Mergers and the Value of Antitrust Deterrence
Author(s): B. Espen Eckbo
Source: *The Journal of Finance*, Vol. 47, No. 3, Papers and Proceedings of the Fifty-Second Annual Meeting of the American Finance Association, New Orleans, Louisiana January 3-5, 1992 (Jul., 1992), pp. 1005-1029
Published by: Blackwell Publishing for the American Finance Association
Stable URL: http://www.jstor.org/stable/2328975
Accessed: 25/11/2009 11:48

THE JOURNAL OF FINANCE • VOL. XLVII, NO. 3 • JULY 1992

Mergers and the Value of Antitrust Deterrence

B. ESPEN ECKBO*

ABSTRACT

While the U.S. has pursued a vigorous antitrust policy towards horizontal mergers over the past four decades, mergers in Canada have until recently been permitted to take place in a virtually unrestricted antitrust environment. The absence of an antitrust overhang in Canada presents an interesting opportunity to test the conjecture that the rigid market share and concentration criteria of the U.S. policy effectively deters a significant number of potentially collusive mergers. The effective deterrence hypothesis implies that the probability of a horizontal merger being anticompetitive is higher in Canada than in the U.S. However, parameters in cross-sectional regressions reject the market power hypothesis on samples of both U.S. and Canadian mergers. Judging from the Canadian evidence, there simply isn't much to deter.

THE UNITED STATES, CANADA, and Europe have over the past four decades implemented substantially different antitrust policies towards horizontal mergers. Since 1950, U.S. antitrust authorities have placed strict regulatory limits on permissible levels of industry concentration and market shares of merging firms, effectively constraining firm size ex ante.[1] To support this policy, the government has filed more than 500 antitrust complaints against horizontal mergers, of which more than 80% resulted in forced divestiture or

* From the Faculty of Commerce, University of British Columbia, Vancouver. This research has benefited from the comments and suggestions of Lisa Meulbroek, Aileen Thompson, Clas Wihlborg and, especially, Vojislav Maksimovic and Joseph Williams. I am also grateful for comments made by the participants of the conference on "Financial Regulation and Monetary Arrangements after 1992" organized by the Centre for Financial Economics at the Gothenburgh School of Economics (Gothenburgh 1990), the conference on "Corporate Control and Corporate Restructuring" organized by the Centre for Economic Policy Research at the Stockholm School of Economics (Stockholm, 1991), and the 1991 meetings of the American Economic Association and the American Finance Association. Partial financial support from the Social Sciences and Humanities Research Council of Canada is also gratefully acknowledged.

[1] For example, according to the 1968 *Merger Guidelines* of the Department of Justice, a merger between two firms each having 4% of the sales in a market with a four-firm concentration ratio of 75% or more was likely to be deemed to "substantially lessen competition" and therefore violate Section 7 of the Clayton Act, the principal federal antitrust law regulating corporate mergers. More recent guidelines issued in 1982 and 1984 shift the definition of concentration to one based on the Herfindahl index and are somewhat less restrictive, but retain the basic focus on market structure.

cancellation of the proposed merger transaction.[2] There can be no doubt that this enforcement policy significantly deters horizontal merger activity.[3] At the other extreme, until 1985 mergers in Canada were allowed to take place in an effectively unconstrained legal environment.[4] The policy in most European countries has been the intermediate solution of encouraging firms to adopt the unconstrained optimal size ex ante, with antitrust authorities monitoring an industry's product price ex post.[5] Naturally, these diverging antitrust philosophies have substantially different implications for the amount of resources needed for effective enforcement, as well as for the social cost of the implicit constraints on productive activity.

The relatively vigorous ex ante enforcement policy of the U.S. implicitly assumes that the theoretical models used to separate anticompetitive from socially efficient takeovers are empirically accurate, and that the policy's deterrent effect on merger activity is welfare enhancing. As to the first assumption, although the U.S. antitrust authorities' reliance on structural standards for selection of merger cases has some intuitive appeal (the cost of enforcing a tacit collusive agreement is inversely related to the number of producers in the industry, Stigler (1964)), there is currently no empirical study which suggests that industry concentration is a *reliable* index of an industry's market power. Traditional empirical studies of the "structure-conduct-performance" paradigm also has minimal power to discriminate between this market concentration doctrine and the alternative hypothesis that competitive pressures cause firms to specialize their productive skills and resources, which in turn leads to increased concentration whenever the resulting cost-advantage increases optimal firm size (Peltzman (1977).[6]

[2] The government's success rate is not surprising; under Section 7 of the Clayton Act, a *potential* threat to competition constitute a civil offense, and anticipated economic efficiencies is not a permissible defense against the illegality of a merger that may "substantially lessen competition".

[3] Stigler (1966), Ellert (1976), Wier (1983), and Eckbo and Wier (1985) present evidence of significant costs imposed on challenged firms.

[4] Under the Canadian Combines Investigation Act of 1910, it is a criminal offense to form business mergers, monopolies or other agreements to limit competition, requiring the prosecution to prove the alleged offense beyond reasonable doubt. This has proved difficult; there has never been a convictions on a monopoly charge, and only one conviction on the merger section of the Combines Act. With the introduction of the Competition Act in 1985, undertaking anticompetitive mergers constitute a civil offense also in Canada, and a tribunal now screens merger proposals for such effects prior to approving the merger transaction (see the 1991 *Merger Enforcement Guidelines* of the Consumer and Corporate Affairs Canada for details).

[5] An exception is Germany, where prenotification and approval of potentially anticompetitive mergers are required. Furthermore, it appears that the European Economic Community now is moving more towards a North-American model for antitrust policy.

[6] See, e.g., Scherer (1980) and Schmalensee (1989) for surveys of the empirical evidence in this area of the industrial organization literature. Stigler (1982) also notes that the economics profession has provided little *tested* knowledge to support the market shares and industry concentration criteria which form the basis for U.S. antitrust policy.

In an efficient capital market, merger-induced changes in expected future product and factor prices translate into abnormal stock returns to firms competing in the same industry as the merging firms. A collusive, anticompetitive merger raises the product price and thus benefits the nonmerging rivals as well. Under the market concentration doctrine, this benefit is increasing in both the pre-merger level and merger-induced change in concentration, an implication which I have rejected in earlier empirical work based on capital market data (Eckbo (1985)). Stillman (1983), Eckbo (1983), and Eckbo and Wier (1985) also present stock return evidence which rejects the proposition that the average industrial horizontal mergers challenged by U.S. antitrust authorities over the 1963–1981 period would have had collusive, anticompetitive effects. Extending these tests into the 1980s, the evidence in Schumann (1990) leads to the same conclusion, as do the related results in James and Wier (1987) for horizontal (unchallenged) mergers in the banking industry. Collectively, these studies increase the odds that the mergers challenged under U.S. antitrust laws would have been efficient had they been allowed to go through.

A finding that the enforcement agencies tend to block efficient mergers is not sufficient to conclude that the policy is socially suboptimal. While difficult to measure, the policy's total social benefit also includes the expected value of deterring potentially anticompetitive mergers from being proposed. This paper takes some first steps towards examining the empirical relevance of this deterrence argument through a comparative empirical analysis of the anticompetitive significance of observed mergers in the U.S. and in the Canadian unregulated environment. The Canadian-U.S. comparison is interesting for at least four reasons. First, as noted earlier, there was a virtual absence of an antitrust overhang in Canada prior to 1985. Second, Canadian industrial markets are characterized by fewer firms and greater concentration of output relative to the U.S. markets, increasing the a priori chance of anticompetitive effects of horizontal mergers. Third, prior to the 1988 free trade agreement with the U.S., foreign entry into Canadian industrial markets were restricted by trade barriers. Thus, the strict U.S. antitrust policy most likely had little if any spillover effect on Canadian domestic markets. Fourth, Canadian policy makers were in fact sufficiently concerned about possible anticompetitive effects of Canadian domestic mergers to eventually put in place the new and much stricter antitrust rules in 1985.

Consequently, under the "effective deterrence hypothesis" of this paper, the probability of observing anticompetitive horizontal mergers is predicted to be higher in Canada than in the U.S. prior to 1985. As explained in Section I, I examine this hypothesis by comparing the estimated parameter values in cross-sectional models explaining the announcement returns to merging firms and their nonmerging industry rivals as a function of industry concentration across the two countries. Data sources and sample selection procedures are described in Section II, while Section III discusses the empirical results. Section IV concludes the paper.

I. Concentration and Industry Wealth Effects of Horizontal Merger

Let AR_j denote the abnormal stock return to firm j as a result of the public announcement of a merger proposal between two firms in the same industry. The focus is on the merging firms and, in particular, their nonmerging industry rivals since the effect of the horizontal merger on product prices tend to cause a revaluation of the rival firms in the direction of the product price change. Assume AR_j is determined by the following linear cross-sectional model,

$$AR_j = x_j\phi + \varepsilon_j, \qquad j = 1,\cdots,N, \tag{1}$$

where x_j is a set of explanatory variables, and the residual ε_j is assumed to satisfy

$$E(\varepsilon_j) = 0, \qquad E\left(\varepsilon_j^2\right) = \sigma^2, \qquad E(\varepsilon_j, \varepsilon_{j'}) = 0 \quad \text{for } j \neq j'.$$

In this context, ε_j represents a summary statistic for variables that are omitted by the econometric equation (1) but which are publicly available to participants in the capital market and therefore reflected in the abnormal return AR_j. The explanatory variables included in the model are given by:

$$x_j\phi \equiv \phi_0 + \phi_1 ln(V_B/V_T)_j + \phi_2 R_j + \phi_3 C_j + \phi_4\, dC_j, \tag{2}$$

where V_B and V_T are the market values of the bidder and target firms, R is the number of nonmerging rival firms in the industry of the horizontal merger, and C and dC are the pre-merger level of and merger-induced change in industry concentration (defined below).

In order to arrive at the predictions of the effective deterrence hypothesis, assume that a horizontal merger can be either "efficient" or "anticompetitive."[7] Frequently cited examples of efficient mergers are those which take advantage of technological complementarities or synergies (which may either increase or decrease the combined output of the merging firms); replace inefficient management teams and organizations; reduce agency costs associated with the incentive to appropriate quasi-rents created by large sunk investments; capture unused corporate income tax credits; or internalize bankruptcy costs. Reflecting the antitrust authorities own view, an anticompetitive merger is perhaps best illustrated in the context of the classical dominant firm or price-leader model. In this model, the bidder firm is part of a coalition of producers who determines the industry's product price by

[7] Both categories presume mergers are value-increasing investments for shareholders of the bidder and target firms. Privately inefficient, value-decreasing mergers fall outside the domain of the antitrust policy debate. The merger literature recognizes that such mergers may take place, whether due to shareholder costs of monitoring managers (see, e.g, Jensen and Ruback (1983), Malatesta (1983), Jensen (1986)), or systematic overbidding on the part of acquiring firms' managers (Roll (1986)). As shown below, in the sample studied here, there is substantial evidence that the mergers on average enhanced shareholder value for both firms involved in the transactions.

unilaterally restricting output, letting the rival fringe firms taking full advantage of the output restriction. In the presence of entry barriers, and ignoring the free-rider problem,[8] it may pay the coalition to buy up one or more of the fringe firms in order to restrict output and raise the product price further, thus the anticompetitive effects.

If high industry concentration tends to be the result of high entry barriers, then the value of C in equation (1) is inversely related to the elasticity of industry demand. For example, a high degree of resource specialization may prevent rapid entry of new firms in response to the existence of rents in the industry. Such an industry will tend to have relatively few firms and, given the uniqueness of the product, few substitutes and a low elasticity of industry demand. A low demand elasticity accentuates the product price impact of a given merger-induced change in industry supply. Thus, if a merger is designed to adopt a scale-increasing new production technology, then the resulting product price decrease is predicted to be relatively large in industries with high concentration. Since the private value of the scale increase is decreasing in this product price decrease, one would therefore expect to find $\phi_3 < 0$ if the mergers in the sample are predominantly of the scale-increasing, efficient type. Vice versa, if the effect of the merger is to reduce output, either through a collusive anticompetitive agreement or through scale-reducing new technology, then the value of a given output reduction is higher in industries with relatively inelastic demand. Thus, if the mergers are predominantly of the output-reducing type, one would expect to find $\phi_3 > 0$.

An analogous argument holds for the merger-induced change in concentration dC. While C captures the cross-sectional effect of a given change in industry output, dC is a proxy for the scale of this output change. Thus, if the merger is of the scale-increasing type, dC will reflect the magnitude of the scale increase. Again, for a given elasticity of industry demand, the larger the scale increase, the larger the product price drop and, ceteris paribus, the lower the private value of the scale-increasing technology. Thus, one expects to see $\phi_4 < 0$ if the sample of mergers are predominantly of the efficient, scale-increasing type. On the other hand, in the price-leader model, it is easily shown that, ceteris paribus, the increase in the industry's price that results from the purchase of one of the fringe firms is an increasing function of the market share of the target.[9] Thus, for a given elasticity of industry demand, if the mergers are predominantly anticompetitive, then one expects to find $\phi_4 > 0$.

It follows from the above that when equation (1) is estimated using the abnormal returns to nonmerging industry rivals, evidence of $\phi_3 > 0$ and

[8] The fringe firms capture benefits but bear none of the costs of the output restriction.

[9] In the dominant firm pricing model, $(p - mc)/p = s^{-1}[e_d + (1 - s)e_f]$, 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. See also Stigler (1964) and Schmalensee (1985) for alternative arguments supporting the predictions in the text.

$\phi_4 > 0$ fails to discriminate between the market concentration doctrine and the hypothesis that the typical merger is of the scale-reducing efficient type. On the other hand, evidence of $\phi_3 < 0$ and $\phi_4 < 0$ rejects the market concentration doctrine in favor of the hypothesis that the typical merger is of the scale-increasing, efficient type. In other words, evidence of $\phi_3 > 0$ and $\phi_4 > 0$ is necessary to conclude in favor of the market concentration doctrine. Furthermore, under the same doctrine, the fewer the effective antitrust constraints on horizontal merger activity, the more likely one is to observe $\phi_3 > 0$ and $\phi_4 > 0$. This is the essential empirical content of the effective deterrence hypothesis of this paper.

Equation (1) also contains the number of identified rival firms, R. If industries with a relatively large R tend to use resources commonly used in other industries as well, then the total merger gains and the industry wealth effect are expected to decrease with R, reflecting the generally low level of infra-marginal rents (including those arising from merger) in competitive industries.[10] Finally, the size ratio $ln(V_B/V_T)$ is included to capture scale effects not already captured by C and dC. Ceteris paribus, the larger the size ratio, the smaller the potential output reduction under the dominant firm model. Also, the smaller the relative size of the (presumably optimally selected) target, the fewer the number of potential alternative targets among the remaining non-merging rivals and, ceteris paribus, the less the scope for industry-wide gains due to anticipated future mergers between rival firms attempting to copy the change in operating strategy signalled by the first merger. Thus, I expect a negative value for ϕ_1.[11]

II. Data

Table I lists the total sample of 471 U.S. and Canadian domestic mergers in the sample, classified by the major two-digit Standard Industrial Classification (SIC) industry of the target firm. The 266 U.S. mergers are from Eckbo (1985) and covers the sample period 1963–1981. The 205 Canadian mergers is a subsample of the mergers and acquisitions in Eckbo (1986) and covers the sample period 1964–1982. Selection procedures and data sources are as follows.

[10] If the sources of merger benefits are industry specific, R also dictates the division of the merger gains between the bidder and the target. In this case, if the non-merging rival firms are predominantly substitute targets, then one would expect ϕ_2 to have a positive value for bidder firms and negative for target firms, with the opposite prediction if the rivals are predominantly rival bidders. This result holds whether the merger is predominantly of the efficient or of the anticompetitive kind.

[11] As discussed by Asquith, Bruner, and Mullins (1983), when AR_j represents the bidder firm, the relative size variable will possibly reflect the effects of a measurement problem as well: a given dollar gain translates into a relatively small percentage return to relatively large bidders. Thus, the precision of the estimated gain is also lower for relatively large bidders since the normal variation in equity value is greater, relative to a given dollar gain. These effects also tend to produce a negative value for ϕ_1.

Table I

Industry Classification of the 471 U.S. and Canadian Mergers in the Sample

A horizontal merger is one where the major four-digit SIC code of the bidder and target firms overlap. The table sorts the mergers based on the first two digits of the target's major four-digit code. The total sample period is 1963–1982.[a]

	Target's Major Two-Digit SIC Industry	266 U.S. Mergers			205 Canadian Mergers	
		Horizontal Unchallenged	Horizontal Challenged	Non-Horizontal	Horizontal	Non-Horizontal
10	Metal mining	2	1	1	6	4
13	Oil and gas extraction	30	1	7	24	11
14	Mining of non-metallic minerals	0	1	1	0	0
20	Food and kindred products	10	7	5	22	10
21	Tobacco manufacturers	0	1	0	0	0
22	Textile mill products	7	0	4	3	2
23	Apparel and other textile products	5	1	0	1	0
24	Lumber and wood products	4	0	1	9	6
25	Furniture and fixtures	0	0	4	4	4
26	Paper and allied products	7	4	6	6	2
27	Printing and publishing	5	2	1	13	13
28	Chemicals and allied products	4	3	3	5	12
29	Petroleum and coal products	2	1	1	1	1
30	Rubber products	1	4	1	1	2
31	Leather and leather products	2	0	1	0	0
32	Stone, clay and concrete products	4	4	5	6	3
33	Primary metal industries	7	4	10	1	3
34	Fabricated metal industries	3	6	4	6	10
35	Machinery, except electrical	8	12	6	0	0
36	Electronic machinery	6	5	5	5	4
37	Transportation equipment	3	6	4	1	2
38	Measuring and controlling instruments	6	0	3	0	0
39	Miscellaneous manufacturing	0	2	0	1	0
	Total	116	80	70	116	89
	Average number of rivals per merger[b]	15	5	10	9	7

[a] The 266 mergers between U.S. firms is from Eckbo (1985) and covers the period 1963–1981. The 205 mergers between Canadian firms are from the period 1964–1982 and represent a subsample of the 247 horizontal and 626 nonhorizontal cases in Eckbo (1986), selected subject to the requirement that the merger was announced in the *Globe and Mail* and that it was possible to identify at least one nonmerging industry rival whose shares were trading at the Toronto Stock Exchange at the time of the merger announcement.

[b] Nonmerging industry rivals are identified using the major four-digit SIC code of the target firm, augmented to a five-digit code based on information on the target's product line, subject to the requirement that the rival is traded on the stock exchanges in New York or Toronto. See the text for data sources and selection procedures.

1. The merger proposal was announced in the *Wall Street Journal* or the *Globe and Mail*, and was approved by shareholders of the merging firms.
2. The bidder or the target firm was listed on the New York Stock Exchange, the American Stock Exchange, or the Toronto Stock Exchange (TSE) at the time of the merger proposal.
3. The major productive activity of the target firm is in a mining or manufacturing industry. For unchallenged mergers in the U.S., this information is based on the four-digit SIC codes listed in Standard & Poor's *Registry of Corporations* and the Federal Trade Commission's *Statistical Report on Mergers and Acquisitions* for the year prior to the year of the merger proposal. If the merger was challenged by U.S. antitrust authorities, the relevant product market is the one listed in the American Bar Association's *Merger Case Digest* as being "endangered" by the alleged anticompetitive effects of the merger.[12] For Canadian mergers, the SIC codes are from Scott's *Industrial Index*, Dun and Bradstreet's *Canadian Key Business Index*, and Consumer and Corporate Affairs Canada's *Merger Registry*.[13] The merger is classified as "horizontal" if the four-digit SIC code representing this primary industry overlaps with one of possibly four primary SIC code industries of the bidder, or if the government in its challenge of the merger classifies the combination as horizontal.
4. The pre-merger four-firm concentration ratio (i.e., the four largest firms' share of total industry output) is available for the target's major four-digit SIC industry. As in Eckbo (1985), data sources for the U.S. concentration ratios are *Census of Manufactures* (volume 1), FTC's *Concentration Levels and Trends in the Energy Sector of the U.S. Economy*, and Rogowsky (1982). Four-firm concentration ratios for the Canadian industries are from the Statistics Canada's *Industrial Organization and Concentration in the Manufacturing, Mining and Logging Industries*.
5. There was at least one nonmerging industry rival firm in the target's major "*five*-digit SIC industry" that was also listed on the NYSE, the AMEX or, for Canadian mergers, on the TSE at the time of the merger proposal. The "five-digit SIC industry" code was generated by augmenting the target's major four-digit SIC code using company-specific information on product lines given in the above data sources.[14] The University of Chicago's Center for Research in Security Prices (CRSP), and the University of Laval (for the TSE), provide four-digit SIC codes for exchange-listed firms. Again using company-specific information on product lines, non-merging rival firms are all firms whose "five-digit SIC industry" matches that of the target.

[12] See Eckbo and Wier (1985) for a list of these industries.

[13] See Eckbo (1986) for details on the *Merger Registry*. To make the SIC code information in the two samples comparable, differences between the Canadian and the U.S. SIC codes was corrected using Statistics Canada's *Standard Industrial Classification Convertibility Index*.

[14] See Eckbo (1983) for examples.

As shown in Table I, the industries with the largest number of sample mergers in the U.S. sample are SIC 13 (oil and gas extraction) with 38 cases, and SIC 20 (food and kindred products) with 22 cases, and SIC 35 (machinery, except electrical) with 26 cases. In Canada, the most merger-intensive industries in the sample also include SIC 13 (35 cases) and SIC 20 (32 cases). Furthermore, there are 26 Canadian cases in industry SIC 27 (printing and publishing).

Table II lists the mean, median and range of the pre-merger four-firm concentration ratio of the target's major four-digit SIC industry, which is the proxy for C used in the subsequent cross-sectional analysis. While there is no generally "best" concentration measure, the four-firm concentration ratio is widely used in the industrial organization literature, as well by antitrust enforcement agencies in determining industry concentration. Another popu-

Table II

Pre-Merger Industry Concentration and Market Shares of Merging Firms

Merger Category	Sample Size	Mean	Median	Minimum	Maximum
I. Four-Firm Concentration Ratio of Target's Major Four-Digit SIC Industry (%)[a]					
Canadian nonhorizontal	89	44	42	4	93
U.S. nonhorizontal	70	41	38	5	90
Canadian horizontal	116	48	45	4	99
U.S. horizontal unchallenged	116	33	31	6	94
U.S. horizontal challenged	80	58	60	5	99
II. Market Shares of Bidder and Target Firms (%)[b]					
Canadian horizontal					
Bidders	62	21	19	1	82
Targets	41	16	15	1	51
U.S. horizontal challenged					
Bidders	72	15	12	1	60
Targets	64	13	9	1	47

[a] As explained in Eckbo (1985), data sources for the U.S. concentration ratios are *Census of Manufactures* (volume 1), FTC's *Concentration Levels and Trends in the Energy Sector of the U.S. Economy*, and Rogowsky (1982). Four-firm concentration ratios for the Canadian indsutries are from from Statistics Canada's *Industrial Organization and Concentration in the Manufacturing, Mining and Logging Industries*.

[b] The market share data for the U.S. horizontal challenged mergers are taken from Eckbo (1985), and are based on case-related court documents for the U.S. challenged mergers (in 80% of the cases), company statistics developed by Economic Information Systems, Inc., and Rogowsky (1982). Market shares for bidders and targets in Canadian horizontal mergers are computed using information in company annual reports, industry Trade Reports, and import/export tables published by Statistics Canada.

lar concentration measure, the Herfindahl index, $H \equiv \sum_j s_j^2$ where s_j is the market share of the jth firm in the industry, is not available on a sufficiently large-scale basis to serve as a useful proxy for C. However,with the information on market shares of the bidder and target firms shown in Panel II of Table II, the merger-induced change in the industry's Herfindahl index is conveniently computed as twice the product of the market shares of the bidder and the target ($dH = 2s_B s_T$), which is the proxy for dC used below. This proxy is arguably superior to the change in the four-firm concentration ratio, because the latter requires the bidder to be among the industry's four largest firms. Following the rival firm selection procedure explained above, the market share data is at the five-digit SIC industry level. For the U.S. horizontal challenged mergers, this data was compiled by Eckbo (1985), and are based on case-related court documents for the U.S. challenged mergers,[15] company statistics developed by Economic Information Systems, Inc., and Rogowsky (1982). Market shares for bidders and targets in Canadian horizontal mergers are computed using information in company annual reports, industry Trade Reports, and import/export-tables published by Statistics Canada.[16]

The mean four-firm concentration ratio is 44% and 41% in the samples of Canadian and U.S. nonhorizontal mergers, respectively. In the sample of horizontal mergers, the mean concentration ratio is 48% in Canada and 33% in the U.S., with 58% in the subsample of U.S. challenged cases. The average market share of bidder firms is 21% in the Canadian subsample and 15% in the U.S. subsample with available data. The corresponding average market shares of targets are 16% and 13%, respectively, which imply average changes in the Herfindahl index of 4% in the U.S. sample and 7% in Canada.

III. Empirical Results

A. Average Industry Wealth Effects

Table III and Table IV show the average event-period abnormal returns (AR) to bidders, targets and rival firms. Assuming stock returns are distributed multivariate normal, abnormal returns are estimated using the following market model,

$$r_{jt} = \alpha_j + \beta_j r_{mt} + \gamma_j d_{jt} + \varepsilon_{jt}, \qquad t = 1, \cdots, T, \tag{3}$$

where r_{jt} and r_{mt} are the continuously compounded rates of return on firm (or portfolio) j and the value-weighted market portfolio over period t,[17] and the binary variable d_{jt} is 1 if t is within the predefined event period and 0 otherwise. For the U.S. mergers in the sample, the estimation is performed

[15] Eighty percent of the market share information comes from this source.

[16] The computation of market shares is subject to the usual caveats concerning the difficulty of identifying and delineating the economically relevant product markets.

[17] Data sources are CRSP (daily returns) and the University of Laval TSE (monthly) stock return files. Daily stock returns for the Canadian sample are not available in machine-readable form for the full sample period.

Table III

Percent Average Event-Period Abnormal Stock Returns (AR) to Bidders and Targets

Abnormal returns are computed based on OLS-estimates of the event parameter γ_j in the market model

$$r_{jt} = \alpha_j + \beta_j r_{mt} + \gamma_j d_{jt} + \varepsilon_{jt},$$

where r_{jt} and r_{mt} are the continuously compounded rates of return to firm j and the value-weighted market portfolio over period t, and d_{jt} is a binary variable which takes on a value of 1 if t is in the pre-specified event period and zero otherwise. For U.S. mergers, the estimation period is Day −200 through Day −21, and the event period Day −20 through Day 10 relative to the *Wall Street Journal* announcement of the merger proposal. For Canadian mergers, the estimation period is Month −60 through Month −13, and the event period Month 0, relative to the *Globe and Mail* announcement of the merger proposal. (z-values and sample size N in parentheses).[a]

Merger Category	U.S.: Day −20 through Day 10 $AR = \frac{31}{N}\Sigma_{j=1}^{N}\hat{\gamma}_j$ (%)		Canada: Month 0 $AR = \frac{1}{N}\Sigma_{j=1}^{N}\hat{\gamma}_j$ (%)	
	Bidders	Targets	Bidders	Targets
Nonhorizontal	−0.41 (−0.10; 59)	19.80 (7.32; 31)	.62 (1.78; 62)	3.17 (3.78; 36)
Horizontal	1.64 (1.87; 160)	18.69 (12.05; 104)	1.10 (2.05; 78)	5.12 (6.32; 48)
Horizontal challenged	3.20 (2.02; 67)	21.70 (10.14; 43)	NA	

NA = not applicable.

[a] Assuming the mergers represent independent events, it follows that $z_j = (1/\sqrt{N})\Sigma_{j=1}^{N}\hat{\gamma}_j/\hat{\sigma}_{ij}$ is distributed approximately standard normal under the hypothesis of zero abnormal return, where $\hat{\sigma}_{ij}$ is the OLS-estimate of the standard error of γ_j.

on $T = 211$ daily return observations, starting on Day −200 and ending on Day 10 relative to the day of the announcement of the merger proposal in the *Wall Street Journal*. The event dummy d_{jt} takes on a value of one over the 31-day period from Day −20 through Day 10. Consequently, the total event-period abnormal return is estimated as $AR_j = 31\gamma_j$. For the Canadian mergers, the estimation period covers 48 observations from Month −60 through Month −13, plus the month of the merger announcement in the *Globe and Mail*. The event dummy takes on a value of one in Month 0 only, making the event period comparable to the total event period for the U.S. mergers in the sample. As a result, the total event-period abnormal return for Canadian firms is estimated as $AR_j = \gamma_j$. For both U.S. and Canadian mergers, the hypothesis that the average total event-period abnormal return equals zero is tested using the following z-statistic,

$$z \equiv \frac{1}{\sqrt{N}} \sum_{j=1}^{N} \frac{\gamma_j}{\sigma_{\gamma j}} \sim N(0,1), \tag{4}$$

Table IV

Percent Average Event-Period Abnormal Stock Returns (AR) to Non-Merging Industry Rivals

Abnormal returns are computed based on OLS-estimates of the event parameter γ_j in the market model

$$r_{jt} = \alpha_j + \beta_j r_{mt} + \gamma_j d_{jt} + \varepsilon_{jt},$$

where r_{jt} and r_{mt} are the continuously compounded rates of return to rival portfolio j and the value-weighted market portfolio over period t, and d_{jt} is a binary variable which takes on a value of 1 if t is in the prespecified event period and zero otherwise. For U.S. mergers, the estimation period is Day −200 through Day −21, and the event period Day −20 through Day 10 relative to the *Wall Street Journal* announcement of the merger proposal. For Canadian mergers, the estimation period is Month −60 through Month −13, and the event period Month 0, relative to the *Globe and Mail* announcement of the merger proposal. The publicly traded industry rivals are identified using the major four-digit industry of the target, augmented by product-specific information, or using the industry which the antitrust authorities claimed was endangered by potential anticompetitive effects of the horizontal merger. (z-values and sample size N in parentheses).[a]

	U.S.: Day −20 through Day 10 $AR = \frac{31}{N}\Sigma_{j=1}^{N}\hat{\gamma}_j$ (%)		Canada: Month 0 $AR = \frac{1}{N}\Sigma_{j=1}^{N}\hat{\gamma}_j$ (%)	
Merger Category	Equal-Weighted Portfolios	Value-Weighted Portfolios	Equal-Weighted Portfolios	Value-Weighted Portfolios
Nonhorizontal	−0.87 (−0.20; 70)	−0.61 (−.11; 70)	2.42 (1.84; 89)	2.04 (1.63; 89)
Horizontal	1.26 (2.50; 196)	.94 (1.97; 196)	−1.51 (−2.14; 116)	−1.36 (−2.01; 116)
Horizontal challenged	2.80 (2.77; 80)	2.01 (2.43; 80)	NA	
Horizontal challenged-A[b]	3.30 (2.90; 36)	3.02 (2.65; 36)	NA	

NA = not applicable.

[a] Assuming the mergers represent independent events, it follows that $z_j = (1/\sqrt{N})\Sigma_{j=1}^{N}\hat{\gamma}_j/\hat{\sigma}_{ij}$ is distributed approximately standard normal under the hypothesis of zero abnormal return, where $\hat{\sigma}_{ij}$ is the OLS-estimate of the standard error of γ_j.

[b] Results based on the nonmerging industry rivals identified by the government agencies challenging the mergers.

where $\sigma_{\gamma j}$ is the standard error of γ_j. Assuming the N sample mergers represent independent events, and replacing the true values of γ_j and $\sigma_{\gamma j}$ with their OLS estimates, this z-statistic is distributed approximately standard normal for large N.

As shown in Table III, bidders and targets in horizontal mergers in both U.S. and in Canada on average earn statistically significant, positive abnormal returns over the event period. In the U.S. sample, the average abnormal return to bidders is 1.64% across all horizontal cases and 3.20% in the subsample of challenged mergers. The corresponding average abnormal return to targets is 18.69% and 21.70%, respectively. In Canada, the average abnormal return to bidders in horizontal mergers is 1.10%, while the average target earns 5.12%. These results confirm that the mergers on average were value-increasing events, a necessary condition for the collusion hypothesis. In contrast, there is much less evidence that bidders in nonhorizontal mergers on average gain from the merger proposal announcements.[18]

The abnormal returns to the average portfolio of non-merging industry rivals shown in Table IV are interesting. In the sample of Canadian horizontal mergers, the average portfolio of industry rivals (both equal- and value-weighted) is a negative -1.51%. This negative rival firm performance rejects the collusion or dominant firm hypothesis under which the merger causes a marginal increase in the industry's product price. On the other hand, the observed negative industry wealth effect is consistent with the hypothesis that the market expects the typical horizontal merger in Canada to be efficient, thus placing rival firms at a competitive disadvantage.

Furthermore, the rival firm performance in the sample of 89 nonhorizontal Canadian mergers is positive and marginally significant. Since nonhorizontal mergers do not have anticompetitive effects, this positive rival performance most likely reflect dissemination of valuable information caused by the merger announcement. For example, the merger may signal that a particular industry-specific resource has increased in value, increasing the probability that other industry members will become future targets.

In the sample of U.S. mergers, the evidence for rival firms is reversed. As documented in earlier papers, rivals earn significantly positive abnormal return in horizontal mergers, with 2.80% in the subsample of challenged mergers, and zero abnormal returns in nonhorizontal mergers. The positive rival firm performance is particularly strong in the subsample denoted "Horizontal Challenged-A," which contains 36 cases for which Eckbo and Wier (1985) obtained a set of rival firms identified by the government agencies responsible for selecting the horizontal mergers for prosecution. In other words, the results appear to be robust with respect to the procedure for identifying economically relevant, nonmerging industry rivals.

The significance of the rival firm abnormal returns are also confirmed by cross-equation tests of the abnormal returns to individual rival firms associated with a given merger announcement. This additional test is of interest since the portfolio approach underlying Table IV fails to reveal abnormal returns if the merger has a positive impact on some individual rivals and a negative and offsetting impact on others. Thus, to complement the z-test, Table V report tests results for the following null hypothesis,

$$H_0\colon \gamma_{ij} = 0, \qquad i = 1,\cdots, R_j. \tag{5}$$

H_0 states that the event parameters across the R_j individual rival firms associated with the j'th merger in the sample are jointly equal to zero. The

[18] See Eckbo (1986) for additional information on the abnormal performance of merging firms in Canada.

Table V

Cross-Equation Tests of Event Parameter Restrictions for Non-Merging Industry Rivals

In this table, the null hypothesis is that the abnormal returns across the rival firms associated with a given merger are jointly equal to zero, i.e., H_0: $\gamma_{ij} = 0$, $i = 1, \cdots, R_j$. The test statistic for H_0 is

$$\frac{1}{(\bar{X}'\bar{X})_{33}^{-1}} \hat{\underline{\gamma}}'\hat{\Sigma}^{-1}\underline{\gamma} \sim \chi^2(R_j),$$

where $\bar{X} = [\underline{1}\underline{r}_m\underline{d}]$ is the $T \times 3$ matrix of market model regressors for merger j, $(\bar{X}'\bar{X})_{33}^{-1}$ is the (3, 3) element of $(\bar{X}'\bar{X})^{-1}$, Σ is the $R_j \times R_j$ contemporaneous residual covariance matrix estimated from the first-pass OLS regression of the R_j individual market model equations, and $\hat{\underline{\gamma}}$ is the $R_j \times 1$ vector of event parameters. The table reports the *sum* of the N values of this χ^2 statistic across a sample of N mergers (which itself has a limiting χ^2 distribution with $\Sigma_j^N R_j$ degrees of freedom), and in parenthesis the percent of the n individual χ^2 values that reject H_0 at a 5% level of significance.[a]

Merger Category	Industry Rivals of U.S. Mergers		Industry Rivals of Canadian Mergers	
	$\Sigma_j^N\chi^2(R_j)$	*df*	$\Sigma_j^N\chi^2(R_j)$	*df*
Nonhorizontal	630 (4.3)	700	602 (6.1)	623
Horizontal	2910[b] (11.3)	2141	1316[b] (8.1)	1044
Horizontal challenged	510[b] (16.4)	401	NA	
Horizontal challenged-A[c]	206[b] (20.3)	144	NA	

[a] For rivals of Canadian mergers, the event parameter γ_{ij} used in the test statistic of this table corresponds to the one-month event parameter in the market model (equation (3) in the text). For rivals of U.S. mergers, the event parameter has been scaled up by a factor of 31 to reflect the total abnormal return over the 31-day event period.

[b] Statistically significant at the 5% level.

[c] Results based on the nonmerging industry rivals identified by the government agencies challenging the mergers and compiled by Eckbo and Wier (1985).

system of R_j rival firm market model equations can be written as (dropping, for convenience, subscript j)

$$\underline{r} = X\underline{\delta} + \underline{\varepsilon}, \tag{6}$$

$\underline{r}' = [\underline{r}_1' \cdots \underline{r}_R']$, the $1 \times TR_j$ vector of returns

$X = [I \otimes \bar{X}]$, a $TR_j \times 3R_j$ matrix

$\bar{X} = [\underline{1}\underline{r}_m\underline{d}]$, the $T \times 3$ matrix of regressors

$\underline{\delta}' = [\alpha_1 \beta_1 \gamma_1 \cdots \alpha_R \beta_R \gamma_R]$, the $1 \times 3R_j$ vector of coefficients,

$\underline{\varepsilon} = [\underline{\varepsilon}_1' \cdots \underline{\varepsilon}_R']$, a $1 \times TR_j$ vector of error terms.

In this notation, I is the $R_j \times R_j$ identity matrix, $\underline{1}$ is the $T \times 1$ unity vector, $\otimes$ indicates the Kronecker or direct product operator, and $\underline{\varepsilon} \sim MVN(0; \Sigma \times I)$. The $R_j \times R_j$ contemporaneous covariance matrix Σ is estimated from the residuals produced by the first-pass OLS regression of system (6).

Since (6) is a system of R_j "seemingly unrelated" regressions with identical regressors, OLS estimation of each equation is asymptotically efficient and provides identical parameter estimates to a generalized least squares procedure (Zellner (1962)). However, significance tests of cross-equation constraints on the R_j event parameters require a procedure which accounts for the contemporaneous cross-correlation of the error terms $\underline{\varepsilon}$. This is particularly important in the present application where the R_j rival firms are operating in the same product market and are simultaneously affected by the merger announcement. Thus, restate H_0 as

$$H_0 : \underline{0} = C_3 \underline{\delta}, \tag{7}$$

where $\underline{0}$ is a $R_j \times 1$ vector of zeros; $C_3 = [I \otimes \underline{c}_3']$ is a $R_j \times 3R_j$ matrix; and $\underline{c}_3$ is a 3×1 vector where the third element equals one and the remaining two elements equal zero. Replacing the true values of $\underline{\delta}$ and Σ with their OLS estimates, the following quadratic form has a limiting χ^2 distribution with R_j degrees of freedom under H_0 (Theil (1971, pp. 312–317)):

$$\underline{\hat{\delta}}' C_3' \left\{ C_3 \left[X' \left(\hat{\Sigma}^{-1} \otimes I_T \right) X \right]^{-1} C_3' \right\}^{-1} C_3 \underline{\hat{\delta}} \sim \chi^2(R_j). \tag{8}$$

This quadratic form can be simplified to[19]

$$\frac{1}{(\bar{X}'\bar{X})_{33}^{-1}} \underline{\hat{\gamma}}' \hat{\Sigma}^{-1} \underline{\hat{\gamma}}, \tag{9}$$

[19] The simplification makes use of the following results:

$$[X'(\hat{\Sigma}^{-1} \otimes I_T) X]^{-1} = \hat{\Sigma} \otimes (\bar{X}'\bar{X})^{-1},$$

and, in our application,

$$\left\{ C_3 \left[\hat{\Sigma} \otimes (\bar{X}'\bar{X})^{-1} \right] C_3' \right\}^{-1} = \left\{ \hat{\Sigma} \otimes \underline{c}_3' (\bar{X}'\bar{X})^{-1} \underline{c}_3 \right\}^{-1} = \frac{1}{(\bar{X}'\bar{X})_{33}^{-1}} \hat{\Sigma}^{-1},$$

where $(\bar{X}'\bar{X})_{33}^{-1}$ is the (3, 3) element of $(\bar{X}'\bar{X})^{-1}$. Finally,

$$C_3' \left\{ \frac{1}{(\bar{X}'\bar{X})_{33}^{-1}} \hat{\Sigma}^{-1} \right\} C_3 = \frac{1}{(\bar{X}'\bar{X})_{33}^{-1}} [\hat{\Sigma}^{-1} \otimes \underline{c}_3 \underline{c}_3'],$$

where the 3×3 matrix $\underline{c}_3 \underline{c}_3'$ contains a value of one in element (3, 3) and zeros otherwise.

where $(\bar{X}'\bar{X})^{-1}_{33}$ is the (3, 3) element of $(\bar{X}'\bar{X})^{-1}$ and $\hat{\underline{\gamma}}$ is the $R_j \times 1$ vector of event parameters.

Table IV shows the value of this test statistic summed over each of the N takeovers in the sample, which itself is distributed χ^2 with $\Sigma_j^N R_j$ degrees of freedom, as well as the percentage of the N individual χ^2 values which reject H_0 at a 5% level of confidence. There is strong evidence that a significant portion of the rival firms across individual mergers indeed earn non-zero abnormal performance as a result of the merger announcements, particularly in the sample of U.S. horizontal challenged cases and when using the rival firms identified by the enforcement agencies.

As discussed earlier, positive rival firm performance in response to the merger proposal announcement does not discriminate between the market power hypothesis (where the dominant coalition reduces industry output and increases the product price post-merger) and the hypothesis that the merger signals opportunities for efficiency gains available to the non-merging industry rivals as well. However, there are several pieces of evidence which tend to favor the "information signalling hypothesis" over the market power hypothesis. First, Eckbo (1983) and Eckbo and Wier (1985) report that the initial gains to rival firms are not reversed in response to the subsequent (successful) government challenge of the merger. Second, Schumann (1990), using a sample of challenged mergers in the 1980s, reports that 12 of 48 "small" rival firms were acquired or restructured within three years of the antitrust challenge of the proposed merger. In other words, conditional upon a horizontal (and subsequently challenged) merger proposal in the industry, there is a 25% chance that the industry's relatively small firms become takeover targets. This surprisingly high takeover frequency is consistent with the information signalling hypothesis, suggesting that the opportunities for efficiency gains require merger.[20] Third, as discussed above, rivals of nonhorizontal mergers in Canada earn positive abnormal returns when no market power effects are possible.

Fourth, Eckbo (1985) provides cross-sectional tests which reject the hypothesis that the industry wealth effect is increasing in either the pre-merger level of, or merger-induced change in, industry concentration, a result which is inconsistent with the market power hypothesis. However, it remains possible that this test result reflects an effective antitrust deterrence policy rather than the absence of anticompetitive mergers per se. I now turn to a comparative cross-sectional analysis of Canadian and U.S. mergers to examine the effective deterrence hypothesis. In the process, I also introduce recently developed econometric procedures designed to produce consistent estimates of coefficients in cross-sectional models with announcement returns as dependent variable.

[20] Schumann (1990) also report that the initial announcement gains to the non-merging industry rivals dissipate within six months unless the rivals in fact become involved in merger or are restructured.

B. Cross-Sectional Analysis

A cross-sectional model such as (1) has obvious econometric advantages over the traditional concentration-profits regressions performed in the industrial organization literature, including a relatively precise, risk-adjusted measurement of the change in industry rents and a direct *causal* relationship running from the merger-induced change in concentration to the industry wealth effect. However, due to the anticipatory nature of prices in an efficient market, the use of stock price changes at the time of the initial merger proposal announcement to measure the merger-induced change in industry rents poses certain econometric challenges of its own. These include adjustments for (a) the effect on announcement returns of managerial private information concerning the value of the voluntary corporate event (Acharya (1988), Lanen and Thompson (1988), Eckbo, Maksimovic, and Williams (1990)), (b) prior market anticipation of the event (Malatesta and Thompson (1985)), and (c) the attenuation effect of the antitrust regulatory overhang in the U.S. (Eckbo (1985), Eckbo, Maksimovic, and Williams (1990)). In the following, I explain these adjustments within the Eckbo-Maksimovic-Williams framework before presenting the parameter estimates for the cross-sectional model.

B.1. Managerial Private Information

Suppose capital market participants know that managers of the bidder firm will initiate a merger only after receiving a private signal η_j which indicates that the event has a positive value, i.e., which is such that:

$$x_j\phi + \eta_j > 0.$$

As a result, when the (unanticipated) event occurs, the market uses its knowledge of managers' incentives to impound into the firm's stock price the following expected return,

$$E\left(AR_j \mid \eta_j > -x_j\phi\right) = x_j\phi + E\left(\eta_j \mid \eta_j > -x_j\phi\right). \tag{10}$$

In effect, the market's inference truncates the residual term η_j that measures the value of managers' private information. This truncation is ignored if the econometrician elects to estimate equation (1) as is, since the nonlinear expectational term in equation (10) ends up in the error term ε_j, causing the estimated parameter values to be inconsistent.

Assume that the private signal η_j is normally distributed and independent across sample firms, i.e., $\eta_j \sim N(0, \omega^2)$. A consistent specification of the cross-sectional model is then given by

$$AR_j = x_j\phi + \omega\frac{n(x_j\phi/\omega)}{N(x_j\phi/\omega)} + \zeta_j, \tag{11}$$

where ζ_j is assumed to satisfy

$$E(\zeta_j) = 0, \qquad E\left(\zeta_j^2\right) = v^2, \qquad E(\zeta_j, \zeta_{j'}) = 0 \quad j \neq j',$$

and where $n(\cdot)$ represents the standard normal density function.[21]

While, in this framework, the *bidder* is the one that receives the private signal η_j and therefore initiates the merger, I also use the nonlinear equation (11) when the dependent variable is the abnormal return to the nonmerging rivals. The purpose of this is twofold. First, under the collusion hypothesis, the bidder's private information, which is partly revealed through the merger proposal, have direct value implications for the rivals as well, in which case OLS or GLS estimation of equation (1) again produces a truncation bias. Second, since the parameter ω represents the standard error of the private information, its estimated value should be nonnegative. Thus, evidence of a significantly negative value of ω would suggest that the model is misspecified.

B.2. Market Anticipation of the Event

Equation (11) assumes that market participants do not recognize that bidding firm managers have received private information until the bid is announced, i.e., the merger announcement is completely unanticipated. However, bidder specific information released prior to the event may cause the market to infer with certainty that the bidder has received a potential target, raising the probability of a subsequent merger announcement from 0 to $\Pr(\eta_j \geq -x_j\phi) = N(x_j\phi/\omega)$. In this case, the conditional expected abnormal return in response to the merger proposal becomes $E(AR_j \mid \eta_j > -x_j\phi)[1 - N(x_j\phi/\omega)]$, which is the specification used below.[22]

B.3. The Antitrust Overhang

The analysis so far ignores the attenuating effect on announcement returns of the possibility that the proposed horizontal merger may be challenged—and possibly blocked—by antitrust authorities. Again using the Eckbo-Maksimovic-Williams framework, let p_{rj} be the probability that the antitrust regulator decides to challenge the case, and let p_{cj} be the probability that the (independent) court decides in favor of the regulator. Both the regulator and the court have access to the public information x_j. Suppose the regulator decides to challenge the merger if the value of $x_j\phi_r + \eta_r \geq 0$. Furthermore,

[21] To arrive at the nonlinear term in equation (11), define $z_j \equiv \eta_j/\omega$ and $\theta_j \equiv -x_j\phi/\omega$. Since $\partial n(z_j)/\partial z_j = -z_j n(z_j)$, the conditional expected value of η_j in (10) can be written

$$\omega E(z_j \mid z_j \geq \theta_j) = \frac{\omega}{1 - N(\theta_j)} \int_{\theta_j}^{\infty} z_j n(z_j)\, dz_j = -\frac{\omega}{N(-\theta_j)} \int_{\theta_j}^{\infty} \frac{\partial n(z_j)}{\partial z_j}\, dz_j = \omega \frac{n(-\theta_j)}{N(-\theta_j)}.$$

See also Maddala (1983) and Hausman and Wise (1977) for discussions of truncated dependent variables.

[22] See Malatesta and Thompson (1985) for an alternative specification of the effect of pre-announcement rumors.

suppose that the court will block the merger if $x_j\phi_c + \eta_c \geq 0$. If the private information received by the merging firm, the regulator and the court (i.e., η, η_r, and η_c) all are normally and independently distributed random variables, then $p_{rj} = N(-x_j\phi_r/\omega_r)$, and $p_{cj} = N(-x_j\phi_c/\omega_c)$, where ω_r and ω_c are the standard errors of the regulator's and the court's private information, respectively. These probabilities are estimated using probit analysis.

The merger proposal now has three possible outcomes: (1) No challenge with probability $(1 - p_{rj})$, (2) an unsuccessful challenge (no court conviction) with probability $p_{rj}(1 - p_{cj})$, and (3) a successful challenge (court conviction) with probability $p_{rj}p_{cj}$. Assume a challenge imposes a legal cost on the merging firms equal to a proportion c of the merger gains. As a result, conditional on the public information x_j, the market reaction to the merger proposal can be written as

$$E\left(AR_j^* \mid \eta_j > -x_j\phi\right) \equiv \left[(1 - p_{rj}p_{cj})\left(x_j\phi + \omega\frac{n(x_j\phi/\omega)}{N(x_j\phi/\omega)}\right) - p_{rj}c\right]\left[1 - N(x_j\phi/\omega)\right], \quad (12)$$

The conditional expectation in (12) is estimated for U.S. mergers using non-linear maximum likelihood, while the estimation for Canadian mergers is performed without the above adjustment for antitrust overhang.

B.4. Parameter Estimates

The probability of a government challenge (p_{rj}), and the conditional probability that the challenge will be successful (p_{cj}), are computed based on the cross-sectional (probit) estimates for the government and the court shown in Table VI. Table VII shows the estimated parameter values for the cross-sectional model for pairs of bidder and target firms with and without the adjustment for antitrust overhang. These estimates, which are identical to those used in Eckbo, Maksimovic and Williams (1990), imply that p_{rj} increases with the pre-merger level of concentration C and decreases with the number of non-merging rival firms R. On the other hand, since the estimates in the second row of Table VI are all insignificant, I use a constant value 59/80 (the sample proportion of successful challenges) as a proxy for p_{cj}. In the nonlinear, maximum-likelihood estimation underlying Table VII and Table VIII these probabilities are treated as constants, as is the proportional court cost c which is assumed to have a value of 0.01.

The sample without antitrust overhang includes all Canadian as well as U.S. nonhorizontal mergers. The regression for the U.S. nonhorizontal mergers in Table VII is statistically insignificant. In contrast, the regressions for the Canadian horizontal and nonhorizontal mergers are statistically significant and indicate that the equal-weighted sum of the gains to bidders and targets decreases with the size ratio $ln(V_B/V_T)$ and the number of rivals R. Furthermore, the gains to the merging firms increase with the pre-merger level of industry concentration C and decreases with the change in concentra-

Table VI

Cross-Sectional Estimates for the U.S. Government and the Court

Maximum likelihood estimates of the coefficients ϕ_r for the probability of a government challenge (p_{rj}) and ϕ_c for the probability of a pro-government court decision (p_{cj}). The binary dependent variable is set equal to one if the merger is challenged (row 1) or if the challenge is successful (row 2) and 0 otherwise. Sample of 196 U.S. horizontal mergers, 1963–1981.[a]

(Asymptotic t-values in parentheses)

Event	Yes	No	Constant ϕ_0	$ln(V_B/V_T)$ ϕ_1	R ϕ_2	C ϕ_3	χ^2
Government challenges a proposed merger	80	116	−1.170	−0.068	−0.040	0.031	77.2[b]
			(−3.67)	(−1.04)	(−2.77)	(6.01)	(3 *df*)
Challenge successful	59	21	0.349	−0.038	0.058	0.003	1.52
			(0.55)	(−0.33)	(1.05)	(0.33)	(3 *df*)

[a] See Table VII for variable definitions. These estimates are identical to those in Eckbo, Maksimovic, and Williams (1990). A successful merger is defined as a government complaint against the merger with one of two results: Either the merger proposal was dropped or, if the merger was consummated before the complaint, the merged firm was ordered to divest. The outcomes of each of the 80 challenged cases are discussed in the Appendix of Eckbo and Wier (1985).

[b] Significant at the 1% level.

tion dC. The positive effect of C is present in both the horizontal and nonhorizontal samples, which suggests that the effect of C is *not* related to market power. The negative impact of dC in the horizontal subsample also fails to support the market power hypothesis. Notice also that the Canadian sample produces a significant value of ω (the standard error of the distribution of managers' private information), which is consistent with the assumption underlying the econometric procedure.

Panel II of Table VII shows the results of the estimation with the adjustment for antitrust overhang given by the conditional expectation in equation (12) above. The relevant sample here consists of U.S. horizontal mergers only, with particular focus on the subsample of challenged cases. The parameter values are similar to the corresponding values in Panel I, with R and dC again producing a negative effect on merger gains, and C having a positive effect. Both regressions are statistically significant at the 1% level, with positive and significant values of the information parameter ω.[23]

Table VIII shows the parameter estimates for the equal-weighted portfolios of rival firms. In Panel I (sample without antitrust overhang), the regressions are statistically insignificant with the exception of the subsample of 37

[23] See Eckbo, Maksimovic, and Williams (1990) for evidence on the distribution of the merger gains between the bidder and the target firms. According to their results, target gains increase with C and decrease with R, while bidder gains decrease in both C and R. Their results also show that the non-linear estimation procedure used here produces parameter estimates for bidder and target firms which differ significantly from the standard OLS estimates.

Table VII

Cross-Sectional Estimates for Pairs of Bidder and Target Firms

Maximum likelihood estimates of the coefficients ϕ, ω in nonlinear cross-sectional models with the total announcement-induced abnormal stock returns (AR_j) to equal-weighted pairs of bidder and target firms as dependent variable, where:

$$AR_j = H(x_j\phi, \omega) + \zeta_j, \qquad j = 1 \cdots N.$$

(Asymptotic t-values in parentheses).[a]

Merger Category	Constant ϕ_0	$ln(V_B/V_T)$ ϕ_1	R ϕ_2	C ϕ_3	dC ϕ_4	ω	χ^2
I. Sample without Antitrust Overhang							
$H(\cdot) = \left[x_j\phi + \omega\frac{n(x_j\phi/\omega)}{N(x_j\phi/\omega)}\right][1 - N(x_j\phi/\omega)]$							
25 U.S. nonhorizontal	−0.024	−0.002	−0.001	0.004	—	0.001	3.3
	(-0.92)	(−0.53)	(−1.55)	(1.98)		(0.33)	(3 *df*)
31 Canadian nonhorizontal	0.093	−0.006	−0.007	0.013	—	0.034	12.2[b]
	(1.01)	(−0.77)	(−1.85)	(2.42)		(2.66)	(3 *df*)
42 Canadian horizontal	−0.048	−0.010	−0.012	0.023	−0.066	0.021	20.1[b]
	(−0.97)	(−1.67)	(−2.05)	(3.41)	(−2.95)	(2.52)	(4 *df*)
II. Sample with Antitrust Overhang							
$H(\cdot) = \left[(1 - p_{rj}p_{cj})\left(x_j\phi + \omega\frac{n(x_j\phi/\omega)}{N(x_j\phi/\omega)}\right) - p_{rj}c\right][1 - N(x_j\phi/\omega)]$							
81 U.S. horizontal	0.091	−0.014	−0.004	0.006	—	0.021	16.1[b]
	(1.01)	(−1.42)	(−2.24)	(2.41)		(2.01)	(3 df)
46 U.S. challenged	0.241	−0.021	−0.003	0.015	−0.065	0.042	23.3[b]
	(0.60)	(−1.98)	(−1.81)	(3.66)	(−2.38)	(3.12)	(4 *df*)

[a] The dependent variable AR_j for Canadian mergers equals the event parameter γ_j estimated from the market model (3), while for U.S. mergers it is equal to the market model estimate of γ_j times 31. The explanatory variables x_j are the log of the ratio of the total equity size of the bidder and the target ($lv(V_B/V_T)$, the number of identified nonmerging industry rivals (R), the pre-merger four-firm industry concentration ratio (C), and the merger-induced change in the industry's Herfindahl Index (dC), computed as $2s_Bs_T$ where s_i is the pre-merger market share of firm i. OLS estimates are used as initial (starting) parameter values in the nonlinear estimation, with the OLS regression standard error as the initial value for ω. The probabilities $\hat{p}_{rj}$ are computed from the parameter estimates in Table VI, while the value of $\hat{p}_{cj}$ is set equal to 52/80 for all j. These probabilities, and the cost of going to court $c = 0.01$, are treated as constants in the estimation above. The normal distribution is approximated to five digits.

[b] Significant at the 1% level.

Canadian horizontal cases with data on market shares. In this regression, dC enters with a statistically significant, negative coefficient, indicating that the higher the increase in industry concentration caused by the merger, the *lower* the wealth impact on the nonmerging industry rivals. The value of C enters with a positive, but statistically insignificant value in the samples of

Table VIII

Cross-Sectional Estimates for Portfolios of Non-Merging Industry Rivals

Maximum likelihood estimates of the coefficients ϕ, ω in nonlinear cross-sectional models with the total announcement-induced abnormal stock returns (AR_j) to equal-weighted portfolios of rival firms as dependent variable, where:

$$AR_j = H(x_j\phi, \omega) + \zeta_j, \qquad j = 1 \cdots N.$$

(Asymptotic t-values in parentheses).[a]

Merger Category	Constant ϕ_0	$ln(V_B/V_T)$ ϕ_1	R ϕ_2	C ϕ_3	dC ϕ_4	ω	χ^2
I. Sample without Antitrust Overhang							
$H(\cdot) = \left[x_j\phi + \omega\frac{n(x_j\phi/\omega)}{N(x_j\phi/\omega)}\right][1 - N(x_j\phi/\omega)]$							
70 U.S. nonhorizontal	−0.003	0.001	−0.004	0.036	—	−0.002	4.2
	(−0.19)	(0.14)	(−0.32)	(1.32)		(−0.16)	(3 df)
89 Canadian nonhorizontal	0.077	−0.023	−0.005	0.015	—	0.001	3.9
	(1.66)	(−1.18)	(−0.24)	(1.07)		(0.03)	(3 df)
116 Canadian horizontal	−0.089	−0.038	0.002	−0.008	—	0.004	5.6
	(−1.99)	(−1.47)	(0.76)	(−0.99)		(0.10)	(3 df)
37 Canadian horizontal	−0.046	−0.030	−0.012	0.010	−0.009	0.006	13.9[b]
	(−1.56)	(−1.36)	(−1.03)	(1.21)	(−2.32)	(0.57)	(4 df)
II. Sample with Antitrust Overhang							
$H(\cdot) = \left[(1 - p_{rj}p_{cj})\left(x_j\phi + \omega\frac{n(x_j\phi/\omega)}{N(x_j\phi/\omega)}\right) - p_{rj}c\right][1 - N(x_j\phi/\omega)]$							
196 U.S. horizontal	0.096	−0.004	−0.010	0.008	—	0.002	3.3
	(1.10)	(−0.01)	(−1.08)	(0.94)		(0.13)	(3 df)
64 U.S. challenged	0.135	−0.032	−0.009	0.001	−0.002	−0.001	8.3
	(1.62)	(−1.05)	(−0.99)	(0.69)	(−2.08)	(−0.68)	(4 df)
27 U.S. challenged-A[c]	0.231	−0.005	−0.017	0.035	−0.003	0.001	16.4[b]
	(2.12)	(−0.34)	(−2.31)	(1.93)	(−2.13)	(0.09)	(4 df)

[a] The dependent variable AR_j for Canadian mergers equals the event parameter γ_j estimated from the market model (3), while for U.S. mergers it is equal to the market model estimate of γ_j times 31. The explanatory variables x_j are the log of the ratio of the total equity size of the bidder and the target ($lv(V_B/V_T)$), the number of identified non-merging industry rivals (R), the pre-merger four-firm industry concentration ratio (C), and the merger-induced change in the industry's Herfindahl Index (dC), computed as $2s_B s_T$ where s_i is the pre-merger market share of firm i. OLS estimates are used as initial (starting) parameter values in the non-linear estimation, with the OLS regression standard error as the initial value for ω. The probabilities $\hat{p}_{rj}$ are computed from the parameter estimates in Table VI, while the value of $\hat{p}_{cj}$ is set equal to 52/80 for all j. These probabilities, and the cost of going to court $c = 0.01$, are treated as constants in the estimation above. The normal distribution is approximated to five digits.

[b] Significant at the 1% level.

[c] Results based on the non-merging industry rivals identified by the government agencies challenging the merges and compiled by Eckbo and Wier (1985).

nonhorizontal as well as in the sample of 37 horizontal Canadian mergers. Furthermore, the parameter value of C is negative (and insignificant) in the full sample of 116 Canadian horizontal mergers. These results are inconsistent with the market concentration doctrine.

In Panel II of Table VIII, the merger-induced change in concentration again enters with a significantly negative coefficient. This is particularly true for the subsample of 27 horizontal "Challenged-A" cases with data on market shares of the merging firms. Since the rival firms in this sample were identified by the enforcement agencies themselves as economically relevant industry rivals, this sample presumably removes some of the possible shortcomings of the rival firm selection procedure described in Section II above. It is therefore particularly interesting to find that the value of ϕ_4 is significantly negative also for this sample. Overall, the results in Table VIII lend no support to the market power hypothesis whether one uses the U.S. or the Canadian merger samples. Consequently, the evidence also rejects the effective deterrence hypothesis set out in Section I.[24]

IV. Conclusions

While the U.S. has pursued a vigorous antitrust policy towards horizontal mergers over the past four decades, mergers in Canada have until recently been permitted to take place in a virtually unrestricted antitrust environment. The absence of an antitrust overhang in Canada presents an interesting opportunity for testing the conjecture that the rigid market share and concentration criteria of the U.S. policy deters a significant number of potentially collusive mergers. This effective deterrence hypothesis implies that the probability of a horizontal merger being anticompetitive is higher in Canada than in the U.S.

A collusive, anticompetitive merger increases the industry's monopoly rents, and thus the market value of the rivals of the merging firms. Under the market concentration doctrine, this rent-increase is necessarily larger the larger the pre-merger level of industry concentraton and market shares of the merging firms. The paper examines this implication by means of cross-sectional regressions with merger-induced abnormal stock returns to non-merging industry rivals as dependent varible. Using an econometric procedure which explicitly accounts for managers' private information concerning the true value of the merger, as well as the effect of the antitrust overhang in the U.S., the paper rejects the market concentration doctrine on samples of

[24] The regressions in Table VIII were recomputed with value-weighted rather than equal-weighted portfolios of rival firms without changing the main conclusions above. Furthermore, the estimation was repeated without the adjustment for pre-announcement rumors, again without a material effect on the results. In fact, the likelihood ratio test of Vuong (1989) fails to reject the hypothesis that the models with and without rumors fit the data equally well, suggesting that pre-announcement anticipation of the event is not important in this sample.

both U.S. and Canadian mergers. By implication, the results also reject the effective deterrence hypothesis. The evidence is, however, consistent with the alternative hypothesis that the horizontal mergers in either of the two countries were expected to generate productive efficiencies.

Since a significant part of the data on market shares and concentration in the U.S. merger sample was collected using sources related to the challenged cases, the rejection of the market concentration doctrine is unlikely to be explained by an excessive level of aggregation in the measures of these industry characteristics, or by a failure to identify economically relevant industry rivals. The enforcement agencies frequently use a definition of the relevant "endangered" product market that contains only a few closely related products, and the evidence reject the market power hypothesis also when using a set of rival firms identified by the enforcement agencies as economically relevant. Furthermore, since the value of industry concentration ranges from a low of 5% to a high of 99% in both the Canadian and the U.S. samples, the rejection of the market concentration doctrine hardly reflects the possibility that the respective product markets were monopolized *prior* to the merger (despite the government challenge). Rather, the evidence systematically rejects the antitrust doctrine even for values of industry concentration and market shares which, during the past four decades, have been considered critical in determining the probability that a horizontal merger will have anticompetitive effects. Judging from the evidence, there simply isn't much to deter.

REFERENCES

Acharya, S., 1988, An empirical test of managerial self-interest on corporate capital structure, *Journal of Finance* 43, 413–429.

Asquith, P., R. F. Bruner, and D. W. Mullins, Jr., 1983, The gains to bidding firms from merger, *Journal of Financial Economics* 11, 121–139.

Eckbo, B. E., 1983, Horizontal mergers, collusion, and stockholder wealth, *Journal of Financial Economics* 11, 241–273.

———, 1985, Mergers and the market concentration doctrine: Evidence from the capital market, *Journal of Business* 58, 325–349.

———, 1986, Mergers and the market for corporate control: The Canadian evidence, *Canadian Journal of Economics* 19, 236–260.

———, V. Maksimovic and J. Williams, 1990, Consistent estimation of cross-sectional models in event studies, *Review of Financial Studies* 3, 343–365.

——— and P. Wier, 1985, Antimerger policy under the Hart-Scott-Rodino Act: A reexamination of the market power hypothesis, *Journal of Law and Economics* 28, 119–149.

Ellert, J. C., 1976, Mergers, antitrust law enforcement and stockholder returns, *Journal of Finance* 31, 715–732.

Hausman, J. and D. Wise, 1977, Social experimentation, truncated distributions and efficient estimation, *Econometrica* 45, 919–939.

James, C. M. and P. Wier, 1987, Returns to acquirers and competition in the acquisition market: The case of banking, *Journal of Political Economy* 95, 355–370.

Jensen, M. C., 1986, Agency costs of free cash flow, corporate finance and takeovers, *American Economic Review* 76, 323–329.

——— and R. Ruback, 1983, The market for corporate control: The scientific evidence, *Journal of Financial Economics* 11, 5–50.

Lanen, W. and R. Thompson, 1988, Stock price reactions as surrogates for the net cash flow effects of corporate policy decisions in cross-sectional studies, *Journal of Accounting and Economics* 10, 311–334.

Maddala, G., 1983, *Limited-Dependent and Qualitative Variables in Econometrics*, (Cambridge University Press, Cambridge, England).

Malatesta, P. H., 1983, The wealth effect of merger activity and the objective functions of the merging firms, *Journal of Financial Economics* 11, 155–181.

——— and R. Thompson, 1985, Partially anticipated events: A model of stock price reactions with an application to corporate acquisitions, *Journal of Financial Economics* 14, 237–250.

Peltzman, S., 1977, The gains and losses from industrial concentration, *Journal of Law and Economics* 20, 229–263.

Rogowski, R. A., 1982, The Justice Department's merger guidelines: A study in the application of the rule, Unpublished paper, Federal Trade Commission.

Roll, R., 1986, The hubris hypothesis of corporate takeovers, *Journal of Business* 59, 197–216.

Scherer, F. M., 1980, *Industrial Market Structure and Economic Performance*, 2nd ed., (Rand McNally, Chicago, IL).

Schipper, K. and R. Thompson, 1983, The impact of merger-related regulations on the shareholders of acquiring firms, *Journal of Accounting Research* 21, 184–221.

Schmalensee, R., 1985, Competitive advantage and collusion, Working paper 8543, CORE, Université Catholique de Louvain.

———, 1989, Inter-industry studies of structure and performance, in R. Schmalensee and R. Willig, eds.: *Handbook of Industrial Organization* (North-Holland, Amsterdam), 951–1009.

Schumann, L., 1990, An analysis of the use of stock returns to determine the competitive effects of horizontal merger, Unpublished paper, Federal Trade Commission, Washington, DC.

Stigler, G. J., 1964, A theory of oligopoly, *Journal of Political Economy* 72, 44–61.

———, 1966, The economic effects of the antitrust laws, *Journal of Law and Economics* 9, 225–258.

———, 1982. The economist and the problem of monopoly, *American Economic Review Papers and Proceedings* 72, 1–11.

Stillman, R., 1983, Examining antitrust policy towards horizontal mergers, *Journal of Financial Economics* 11, 224–240.

Theil, H., 1971, *Principles of Econometrics* (John Wiley, New York).

Vuong, Q., 1989, Likelihood ratio tests for model selection and non-nested hypotheses, *Econometrica* 67, 307–333.

Wier, P., 1983, The costs of antimerger lawsuits: Evidence from the stock market, *Journal of Financial Economics* 11, 207–224.

Zellner, A., 1962. An efficient method of estimating seemingly unrelated regressions and tests for aggregation bias, *Journal of the American Statistical Association* 67, 348–368.