Industrial De-Diversification and Its Consequences for Productivity

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ABSTRACT

Due in large part to intense takeover activity during the 1980s, the extent of American firms' industrial diversification declined significantly during the second half of the decade. The mean number of industries in which firms operated declined 14 percent, and the fraction of single-industry firms increased 54 percent. Firms that were "born" during the period were much less diversified than those that "died", and "continuing" firms reduced the number of industries in which they operated. Using plant-level Census Bureau data, we show that productivity is inversely related to the degree of diversification: holding constant the number of the parent firm's plants, the greater the number of industries in which the parent operates, the lower the productivity of its plants. Hence de-diversification is one of the means by which recent takeovers have contributed to U.S. productivity growth. We also find that the effectiveness of regulations governing disclosure by companies of financial information for their industry segments was low when they were introduced in the 1970s and has been declining ever since.

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In previous research (Lichtenberg and Siegel, 1987, 1989a, 1989b; Lichtenberg and Kim, 1989), we presented evidence that certain types of corporate control transactions during the 1970s and 1980s tended to increase the efficiency of U.S. enterprises. In particular, we showed that the relative (to industry mean) total-factor productivity (TFP) of (1) manufacturing plants involved in ownership changes in the 1970s, (2) plants involved in leveraged buyouts (LBOs) in the 1980s, and (3) airlines involved in mergers during 1970-84, tended to increase in the years following the transaction. We provided a number of reasons why these changes in corporate control resulted in improvements in efficiency. First, we argued that "re-matching" of owners and plants may yield efficiency gains if the "quality of the match" between an owner and plant is heterogeneous and cannot be known with certainty unless the match is made. Second, we demonstrated that ownership changes are associated with substantial reductions in corporate overhead (e.g., the ratio of administrative employment to total employment), and that this represents an important source of productivity gains. Third, both managers' incentives and their opportunities to engage in inefficient behavior may be much lower under an LBO partnership arrangement than they are in a typical publicly-held corporation. Fourth, airline mergers tended to result in significant improvements in capacity utilization (load factor).

This paper examines another means by which changes in corporate control may bring about improvements in operating efficiency: by reducing the extent of industrial diversification, i.e. the number of industries in which a firm operates. Our previous research suggested the existence of the following causal relationship:

$$\text{Control changes} \xrightarrow{(+)} \text{Productivity} \qquad (1)$$

where the (+) above the arrow denotes a positive relationship. We will attempt to establish that the sign of this "reduced form" relationship is positive in part because of the (negative) signs of the "structural" relationships between these two variables and a mediating variable:

$$\text{Control changes} \xrightarrow{(-)} \text{Diversification} \xrightarrow{(-)} \text{Productivity} \qquad (2)$$

In other words, control changes of the 1970s and 1980s led to increases in productivity in part because these changes (unlike the control changes of the earlier postwar era, particularly the late 1960s) reduced the extent of industrial diversification, and diversification is inversely related to productivity.

Our first objective will be to provide empirical support for the second of the two hypotheses indicated in (2) above, the one concerning the effect of diversification on productivity. Several previous papers have examined the effect of diversification on other measures of firm performance, such as profitability, Tobin's q and shareholder wealth. Ravenscraft and Scherer found that "unrelated" lines of business acquired during the conglomerate merger boom of the late 1960s experienced below-average profitability in the 1970s and were often subsequently divested. Wernerfelt and Montgomery found that "narrowly diversified firms do better [i.e., have higher values of q, _ceteris paribus_] than widely diversified firms." Morck, Shleifer, and Vishny found that diversification reduced bidding firms' shareholder wealth in the 1980s, although it failed to do so in the 1970s. However we are not aware of any previous research on the effect of diversification on TFP--output per unit of total input--which is generally regarded by economists as the purest measure of technical efficiency. We will estimate this effect using rich and detailed Census Bureau data on over 17 thousand manufacturing establishments in the year 1980.

Our investigation of the first hypothesis indicated in (2), concerning the effects of (recent) control changes on the extent of diversification, will be based on a different data set, and will be less direct. Using Compustat data, we will describe and analyze changes between January 1985 and November 1989 in the distribution of companies by the number of industries in which they operate. Due to data limitations, control changes won't be explicitly accounted for in this analysis. But because the 1980s was a period of high and accelerating takeover activity--the value of takeover transactions as a fraction of GNP increased from 1.5 percent in 1979 to 4.5 percent in 1986--takeovers are probably responsible for much of the change in the extent of diversification.

In addition to analyzing one of the causes (control changes) and effects (productivity) of diversification, we will also investigate the issue of segmented financial reporting by diversified companies. In the mid-1970s the Financial Accounting Standards Board and the Securities and Exchange Commission began requiring firms to disclose financial data for individual business segments. Again using Compustat data, we will assess the effectiveness of these regulations by examining the time-series of distributions of companies by number of reported segments, and comparing it to the distributions of companies by the "true" number of industries in which they operate.

I. Industrial Diversification and Productivity of Manufacturing Plants

The measure of productivity that we will use is the same as the one employed in our previous analyses of the effects of takeovers and leveraged buyouts on productivity (Lichtenberg and Siegel 1987, 1989b). It is a residual from a production function of the following form, estimated separately by 4-digit SIC industry:

$$\ln VQ_{ij} = \beta_{0j} + \beta_{Lj} \ln L_{ij} + \beta_{Kj} \ln K_{ij} + \beta_{Mj} \ln VM_{ij} + u_{ij} \qquad (3)$$

where VQ denotes the value of production (the value of shipments adjusted for changes in finished-goods and work-in-process inventories); L denotes labor input ("production-worker-equivalent" manhours); K denotes capital input (the "perpetual inventory" estimate of the net stock of plant and equipment); VM denotes the value of materials consumed (materials purchased adjusted for changes in raw-materials inventories): u is a disturbance term; and the subscript ij refers to establishment i in 4-digit industry j.[1] All of the data (with one exception noted below) for this study are for the year 1980. Output and materials are measured in nominal terms because the Census database does not include establishment-specific deflators. It is conventional to assume that output and materials prices do not vary across establishments within an industry, which would imply that the nominal measures are proportional to their real counterparts, although there is some evidence inconsistent with this hypothesis (see Abbott (1988)). Thus the computed residual may be capturing price differences as well as productivity differences. Because eq. (3) was estimated separately by industry, the residual for a given observation measures the percentage deviation of that establishment's TFP from the mean TFP of all establishments in the same industry. By construction, of course, the residuals have a mean value of zero.

A basic premise of our research design is that the industrial <u>structure</u> of a plant's parent firm--measured in terms of the number and industry-distribution of its plants--determines

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[1]This 3-factor Cobb-Douglas production function may be regarded as a local first-order logarithmic approximation to any arbitrary production function. Maddala (1979, p. 309) has shown that, at least within a "limited class of functions...(viz. Cobb-Douglas, generalized Leontief, homogeneous translog, and homogeneous quadratic) differences in the functional form produce negligible differences in measures of multi-factor productivity." This is because these different functional forms differ in their elasticities of substitution (which depends on the second derivatives of the production function) whereas productivity depends primarily on the first derivatives.

the plant's performance (productivity). We assume that the parent's structure is exogenous with respect to the plant's performance. It is possible, however, that the (average) performance of a firm's plants may in the long run influence the firm's industrial structure. Some observers have suggested that it is very profitable firms with large free cash flows that are most likely to engage in diversifying acquisitions. These profitable firms are likely to own plants that are efficient relative to their respective industries (although they may merely own "average" plants in industries with above-average profitability). Thus feedback from plant performance to firm structure might be expected to bias upward the coefficient on a diversification index in a productivity equation.

Our research strategy is to estimate cross-sectional regressions of the plant's productivity residual (RESIDUAL) on several different measures of its parent firm's industrial structure. This may be repesented algebraically by

$$\text{RESIDUAL} = f(\text{STRUCTURE}) \qquad (4)$$

The measures included in the STRUCTURE vector are (1) SINGLE, a dummy variable equal to one if the firm operates only one plant, and otherwise equal to zero; (2) NPLANTS, the total number of manufacturing plants owned and operated by the firm; (3) NINDS, the total number of 4-digit SIC manufacturing industries in which the firm operates; and (4) SAMEIND, the fraction of the firm's plants that operate in the same industry as this plant. Due to the way in which our sample was constructed, there are some problems associated with the measurement of the last three variables. The ultimate source of the data is the 1980 Annual Survey of Manufactures (ASM), which collected data from a sample of approximately 50 thousand manufacturing establishments, out of a population of roughly 350 thousand establishments.[2] Our analysis is based on a nonrandom subset (constructed for our earlier (Lichtenberg and Siegel, 1987) research project) of about 18 thousand of the ASM establishments. All of the establishments in the subset we examined had been in continuous operation and had been included in the ASM sample since at least 1972. Thus the sample is biased towards mature establishments that are themselves large or that are owned by large firms. We calculated NPLANTS simply by counting the number of plants within the subset of 18 thousand with the same parent company identification code as a given plant. We calculated NINDS by counting the number of industries in which these plants primarily operated. Because these counts were based on the subset of 18 thousand establishments rather than on the entire population of 350 thousand establishments, they are subject to measurement error. In particular, they are <u>lower bounds</u>.[3] Although the measurement

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[2]Large establishments (those with greater than 250 employees) are sampled with certainty, and smaller establishments are sampled with probability inversely related to their size.

[3]The downward bias in NPLANTS and NINDS would perhaps have been reduced if we had used data for an ASM year prior to 1978. Beginning in 1978, to reduce the cost of the ASM the Census Bureau switched from sampling with certainty all establishments of large firms to only sampling large establishments with certainty.

Even if they were based on the entire Census of Manufactures, NPLANTS and NINDS would still be truncated due to the omission of error is not of the classical (e.g. normal, i.i.d.) form, one suspects that it would bias the coefficients and t-statistics on these variables towards zero. The variable SINGLE is not subject to measurement error (at least of this kind), since for administrative purposes the Census Bureau records this attribute in the establishment data files. Even if NINDS were not subject to truncation, it would still undoubtedly be a cruder (noisier) measure of firm diversification than the standard Gort-Herfindahl index or the concentric index of Caves, Porter, and Spence (1980).

The performance measure we have chosen--the residual from the production function (3)--is output produced by the plant per unit of total input employed in the plant. Some of the inputs that contribute to the production of a plant's output, however, may not be employed in the plant itself; they may be employed in what the Census Bureau calls "auxiliary establishments." These are establishments

> whose employees are primarily engaged in general and business administration; research, development, and testing; warehousing; electronic data processing; and other supporting services performed centrally for other establishments of the same company rather than for other companies or the general public.[4]
>
> The primary functions of these establishments are to manage, administer, service, or support the activities of the other establishments of the company.[5]

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nonmanufacturing establishments.

[4]U.S. Bureau of the Census (1986, p. A-1).

[5]U.S. Bureau of the Census (1986, p. 2).

Although only 0.4 percent of the entire 3.4 million companies (in all industries) recorded in Census data had at least one auxiliary establishment, in 1982 these establishments accounted for about 7 percent of employment and 10 percent of payroll in the U.S. manufacturing sector.[6] Hence failure to account for auxiliary inputs could result in seriously distorted estimates of plant productivity, and these distortions are likely to be strongly correlated with our measures of firm industrial structure. Fortunately, because we had access to firm-level data on both total employment (TE) and employment in auxiliary establishments (AE), we can control (imperfectly, perhaps) for inputs employed in auxiliaries.[7] Because a given auxiliary establishment typically provides services to a number of production establishments (plants), there is a problem of allocating the auxiliary's inputs across plants. We assume that the ratio of auxiliary inputs dedicated to a plant to the plant's own employed inputs is the same for all of the firm's plants, which implies that inputs employed within the plant understate the plant's "true" total input (including allocated auxiliary inputs) by the ratio AUXSHARE≡AE/TE. Given this assumption,

[6]See Lichtenberg and Siegel (1989a) for a detailed discussion of the role of auxiliary establishments.

[7]Since auxiliary employment data are collected only in Census years, we used values of AE and TE for 1982, the Census year closest to 1980. The time misalignment clearly introduces some error into the correction for auxiliary inputs, although we suspect that firms' relative values of the ratio AE/TE are fairly stable over time.

there are two alternative ways of accounting for auxiliary inputs in our analysis. The first is to "inflate" (some or all of) the plant's recorded input values (e.g., L and K) by multiplying them by (1 + AUXSHARE) prior to calculating the productivity residual via eq. (1). The second is, instead of inflating the input values, to include AUXSHARE as an explanatory variable in the productivity equation. Because the first approach is much more restrictive, and because it isn't clear which inputs should be inflated by (1 + AUXSHARE), we've adopted the latter procedure.

One additional econometric issue deserves our attention. As noted above, we will use a two-step estimation procedure to analyze the effect of diversification on productivity. The first step is to estimate the production function (3) by industry, and to compute the residuals. The second step is to regress these residuals on a vector of explanatory variables. The formulas derived by Neter et al (1985, p. 402) imply that the variance of $RESIDUAL_{ij}$ is $V_{ij} \equiv S_j^2(1 - X'_{ij}(X'_jX_j)^{-1}X_{ij})$ where S_j is the standard error of the residual for industry j; X_j is the design matrix from eq. (3) for industry j; and X_{ij} is the i^{th} row of this matrix (i.e., the row corresponding to the i^{th} plant). Because this variance differs both within and between industries, the disturbances of eq. (4) are heteroskedastic. We will therefore estimate eq. (4) using weighted least-squares (WLS), with weights equal to $V_{ij}^{-1/2}$.

Descriptive statistics for our sample of 17,664 plants are provided in Table 1. Means, standard deviations, and selected quantiles of the variables are presented in the top part of the table. Judging from the quantiles, the distribution of the RESIDUAL appears to be quite symmetric, as we would hope. Only 7 percent of the plants in our sample are "single-unit" plants, i.e. the only plants owned by their parent firms. The mean and median number of plants owned by the parents of our sample plants are 23 and 11, respectively. The mean and median number of manufacturing industries in which the parents operated are 9 and 5, respectively. The distributions of both of these variables are obviously quite skewed, so we will use the logarithms of these variables, rather than their levels, in the remainder of the empirical analysis. The mean (median) value of the fraction of the parent's plants operating in the same industry as a given plant is 45 (31) percent. The sample mean value of AUXSHARE, the ratio of auxiliary employment to total firm employment, is virtually identical to the population (weighted) mean value of 7 percent cited earlier. Although the production function (3) allows for non-constant returns to scale (since the input coefficients aren't constrained to sum to one), we will also control for possible scale effects by including in the regressions a measure of establishment size--total plant employment (PLANTEMP). As Table 1 indicates, this variable is also highly skewed (the mean of 525 is almost double the median), so the log transformation will also be applied to it.

Sample correlation coefficients are shown in the bottom part of Table 1. The absolute values of the correlations among three variables--NPLANTS, NINDS, and SAMEIND--are very high (above 0.8). As we shall now see, this fact is of crucial importance in interpreting our estimates of the effects of parent firm industrial structure on plant productivity.

Weighted least-squares regressions of the plant productivity RESIDUAL on plant and parent firm characteristics are displayed in Table 2. Each column of the table represents a separate regression. The only regressor in the first equation is the variable SINGLE. The coefficient on it indicates that single-unit establishments are, on average, 5.6 percent less productive than multi-unit establishments in the same industry. The difference is highly statistically significant. Part of this difference may be due to the fact that some multi-unit establishments are serviced by auxiliary establishments, whereas (by definition) no single-unit establishments are. In column 2 we attempt to control for auxiliary inputs by including AUXSHARE in the equation; we also include log(PLANTEMP) to allow for scale effects. Including these regressors reduces the magnitude of the SINGLE coefficient, but by only 16 percent, and it remains highly significant. As expected, the coefficient on AUXSHARE is positive and significant, consistent with the view that auxiliary establishment inputs contribute to production establishment output. The positive coefficient on log plant employment is significant but very small, suggesting that there may be very modest economies of scale.

Because the mean values of NINDS for single- and multi-unit establishments are 1 and 9.6, respectively[8], the negative coefficient on SINGLE might give the impression that diversification has a positive effect on productivity: single-unit plants are both less efficient and owned by less-diversified firms than multi-unit plants. This impression is reinforced by the regression in column 3, which includes log(NINDS) as a regressor. Its positive and significant coefficient implies that, among multi-unit establishments, the greater the number of industries in which the parent operates, the higher is plant productivity. So there is an apparent positive relationship between diversification and productivity both <u>between</u> establishment categories (single- vs. multi-unit) and <u>within</u> the multi-unit category.

This apparently positive relationship, however, is completely <u>spurious</u>: it results from failing to control for the number of <u>plants</u> owned by the firm NPLANTS and from the high positive correlation noted above between (the logarithms of) NINDS and NPLANTS. In column 4 we replace log(NINDS) by log(NPLANTS). Its coefficient indicates that productivity increases with the number of plants owned by the firm. The low productivity of single-unit plants may therefore be due to their low value of NPLANTS, not their low value of NINDS. In column 5 we include <u>both</u> of these regressors. The coefficients on both

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[8]This may be inferred from the top of Table 1, since the value of NINDS for all single-unit plants is 1 by definition, and 7 percent of sample plants are single-unit plants.

of these variables are very different from what they were when they were included separately. The coefficient on log(NINDS) becomes negative, almost triples in magnitude and becomes more significant. The coefficient on log(NPLANTS) more than doubles and also becomes more significant. The equation in column 5 reveals that holding constant the number of the parent firm's plants, the greater the number of industries in which the parent operates, the lower the productivity of its plants.

We can get a feeling for the magnitude of these effects by considering the implications of moving "halfway across"--from the .25 quantile to the .75 quantile of--the distributions of these variables. The difference between the .25 and .75 quantile values of the RESIDUAL is .22(=.10-(-.12)). The effect of a *ceteris paribus* decrease in NINDS from its .75 to its .25 quantile value is .049 = -.019 * log(1 / 13), or 22 percent of this productivity difference. The effect of an increase in NPLANTS from its .25 to its .75 quantile value is .065 = .023 * log(34 / 2), or 30 percent of the productivity difference. Of course, in view of the high correlation between NINDS and NPLANTS, the effects of *ceteris paribus* changes in these variables may not be of great practical significance.

It may be useful to offer a slightly different interpretation of the equation in column 5. The equation may be represented as follows: $r = \beta_1 i + \beta_2 p$, where $r \equiv$RESIDUAL, $i \equiv$log(NINDS), and $p \equiv$log(NPLANTS), and we have ignored other terms on the right-hand side for simplicity. This equation may be rewritten in two alternative ways: $r = (\beta_1 + \beta_2)\ i - \beta_2\ (i - p)$, and $r = (\beta_1 + \beta_2)\ p + \beta_1\ (i - p)$. Thus $-\beta_2 = -.023$ may be interpreted as the effect of increasing the number of industries per plant (i - p), holding constant the total number of industries, and $\beta_1 = -.019$ may be interpreted as the effect of increasing this ratio, holding constant the total number of plants.

Above we characterized the parent firm's industrial distribution of plants by NINDS, the number of industries in which it operates. Another attribute of this distribution that may influence a plant's productivity (conditional on NPLANTS) is the fraction of the parent's plants in the same industry. Let NSAME denote the number of parent's plants in the same industry and NOTHER (≡ NPLANTS - NSAME) denote the number in other industries. The productivity-determination equation might be hypothesized to be RESIDUAL = β log (NOTHER + (1 + π) NSAME) + other regressors, where π is the percentage difference between the productivity effect of NOTHER and NSAME. The preceeding equation is nonlinear, but it can be approximated by the linear equation RESIDUAL ≈ β NPLANTS + $\beta\pi$ SAMEIND + other regressors, where SAMEIND ≡ NSAME/NPLANTS is the fraction of plants in the same industry. The ratio of the SAMEIND coefficient to the NPLANTS coefficient may be interpreted as an estimate of π, and the significance of π may be inferred from the t-statistic on the SAMEIND coefficient.

The equation shown in column 6 of Table 2 includes SAMEIND *instead of* log(NINDS). The coefficient on SAMEIND is positive and highly significant. The implied estimate of π is 1.86 (= .026/.014). This implies that a unit increase in the number of plants in the same industry raises a plant's productivity almost three times as much as a unit increase in the number of plants in other industries. The regression in column 7 includes both SAMEIND and log(NINDS) as explanatory variables. The coefficient on SAMEIND is very small and insignificant, and the coefficients on the other regressors are essentially identical to their counterparts in column 5 (although the standard error on the log(NINDS) coefficient increases by a third). It is not surprising that the SAMEIND and log(NINDS) coefficients are not both significant, given the high inverse correlation (-.90) between these variables. The fact that log(NINDS) dominates SAMEIND perhaps signifies that plant productivity depends more on the general extent of parent-firm diversification than it does on the fraction of firm activity in the plant's specific line of business.

II. Changes in the Extent of Industrial Diversification, January 1985 to November 1989

In this section we describe and analyze changes in the extent of industrial diversification between January 1985 and November 1989, the earliest and most recent dates for which this kind of information was available. The data for this section were derived from two editions (corresponding to those dates) of the Standard Industrial Classification (SIC) File, a subset of the Business Information Compustat II file produced by Standard & Poor's Compustat Services, Inc. The SIC file identifies firms' principal products and services by listing up to 90 SIC codes for each company. The SIC codes are derived by Compustat from Annual Reports to Shareholders and from 10-K Reports to the SEC. Our index of diversification will be the same (admittedly crude) one we used in our analysis of the Census data: a simple <u>count</u> of the SIC codes reported for the firm.

Table 3 displays mean values of NSIC (the number of SIC codes) and the number of observations in 1985 and 1989. There were 6505 firms included in the 1985 SIC file, and 7541 firms in the 1989 file. The number of ("continuing") firms present in <u>both</u> files (with a common firm identification (CUSIP) number) was 3829. Thus there were 2676 "deaths" and 3712 "births" between 1985 and 1989. The mean value in 1985 of NSIC for all firms present in that year was 5.46, and the corresponding mean for 1989 was 4.70.[9] Hence the mean declined by .76 (about 14 percent), and this decline is highly statistically significant.

It is interesting to note that the number of <u>firms</u> in the SIC file increased about 16 percent (from 6505 to 7541) between 1985 and 1989, so that the total number of "divisions" (industry-<u>cum</u>-firms) remained almost unchanged (it increased by 2 percent).

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[9] Because the unit of observation here is the firm, whereas in the previous section it was the plant, one would expect the mean value of NSIC to be lower than the previously-reported mean value of NINDS (since firms with higher values of NINDS tend to have more plants); this is indeed the case.

over the course of this period, markets replaced hierarchies as the medium of interaction and exchange among a relatively stable number of divisions.[10]

The last three rows of the table indicate that three distinct factors contributed to the decline in the mean value of NSIC. First, the mean value for continuing firms declined; the decline was only about one-third as great as for all firms (-.27) but was still highly significant. Second, the mean value in 1985 for deaths was substantially higher than the mean value in 1989 for births--4.78 compared to 3.70. Entering firms were much less diversified than exiting firms. Finally, the number of births exceeded the number of deaths.

Because the distributions of companies by NSIC are highly skewed in both years, it may be appropriate to consider changes in the distribution of the logarithm of NSIC rather than in NSIC itself. The mean of ln(NSIC) also declined about 14 percent from 1985 to 1989, from 1.29 to 1.12.

Table 4 provides further evidence of the decline in the extent of diversification, by reporting percentages of companies in 1985 and 1989 with values of NSIC in selected ranges. The fraction of "single-industry" companies--those with only one SIC code--increased by 54 percent, from 16.5 to 25.4 percent. The fraction of companies that were highly diversified--those with values of NSIC in excess of 20, say--declined by 37 percent, from

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[10] The distinction between markets and hierarchies was developed by Williamson (1975).

3.5 to 2.2 percent.

The results of the previous section imply that the reduction in the extent of diversification between 1985 and 1989 was a source of productivity growth during that period. One might attempt to estimate the productivity contribution of de-diversification simply by multiplying the change in the mean value of ln(NSIC) by the coefficient on ln(NINDS) in the productivity equation. This yields an estimate of (1.12 - 1.29) * -0.19 = .0032, or 0.3 percentage points. This does not appear to be very large, but the estimate may be distorted for several reasons. First, as noted above, due to errors in measuring NINDS, the coefficient on ln(NINDS) is probably biased towards zero. Second, the unit of observation in the regression analysis was the plant, whereas our estimate of the mean change in diversification between 1985 and 1989 was based on firm-level data. Third, the productivity regressions were based on manufacturing establishments only, while nonmanufacturing companies were also included in the NSIC calculations. The estimate of 0.3 percentage points is much larger relative to typical nonmanufacturing productivity growth rates than it is relative to manufacturing growth rates.

As shown in the previous section, the cross-sectional correlation between a plant's parent's number of industries (NINDS) and its number of plants (NPLANTS) is positive and very high (.94). One might therefore expect that the mean value of NPLANTS would have declined, along with the mean value of NSIC, between 1985 and 1989: firms became smaller as they de-diversified. If so, then de-diversification might not have increased productivity, since NPLANTS has a positive partial effect on plant productivity. Since we lack time-series data on NPLANTS, to investigate this possibility we will use an alternative measure of firm size, total firm employment (FIRMEMP), which is available for the subset of firms included in the Compustat Annual Industrial File. We calculated the logarithm of the ratio of the value of FIRMEMP in 1987 (the most recent year for which fairly complete data were available) to its value in 1984, for a sample of 1562 continuing firms with nonmissing values in both years. The mean value of this variable was positive (=.047) and significantly different from zero (t = 5.9).[11] Thus although, as shown above, continuing firms became decreasingly diversified (albeit less so than all firms), such firms were apparently not shrinking during roughly the same period. Due to data limitations, this test is not conclusive, but it does suggest that the productivity impact of declining diversification was not offset by the impact of declining firm size.

III. Evaluation of the Effectiveness of FASB/SEC Regulations Concerning the Disclosure of Financial Information by Industry Segment

In the mid-1970s, the Financial Accounting Standards Board

[11] We eliminated 331 "outliers" with absolute values of this variable greater than 1. Including them raised the mean to .110.

(FASB) issued Statement of Financial Accounting Standards (SFAS) No. 14, "Financial Reporting for Segments of a Business Enterprise". This Statement required firms to report financial data (for fiscal years ending after December 15, 1977) for industry segments which accounted for 10 percent or more of the consolidated firm's sales, operating profits, or assets. SFAS No. 14 defined an industry segment as "a component of an enterprise engaged in providing a product or service, or a group of related products or services primarily to unaffiliated customers (i.e., customers outside the enterprise) for a profit." Since this definition is quite general and perhaps vague, firms had considerable latitude in the extent and nature of segmentation in their financial reporting. When, or soon after, SFAS No. 14 was issued, the Securities and Exchange Commission (SEC) issued Regulation S-K, "Instructions Regarding Disclosure," which required that the information prescribed by SFAS No. 14 be included in SEC Form 10-K.[12]

The SEC was not the only government agency to respond (with a lag) to the increase in industrial diversification that

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[12] Both SFAS No. 14 and Regulation S-K required disclosure of: sales net, operating profit (loss), and identifiable assets. FASB No. 14 also required disclosure of: depreciation, depletion, and amortization; capital expenditures; equity in earnings; investments in equity; the name and amount of sales to each customer, and identification of each industry segment or segments selling to principal customers. Regulation S-K also required disclosure of: order backlog; research and development (company- and customer-sponsored); employees; the amount of revenue accounted for by major products or groups of related products or services; and the names of customers from whom more than 10 percent of consolidated revenues are derived.

occurred in the late 1960s by requiring firms to disclose financial data for industry segments. The Federal Trade Commission (FTC) also did so (for very different reasons) by instituting its Line of Business (LB) Program[13]. This program, authorized by Section 6 of the Federal Trade Commission Act (15 U.S.C. 46), required firms to disaggregate their financial performance statistics into a maximum of 261 three- or four-digit Standard Industrial Classification (SIC) manufacturing industry categories. However, whereas SFAS No. 14/Regulation S-K required firms to disclose segmented financial data to the public, only sworn officers and employees of the FTC were allowed access to the LB reports.[14] For a number of reasons, including reluctance of firms to respond to the survey[15] and budgetary pressures at the FTC, the survey was administered in only four years, 1974-1977.

In contrast, Regulation S-K and SFAS No. 14 remain in effect to this day. The purpose of this section is to assess the effectiveness of these regulations by examining data on the extent of industry segmentation in company financial reporting, and by comparing these to data on the "true" extent of industrial

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[13] See Federal Trade Commission, Bureau of Economics, <u>Report on the Line of Business Program</u>.

[14] Also, Regulation S-K applied to all publicly-held corporations, while fewer than 500 of the nation's largest manufacturing corporations were required to file LB reports.

[15] About one-third of the 345 companies ordered to file the first survey were parties to litigation challenging the legality of the survey.

diversification. First we will describe the data upon which our analysis is based. Next we will present time-series evidence on the extent of segmentation in reporting. We will then consider alternative potential explanations for this evidence.

The data for this section are derived from the Industry Segment file, another subset of the Business Information Compustat II file used in the previous section. The ultimate sources of the data in the Industry Segment file are also Annual Reports to Shareholders and 10-K Reports to the SEC. Data for up to 10 segments per company are reported in the Industry Segment file, although as we shall see below the fraction of firms with 10 (or more) reported segments never exceeds 0.2 percent. The file is longitudinal, containing up to seven fiscal years of information for each company. If data for a particular fiscal year are missing, no data for previous fiscal years are reported. Thus the file is subject to a kind of censoring: past data are not available for firms that have dropped out of the sample.

As part of the file documentation, Compustat provides a Data Availability Report (DAR). Among other things, the DAR reveals how many of the companies present in the file have N reported industry segments (N = 1,...,10) in each fiscal year. Thus, one can generate for each year a frequency distribution of companies, by number of reported industry segments.

We had access to DARs corresponding to two different "editions" of the Industry Segment file. The first DAR is for the January 1984 edition of the file, and contains fairly complete data for fiscal years 1977-82; the second is for the August 1988 edition, and contains fairly complete data for 1981-87. Thus we can generate an annual time series of distributions of companies by N, beginning in 1977--about the time the regulations went into effect--and ending in 1987.[16]

Data on the percent of companies in the Industry Segment file reporting at least N industry segments, by year, are presented in Table 5. A comparison of the data for the year 1985 in Tables 4 and 5 reveals that the extent of segmentation in reporting is very low, relative to the true extent of industrial diversification. In 1985, the fraction of companies with more than one SIC code was 83.5 percent, whereas the fraction of companies with more than one reported segment was only 29.7 percent.

The data in Table 5 also reveal a sharp, steady, virtually monotonic decline over time in the percent of firms with at least N reported segments, for every value of N. In 1977, about half of the included companies reported at least two industry segments, and a third reported at least three. By 1987, these fractions had declined to about one-quarter and one-seventh, respectively. Moreover, the ratio of the 1987 to the 1977 percentage tends to decline as N increases: the relative decline in segmentation is greatest at the "upper tail" of the distribution.

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[16] Unfortunately, we lack data for the "pre-regulatory" years prior to 1977.

We will consider three alternative potential explanations for the steady decline in segmentation in reporting: (1) a decline in the true extent of diversification; (2) data censoring; and (3) declining enforcement of, and compliance with, the spirit (although not the letter) of the disclosure regulations.

In the previous section we established that the true extent of industrial diversification, as measured by the number of SIC codes, declined significantly from 1985 to 1989. The data suggest that a decline in true diversification explains part of the decline in the reported number of segments between 1977 and 1987, but only a small part. Table 6 juxtaposes some of the data for 1985 and 1989 from Table 4 and some of the data for 1982 and 1987 from Table 5; the latter two years span the five-year period closest to the 1985-89 period analyzed in the previous section for which we have segment data. Each line of the table shows the percent of companies with values of NSIC greater than X in 1985 and 1989, and the percent of companies with NSEG (the number of reported industry segments) greater than Y in 1982 and 1987. The values of X and Y were chosen so that the 1985 percentage for NSIC was roughly equal to the 1982 percentage for NSEG. The table reveals that there were much greater relative declines in the NSEG percentages than there were in the NSIC percentages. For example, the percent of companies with values of NSIC > 5 declined 20 percent from 1985 to 1989, from 12.3 to 9.9, while the percent of companies with values of NSEG > 3 declined 46 percent from 1982 to 1987, from 12.5 to 6.8.

As noted above, due to Compustat's procedures for processing the file (i.e., including companies with missing data for early years but not those with missing data for middle and late years), the data are subject to censoring. The apparent decline in the extent of segmentation might be an artifact of this censoring. For example, firms for which only recent years' data are available might be hypothesized to be newer, smaller firms, with fewer industry segments than large, established, continuing firms. (On the other hand, firms that are entirely absent from the file because they have "dropped out" of the sample also probably had few segments; this would tend to offset the bias.) Fortunately, because we have two different "snapshots" (DARs), taken almost five years apart, of two fiscal years (1981 and 1982), we can assess the extent of censoring-induced bias simply by comparing the two snapshots of the same year. Substantial differences between the two snapshots of the same year would suggest that the bias issue is an important one. Table 5 shows distributions of companies by NSEG in 1982, as reported in both the 1984 and 1988 DARs.[17] Although the distribution from the later DAR lies everywhere above that from the earlier DAR (consistent with the presence of censoring-induced bias), the two distributions are very similar. Moreover, the later 1982

[17] The extent of sample attrition is suggested by the fact that the 1984 report contained 1982 data for 5651 companies, whereas the 1988 report contained 1982 data for only 4313 companies.

distribution is almost uniformly below the 1981 distribution, and the earlier 1982 distribution is almost uniformly above the 1983 distribution.[18] Data censoring therefore appears to be responsible for a negligible fraction of the total estimated decline in the extent of reported segmentation.

Instead, it appears that the change in reporting reflects a decline in enforcement of, and compliance with, the spirit, if not the letter, of Regulation S-K. That there may have been a decline in enforcement during the 1980s is not too surprising, since it is well known that the enforcement staffs of many federal regulatory agencies were drastically reduced during the Reagan Administration. However SEC expenditures increased in real terms during the 1980s, from $84 million in 1980 to $94 million in 1987 and $111 million in 1988.[19] Moreover the decline in segmentation clearly preceded the Reagan Administration; it began, in fact, as soon as the regulation went into effect (if not before). The immediate and uninterrupted decline in segmentation may simply reflect the normal time-path of response of economic agents to the issuance of poorly-defined regulations. In this context, one might interpret the time-series data of Table 5 as being generated by a process of diffusion of noncompliance (and nonenforcement) behavior across the population

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[18] A comparison of the "early" and "late" distributions for 1981 yields similar results.

[19] All figures are in constant 1982 dollars and are reported in <u>Regulation</u>, 1988 No. 3, p. 12. The 1988 figure is estimated.

of firms.

IV. Summary and Conclusions

During the quarter century following the Second World War, U.S. industrial enterprises became increasingly diversified. Rumelt[20] has estimated that the percentage of diversified companies in the Fortune 500 more than doubled from 1949 to 1974, from under 30 percent to over 60 percent. The greatest increase in the extent of diversification apparently occurred during the conglomerate merger wave of the late 1960s, which Golbe and White (1988) have shown to be the most intense period of merger and acquisition (M&A) activity between 1940 and 1985.

Previous studies have demonstrated that diversification tends to have a negative impact on financial variables such as profitability, Tobin's q and (in recent years) stock prices. We have provided evidence consistent with the view that diversification has a negative effect on technical efficiency, i.e. on total-factor productivity. The effect of diversification on efficiency might be regarded as an important, if not the main, underlying mechanism by which diversification influences financial variables.

Our analysis, based on Census Bureau data for over 17 thousand plants in 1980, indicated that holding constant the number of the parent firm's plants (and other variables), the greater the number of industries in which the parent firm

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[20] Cited by Bhide (1989, p. 53).

operates, the lower the productivity of its plants. This suggests that the conglomerate merger boom of the late 1960s may have contributed to the slowdown in U.S. productivity growth which began at or slightly after that time.

If diversification is bad for productivity, and therefore for profitability, why did managers pursue aggressive diversification strategies in the late 1960s? One possible explanation is that managers were interested in maximizing shareholder wealth but that they miscalculated, and expected diversifying acquisitions to yield profitable synergies. An alternative explanation is in the spirit of Jensen's free cash-flow theory. Firms were generating large cash flows, their managers preferred using these cash flows to finance acquisitions to paying dividends to shareholders, and the latter were unable to force managers to do so. Due to vigorous antitrust enforcement, managers were unable to acquire firms in the same line of business, which would have been both technically efficient and highly profitable, although not necessarily socially desirable. Therefore firms acquired business units in unrelated industries, even though they knew little about these businesses and were unlikely to be able to manage them efficiently.[21]

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[21] Wernerfelt and Montgomery offer another explanation of why firms may be prompted to diversify, even if diversification reduces the firm's profitability. They argue that firms may have excess capacity of less-than-perfectly marketable factors, and that the marginal returns to these factors declines as the firm diversifies beyond the first industry chosen.

The extent of industrial diversification probably peaked in the early 1970s. As Ravenscraft and Scherer have documented, by the mid-1970s conglomerate firms began to divest the unrelated (to their primary industry) and unprofitable lines of business they had acquired during the 1960s. A substantial fraction of the corporate control transactions of the 1970s were divestitures of previously-acquired units.

But much larger declines in the extent of diversification probably occurred in the 1980s. The rate of business ownership change was much higher in the 1980s than it had been in the 1970s.[22] Deregulation, intensified foreign competition, junk-bond financing, and relaxed antitrust enforcement may have contributed to this increase in takeover activity. Moreover, the <u>nature</u> of corporate control transactions changed in the 1980s. Hostile, "bust-up" takeovers undertaken by "corporate raiders", along with leveraged buyouts (which are frequently followed by asset sales), accounted for a rapidly growing share of overall takeover activity. The size of the average and largest takeover targets also increased dramatically during the 1980s.

Using Compustat data, we have shown that the extent of industrial diversification declined significantly during the

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[22] Unpublished Census Bureau data indicate that the average annual rate of ownership change among fairly large manufacturing plants increased from 2.3 percent during 1973-79 to 4.2 percent--an 80 percent increase--during 1979-86. Moreover, the <u>lowest</u> annual rate in the second period (3.3 percent in 1979-80) was greater than the <u>highest</u> annual rate in the first period (3.2 percent in 1973-74).

second half of the 1980s. The mean number of industries in which firms operated declined by 14 percent from January 1985 to November 1989. Two factors contributed to this decline: firms that were "born" during this period were much less diversified than those that "died", and "continuing" firms reduced the number of industries in which they operated. The fraction of companies that were highly diversified--operating in more than 20 industries--declined 37 percent, and the fraction of single-industry companies increased 54 percent. The apparent acceleration in the rate of de-diversification from the 1970s to the 1980s contributed to the acceleration in the rate of U.S. productivity growth, but it is difficult to determine the magnitude of this contribution.

We have also examined another issue related to industrial diversification: the effectiveness of FASB and SEC regulations concerning company disclosure of financial information for its industry segments. Our findings indicate that, because firms are free to define industry segments as they see fit, the effectiveness of these regulations was low when they were introduced in the 1970s and has been declining ever since. In 1985, only 30 percent of the companies in Compustat's Industry Segment file reported data for more than one industry segment, whereas 84 percent were truly multi-industry firms. Moreover the extent of industry segmentation in financial reporting has declined much faster than the extent of true diversification: between 1977 and 1987, the fraction of companies reporting data for at least two industry segments declined from one-half to one-quarter. This is unfortunate because appropriately segmented financial data for diversified firms are necessary, or at least highly useful, for both economic policymaking and for economic and financial research.[23]

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[23] Lichtenberg and Siegel (1989c) have shown that segmented data permit more efficient estimates of companies' total-factor productivity growth and of the rate of return to research and development investment than consolidated company data.

TABLE 1

Descriptive Statistics for Sample of 17,664 Plants

STATISTIC	RESIDUAL	SINGLE	NPLANTS	NINDS	SAMEIND	AUXSHARE	PLANTEMP
Mean	0	.07	23	9	.45	.07	525
Std. dev.	.19	--	29	11	.38	.14	1107
Quantiles:							
.05	-.31	--	1	1	.03	0	45
.25	-.12	--	2	1	.11	0	144
.50	-.01	--	11	5	.31	.04	284
.75	.10	--	34	13	1	.10	517
.95	.31	--	82	28	1	.28	4562
Correl. Coeffs. *							
SINGLE	-.06						
NPLANTS	.08	-.33					
NINDS	.07	-.35	.94				
SAMEIND	-.06	.39	-.81	-.90			
AUXSHARE	.06	-.15	.28	.25	-.25		
PLANTEMP	.03	-.05	.04	.07	-.06	-.01	

* Log transformation applied to NPLANTS, NINDS, and PLANTEMP.

TABLE 2

Weighted Least-Squares Regressions of
Plant Productivity Residual on
Plant and Parent-Firm Characteristics
(t-statistics in parentheses)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
SINGLE	-.056 (8.66)	-.047 (7.23)	-.037 (5.42)	-.033 (4.85)	-.037 (5.43)	-.038 (5.51)	-.038 (5.39)
AUXSHARE		.094 (7.95)	.082 (6.75)	.072 (5.88)	.069 (5.67)	.073 (5.94)	.069 (5.67)
log(PLANTEMP)		.005 (3.60)	.005 (3.33)	.005 (3.44)	.006 (3.90)	.005 (3.68)	.006 (3.91)
log(NINDS)			.007 (4.61)		-.019 (5.10)		-.018 (3.62)
log(NPLANTS)				.009 (7.14)	.023 (7.45)	.014 (7.35)	.023 (7.25)
SAMEIND						.026 (3.59)	.003 (0.28)
Intercept	-.003 (1.94)	-.040 (4.80)	-.048 (5.64)	-.057 (6.61)	-.064 (7.30)	-.082 (7.40)	-.066 (5.54)

TABLE 3

MEAN NUMBER OF SIC CODES IN 1985 AND 1989
(Standard error of mean in parentheses)

Companies included	1985	1989	Change
All companies	5.46 (.075)	4.70 (.061)	-0.76 (.048)
Continuing companies	5.94 (.103)	5.67 (.097)	-0.27 (.063)
Births	---	3.70 (.069)	---
Deaths	4.78 (.104)	---	---

Note: There were 6505 companies in 1985 and 7541 companies in 1989, 3829 continuing companies, 3712 births, and 2676 deaths.

Source: Author's calculations based on January 1985 and November 1989 Business Information Compustat II SIC files.

TABLE 4

PERCENT OF COMPANIES WITH 1985 AND 1989 VALUES OF NSIC IN SELECTED RANGES

	Percent of companies with NSIC in range in:	
Range	1985	1989
NSIC = 1	16.5	25.4
NSIC LE 2	35.4	43.6
NSIC LE 3	50.3	57.4
NSIC GT 5	31.1	26.0
NSIC GT 10	12.3	9.9
NSIC GT 20	3.5	2.2
NSIC GT 30	1.3	0.8

Note: These calculations are based on all 6505 observations in 1985 SIC file and all 7541 observations in 1989 SIC file.

TABLE 5

PERCENT OF COMPANIES IN COMPUSTAT INDUSTRY SEGMENT FILE
REPORTING AT LEAST N INDUSTRY SEGMENTS, BY YEAR, 1977-1986

	[-----Jan. 1984 Report------]						[-----Aug. 1988 Report------]					
N	77	78	79	80	81	82	82	83	84	85	86	87
2	49.7	46.0	44.9	42.6	40.3	37.4	37.9	35.3	33.2	29.7	27.4	27.3
3	32.5	29.5	27.9	26.3	24.3	22.0	23.0	21.0	19.3	16.7	15.0	14.8
4	17.3	15.6	14.9	13.6	12.9	11.6	12.5	11.4	9.9	8.0	7.1	6.8
5	8.2	7.4	7.0	6.5	5.9	5.3	5.8	5.0	4.1	3.2	2.8	2.8
6	3.2	3.0	3.0	2.8	2.5	2.2	2.5	2.2	1.6	1.3	1.2	1.1
7	1.4	1.3	1.4	1.3	1.1	1.0	1.1	0.9	0.8	0.6	0.5	0.5
8	0.6	0.5	0.6	0.7	0.6	0.4	0.5	0.4	0.4	0.3	0.2	0.2
9	0.3	0.3	0.3	0.3	0.2	0.2	0.3	0.2	0.2	0.1	0.1	0.1
10	0.2	0.2	0.2	0.2	0.1	0.1	0.1	0.1	0.1	0.0	0.0	0.0
Total Number of Firms	3811	4460	4660	5034	5388	5651	4313	4775	5224	6100	6735	6135

TABLE 6

PERCENT OF FIRMS WITH 1985 AND 1989 VALUES OF NSIC GREATER THAN X (X = 5, 10, 20, 30) AND WITH 1982 AND 1987 NSEG VALUES GREATER THAN Y (Y = 1, 3, 5, 6)

	(1)	(2)	(3)		(4)	(5)	(6)
NSIC range	Percent of companies in NSIC range: 1985	1989	Ratio (2)/(1)	NSEG range	Percent of companies in NSEG range: 1982	1987	Ratio (5)/(4
NSIC > 5	31.1	26.0	0.84	NSEG > 1	37.9	27.3	0.72
NSIC > 10	12.3	9.9	0.80	NSEG > 3	12.5	6.8	0.54
NSIC > 20	3.5	2.2	0.63	NSEG > 5	2.5	1.1	0.44
NSIC > 30	1.3	0.8	0.62	NSEG > 6	1.1	0.5	0.45

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