., a classical dominant-firm or price-leader framework). In the presence of entry barriers, it may pay the coalition to buy up one or more of the fringe firms in order to restrict industry output and raise the product price further. In the dominant-firm model it is easily shown that $(p - mc)/p = s/[e_d + (1 - s)e_f]$, where p is the product price, mc is the sum of the marginal costs of the firms in the coalition, s is the market share of the price leader, and e_d and e_f are the numerical values of the price elasticities of industry demand and fringe firm supply. Thus, ceteris paribus, the increase in the industry's price that results from the dominant coalition's purchase of one of the fringe firms is an increasing function of the market share of the target. Since *all* firms in the industry benefit from this product price increase, it also follows that AR will be increasing with the increase in concentration. Alternatively, one could use Stigler's (1964) oligopoly model, which is not based on ad hoc conjectures concerning firm behavior, to support the prediction in the text. In Stigler's framework, the success probability of a collusive agreement is inversely related to the number of firms in the industry. The fewer the number of firms, the lower the costs of policing the cartel agreement and the closer the cartel price will be to the single-firm monopoly level. Note that because my empirical analysis focuses on stock returns, and since stock prices permit any (price or nonprice) source of monopoly rents, the empirical results of this paper are not necessarily restricted to the product price prediction of any particular oligopoly model.

orthogonal to both CR_j and dCR_j. Under the market concentration doctrine we should find that $\alpha_2 > 0$. Furthermore, under the stronger proposition embedded in antimerger policy, which holds that a merger is more likely to have anticompetitive effects the larger the premerger value of CR_j, we should also find evidence of $\alpha_1 > 0$.

The form of equation (1) is similar in spirit to the regression models typically estimated in the "structure-conduct-performance" literature. However, while my dependent variable *AR* measures directly the capitalized value of the increased profits expected to follow from the change in the industry's level of concentration, the tradition has been to regress an accounting measure of the level of industry profits on the level of concentration. The traditional approach has been criticized on the grounds that accounting profits are a poor proxy for economic profits, and that any cross-sectional variation in the level of industry profits can reflect differences in risk. This criticism does not apply here, since AR_j is measured using market values and represents a *risk-adjusted* change in the level of industry rents. Furthermore, since equation (1) is in the form of changes in the central variables, α_2 can be meaningfully interpreted without specifying a structural model relating the level of industry profits to concentration.

On the other hand, my test suffers to some extent from the ambiguity common to traditional studies that correlate profits with concentration without controlling for the cross-sectional variation in costs of production. However, the analogy is not complete. Recall that *AR* represents the merger-induced abnormal returns to the representative rival of the bidder and target firms. Although it is possible that assumption 2 is wrong—and that the merger in fact signals opportunities for efficiency gains available to other industry members—there is no theory that predicts such signaling effects to be systematically correlated with concentration. In other words, evidence of a positive α_2 is not easily interpreted in terms of an efficiency argument. The analogous test based on the level of industry profits is weaker also in this respect since one can reasonably argue that relatively high industry profits are likely to be found in relatively concentrated industries because of the expansion of relatively cost-efficient, competitive firms (see, e.g., McGee 1971; Demsetz 1973; Peltzman 1977; Brozen 1982).

B. Abnormal Returns and the Hypothetical Product Price Change

If *AR* hypothetically reflects the present value of a merger-induced expected change in the product price, what is the magnitude of this hypothetical product price change? Let $\theta \equiv dp/p$ be a constant percentage change in the product price that would generate merger-induced abnormal returns of *AR* to the representative rival firm. Furthermore, suppose (assumption 4) the rival firm operates exclusively in the industry where the merger takes place; (5) θ is not expected to affect the

output rate of the rival;[9] and (6) the rival firm's total sales follow a geometric random walk with a constant growth rate g. It then follows that

$$dv = \theta S_0(1 - \tau) \int_0^{\infty} e^{(g - r)t} dt, \tag{2}$$

where τ is the firm's (constant) tax rate, r is the (constant) risk-adjusted discount rate, and S_0 is the current (premerger) level of sales.[10] Finally, if (assumption 7) the firm's net earnings available to stockholders, X, follow the same stochastic process as S, the integral in equation (2) equals the firm's current price-earnings ratio (v/X_0). Thus, we have

$$\theta = \frac{X_0}{S_0(1 - \tau)} AR. \tag{3}$$

In this constant-growth valuation framework, an estimate of θ given by (3) represents a hypothetical expected permanent change in the product price, consistent with merger-induced abnormal returns of AR to the representative rival of the merging firms.

Of course, θ is only hypothetically a summary of an underlying product price change: under the alternative hypothesis where AR is interpreted as the present value of expected future cost savings available to other industry members as a result of the information in the merger announcement, θ summarizes the percentage expected decrease in the rival firm's average costs of production. Thus, while evidence of $\alpha_2 \leq 0$ refutes the market concentration doctrine whether one uses AR or θ as the dependent variable in the regression of equation (1), evidence of $\alpha_2 > 0$ does not discriminate between the market power argument and alternative productive efficiency (cost-advantage) theories.

III. Data

The primary data base of this study consists of 266 merger proposals that were made during the period July 1963 through December 1981. Two hundred thirty-two of these merger proposals, and the associated rival firms, were identified using the data base compiled by Eckbo (1983), which covers the period 1963–78. Additionally, 34 proposals were collected covering the period 1979–81, of which 16 were taken

9. This assumption precludes a scenario in which the merger triggers a collusive agreement and the rival firm participates by restricting his output to support a product price increase. Thus, a dominant-firm scenario, in which the rival free rides on the price increase supported by the price leader (or a coalition of other firms), is more appropriate for the derivation of the hypothetical product price change.

10. Under the hypothesis that AR derives solely from θ, the change in the rival's expected net earnings in period t is given by $\theta E(S_t)(1 - r)$. Furthermore, under assumption 6, $E(S_t) = (1 + g)^t S_0$.

from the sample compiled by Eckbo and Wier (1985). Each merger in my sample satisfies the following five sampling criteria: (1) the merger proposal was announced in the *Wall Street Journal* and was approved by the shareholders of the merging firms; (2) the bidder or the target firm was listed on the New York (NYSE) or the American (ASE) Stock Exchange; (3) the major productive activity of the target firm was in a mining or manufacturing industry; (4) the premerger four-firm concentration ratio for this particular industry could be identified; and (5) there was at least one NYSE-or ASE-listed rival firm in this industry at the time of the proposal announcement.[11]

The sample consists of 196 horizontal and 70 nonhorizontal (vertical or conglomerate) mergers. Ninety-eight mergers (80 horizontal) were challenged by the government as violating Section 7 of the Clayton Act.[12] Counting the major four-digit SIC code of the target firm only, the total data base covers more than 136 different four-digit SIC industries, typically with one or two mergers per industry. On the two-digit SIC level, Oil and Gas Extraction (SIC 13) and Machinery, except Electrical (SIC 35) alone account for 62 (30 + 32) cases. Other heavily represented 2-digit industries are Food and Kindred Products (SIC 20; 27 cases), Paper and Allied Products (SIC 26; 18 cases), Chemicals and Allied Products (SIC 28; 18 cases), Primary Metal Industries (SIC 33; 21 cases), and Electric and Electronic Equipment (SIC 36; 19 cases).

I use the four-firm concentration ratio (CR_4) of the major four-digit

11. The rival firm selection procedure can be summarized as follows: *Step 1:* Identify the four major four-digit SIC codes of the target firm as listed in the Standard & Poor's *Registry of Corporations* for the year prior to the merger proposal. Following the *SIC Manual 1972* and the company-specific information on product lines given in the Standard & Poor's *Registry*, transform the four-digit SIC code into a five-digit industry code. If the merger was challenged, define the five-digit code by the product line listed in the data sources for Section 7 cases. *Step 2:* Produce a list of all firms on the University of Chicago Center for Research in Security Prices (CRSP) daily stock returns tape that are associated with the first four digits of the target's five-digit code(s). Finally, use information in the Standard & Poor's *Registry* to transform the four-digit codes of the firms on this CRSP list to five-digit codes, as under step 1. Label the firms with at least one five-digit industry code that overlaps the target firm "product market rivals."

12. Information on these challenged mergers is found in the American Bar Association's *Merger Case Digest* and the Commerce Clearing House's *Trade Regulation Reporter*, which also classify each case as horizontal, vertical, or conglomerate. The Section 7 complaint (and subsequent litigation) resulted in cancellation of 19 merger proposals, and divestiture was ordered in another 50 cases. Sixteen cases were dismissed by the court; in eight cases the defendant firm was convicted but no divestiture was ordered; and five cases were still pending as of December 1983. For additional information on the case outcome and relevant product markets for the 80 horizontal challenged cases, see Eckbo and Wier (1985, app.). The 168 unchallenged mergers are listed in the FTC's *Statistical Report on Mergers and Acquisitions* 1979 (the Large Merger series), which also provides a classification of the cases as horizontal, vertical, or conglomerate. For all the horizontal cases in our data base, the bidder and target firms have at least one overlapping four-digit SIC code according to the information listed in the Standard & Poor's *Registry of Corporations* for the year prior to the merger proposal.

SIC industry of the target firm to represent *CR* in equation (1), and the change in the industry's Herfindahl index (*dH*) as a measure of *dCR*.[13] Data on CR_4 were collected using published sources related to the challenged cases,[14] volume 1 of the 1977 *Census of Manufactures*, the FTC's *Concentration Levels and Trends in the Energy Sector of the U.S. Economy* for 1974, and Rogowsky (1982). For the unchallenged cases the relevant industry is the one represented by the first four digits of the five-digit industry code used in the rival firm selection procedure.[15] For the challenged cases the industry is the one listed in the sources for Section 7 cases, that is, the five-digit industry itself.[16]

The market shares of the bidder and target firms, which yield *dH*, were collected using the case-related publications cited above; Rogowsky (1982); and company statistics developed by Economic Information Systems, (EIS).[17] A measure of the adjustment factor needed to estimate the product price change θ in equation (3) was obtained by searching Standard & Poor's Annual Industrial Compustat tape for the rival firms in the sample. Using these Compustat-listed rivals I formed an equal-weighted industry average adjustment factor based on each company's ratio of "Net Sales" (a proxy for S_0) to "Income before Extraordinary Items and Discontinued Operations" (a proxy for X_0),

13. $CR_4 \equiv \Sigma_{i=1}^{4} s_i$, and $H \equiv \Sigma_{i=1}^{n} s_i^2$, where s_i is the market share of firm i, (in CR_4 the sum is over the four firms with the largest market shares), and n is the total number of firms in the industry. The change in the Herfindahl index caused by a merger between firms i and k is therefore given by $dH = 2s_i s_k$.

14. The sources include the *Merger Case Digest, Trade Regulation Reporter, Federal Trade Commission Decisions, Federal Reporter 2d Series*, and *Federal Supplements* (various issues).

15. Although step 1 of the rival firm selection procedure allows use of up to four industry codes per target firm (see n. 11, above), in 97% of the cases step 2 resulted in rival firms representing only *one* (usually the first) of the four industry codes, which is the one used for the purpose of collecting CR_4. The census lists CR_4 for most four-digit SIC manufacturing industries for the years 1963, 1966, 1967, 1970, 1972, and 1977. If the merger took place in a year not listed by the census, I extrapolated linearly the concentration ratio between the two closest census years on each side of the merger year.

16. If the challenged cases are truly anticompetitive, and since the four-firm concentration ratio tends to increase as one is narrowing the definition of the industry from a four-digit to a five-digit code, the selection of concentration data for the challenged cases may seem to introduce a bias in favor of the market concentration doctrine. However, as reported in an earlier version of the paper, regression results using Census-based concentration data for the challenged cases as well were indistinguishable from the results obtained with concentration data selected using the Section 7 information.

17. The FTC supplied me with EIS data for the years 1974, 1976, 1978, and 1979, covering exchange-listed firms that operate in the four-digit SIC industries represented by the first four digits of the five-digit code used in the rival firm selection procedure. To match a bidder or target firm in my sample with this EIS population, I used the year prior to the year of the merger proposal. The EIS data provide market shares at the level of the four-digit SIC industry of the firm. A more disaggregated analysis of economically relevant markets would almost certainly provide higher levels of market shares. Such a disaggregated analysis underlies the market shares in the publications related to the Section 7 cases, and, as it turned out, more than 80% of the market shares used in this paper come from the Section 7 publications.

and a flat tax rate of 48% for all companies. This procedure generated an industry adjustment factor for 186 of the 196 horizontal and 63 of the 70 nonhorizontal mergers in the data base.

Table 1 details the range, mean and median values for several of the central characteristics of the data base. As expected (given the law enforcement agencies' selection rule), the average market values of the bidder and target firms, the four-firm concentration ratio, and the market shares are higher in the challenged sample than in the unchallenged sample.[18] For example, in the sample of challenged horizontal mergers, the average level of concentration is 58%, while the average market share of the bidder (target) is 14.8% (12.8%). In the sample of unchallenged horizontal mergers, the corresponding numbers are 33% for the concentration level and 2.2% (2.8%) for the market shares of the bidder (target) firms. The change in the industry's Herfindahl index ranges from 0.02% to 24.18% (mean 3.33% over 64 observations) in the challenged sample compared to a range of 0.02%–0.33% (mean 0.15% over five observations) in the unchallenged sample.

IV. Estimation Procedure

A. *Abnormal Returns and the Probability of a Government Challenge*

In equation (1), *AR* is defined as the abnormal returns to the rival firm resulting from a merger announcement that drives the probability the merger will be successfully consummated from zero to one. To proxy this type of announcement, I use the first *Wall Street Journal* reference to the merger proposal. There is substantial evidence that the proposal announcement resolves a significant portion of the uncertainty concerning the consummation of the merger.[19] However, since a merger proposal can fail if stockholders of the merging firms reject the proposal,[20] or if the government prevents the merger from taking place, this announcement almost certainly does not drive the probability of

18. To capture the "size" of the merger, V_B and V_T in table 1 are computed as the maximum of the firm's book value and (if listed) equity value. The book values are from the Compustat tape and from Moody's *Industrial Manual*, while the equity values are from the CRSP tape, all for the year prior to the year of the merger proposal.

19. For example, the evidence in Dodd (1980) and Asquith (1983) indicates that the merger proposal announcement is a more significant news event than the subsequent announcement of the final outcome of the merger attempt. In Dodd's sample of successful merger proposals, target stockholders earn 24% abnormal return over the year leading up to the "effective merger date." However, more than 80% of this abnormal performance (19.6% abnormal return) occurs over the 10 trading days up to and including the proposal announcement date.

20. Generally, a merger proposal (which is an agreement between the managements of the bidder and target firms) must be approved by a two-thirds majority of the outstanding shares of the target firm.

TABLE 1 **The Number of Rival Firms per Merger (NR), Values (V) and Market Shares (S) of the Bidder and Target Firms, Four-Firm Concentration (CR_4) and Merger-induced Change in the Herfindahl Index (dH) across Subsamples of the Data Base, 1963–81**

	168 Unchallenged Mergers						98 Challenged Mergers					
	116 Horizontal			52 Nonhorizontal			80 Horizontal			18 Nonhorizontal		
Sample Characteristic	Range	Mean	Median	Range	Mean	Median	Range	Mean	Median	Range	Mean	Median
NR	1–55	15	6	1–51	10	5	1–29	5	3	1–18	5	3
V_B	9–8,557	847	325	7–9,896	1,017	450	10–23,555	1,544	452	101–4,520	1,577	845
V_T	10–6,480	151	27	10–1,130	80	22	1–9,470	471	101	10–1,412	313	100
CR_4	6–94	33	31	5–90	34	31	5–99	58	60	7–83	47	50
S_B*	.8–7.1	2.2	1.3		NA		.1–60.0	14.8	12.0		NA	
S_T†	.4–5.1	2.8	1.3	.6–20.0	10.0	5.1	.2–47.0	12.8	9.0	8.0–40.0	18.3	15.0
dH‡	.02–0.33	0.15	.05		NA		.02–24.18	3.33	1.00		NA	

NOTE.—NA = not available; subscript B and T denote bidder and target; CR_4, S, and dH are in percentages and refer to the major industry of the target firm; V is measured in millions of dollars and represents the *maximum* of the book value of the firm's assets and (if listed) the market value of its equity. The median, as reported, is such that 50% of the sample has a numerical value less than or equal to the numerical value listed.

* N = 37 observations (of the 116 unchallenged mergers) and 72 (of the 80 challenged mergers).

† N = nine observations (of the 116), six (of the 52), 64 (of the 80), and 13 (of the 18).

‡ N = five observations (of the 116) and 64 (of the 80).

merger consummation to one. In our particular sample, where each merger proposal (by selection) was approved by stockholders, the former source of uncertainty is unlikely to significantly affect the cross-sectional estimates of α_1 and α_2 in equation (1). On the other hand, there is a substantial cross-sectional variation in our sample with respect to various merger-specific characteristics that determine the likelihood the merger will be successfully challenged by the government. If the mergers with the largest anticompetitive effects are also the ones with the largest probability of being prevented by the government, then this antitrust "overhang" can bias our cross-sectional tests against the market concentration doctrine.

To illustrate, suppose the merger is completely unanticipated prior to the proposal announcement, and that the only source of uncertainty remaining after the proposal concerns the response by the law enforcement agencies. Furthermore, define γ as the *daily* abnormal return to the rival firm measured over w days relative to the proposal announcement, and suppose we use γw as a proxy for AR in the regression of equation (1). If π denotes the probability that the merger will be prevented by the government, we have that

$$AR = \frac{\gamma w}{1 - \pi}, \tag{4}$$

where $1 - \pi$ is the probability, conditional on the merger proposal, that the merger will eventually be consummated. If AR and π are positively correlated (i.e., if the firms with the largest AR are most likely to be successfully challenged), then the correlation between γw and dCR underestimates the correlation between AR and dCR, biasing the test against the market concentration doctrine. Since it is impossible to assess a priori the importance of this measurement problem, I report empirical results where both AR (estimated using eq. [4]) and γw are used to proxy the industry wealth effect of the merger.

B. Proposal-induced Abnormal Returns

For the jth merger in the sample we estimate the impact of the merger proposal announcement on the returns to an equal-weighted portfolio of the NR_j rival firms, using the following "market model" conditioned on the merger event:[21]

$$r_{jt} = \beta_{0j} + \beta_{1j} r_{mt} + \gamma_j d_t + \epsilon_{jt} \quad t = 1, \ldots, T, \tag{5}$$

where r_{jt} is the continuously compounded return to portfolio j over day t; r_{mt} is the return to the value-weighted market portfolio over day t; d_t

21. For a discussion of the market model, which represents the equilibrium-return-generating process when security returns follow a multivariate normal distribution, see Fama (1976).

is a dummy variable with a value of one if t is an event day (defined below) and zero otherwise; and ϵ_{jt} is assumed to be a mean-zero, normally distributed error term independent of r_{mt} and d_t. While the coefficient β_{1j} extracts marketwide movements from the return series, the abnormal return parameter γ_j directly isolates the component of the security's daily return that is due to the merger proposal itself.[22]

The abnormal return parameter γ_j is estimated for 2 different event periods around the day of the *Wall Street Journal* publication of the merger proposal (day 0). The estimation period (T) is 200 trading days prior to the day 0 through the tenth trading day after this date. The first event period extends from day -20 through day 10, covering any leakage of merger-related information prior to the announcement as well as any immediate postannouncement developments related to the merger. The dummy variable d_t therefore takes on a value of one during this 31-day interval. The resulting estimate of γ_j multiplied by 31 ($w = 31$) measures the total abnormal return to security j over that period. The second event period extends from day -3 through day 3 and is used to single out more precisely the effect of information released during the days immediately surrounding the proposal announcement. When estimating γ_j over the -3, 3 period, I delete days -20 through -4 and days 4 through 10 from the return series. Thus, β_{0j} and β_{1j} are estimated over the same "normal" period (extending from day -200 through day -21) for both event parameters.

Table 2 summarizes the sample averages of the OLS-estimates of $\gamma_j w$.[23] Target firms earn on average large and significant abnormal re-

22. Since my estimate of γ_j is based on an equal-weighted portfolio of rival firms, it measures the impact of the merger proposal on the average rival in the industry of the jth merger, fully incorporating any normal contemporaneous correlation of returns across firms in the same industry. Altering the portfolio weights generally changes both the interpretation and the estimated value of γ_j. For example, if the rival firms have their assets invested exclusively in the industry of the merging firms, then weights based on the relative market values of the rival firms would lead to an interpretation of γ_j as the abnormal return to a dollar invested in the industry. Alternatively, the set of portfolio weights that minimizes the residual variance in eq. (5) generates an estimate of γ_j identical to the estimate one would obtain by running a pooled time-series cross-sectional regression on the rival firms and constraining γ_j to be equal across rivals (Schipper and Thompson 1983). The weights of this particular portfolio are given by $(\mathbf{1}'\hat{\Sigma}^{-1}\mathbf{1})^{-1}\hat{\Sigma}^{-1}\mathbf{1}$, where $\hat{\Sigma}$ is the $NR_j \times NR_j$ contemporaneous covariance matrix of the OLS-residuals from the first-pass regression of eq. (5) across the NR_j rival firms associated with merger j, and $\mathbf{1}$ is a $NR_j \times 1$ vector of ones. Results based on these alternative portfolio weights were reported in an earlier version of this paper but have been omitted, as none of the basic conclusions were altered.

23. The Z-value in table 2 tests the hypothesis that the average abnormal return (AAR_J) is equal to zero. That is,

$$H_0: \quad AAR_J \equiv \frac{w}{J} \sum_{j=1}^{J} \gamma_j = 0,$$

TABLE 2 **Percentage Average Abnormal Returns (*AAR*) to Bidder, Target, and Rival Firms over 2 Event Periods relative to the Day of the *Wall Street Journal* Announcement of the Merger Proposal, 1963–81** (*Z*-Values in Parentheses)[a]

Sample	*N* Observations	*AAR*(−20, 10)	*AAR*(−3, 3)
I. 70 nonhorizontal mergers:			
Target firms	31	19.80***	14.88***
		(7.32)	(13.15)
Bidder firms	59	−.41	.37
		(−.10)	(.96)
Equal-weighted pairs of bidder and target firms	24	8.26***	6.96***
		(4.59)	(9.32)
Equal-weighted portfolios of rival firms	70[b]	−.87	−.07
		(−.20)	(1.02)
II. 196 horizontal mergers:			
Target firms	104	18.69***	11.19***
		(12.05)	(15.87)
Bidder firms	160	1.64*	.64*
		(1.87)	(1.77)
Equal-weighted pairs of bidder and target firms	81	10.18***	5.49***
		(9.84)	(12.11)
Equal-weighted portfolios of rival firms	196[c]	1.26***	.58***
		(2.50)	(2.86)
III. 80 horizontal challenged mergers:			
Target firms	43	21.70***	11.46***
		(10.14)	(12.55)
Bidder firms	67	3.20**	1.01*
		(2.02)	(1.85)
Equal-weighted pairs of bidder and target firms	37	13.43***	6.41***
		(9.31)	(10.71)
Equal-weighted portfolios of rival firms	80[d]	2.80***	.48**
		(2.77)	(2.01)

[a] *AAR* and *Z* are defined in n. 23.
[b] 70 portfolios containing 553 individual rival firms.
[c] 196 portfolios containing 2,062 individual rival firms.
[d] 80 portfolios containing 384 individual rival firms.
* Significantly different from zero at a 10% level of significance.
** Significantly different from zero at a 5% level of significance.
*** Significantly different from zero at a 1% level of significance.

turns whether the merger is horizontal or not. For example, for the period −20 through 10, the average $\gamma_j w$ is 19.80% for the nonhorizontal sample, 18.69% for the horizontal sample, and 21.70% for the horizon-

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where J is the number of mergers in the sample. Assuming the mergers are independent events, we have

$$Z \equiv \frac{1}{\sqrt{J}} \sum_{j=1}^{J} \frac{\gamma j}{\sigma_{\gamma j}} \sim N(0, 1),$$

where $\sigma_{\gamma j}$ is the standard error of γ_j. Under H_0, and replacing the true values of γ_j and $\sigma_{\gamma j}$ with their OLS estimates, this Z-statistic is approximately standard normal for large J.

tal challenged cases. The corresponding Z-values are large. Perhaps more surprising is the significant abnormal returns to *bidder* firms in horizontal mergers.[24] Whereas bidder firms in nonhorizontal mergers earn virtually zero average abnormal returns, in the two other groups the average abnormal returns is 1.64% and 3.20% relative to the period −20 through 10. In no group do the bidder or target firms on average lose from the investment, confirming the assumption of managerial value maximization underlying the market power (as well as the alternative cost-advantage) theory. Note also that the strong performance of the target firms is consistent with my assumption that the merger proposal is a significant news event. This is an important verification, particularly since no explicit *Wall Street Journal* news announcement exists for the rival firms.

The rivals of the 70 nonhorizontal mergers in the data base earn on average zero abnormal performance over either one of the 2 event periods. This contrasts with the significantly positive abnormal performance by the average portfolio of rivals associated with the 196 horizontal mergers. Over the period −20 through 10, these rival firms (the total number of rivals across the 196 portfolios is 2,062) earn abnormal returns of 1.26% ($Z = 2.50$) while the corresponding performance is 2.80% ($Z = 2.77$) for the subgroup of 80 challenged horizontal mergers. For all the groups, the average abnormal return to the rival firm appears to be slightly lower than the corresponding abnormal return to the average bidder firm.

As discussed in Section II, the positive average industry wealth effect cannot be used to discriminate between the market concentration doctrine and the competing cost-advantage theory of mergers. Notice, however, that the same evidence is somewhat inconsistent with the notion that the average merger was intended to implement substantial scale economies, as this would have caused expectations of a substantial fall in the product price (or rise in factor prices) and capitalized losses to rival firms.[25]

24. As reviewed by Jensen and Ruback (1983), the bulk of the evidence in the merger literature indicates that stockholders of the acquiring firm earn on average positive but statistically insignificant abnormal returns. The results for the bidder firm in table 2 is consistent with the proposition that a horizontal merger is on average more profitable for the bidder than is a nonhorizontal merger.

25. To complement the Z-test, which can fail to reject the hypothesis of zero abnormal returns if some mergers have a positive impact while others have a negative and offsetting impact, I also performed a test of the null hypothesis H_0': $\gamma_{ij} = 0$, $i = 1, \ldots, NR_j$. H_0' states that the event parameters are jointly equal to zero across each of the NR_j *individual* rivals of the jth merger and represents a set of NR_j linear constraints on the NR_j event parameters γ_{ij}. Under the assumption of normality, recognizing the structure of eq. (5), the following statistic has a χ^2 distribution with NR_j degrees of freedom:

$$(X'X)_{33}^{-1}\,\boldsymbol{\gamma}_j'\,\boldsymbol{\Sigma}^{-1}\,\boldsymbol{\gamma}_j \sim \chi^2(NR_j).$$

$X \equiv [\mathbf{1}\ \mathbf{r}_m\ \mathbf{d}]$ is the $T \times 3$ matrix of the explanatory variables in eq. (5), $(X'X)_{33}^{-1}$ is the (3, 3) element of the 3×3 matrix $(X'X)^{-1}$, $\boldsymbol{\gamma}_j$ is the $NR_j \times 1$ vector of the event parameters

C. *The Probability of a Successful Government Challenge*

Let Δ_j be a vector of firm- and industry-specific characteristics associated with the jth merger at the time of the merger proposal, and let y_j be a binary variable such that

$$y_j = \begin{cases} 1 & \text{if } \delta'\Delta_j + u_j > 0 \\ 0 & \text{otherwise,} \end{cases} \tag{6}$$

where δ is a parameter vector and u_j is a random variable. Defining $y_j = 1$ when the merger is successfully challenged by the government, and assuming u_j has the logistic density function, we have that

$$\pi_j = \text{prob}(y_j = 1) = \text{prob}(u_j > -\delta'\Delta_j) = \frac{1}{1 + e^{-\delta'\Delta_j}}. \tag{7}$$

Equivalently, since

$$\frac{\pi_j}{1 - \pi_j} = e^{\delta'\Delta_j}, \tag{8}$$

this logit model implies that the logarithm of the odds in favor of a successful government challenge is a linear function of the merger characteristics Δ_j.

We define a government challenge as "successful" if it leads to divestiture or to cancellation of the merger proposal and estimate δ using maximum likelihood procedures.[26] Since, at the time of the merger proposal announcement, there is substantial uncertainty regarding which mergers will be selected for prosecution by the law enforcement agencies, we include both unchallenged and challenged mergers in the estimation of δ.[27] We include the asset sizes of the bidder and target

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associated with the NR_j rival firms of merger j, and Σ is defined as in n. 22. This test statistic, which has a limiting χ^2 distribution when γ_j and Σ are replaced with their OLS estimates, is a straightforward simplification of the general quadratic form given in Theil (1971, p. 313). I estimated the value of this χ^2 statistic for each of the J mergers in the sample, and made statistical inferences on the basis of the percentage of these J values that rejected H_0'. The results, which were reported in an earlier draft, were generally consistent with and reinforced the results of the Z-test in table 2. Furthermore, although the estimation methodology used here differs from the one used in Eckbo (1983), the magnitudes of the abnormal returns are very similar. Thus the estimates in table 2 appears to be robust also with respect to the estimation methodology. In addition, Eckbo and Wier (1985) fail to reject the proposition that the abnormal performance of the rivals of the challenged mergers reported in table 2 is different from the corresponding performance of a set of rival firms selected using case-related publications and internal government memoranda prepared for the Section 7 cases. In this sense, the estimates in table 2 are robust with respect to the procedure for selecting rival firms as well.

26. With this definition, 59 of the 80 horizontal and 10 of the 18 nonhorizontal challenged mergers were successfully prevented by the government. My estimation procedure uses the Newton-Raphson method, which, in the case of the logit distribution, converges to the maximum of the likelihood function regardless of the setting of the initial parameter values (see, e.g., Maddala 1983, p. 25).

27. If the enforcement agencies adhered strictly to their own merger guidelines, π_j could a priori be set equal to zero for the mergers with values of CR_4 falling below the

firms, the number of identified rival firms, and the four-firm concentration ratio in Δ_j, all of which are available for the unchallenged as well as the challenged cases in the data base (see table 1). The regressions are run with a constant term, constraining the sum of the estimated probabilities to be equal to the number of observations in the sample for which $y_j = 1$. Since the law enforcement agencies treat horizontal and nonhorizontal merger differently, the parameter vector $\boldsymbol{\delta}$ is estimated for each of the two merger categories separately.

The results of the logit estimation, which are listed in table 3, indicate that the number of rival firms and the four-firm concentration ratio are both useful in explaining the probability of a successful government challenge of a horizontal merger. The first regression in table 3 yields a maximum likelihood estimate of -0.054 (assymptotic t-value of -2.13) for the coefficient multiplying NR, while the estimated value of the coefficient multiplying CR_4 is 0.036 (t-value of 4.62). Thus, as might be expected, the estimate of π_j increases with CR_4 and decreases with NR. While the asset sizes of the merging firms have virtually zero explanatory power, the likelihood ratio test strongly rejects the hypothesis that all the slope coefficients are equal to zero.[28] The second regression in table 3 indicates that the same independent variables have zero power to explain the probability of a successful government challenge of a nonhorizontal merger. Consequently, the proper assumption is $\pi_j = 0$ for this category of mergers.[29] The value of $\hat{\pi}_j$ for the unchallenged horizontal cases, implied by the parameter estimates in table 3, ranges from 0.01 to 0.78, with a mean of 0.20 and standard deviation

guidelines. However, Rogowsky (1982) finds that over 20% of the mergers challenged under the 1968 guidelines actually fell *below* the guidelines, and one-third of these were found in violation of Section 7 of the Clayton Act. Furthermore, Eckbo (1983) and Eckbo and Wier (1985) find that the merging firms earn significantly negative abnormal returns at the time the challenge is announced, which strongly indicates that the challenge itself was not fully anticipated.

28. The likelihood ratio reported in table 3 is given by $-2\log_e\lambda$, where $\lambda \equiv \max L(\boldsymbol{\delta}, H_0)/\max L(\boldsymbol{\delta})$, where L is the likelihood function and H_0 is the hypothesis that the slope coefficients in $\boldsymbol{\delta}$ are jointly equal to zero. This ratio approximates a χ^2 distribution with degrees of freedom equal to the number of parameters in $\boldsymbol{\delta}$ for which H_0 specifies given values (Theil 1971, p. 396), in this case 4. Reestimating $\boldsymbol{\delta}$ for the challenged mergers only, and including the change in the industry's Herfindahl index as an explanatory variable, I found that the likelihood ratio test rejected the proposition that the slope coefficients in $\boldsymbol{\delta}$ were different from zero. Thus most of the explanatory power of the first regression in table 3 apparently comes from predicting the challenge itself rather than the outcome of the challenge.

29. Since some nonhorizontal mergers were in fact challenged, this assumption is obviously wrong. However, both the value of and cross-sectional variation in the true π_j for this category of mergers are likely to be very small. The proportion of the nonhorizontal mergers in the sample that were successfully challenged (10 in 70) is substantially larger than the corresponding proportion based on the total population of all nonhorizontal mergers. Thus, this sample proportion would substantially overestimate the unconditional, true value of π_j.

TABLE 3 **Maximum Likelihood Estimates of the Coefficients of the Logit Model Generating the Probability of a Successful Government Challenge of a Proposed Merger** (Assymptotic *t*-Values in Parentheses)

Sample of Mergers	*N* Observations		Independent Variables					Likelihood Ratio
	$y = 1$	$y = 0$	Constant	V_B	V_T	NR	CR_4	
All horizontal	59	137	−2.175	−.001	.001	−.054	.036	43.78
			(−4.49)	(.43)	(.25)	(−2.13)	(4.62)	(4*df*)
All non-horizontal	10	60	−2.232	.001	.002	−.015	.001	4.16
			(−2.53)	(.76)	(1.63)	(−.38)	(.07)	(4*df*)

NOTE.—The dependent variable, y, takes on a value of one if the merger was ordered to divest or if the merger proposal was canceled as a result of the antitrust complaint and zero otherwise. V_B = asset size of bidder firm, V_T = asset size of target firm, NR = number of identified rival firms in the major industry of the target, CR_4 = four-firm concentration ratio in the major industry of the target. The likelihood ratio is described in n. 28.

0.14. For the challenged horizontal cases, the range is 0.06 to 0.81, with a mean of 0.45 and standard deviation 0.21.

V. Abnormal Returns and Concentration

A. Regressions Based on the Premerger Level of Concentration

Table 4 and table 5 list the OLS regression results based on "low," "medium," and "high" levels of four-firm concentration as independent variables.[30] The medium level of concentration is defined by $30\% \leq CR_4 < 60\%$. All variables are expressed in percentages. In table 4 the basic dependent variable is the abnormal returns to equal-weighted pairs of bidder and target firms, while table 5 gives the corresponding results for the equal-weighted portfolios of rival firms. I focus on the bidder and target firms jointly since I am interested in the behavior of the total gains from the merger as opposed to the division of the gains between the merging parties.

30. Statistical inferences based on the OLS estimates of the parameters in the cross-sectional regressions are valid under the assumptions that the error terms in the regressions are independent of the explanatory variables, homoscedastic, and not cross-correlated. The assumption of zero contemporaneous correlation of the error terms is justified, since the merger proposals in the data base take place at different points in calendar time, and since portfolios are formed to estimate the abnormal performance of the bidder, target, and rival firms in a given merger. To gauge the importance of the assumption of homoscedasticity, all regressions were repeated using weighted least squares (WLS), where the estimated standard error of the abnormal return was used as the weight. This linear transformation of the data preserves the full column rank of the matrix of explanatory variables and therefore the unbiasedness of the parameter estimates. Furthermore, under the usual WLS assumptions, this procedure also produces unbiased estimates of the standard errors of the parameters. As it turned out, the results based on the WLS procedure were indistinguishable from the OLS results, and only the latter are reported in the paper.

TABLE 4 **OLS Estimates of the Coefficients in the Cross-sectional Regressions of the Merger-induced Abnormal Returns to Pairs of Bidder and Target Firms on the Level of Concentration, 1963–81** (*t*-Values in Parentheses)

Sample	Dependent Variable	*N* Observations	Event Period[a]	Independent Variables[b]				R^2	*F*
				Constant	Low CR_4	Medium CR_4	High CR_4		
I. Equal-weighted pairs of bidder and target firms in horizontal mergers	γw	81	−20, 10	6.147	.018	.100	.098	.06	1.73
				(1.13)	(.07)	(.76)	(1.30)		
			−3, 3	3.654	.011	.051	.039	.02	.54
				(.92)	(.06)	(.53)	(.71)		
	AR	81	−20, 10	−3.537	.465	.409*	.571***	.371	15.15
				(−.38)	(1.05)	(1.83)	(4.40)		
			−3, 3	2.715	.079	.120	.191**	.129	3.82
				(.40)	(.25)	(.73)	(2.00)		
	θ	79	−20, 10	1.570	−.009	.011	.039**	.196	6.08
				(1.16)	(−.13)	(.34)	(2.11)		
			−3, 3	.979	−.013	.002	.018	.084	2.30
				(.86)	(−.24)	(.08)	(1.16)		
II. Equal-weighted pairs of bidder and target firms in nonhorizontal mergers	AR	24	−20, 10	10.057	.019*	−.045	−.062	.043	.30
				(.86)	(.03)	(−.18)	(−.40)		
			−3, 3	8.765	.642	−.289	−.044	.013	.09
				(0.99)	(.16)	(−.16)	(−.33)		
	θ	21	−20, 10	−2.905	.251*	.078	.044	.013	.09
				(−.97)	(1.80)	(1.22)	(1.00)		
			−3, 3	−1.463	.141	.046	.026	.178	1.22
				(−.71)	(1.59)	(1.14)	(.95)		

NOTE.—γw is the abnormal return measured over ω days relative to the day of the *Wall Street Journal* announcement of the merger proposal (see eq. [5]). $AR = \gamma w/(1 - \pi)$, where π is the probability that the proposed merger will be successfully challenged by the government. π is estimated using maximum likelihood procedures, which yielded $\pi = 0$ for nonhorizontal mergers. $\theta = AR[X/S(1 - \tau)]$, where X = net earnings, S = sales, τ = corporate tax rate.

[a] Day 0 is the day of the *Wall Street Journal* announcement of the merger proposal.

[b] The three variables low CR_4, medium CR_4, and high CR_4 are obtained by multiplying the four-firm concentration ratio by a dummy variable that takes on a value of one if $CR_4 < 30\%$, $30\% \leq CR_4 < 60\%$, $CR_4 \geq 60\%$, respectively, and zero otherwise.

* Significantly different from zero at a 10% level of significance.

** Significantly different from zero at a 5% level of significance.

*** Significantly different from zero at a 1% level of significance.

TABLE 5 **OLS Estimates of the Coefficients in the Cross-sectional Regressions of the Merger-induced Abnormal Returns to Rival Firms on the Level of Concentration, 1963–81** (*t*-Values in Parentheses)

Sample	Dependent Variable	*N* Observations	Event Period	Independent Variables				R^2	*F*
				Constant	Low CR_4	Medium CR_4	High CR_4		
I. Equal-weighted portfolios of rivals of horizontal mergers	γw	196	−20, 10	1.542	−.070	.001	.009	.011	.70
				(.60)	(−.56)	(.01)	(.26)		
			−3, 3	2.783**	−.085	−.047*	−.042***	.036	2.37
				(2.29)	(−1.44)	(−1.65)	(−2.45)		
	AR	196	−20, 10	−.993	.036	.074	.115	.030	1.95
				(−.18)	(.14)	(.57)	(1.48)		
			−3, 3	3.505*	−.112	−.056	−.067**	.037	2.44
				(1.67)	(−1.08)	(−1.12)	(−2.22)		
	θ	186	−20, 10	.279	−.004	−.001	.013	.036	2.25
				(.38)	(−.12)	(−.06)	(1.22)		
			−3, 3	.333	−.004	−.007	−.002	.011	.67
				(.94)	(−.25)	(−.89)	(−.45)		
II. Equal-weighted portfolios of rivals of nonhorizontal mergers	AR	70	−20, 10	1.604	−.138	−.072	−.002	.019	.44
				(.36)	(−.67)	(−.67)	(−.031)		
			−3, 3	−.265	−.014	.007	.018	.022	.50
				(−.15)	(−.16)	(.16)	(.62)		
	θ	63	−20, 10	.182	−.015	−.008	−.003	.005	.09
				(.26)	(−.46)	(−.44)	(−.22)		
			−3, 3	.097	−.003	−.001	−.002	.005	.10
				(.44)	(−.30)	(−.19)	(−.48)		

NOTE.—See table 4 nn. for definitions of variables and significance levels.

Adjusted for the probability that the merger will be successfully challenged by the government, the 81 equal-weighted pairs of bidder and target firms involved in horizontal mergers earn significantly higher abnormal returns in highly concentrated industries than in industries with relatively low premerger level of concentration. In table 4, the regression of AR over the -20 through 10 event period produces a coefficient multiplying the variable low CR_4 of .465 ($t = 1.05$), while the coefficient of the high-concentration variable is .5711 ($t = 4.40$). The regression R^2 is .371 with a significant F-statistic of 15.15. The interpretation of the fourth coefficient of this particular regression is that, ceteris paribus, an increase by one percentage point in the level of concentration in an already highly concentrated industry on average results in an increase in the abnormal returns to the merging firms by 0.571%. A similar (though somewhat weaker) relationship emerges from the regression of the product price change estimator θ on concentration, where the coefficient multiplying high CR_4 is .039 ($t = 2.11$, $R^2 = .196$, $F = 6.08$). Furthermore, a comparison with the results for the bidder and target firms involved in nonhorizontal mergers reveals that the positive correlation with concentration is present only for horizontal mergers. There is no evidence in table 4 that the gains to firms involved in nonhorizontal mergers are significantly positively related to medium or high levels of concentration.

These results are consistent with the antitrust doctrine, where the abnormal returns to the merging firms are viewed as the capitalized value of increased monopoly rents. On the other hand, the same results are also consistent with the proposition that horizontal mergers generate a cost advantage for the merging firms that is larger the higher the premerger level of concentration. Of the two competing hypotheses, however, only the former implies that also the abnormal returns to the *rivals* should increase with concentration. The evidence in table 5 clearly rejects that prediction: rival firms of horizontal mergers earn significantly *lower* abnormal returns over the -3, 3 period the higher the level of concentration, with the largest reductions in AR (as well as in γw) for the highest concentration levels. Increasing the level of the high CR_4 variable by 1% implies a 0.042% ($t = -2.4$) decrease in γw and a 0.067% ($t = -2.22$) decrease in AR, measured over the -3, 3 event period. The product price change estimator θ shows no significant relationship to the premerger level of concentration, and neither do any of the three dependent variables when measured over the longer event period -20, 10. Finally, the abnormal performance of rivals of *non*horizontal mergers is uncorrelated with the level of concentration, regardless of the event period or the transformation of the abnormal returns. In other words, increasing the level of concentration has a greater negative effect on the rivals of horizontal mergers than on

the rivals of nonhorizontal mergers, a result which is not predicted by the antitrust doctrine.

B. *Regressions Based on the Merger-induced Change in Concentration*

Table 6 contains the results of regressing abnormal returns on the merger-induced change in the Herfindahl index (dH) of the major industry of the target firm, controlling for the premerger level of CR_4. Since a nonhorizontal merger does not change the level of concentration, this table is based on the horizontal mergers only. Furthermore, with data on dH available for only five of the 116 horizontal unchallenged mergers in our sample (see table 1), I restrict the analysis to challenged mergers. If anything, this restriction possibly introduces a sample bias in favor of the market concentration doctrine since the challenged mergers were all accused by the government of "monopolizing" product markets.

Despite this potential bias, the regression results in table 6 do not support the market concentration doctrine. The coefficient multiplying the change in the Herfindahl index is uniformly negative across all the reported regressions. For example, increasing dH by 1% implies a reduction of 0.570% in the abnormal returns (γw) to the merging firms over the -20, 10 period. With a t-value of -1.79, this coefficient is statistically different from zero at the 10% level of significance. Similarly, the abnormal returns adjusted for the probability of a successful government challenge (AR) are reduced by 1.299% ($t = -1.79$) as a result of a 1% increase in dH.

The results for the rival firms are similar, although the significance levels are somewhat lower. The proposal-induced abnormal return (γw) to the average portfolio of rival firms declines by 0.420% as the merger-induced change in the industry's Herfindahl index increases by 1%. This coefficient is statistically different from zero at a 10% level of significance, with a t-value of -1.70. For both AR and θ, the coefficient multiplying dH appears with a negative sign, although the t-values are too low to conclude that they are statistically different from zero.[31] Thus, on balance, the correlation between the merger-induced change in concentration and the abnormal performance of both the merging firms and their rivals appears to be negative. Adjusting the abnormal performance by a maximum likelihood estimate of the proba-

31. The results in table 6 based on the product-price-change estimation θ are generally consistent, although not closely comparable, with results reported by Stillman (1980). Using monthly security returns and 15 horizontal mergers challenged between 1954 and 1970, Stillman finds no correlation between what he calls the "effective price change" (measured over 7 and 13 months preceding and including the month of the merger announcement) and the merger-induced change in concentration.

TABLE 6 **OLS Estimates of the Coefficients in the Cross-sectional Regressions of the Merger-induced Abnormal Returns on the Level of and Change in Concentration: Sample of Horizontal Challenged Mergers, 1963–81** (t-Values in Parentheses)

Sample	Dependent Variable	N Observations	Event Period	Independent Variables			R^2	F
				Constant	CR_4	dH		
I. Equal-weighted pairs of bidder and target firms	γw	26	−20, 10	12.127***	.059	−.570*	.136	1.82
				(2.64)	(.82)	(−1.79)		
			−3, 3	4.048	.057	−.190	.056	.68
				(1.09)	(.97)	(−1.74)		
	AR	26	−20, 10	−5.151	.641***	−1.299*	.407	7.89
				(−.47)	(3.72)	(−1.79)		
			−3, 3	−2.275	.289*	−.176	.136	1.82
				(−.24)	(1.91)	(−.27)		
	θ	25	−20, 10	1.584	.048**	−.242	.170	2.25
				(1.03)	(1.96)	(−1.11)		
			−3, 3	.019	.038	−.059	.094	1.13
				(.01)	(1.51)	(−.27)		
II. Equal-weighted portfolios of rival firms	γw	64	−20, 10	3.715	−.004	−.420*	.045	1.43
				(1.13)	(−.67)	(−1.70)		
			−3, 3	4.279***	−.060***	−.075	.115	3.97
				(2.92)	(−2.58)	(−.66)		
	AR	64	−20, 10	1.927	.088	−.849	.037	1.17
				(.25)	(.73)	(−1.45)		
			−3, 3	7.584***	−.119***	−.207	.102	3.45
				(2.32)	(−2.31)	(−.82)		
	θ	58	−20, 10	−.218	.020	−.087	.025	.70
				(−.18)	(1.04)	(−.62)		
			−3, 3	.163	.001	−.038	.006	.16
				(.26)	(.14)	(−.55)		

NOTE.—CR_4 = four-firm concentration ratio, dH = merger-induced change in the Herfindahl index of concentration. Both CR_4 and dH refer to the major industry of the target firm. See table 4 nn. for definitions of variables and significance levels.

bility of a successful government challenge reduces the significance of the regression but does not entirely remove the negative correlation. Furthermore, transforming the abnormal returns into a hypothetical product price change estimator essentially drives the correlation to zero.[32]

Since the evidence fails to support the market concentration doctrine, it follows that γw (and therefore AR) is unlikely to represent monopoly rents, even for high levels of both CR_4 and dH. Consequently, it is also inappropriate to refer to θ as a product price change estimator. One consistent explanation for the positive average abnormal returns to the merging firms and their rivals is that the merger announcement signals the presence of opportunities for realizing productivity increases available to the industry at large.[33] In sum, the evidence in this paper is consistent with the proposition that the merger-induced abnormal returns represent the capitalized value of future cost savings from a more efficient allocation of corporate resources.

VI. Conclusions

The classical market concentration doctrine predicts that a horizontal merger is more likely to have collusive, anticompetitive effects the greater the merger-induced change in industry concentration. Since successful collusion increases the industry's product price (or decreases factor prices), it also follows from this doctrine that the expected benefits to the product market rivals of the merging firms are increasing with the concentration change. I test this proposition by studying the cross-sectional relationship between the industry wealth effect of merger proposal announcements and industry characteristics such as the level of the four-firm concentration in the major industry of the target firm and the merger-induced change in the industry's Herfindahl index.

Although the announcement of a horizontal merger is typically associated with positive abnormal returns to rivals of the target firm, I reject the proposition that this industry wealth effect is positively correlated with the change in concentration implied by the horizontal

32. The significance of the coefficients multiplying the level of concentration (CR_4) in table 6 reflects the results discussed earlier in tables 4 and 5. In other words, the significance of this variable is predominantly associated with values in the upper range CR_4.

33. As discussed by Eckbo (1983), a merger proposal announcement may disseminate information detailing the new production or investment strategy to be implemented by the merged firm. This information can increase the market value of the rivals even if implementation of the new strategy requires merger with some other firm. In the latter case, the abnormal returns to rival firms can be interpreted as capitalized gains from future merger activity, a proposition that, at least in principle, is testable.

merger. This conclusion is based on a sample of 80 horizontal mergers which the antitrust law enforcement agencies alleged would "monopolize" product markets. Furthermore, including a larger sample of horizontal unchallenged mergers, and a comparison group of nonhorizontal mergers, I also reject the proposition that the industry wealth effect is positively correlated with the premerger level of concentration.

Since a significant part of the data on market shares and concentration were collected using sources related to the challenged cases, the rejection of the market concentration doctrine is not likely to be explained by an excessive level of aggregation in my measures of these industry characteristics: the enforcement agencies frequently use a definition of the relevant "endangered" product market that contains only a few closely related products. Second, the lack of evidence in favor of the doctrine hardly reflects the possibility that the relevant product markets were, in spite of the government challenge, already monopolized prior to the merger: the mergers in my data base cover a wide range of concentration levels, and there is some evidence of a significantly *negative* correlation between the industry wealth effect of the merger announcement and the merger-induced change in concentration. Third, the results are not explained by a general antitrust "overhang," that is, a large probability that anticompetitive mergers will be successfully challenged by the government: I adjust the merger proposal-induced industry wealth effect by a maximum likelihood estimate of the probability of a successful antitrust complaint. Finally, I also reject the market concentration doctrine after transforming the industry wealth effect of the merger into a perpetuity that, in a constant-growth valuation framework, can be interpreted as a permanent increase in the industry's product price expected to follow from a hypothetically anticompetitive merger.

On the other hand, the market reaction to the merger proposal announcement is consistent with the proposition that the horizontal mergers in the sample were expected to provide a more efficient allocation of corporate resources *regardless* of the level and change in concentration. The positive industry wealth effect may reflect information concerning opportunities for productivity increases that are available to the nonmerging rival firms. Of course, while this information argument is consistent with the evidence, it is, at this stage, just conjecture. The information signaling argument clearly needs further development; it is a topic that is currently being researched. The conclusive finding of this paper is that my evidence systematically rejects the antitrust doctrine, even for values of the level and merger-induced change in industry concentration which, over the past 3 decades, have been considered "critical" in determining the probability that a horizontal merger will have anticompetitive effects.

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