Chapter 16

INTER-INDUSTRY STUDIES OF STRUCTURE AND PERFORMANCE

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Contents

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1. Introduction

This chapter is concerned with inter-industry studies of the relations among various measures of market structure, conduct, and performance.[1] Prior to the seminal work of Bain (1951, 1956), most empirical research in industrial organization involved detailed case studies of particular industries. These were time-consuming, involved a great deal of subjective judgement, and covered only a small sample of industries, in many of which antitrust litigation had made data available. Bain's inter-industry, cross-section approach seemed to make possible rapid and objective analysis of large samples of markets. Research interest accordingly shifted from industry studies to inter-industry work during the 1960s.

In a comprehensive survey written at the start of the 1970s, Weiss (1974) discussed 46 cross-section studies of the correlates of seller concentration. Ten years later, Gilbert (1984) found 45 such studies of the U.S. banking industry alone. But during this same decade a number of critics effectively challenged the data and methods used in inter-industry research, as well as the conventional interpretation of its findings. Interest shifted to work on the theory of imperfectly competitive markets and, more recently, to econometric industry studies employing formal models of conduct. Inter-industry studies are now out of fashion.

While some feel that fashion is unjust because cross-section research can reveal the structural parameters that determine market conduct and performance, others contend that the cross-section approach is inherently incapable of producing anything useful. In the next section I argue for an intermediate position: cross-section studies rarely if ever yield consistent estimates of structural parameters, but they can produce useful stylized facts to guide theory construction and analysis of particular industries. They typically fail to be persuasive when they attempt to do much more than this. Inter-industry research can complement industry studies by describing robust relations that hold across large samples of markets.

Cross-section studies also fail to be persuasive when they ignore serious measurement problems. Section 3 considers some of these, focusing on the problem of measuring profitability. Again I take an intermediate view: these problems deserve to be taken seriously but, if handled sensibly, they are not so severe as to render cross-section work valueless.

[1] Other chapters in this Handbook discuss related work on the dynamics of pricing behavior (Chapter 15 by Dennis Carlton), technical change (Chapter 18 by Wesley Cohen and Richard Levin), and international comparisons (Chapter 21 by Richard Caves). Chapter 17 by Timothy Bresnahan considers single-industry studies concerned with many of the issues treated here.

Sections 4–6 discuss the main empirical regularities that have been uncovered in inter-industry research. The discussion is organized by dependent variable: Section 4 describes studies that attempt to explain differences in profitability, Section 5 considers studies of prices and costs, and Section 6 discusses studies of concentration, advertising intensity, and conduct-related variables. Section 7 contains a few concluding remarks.

The literature discussed here is enormous, and this chapter is inevitably incomplete despite its length. The reference list at the end is biased toward recent works (and full, book-length presentations) in which earlier contributions are discussed. A number of previous surveys treat some topics in more depth than is possible here.[2]

2. Method and interpretation

2.1. *Long-run equilibria and the endogeneity problem*

The usual presumption in cross-section work in all fields of economics is that observed differences across observations reflect differences in long-run equilibrium positions.[3] Thus, for instance, cross-section studies of demand are usually interpreted as producing estimates of long-run elasticities.

In general, in order to use cross-section data to estimate long-run relations, deviations from long-run equilibrium must be uncorrelated with the independent variables employed. If this strong requirement (discussed further below) is satisfied, and if theoretically sound structural equations can be formulated and identified, simultaneous equations techniques can be employed to yield consistent estimates of long-run structural parameters. In order to estimate any structural equation consistently, one must generally have at least as many available instrumental variables as there are variables on the right-hand side of the equation. Instrumental variables must be exogenous – that is, uncorrelated with the structural residual. If an equation has K endogenous variables and L exogenous variables on the right-hand side, consistent estimation requires at least K additional instrumental variables that (a) can be excluded from the equation on theoretical grounds and (b) are correlated with one or more of the included

[2] Useful surveys of portions of the cross-section literature include Weiss (1971, 1974), Ferguson (1974), Jacquemin and deJong (1977), Comanor and Wilson (1979), Hay and Morris (1979, chs. 7 and 12), Scherer (1980, chs. 4 and 9), Brozen (1982), Geroski (1983), Curry and George (1983), Gilbert (1984), Waterson (1984, ch. 10), and Caves (1985).

[3] For general discussions of the cross-section approach, see Phillips (1976), Cowling (1976), Caves, Porter, and Spence (1980, ch. 1), Sawyer (1982), and Donsimoni, Geroski and Jacquemin (1984), in addition to the surveys cited in footnote 2, above. Caves and Pugel (1980) provide a nice discussion of the long-run equilibrium presumption and use it explicitly to structure their analysis.

endogenous variables. I now argue that the instruments necessary for consistent estimation are rarely available in inter-industry empirical work in industrial organization.

Inter-industry studies in industrial organization are part of an enterprise that seeks tools, based on either deductive or inductive analysis, that permit one to make useful predictions about real markets based on relatively stable, observable variables. These quantities, which together comprise *market structure*, are loosely divided into two sets. *Intrinsic* structural variables [called *basic conditions* by Scherer (1980)] are more or less completely determined by the nature of the product and the available technologies for production and marketing. Other elements of market structure are *derived* in that they may reflect government policy, business strategies, or accidents of history, as well as the relevant intrinsic variables. List of derived structural variables usually include seller concentration, conditions of entry, buyer concentration, and product differentiation.

In any complete market model, such as the textbook models of monopoly and competition, market structure determines *market conduct* – the behavioral rules followed by buyers, sellers, and potential entrants to choose the variables under their control. *Market performance* is assessed by comparing the results of market conduct to first-best ideals, such as perfect competition, or feasible alternatives.

Most of the cross-section literature has been concerned with the effects of intrinsic structural variables on derived structure and with the effects of structure as a whole on conduct and performance. But, except in textbook competitive markets, derived market structure is clearly affected by market conduct in the long run. Mergers and investments alter seller concentration; marketing strategies may affect product differentiation; the attractiveness of entry depends on the actual and expected conduct of established sellers. And, though the linkages may be looser, intrinsic structure is also affected by conduct in the long run. Invention and innovation can change the nature of the product and the available technologies. Industry-specific aspects of government policy are generally affected by industries' lobbying and other political activity and may also be affected by observed performance.

Thus, in the long-run equilibria with which cross-section studies must be primarily concerned, essentially all variables that have been employed in such studies are logically endogenous. This means that there are in general no theoretically exogenous variables that can be used as instruments to identify and estimate any structural equation. (Even if one is willing to argue that intrinsic structure is approximately exogenous, one has only a small number of potential instrumental variables. And these are difficult to observe because actual characteristics of existing firms or plants are not determined only by intrinsic structure.) Moreover, recent theoretical work emphasizes the complexity of market conduct and its determinants and thus makes it difficult to argue strongly for the exclusion of any variable from any structural equation.

Several authors, beginning with Strickland and Weiss (1976), have estimated three-equation models, with profitability, advertising and concentration all treated as endogenous.[4] An examination of some of the specification and exogeneity assumptions made in a leading example of this approach makes the basic endogeneity problem clear. Martin (1979a) needs two instruments to identify and estimate his profitability equation. He employs a durable good dummy variable and lagged values of concentration and profitability. Given the long-run equilibrium presumption and the likelihood that departures from long-run equilibrium are serially correlated, it seems unlikely that lagged endogenous variables are generally valid instruments in cross-section studies. He treats the ratio of imports to sales as exogenous, even though high domestic prices should attract imports [Geroski (1982), Caves (1985)]. And he also treats as exogenous three variables that depend on the technologies actually employed by the industry's firms. Since these depend on seller conduct as well as the menu of available technologies, they are unlikely to be valid instruments.

Martin excludes the three technology-based variables mentioned above from his structural equation for advertising intensity, even though it is unclear in theory why marketing decisions should not depend on production technology. He uses these three variables, along with a regional markets dummy and lagged profitability and concentration, as instruments. The validity of all these seems suspect, with the possible exception of the regional dummy. Similar problems affect Martin's concentration equation.

Some authors have treated additional variables as endogenous.[5] The longer the list of endogenous variables in any model, the more difficult it is to obtain valid instruments from available data. It seems very unlikely that the endogeneity problem posed here can be solved by more elaborate model specifications. Nor can this problem be simply dismissed by the observation that least-squares and simultaneous equations methods generally yield very similar estimates. The relation between these estimates is entirely determined by the set of variables used as instruments.[6] And specification tests of the sort employed by Geroski

[4]Martin (1979a, 1979b) has pointed out that the original Strickland–Weiss (1976) model is not identified, even if one assumes all their judgements about exogeneity are correct.

[5]Studies using simultaneous equations methods include Comanor and Wilson (1974), Geroski (1982), Marvel (1980), Caves, Porter and Spence (1980), Martin (1983), and Connolly and Hirschey (1984).

[6]But the general similarity between the coefficient estimates produced by least squares and simultaneous equations techniques in cross-section studies does present something of a puzzle. (I am indebted to Jerry Hausman for posing this problem and for suggesting the development that follows.) Consider the simplest possible structural equation: $y = \beta x + \varepsilon$, where y, x, and ε have mean zero, and β is a constant. Suppose that x is correlated with ε, as is a proposed instrumental variable, z. Let β_o and β_i be the ordinary least squares and instrumental variables estimates of β, respectively. Then a

(1982), Connolly and Hirschey (1984), and others to test for endogeneity are valid only if one has available a sufficient number of instruments known a priori to be valid.

Consistent structural estimation is possible without instrumental variables in recursive systems,[7] but arguments for recursivity are rarely made in this literature. Similarly, panel data, in which a set of firms or industries is observed over time, can yield consistent structural estimates if an explicit model of disequilibrium behavior is employed. But this has rarely been done; almost all inter-industry studies have had only a cross-section dimension, and studies using panel data have generally had a non-structural, descriptive focus.

2.2. Design and interpretation

Even if cross-section studies in industrial organization generally can only describe relations among long-run equilibrium values of endogenous variables, such studies can make a contribution. But they should be designed, executed, and interpreted with due regard for their limitations.

A simple example will help to structure the discussion. Consider a competitive market in which data are available only on price, P, and quantity, Q, both endogenous variables. Suppose, further, that it is considered reasonable to work with linear approximations to the supply and demand curves:

$$Q^{s} = a + bP + e, \tag{2.1a}$$

$$Q^{d} = \alpha - \beta P + \varepsilon \tag{2.1b}$$

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bit of algebra shows that the probability limit of $(\beta_i - \beta_o)/\beta_o$ is given by

$$\left[(1 - R_o^2)/R_o^2\right]^{1/2}\left[(\rho_{z\varepsilon}/\rho_{zx}) - \rho_{x\varepsilon}\right]/\left[1 - (\rho_{x\varepsilon})^2\right]^{1/2},$$

where R_o^2 is the least-squares R^2, and the ρ's are correlations. In time-series studies with trendy variables, ρ_{xz} and R_o^2 will be close to one, so that even if z is a terrible instrument (so that $\rho_{z\varepsilon}$ and $\rho_{x\varepsilon}$ are approximately equal), the difference between least squares and instrument variables estimates will generally be small. In cross-section data, however, ρ_{xz} and R_o^2 are generally well below one, so that large changes in coefficients would be expected to be the rule even with poor instruments.

[7]Suppose the structural equation for some variable, y, involves a set of independent variables, X, some of which are endogenous, and a disturbance term, ε. This equation can be consistently estimated by least squares as long as y does not enter the structural equations for the endogenous elements of X *and* ε is uncorrelated with the disturbances of those equations. That is, y must not affect X directly, and the unobservable variables determining X and y must be uncorrelated.

where e and ε summarize the effects of the exogenous variables affecting supply and demand, respectively. Working with the reduced form of (2.1), it is easy to see that the variances of price and quantity and the covariance between them implied by this model are as follows:

$$(b+\beta)^2\sigma_{PP} = \sigma_{\varepsilon\varepsilon} + \sigma_{ee} - 2\sigma_{e\varepsilon}, \tag{2.2a}$$

$$(b+\beta)^2\sigma_{QQ} = b^2\sigma_{\varepsilon\varepsilon} + \beta^2\sigma_{ee} + 2b\beta\sigma_{e\varepsilon}, \tag{2.2b}$$

$$(b+\beta)^2\sigma_{PQ} = b\sigma_{\varepsilon\varepsilon} - \beta\sigma_{ee} + (\beta - b)\sigma_{e\varepsilon}. \tag{2.2c}$$

These three equations in five unknowns (b, β, σ_{ee}, $\sigma_{\varepsilon\varepsilon}$, and $\sigma_{e\varepsilon}$) cannot generally be solved for unique estimates of b or β. But this does not mean that data on price and quantity provide no useful information.

In early studies of the demand for agricultural products, for instance, it was argued that σ_{ee} was much larger than $\sigma_{\varepsilon\varepsilon}$. In this case least squares estimation of (2.1b) yields an approximately consistent estimate of β. Alternatively, suppose that one feels that $\beta = 0$ but observes $\sigma_{PQ} < 0$. Equation (2.2c) shows immediately that $\rho_{e\varepsilon} = \sigma_{e\varepsilon}/[\sigma_{ee}\sigma_{\varepsilon\varepsilon}]^{1/2}$ must be positive. The plausibility of this implication can perhaps be evaluated on the basis of theory and evidence from other contexts. More formally, with $\beta = 0$, equations (2.2) can be solved for $\rho_{e\varepsilon}$ as a function of σ_{PP}, σ_{QQ}, σ_{PQ}, and b. Using this relation, and allowing for sampling variability, one can ask whether plausible values of the elasticity of supply are consistent with plausible values of $\rho_{e\varepsilon}$.

The first point of this example is that data on endogenous variables do provide information, though not the sort that can be handled by commonly employed estimation techniques. For models noticeably more complex than (2.1), such information can be quite difficult to interpret. [Explicit latent variable models may be useful here; see Aigner, Hsiao, Kapteyn and Wansberg (1984).] The second point is that even in simple models, the interpretation of relations among endogenous variables requires a good deal of prior information. Since the prior information one brings to any empirical study is derived both from theory and from previous empirical work, this suggests (correctly, I think) that progress is often made by assembling pieces of theory and evidence from a variety of sources, rather than through definitive tests of critical hypotheses.

All this implies that the primary objective in cross-section studies must be to describe the main patterns in the data set employed as clearly and completely as possible. The appropriate mind-set, which some recent work seems clearly to reflect, is accordingly that of descriptive statistics, not structural hypothesis testing. Of course, all correlations are not created equal. Theory and previous empirical work must determine what is worth studying and how it should be measured. Structural hypotheses must inevitably play a key role at the design stage, even when the endogeneity problem prevents structural estimation and testing. On the other hand, strong and robust relations among variables with economic content should always be reported, even if they do not make sense in light of existing theory; they may be central to the development of better theory.

Regression analysis may be an appropriate technique for data description in many cases. But in a world of what Krasker, Kuh and Welsch (1983) describe as "dirty data and flawed models", ordinary least squares (and other methods based on second moments) should be supplemented by the techniques they and others have developed for detecting and dealing with extreme observations. [For a striking example of the effects of outliers in cross-section data, see Cohen, Levin and Mowery (1987).] And if statistical analysis is to be used as a tool to summarize data, rather than to estimate structural models, it is important to let the data speak. This points toward the use of relatively simple specifications and careful treatment of specification uncertainty [Leamer (1983), Bothwell, Cooley and Hall (1984), Connolly and Hirschey (1984)].

The almost exclusive attention paid to t- and F-statistics in much cross-section work is inconsistent with the methodological viewpoint taken here. Such statistics do help the reader sort out the impact of sampling variation, of course. But the relation between structural hypotheses and estimated coefficients is often unclear or controversial. (Indeed, the existence of competing structural explanations for the findings of many cross-section studies is both a reason why this line of research has fallen from favor and a symptom of the general impossibility of structural estimation in this context.) From the descriptive point of view, equal interest usually attaches to the magnitude of estimated coefficients and to the contribution of particular independent variables to explaining the variance of the dependent variable. In the supply–demand example, the t-statistic on the slope coefficient in a least squares estimate of (2.1a) would be of interest, but it would provide only a small fraction of the information in the data. In the present context, one would like to know not just whether concentrated industries are on average more profitable than unconcentrated industries, but also whether the difference (if any) is large or trivial.

In general, descriptive work should be concerned with measurement and data summarization in the broadest sense. Convincing evidence on the validity of structural hypotheses rarely emerges from a single empirical study – here or in other branches of economics. Progress is facilitated if the main features of individual data sets are fully described, so that diverse studies can be compared and contrasted. Improvements in data collection and measurement methods are likely to add more value than refinements in the specification of underidentified structural equations.

Because it is often not clear how best to measure many variables suggested by theory, the most interesting empirical relations are those that are robust to plausible variations in measurement methods as well as to variations in specification. And because different countries often have different accounting conventions and construct official data in different ways, and we seek economic laws that hold across national boundaries, international replications are especially valuable. [The use of matched Canadian and U.S. industries by Caves, Porter and Spence (1980), Baldwin and Gorecki (1985), and a few others is noteworthy in this context.]

Finally, it is important to recall that if departures from long-run equilibrium are correlated with the independent variables, cross-section studies will produce a biased picture of relations among long-run equilibria. And such correlations are often plausible. Capital-intensive industries tend to be concentrated and to have cyclically sensitive profitability, for instance. New industries or those that have been disturbed by major innovations are likely to be farther from equilibrium than others, with the direction of departure from equilibrium dependent on the source and nature of the innovation. All of this points to the desirability of attempting to control for departures from equilibrium in cross section, of using replication to check for robustness with respect to sample selection and period of study, and of employing panel data creatively. Panel data sets make it possible in principle to control for or to study cyclical and secular disequilibria and to analyze directly the long-run differences among industries.[8] Panel data share another very desirable feature with data on geographically-separated markets in a single industry: they make it possible to control for unobservable industry-specific variables by focusing on differences over time or across space.

The descriptive orientation presented here implies that cross-section studies in industrial organization should be modest, both in their goals and in their conclusions, since it is generally impossible to estimate structural models complex enough to be theoretically defensible. Modesty would go a long way toward making cross-section studies persuasive, thus putting them on the same plane as good cross-section work in other fields. Consistent with this orientation, I concentrate in Sections 4–6 on empirical regularities that seem to be robust to variations in specification, time period, country, and plausible changes in variable definition.[9] Theory enters in discussions of measurement and specification choice, but I do not attempt to provide definitive structural interpretations of results.

[8] At the simplest level, industry averages over relatively long time periods shed more light on long-run differences than observations for any single year. More generally, modern econometric techniques make it possible to combine intertemporal (within-unit) and inter-industry (between-unit) information efficiently and to perform revealing specification tests; see Hausman and Taylor (1981).

[9] Assmus, Farley and Lehmann (1984) and the references they cite discuss formal approaches to the analysis of multiple studies, the use of which is effectively precluded here by the breadth of the literature covered.

3. Measuring key variables

Most of the cross-section literature focuses on relations involving one or more of the following variables: profitability, concentration, and barriers to entry. Subsections 3.1 and 3.2 consider measures of profitability that have been employed (and attacked) in this literature. Subsections 3.3 and 3.4 discuss measurement of concentration and barriers to entry, respectively.

3.1. Measures of profitability

The many measures of profitability that have been employed in the cross-section literature fall into four basic classes. First, Bain (1951, 1956) argued that the relevant theory deals with the ability of firms to hold price above long-run average cost, where "cost" is defined as usual to include competitive returns on capital employed. Since most firms (and plants) produce multiple products, this suggests using the ratio of excess profit to sales revenue. Only Qualls (1972) and a few others have used accounting-based estimates of this measure of profitability, perhaps because it requires an estimate of the competitive rate of return on capital employed.

Second, many studies have employed accounting rates of return on assets or equity. Bain (1951, 1956) used the after-tax rate of return on equity because of data limitations, and other authors have employed the before-tax rate of return on equity and the before- and after-tax rates of return on assets. [Returns on assets are most naturally defined to include both interest payments and profits; see Schmalensee (1976).] Before-tax measures are undistorted by peculiarities of tax systems, though long-run (risk-adjusted) after-tax (economic) rates of return should be equalized under the null hypothesis of perfect competition. Increases in leverage make the residual return to equity more variable, and in competitive capital markets investors must generally be paid higher average returns to compensate. Rates of return on assets, on the other hand, will mainly reflect operating results, not capital structure decisions.

Third, Collins and Preston (1968, 1969) introduced and employed the so-called price–cost margin (PCM), which can generally be computed for more narrowly-defined industries than accounting rates of return. Consider a firm with long-run constant returns to scale, and let v = variable cost per unit, δ = depreciation rate of capital, ρ = competitive rate of return, P = price, Q = output, and K = dollar value of capital employed. Then the markup of price over long-run average (and marginal) cost is given by

$$\frac{P - v - (\rho + \delta)(K/Q)}{P} = \frac{PQ - vQ}{PQ} - (\rho + \delta)\frac{K}{PQ}. \tag{3.1}$$

The first quantity on the right, (revenue − variable cost)/revenue, is the PCM.

Under competitive conditions, the PCM should on average equal the second quantity on the right of equation (3.1), the required rental on assets employed per dollar of sales. Many authors have used the PCM as the dependent variable in linear regressions and included the ratio of assets (sometimes depreciable assets) to revenue among the independent variables. In light of (3.1), this procedure amounts to assuming that both the competitive rate of return, ρ, and the rate of depreciation, δ, are the same for all industries in the sample.

Fourth, measures that employ the market value of a firm's securities (often, because of data limitations, only its common stock) are attractive because, under the widely-accepted hypothesis of capital market efficiency, the market value of a firm's securities reflects all available information about its future profitability [Schwert (1981)]. In an early study, Stigler (1963) employed the ratio of the market value of a firm's equity to its inflation-adjusted book value. Two other measures have been widely used. Tobin's q, defined as the ratio of the market value of a firm to the replacement cost of its tangible assets, should on average equal one under the competitive null hypothesis if (and only if) intangible assets are not present [Lindenberg and Ross (1981), Salinger (1984)]. The excess value ratio (EVR), defined as (market value − book value)/revenue, was introduced by Thomadakis (1977) as a measure of the ratio of (capitalized) excess profits to sales. [Smirlock, Gilligan, and Marshall (1984) compare these two measures.]

Are all these measures so highly correlated with one another that debates about their relative merits are pointless? At least for the last three classes of measures, the answer seems to be as follows:[10]

Stylized Fact 3.1

Correlations among accounting rates of return are high, and regression results are usually not sensitive to which measure of this type is employed. Correlations of accounting rates of return with the PCM and with measures based on market values are lower, and regression results often depend on which type of measure is used.

The weak correlation between the PCM and the ratio of accounting profits to sales reported by Liebowitz (1982a) and others suggests important inter-industry differences in rates of depreciation and competitive rates of return.

3.2. Accounting problems

All of the profitability measures mentioned above rely on accounting data, even those also using data on securities prices [Schwert (1981)]. As Benston (1985)

[10]See, for instance, Ornstein (1972), Caves, Porter and Spence (1980), Lindenberg and Ross (1981), Liebowitz (1982a), Martin (1979b), Salinger (1984), and Hirschey (1985).

demonstrates, it is easy to list many reasons why accounting data yield noisy measures of economic variables. (The PCM is particularly easy to criticize because it omits capital costs.) Important problems arise because large firms are generally active in many markets. Firm-level data are thus multi-market aggregates, while data constructed at the plant level do not reflect costs incurred at the firm level, and the allocation of those costs to individual lines of business is inevitably somewhat arbitrary.[11]

On the other hand, it is unlikely that accounting numbers are pure noise: firms use accounting data (though perhaps not the aggregates in published reports) in decision-making, and many studies in the finance and accounting literatures find that the stock market reacts to the publication of accounting reports. While the signal to noise ratio in accounting data is of interest, the more important question is the extent to which errors in accounting data are correlated (positively or negatively) with independent variables used in regression analysis. If such correlations are important, coefficient estimates will be biased, and statistical studies, even with large samples, may miss real relations involving true, economic profitability and report spurious relations that are mere artifacts of accounting practices.

Stigler (1963) noted that owners of small U.S. corporations have an incentive to pay themselves high salaries, and thus to understate their accounting profits, in order to avoid the double taxation of dividend income. His results indicate that small corporations tend to account for a larger share of industry assets the lower is concentration. [See also Kilpatrick (1968) on adjusting for the effects of this incentive.] One can argue on theoretical grounds that managers have strategic and public relations incentives to understate high profits and overstate low profits, though the extent of such behavior has apparently not been systematically studied.

More recently, considerable attention has been focused on capitalization and depreciation practices and inflation as sources of bias. Much of the relevant theoretical literature [see especially Stauffer (1971) and Fisher and McGowan (1983)] has considered a firm composed of a large number of identical investment projects. Each project requires an initial outlay of one dollar and produces a net cash flow of $\pi(\tau)$ dollars when it is τ periods old, with all dollar figures deflated to some base period. Suppose there are no taxes and the following relation holds:

$$1 = \int_0^{\infty} \pi(\tau) e^{-r\tau} \mathrm{d}\tau, \tag{3.2}$$

so that r is the real, economic rate of return on the firm's operations.

[11]These problems apply in the United States to data published by the Internal Revenue Service, the Census of Manufactures, and the Federal Trade Commission's Line-of-Business program, respectively.

To see how r relates to accounting rates of return, let $I(t)$ be the number of projects the firm starts in period t, and let $P(t)$ be the ratio of prices in period t to those in the base period. Thus, $P(t)I(t)$ is the current dollar value of the firm's investment in period t. Let $b(\tau, t)$ be the book value in period t of a dollar invested in period $t-\tau$, and let $d(\tau, t)$ be the accounting depreciation charged against this investment. Then the firm's accounting rate of return on assets in year t is given by cash flow minus depreciation, all over the book value of assets:

$$r_{\mathrm{a}}(t) = \frac{\int_0^{\infty} I(t-\tau)P(t)\pi(\tau)\,\mathrm{d}\tau - \int_0^{\infty} I(t-\tau)P(t-\tau)d(\tau,t)\,\mathrm{d}\tau}{\int_0^{\infty} I(t-\tau)P(t-\tau)b(\tau,t)\,\mathrm{d}\tau}. \tag{3.3}$$

In the simplest case prices are constant, so that $P(t) = 1$ for all t, and that accounting depreciation is not time-dependent, so that $b(\tau, t) = b(\tau)$ and $d(\tau, t) = d(\tau)$ for all τ and t. Then in order for r_{a} to equal r for all possible investment paths, $I(t)$, it follows from (3.3) that the following equation be satisfied for all τ:

$$-d(\tau) = rb(\tau) - \pi(\tau). \tag{3.4}$$

As Hotelling (1925) first demonstrated and many others have independently discovered since, (3.4) will be satisfied with $b(0) = 1$ and $-d(\tau) = b'(\tau)$ for all τ if and only if the asset's net book value is given by

$$b_{\mathrm{e}}(\tau) = \int_{\tau}^{\infty} \pi(x)\mathrm{e}^{-r(x-\tau)}\,\mathrm{d}x. \tag{3.5}$$

That is, depreciation is *exact* if the asset's book value is equal to the net present value, computed at the economic rate of return, of its future net cash flows. Then exact or economic depreciation is just the decline in book value: $d_{\mathrm{e}}(\tau) = -\mathrm{d}[b_{\mathrm{e}}(\tau)]/\mathrm{d}\tau$.

If prices are changing, equation (3.4) is replaced by

$$-P(t-\tau)d(\tau,t) = rP(t-\tau)b(\tau,t) - P(t)\pi(\tau). \tag{3.6}$$

It is easy to see that (3.6) is satisfied for all t and τ and for any price trajectory, $P(t)$, if both Hotelling book values and Hotelling depreciation deductions are adjusted to take into account inflation since the asset was purchased:

$$b(\tau,t) = [P(t)/P(t-\tau)]\,b_{\mathrm{e}}(\tau), \tag{3.7a}$$

$$d(\tau,t) = [P(t)/P(t-\tau)]\,d_{\mathrm{e}}(\tau). \tag{3.7b}$$

[See Shalchi and Smith (1985) for an overview of the accounting literature on methods for handling price changes in practice.]

To see what happens when exact depreciation and inflation adjustments are not employed, it is convenient and traditional (but somewhat unrealistic) to consider steady-state growth paths. Suppose prices rise at rate i and investment grows at rate g, and define the Laplace transform, $f^*(s)$, of any function of time, $f(t)$, by

$$f^*(s) = \int_0^\infty f(t)\mathrm{e}^{-st}\mathrm{d}t, \tag{3.8}$$

where s is a constant [Stauffer (1971)]. Then, if depreciation is not time-dependent, substitution in (3.3) and integration by parts allow the steady-state accounting rate of return to be rewritten as

$$r_\mathrm{a} = g + i + \frac{\pi^*(g) - 1}{b^*(g+i)} = (g+i)\frac{\pi^*(g) - d^*(g+i)}{1 - d^*(g+i)}. \tag{3.9}$$

Since $\pi^*(r) = 1$ by definition, equation (3.9) shows that when $g = r$ the accounting rate of return overstates the economic rate of return by exactly the rate of inflation. It thus provides an unbiased estimate of the firm's *nominal* rate of return, which can be related to observed nominal interest rates, for instance. In the *usual case* in which $r > g$ and $\pi^*(g)$ thus exceeds one, r_a exceeds $g + i$ but may be greater or less than $r + i$ in general [Fisher and McGowan (1983)]. More rapid depreciation, perhaps taking the form of expending some of the initial investment, will reduce $b^*(g+i)$ and $d^*(g+i)$. As long as $\pi^*(g) > 1$, it follows that the steady-state accounting rate of return will be *increased*, even though both profits and assets will be reduced.

In the very special case of exponential decay, $\pi(\tau) = (r+\delta)\mathrm{e}^{-\delta\tau}$. Then if $b(\tau) = \mathrm{e}^{-d\tau}$, where d and δ may differ, equation (3.9) becomes:

$$r_\mathrm{a} = r + i + [i + (d - \delta)]\left[\frac{r-g}{g+\delta}\right]. \tag{3.10}$$

Depreciation is exact if and only if $d = \delta$.[12] If $d = \delta$, r_a is equal to the nominal rate of return, $r + i$, for very short-lived assets ($\delta \to \infty$) and approaches $r + i(r/g)$ as asset longevity increases ($\delta \to 0$). In the usual case in this example, at least, the steady-state bias is thus worse for longer-lived assets. Similarly, as long as $r > g$, the bias is a decreasing function of the firm's growth rate for fixed r.

[12] One should not be misled by this example: in general the time-path of Hotelling depreciation does *not* have the form of the cash-flow profile – nor, necessarily, of standard accounting depreciation schemes.

It is clear that if accounting data are to be used to measure economic profits, an inflation adjustment of the sort described by equations (3.8) is appropriate. Bain (1951) recognized this point, and Stigler (1963) adjusted his data for the effects of inflation. Few later authors have followed suit. The analysis above indicates that failure to adjust for inflation will induce bias if asset lifetimes or firm growth rates are correlated with independent variables employed in profitability regressions.

Outlays for advertising, research, and development are treated as current expenses in conventional accounting, as are costs of producing firm-specific human capital, even though all these outlays are expected to produce future cash flows. The analysis above indicates that in the usual case, these procedures tend to understate firms' capital stocks (by depreciating more rapidly than is economic) and overstate rates of return (see Subsection 4.4 below).

Firms have some discretion over the accounting procedures they employ. Studies of choices of accounting methods [Zmijewski and Hagerman (1981), Holthausen and Leftwich (1983)] consistently support

Stylized Fact 3.2

Large U.S. firms are more likely than small ones to adopt accounting practices (like accelerated depreciation) that lower current profits and increase steady-state accounting rates of return.

Salamon (1985) argues that this phenomenon is the source of the correlation between firm size and accounting profitably detected by Hall and Weiss (1967) and some other authors. On the other hand, there is little support for the existence of a correlation between accounting method choices and industry concentration [Hagerman and Senbet (1976), Zmijewski and Hagerman (1981), Holthausen and Leftwich (1983)].

All this suggests that empirical work on profitability should take accounting biases seriously. In some cases it may be sufficient to use alternative profitability measures that are likely to be biased in different directions. Sometimes controls for accounting distortions can be included among the regressors [Telser (1972, ch. 8), Salinger (1984)]. Alternatively, it may be possible to construct subsamples that differ in the likely direction or importance of accounting biases and to check for stability among the subsamples [Demsetz (1979)]. Inflation-related distortions can be corrected, at least approximately, on a fairly routine basis.[13] One can either exclude small corporations [Kilpatrick (1968)] or attempt to adjust their

[13]For instance, under the assumptions made to derive equation (3.10), if $d = \delta$, then δ equals the observed accounting depreciation/assets ratio, and the inflation-induced bias in the rate of return on assets can be corrected by multiplying both assets and depreciation by $(g + \delta + i)/(g + \delta)$.

accounting profits for excessive salaries [Stigler (1963)]. A number of authors have attempted to adjust accounting data for depreciation-related biases using the steady-state framework presented above [e.g. Weiss (1969), Stauffer (1980), and Salamon (1985)]. Such adjustments require considerable prior information, since the basic cash flow profile, $\pi(\tau)$, cannot be directly estimated from aggregate accounting data.

3.3. Measures of concentration

Two questions are of central importance here: Which measure of concentration should be employed? And how should geographic and product market boundaries be drawn?

A number of authors have presented axiomatic arguments for particular concentration measures; see Hannah and Kay (1977), Curry and George (1983), and Waterson (1984). Ideally, of course, the appropriate measure of concentration should be derived from oligopoly theory. As Cowling and Waterson (1976), among others, have observed, the H-index of concentration, equal to the sum of squared market shares, emerges as an endogenous correlate of industry profitability in a Cournot oligopoly with (exogenous) cost differences. Saving (1970) shows that concentration ratios (the aggregate shares of domestic output or employment of, for instance, the four or eight largest sellers) emerge similarly under alternative behavioral assumptions. But the usual hypotheses of interest involve the effect of concentration on behavior, and this argues against assuming the mode of behavior in advance. Stigler (1964) suggests that the H-index provides a reasonable measure of the ease of detecting cheating on collusive agreements, but his arguments are not fully rigorous. In short, received theory does not dictate the choice of concentration measure.

Most authors use concentration ratios because they are available in government-supplied data and because many studies have found alternative concentration measures to be highly correlated. But the choice among even highly correlated concentration measures can affect the results obtained [Kwoka (1981), Sleuwaegen and Dehandschutter (1986)]. And concentration ratios can be used to develop good estimates of the H-index [Schmalensee (1977), Michelini and Pickford (1985)], so that good estimates of other measures may also be obtainable from published data.

Many authors have also simply used the market boundaries provided by the compilers of official data. As antitrust cases make clear, it is often difficult to choose among market definitions, and the official definitions are often inappropriate. (Geographic market boundaries in official data usually coincide with national boundaries, and product markets boundaries are often based mainly on similarity of production technologies.) Bain (1951) chose to drop from his sample those officially-defined "markets" for which geographic or product boundaries did not seem sensible. This reduced his sample size from 149 to 83. Most subsequent authors have been unwilling to sacrifice so many degrees of freedom to obtain well-defined markets. [But see Mann, Henning, and Meehan (1967).]

Most investigators (but not all) do drop catch-all industries with such terms as "not elsewhere classified" or "miscellaneous" in their descriptions. Some also adjust published concentration ratios for the existence of regional markets [see Shepherd (1974) and, especially, Weiss and Pascoe (1986)]; others include dummy variables for products that are rarely shipped long distances. Going in the other direction, it is common to allow for foreign competition by using the ratio of imports to domestic production [Caves (1985)]. But Leitzinger and Tamor (1983) and others note that if a product is already imported in non-trivial quantities, and if there are no non-tariff barriers preventing an increase in imports, imports can respond to domestic price changes, so that it may be better to work with world markets (and world concentration) rather than domestic markets.

A few studies have considered the relation between *buyer* concentration and seller profitability. The basic Bainian argument here is that buyers who are large relative to the market should be able to destabilize collusion in concentrated industries and push sellers' prices and profits down toward competitive levels. On the other hand, if a seller faces few buyers because he sells to a single concentrated industry and if the concentration–collusion hypothesis is valid, downstream input demand may be less elastic that it would be under competitive conditions, tending to offset increased downstream bargaining power [Waterson (1980)]. Measurement of buyer concentration has proven to be difficult in practice [compare Lustgarten (1975), Guth, Schwartz and Whitcomb (1976), and Waterson (1980)] and tightly constrained by data availability.

3.4. Measures of entry barriers

The cross-section literature has taken three different approaches to measuring the elements of market structure that Bain (1956) argued affected the ability of established firms to prevent supra-normal profits being eroded by entry. First, Bain (1956) performed a detailed structural analysis of each of the industries in his sample and classified them according to the height of the barriers to entry in each. This approach is labor-intensive, and subjective judgement must be used to integrate the information on each industry into an overall estimate of the height of entry barriers. For these reasons, only a few subsequent authors [notably Mann (1966) and Qualls (1972); see also Palmer (1973)] have used this approach.

Second, Orr (1974a), using 1964–67 Canadian data, estimated a model of the following sort:

$$\Delta N = \gamma(r - r^*), \tag{3.11}$$

where ΔN is the gross increase in the number of sellers over the period, γ is a positive constant measuring the speed of adjustment, r is the average observed profit rate, and r^* is the profit rate at which entry would cease. Orr (1974a) substituted for r^* a linear function of variables designed to measure the conditions of entry. In a later study, Orr (1974b) used the estimated coefficients of this function to construct a measure of entry barriers for each industry in his sample. Only Masson and Shaanan (1982) and a few others have adopted this two-step approach.

Third, the most common approach to the treatment of entry barriers in studies of profitability appears to be due to Comanor and Wilson (1967). They investigated regression equations of the following form:

$$r = \beta_0 + \beta_1(CON) + \beta_2(BE_1) + \cdots + \beta_{N+1}(BE_N), \tag{3.12}$$

where r is a measure of profitability, the β's are unknown coefficients, CON is a measure of seller concentration, and the BE's are variables designed to measure the structural determinants of entry barriers. BE variables that appear in the literature generally correspond to three of the four possible sources of entry barriers discussed by Bain (1956): scale economies, capital requirements, and product differentiation advantages of established sellers. [Bain (1956) found the fourth possible source, absolute cost advantages of established sellers, to be the least important in his sample.]

Comanor and Wilson (1967, 1974) and many others have measured the importance of *scale economies* by the ratio of the output of a *plant* of minimum efficient scale (MES) to the output of the market as a whole. [Only Neumann, Bobel and Haid (1979) and a few others have attempted to measure minimum efficient *firm* scale.] MES is most commonly measured as either the average size of the largest plants accounting for half the industry's output or the size of the smallest of these plants. Both measures rest on the assumption that the distribution of observed plant sizes relative to MES does not vary systematically across industries, though Weiss (1976) and Baldwin and Gorecki (1985) find evidence that this assumption is false. And Davies (1980) shows that the differences between MES measures based on the size distribution of existing plants and MES measures computed using either survivorship methods or the interesting approach of Lyons (1980) tend to be positive and to be positively correlated with concentration. [See also Ornstein, Weston, Intriligator and Shrieves (1973) on MES measures.]

Caves, Khalizadeh-Shirazi and Porter (1975) argued that even if MES is large relative to the market, small-scale entry may be attractive unless the cost penalty for operation at suboptimal scale is substantial. They compute a cost disadvantage ratio (CDR) by taking the ratio of value-added per worker in plants below MES to that in larger plants, and they propose multiplying the

MES/market ratio by a zero/one dummy variable that equals one only when CDR is small. [Chappel and Cottle (1985) use firm-level data in this fashion.] This procedure tends to overstate the cost disadvantages of small plants, since capital/labor ratios typically rise with scale [Caves and Pugel (1980)] – but CDRs above unity are not uncommon in U.S. data. It is unclear why a zero/one specification is preferable to some sort of continuous interactive form. Some studies employ a CDR-based dummy variable by itself in linear equations; the theoretical rationale is not apparent.

Bain (1956) argued that a potential entrant might be deterred if the *capital requirements* of entry were large in absolute terms. The hypothesis that capital markets are seriously imperfect, on which Bain rested his argument, does not now command much respect. But recent theory implies that entry will be deterred if a large fraction of entry costs are sunk (i.e. cannot be recovered upon exit), and the relative importance of sunk cost may be correlated with the absolute level of capital requirements [see Kessides (1986)]. Capital requirements are often handled in the Comanor–Wilson (1967) framework by including among the regressors a variable measuring the capital cost of a plant of minimum efficient scale. Some authors multiply this quantity by a CDR-based dummy variable.

Finally, Bain (1956) attributed the greatest importance to *product differentiation advantages of established sellers*. Comanor and Wilson (1967) introduced the idea of using an industry's advertising/sales ratio to measure this structural feature; they and others have also employed advertising spending per firm for this purpose. [Cowling (1976) and Porter (1976a) compare these measures.] Some studies have also used research and development spending as a percentage of sales or patents/sales, and Neumann, Bobel and Haid (1985) used the ratio of registered trademarks to owners' equity. All of these variables are basically measures of conduct and thus clearly endogenous in the long run.

4. Profitability

Bain (1951) began the literature considered in this section by arguing that if high seller concentration facilitates collusion, firms in highly concentrated industries should on average earn supra-competitive profits. He found support for this hypothesis using data on leading U.S. firms in the 1936–40 period.[14] In a second seminal study, Bain (1956, esp. pp. 190–191) argued that both high concentration and high barriers to entry were necessary to produce excess profits in long-run equilibrium. He found support for this interactive hypothesis using data on leading U.S. firms in 1936–40 and 1947–51. Bain (1951, 1956) is still worth

[14]For detailed discussions if Bain's (1951) study, see Brozen's (1970, 1971) critique and the responses by Qualls (1974) and Weiss (1974).

reading today for his careful handling of data and his thoughtful discussion of many of the hypotheses, problems, and results that have dominated the subsequent literature.

Subsection 4.1 discusses some descriptive statistics on differences in measured profitability, and Subsection 4.2 considers control variables that have been employed in cross-section studies of profitability levels. Subsections 4.3–4.5 present the main results that have been obtained in these studies, and Subsection 4.6 discusses studies concerned with the variability of profits.

4.1. *Descriptive statistics on profitability*

Are many firms sufficiently profitable as to suggest large percentage differences between price and average cost? Analysis of U.S. data on accounting rates of return [Alberts (1984)] and Tobin's q's [Salinger (1984)], along with the generally small estimates of monopoly welfare losses based on cross-section differences in profit rates, imply

Stylized Fact 4.1

Differences among observed accounting rates of return and market/book ratios in the U.S. are generally too low to be easily reconciled with the existence of textbook monopolies.

The 1936–40 profitability data reported in Bain (1951) support this observation. The average after-tax rate of return on equity in Bain's 20 unconcentrated industries is 6.9 percent. If this is the competitive rate of return, r_c, any other firm's excess after-tax return on equity is given by

$$r - r_c = \frac{(1-\tau)(R-C)}{E} = \left[\frac{(1-\tau)R}{E}\right]\left[\frac{R-C}{R}\right], \tag{4.1}$$

where r is the actual after-tax rate of return on equity, τ is the corporate tax rate, R is revenue, C is total cost (including normal profit), and E is owners' equity. Data from the U.S. Internal Revenue Service *Statistics of Income, 1938* indicate that $[(1-\tau)R/E]$ averages about 1.12 for manufacturing firms in Bain's sample. Thus an observed r of 16 percent corresponds to a markup over total cost $[(R-C)/R]$ of about 8.1 percent $[= (16.0 - 6.9)/1.12]$, which would be chosen by a monopoly facing a demand elasticity of about 12. A firm with such an elastic demand curve has little monopoly power, and yet only 3 of Bain's 22 concentrated industries had r's above 16 percent. The highest r in Bain's sample was for "aircraft and parts", an industry far out of equilibrium in the 1930s; it corresponds to a demand elasticity of about 8.

On the other hand, Mueller (1977, 1986) and Connolly and Schwartz (1985), using both accounting rates of return and the EVR, find that profit rates of large U.S. firms to not converge over time to a common mean [see also Stigler (1963)], and Odagiri and Yamawaki (1986) and Geroski and Jacquemin (1988) report similar results for large Japanese and U.K. firms, respectively. Some studies of other countries (discussed below) cannot reject convergence in the limit, but none finds rapid convergence.

Stylized Fact 4.2

Accounting profitability differences among large firms tend to persist for long periods.

Connolly and Schwartz (1985) find that highly profitable U.S. firms regress toward the mean noticeably more slowly than others.

There appear to be important international differences in profit dynamics. Geroski and Jacquemin (1988) cannot reject the null hypothesis that profitabilities of large French and German firms converge to a common value in the limit, and Odagiri and Yamawaki (1986) find more dispersion in asymptotic profit rates in the United States than in Japan. These studies also argue that rates of convergence are more rapid in Japan than in the United States and more rapid in France and Germany than in the United Kingdom. Odagiri and Yamawaki (1987) find the United States to have the largest asymptotic differences and the slowest convergence in a larger set of countries.

The importance of industry differences in the determination of profitability has been studied by Gort and Singamsetti (1976), using firm-level data for the United States in 1970, and by Schmalensee (1985), Scott and Pascoe (1986), Ross (1986), and Kessides (1987) using U.S. Federal Trade Commission Line of Business data for the mid-1970s. All employed dummy variables for each industry in the sample, and all support

Stylized Fact 4.3

At the firm or business unit level in the United States, industry characteristics account for only about 10–25 percent of the cross-section variation in accounting rates of return.[15]

[15]Gort and Singamsetti (1976) and Kessides (1987) come up with 10 percent, Schmalensee (1985) with 20 percent, and Ross (1986) with 30 percent. [The Scott–Pascoe (1986) estimates are not strictly comparable.] Mueller (1986, pp. 218–219) argues plausibly that the Schmalensee and Ross estimates may be unusually high because of the extraordinary impact of the first oil shock in 1975, the year covered by their data.

This suggests that Stylized Fact 4.2 reflects more than industry-specific accounting biases and stable mixes of firms' activities, though it does not rule out persistent differences in growth rates or accounting practices as sources of long-lived differences.[16] Recent work by Cubbin and Geroski (1987) with a panel of 217 large U.K. firms over the 1951–77 period finds that changes in firms' profit rates are not well explained by industry averages; their estimates reveal important firm-specific dynamic effects. On the other hand, Schmalensee (1985) found that industry characteristics accounted for about 75 percent of the variation in industry average accounting rates of return, suggesting that the industry is at least a valid unit of analysis.

Gort and Singamsetti (1976) attributed the variation in rates of return not explained by industry dummy variables to firm characteristics. But Schmalensee (1985) found that knowing a firm's profitability in one industry provided no information on the likely profitability of its other businesses. This is consistent with Mueller's (1986) finding that the only detectable impact of merger activity in a sample of large U.S. firms during the 1950–72 period was to hasten the regression of acquiring firms' profitability toward the mean. Recently, however, Scott and Pascoe (1986) and Kessides (1987) have detected significant firm effects in the FTC Line of Business data.

4.2. Control variables

Many control variables have been employed in cross-section studies of profitability. [Ravenscraft (1983) and Bothwell, Cooley and Hall (1984) provide long lists.] Following Comanor and Wilson (1967), many authors have used some measure of recent sales growth in order to control for the effects of disequilibrium.[17] This variable almost always "works" statistically [see, especially, Bothwell, Cooley and Hall (1984)]:

Stylized Fact 4.4

Recent growth in revenue is positively correlated with measured profitability.

Bradburd and Caves (1982) argue that only unanticipated growth should affect profitability, but they find support for this hypothesis only among uncon-

[16] Imel, Behr and Helmberger (1972) demonstrate that heteroskedasticity is almost certainly present in regression analysis of the profitability of diversified firms. Unfortunately, the data needed to estimate the disturbance covariance matrix consistently are rarely available. But even if fully efficient estimation is not possible, White's (1980) techniques can be used to avoid biased inferences.

[17] Mueller (1986) controls for disequilibrium by using time series data to estimate the steady-state profitability of each firm in his sample.

centrated U.S. industries in 1972. Liebowitz (1982b) examines the dynamic effects of several alternative measures of disequilibrium on the rate of return on assets in U.S. data for the 1960s. He concludes that revenue growth over a one- or two-year period if the best available measure. (Growth rates over longer periods may serve as crude controls for growth-related accounting biases.) Liebowitz (1982b) also finds that his measures of disequilibrium are generally uncorrelated with concentration, so that estimates of the concentration–profitability relation may not be biased by failure to control for disequilibrium.

Studies using the PCM to measure profitability generally employ the capital/revenue ratio as a control. The coefficient of this variable is usually plausible and statistically significant. But significant negative estimates have been reported by a number of authors [e.g. Ornstein (1975), Liebowitz (1982a), Domowitz, Hubbard and Petersen (1986a, 1986b)] especially when U.S. data for the 1970s are employed.

Pagoulatos and Sorenson (1981) and Harris (1986) present evidence that PCMs are lower in industries with more elastic demand, and Bradburd (1982) finds that PCMs are lower in producer goods industries selling inputs that are important to downstream buyers (and for which demand elasticities should accordingly be high). But Bradburd finds no support for the plausible hypothesis that demand elasticity is negatively related to PCM only in concentrated industries.

Finally, a number of authors have attempted to control for differences in risk, using a variety of measures and obtaining a variety of significant [Bothwell and Keeler (1976), Neumann, Bobel and Haid (1979, 1985), Harris (1984, 1986)], insignificant [Grabowski and Mueller (1978), Bothwell, Cooley and Hall (1984), Hirschey (1985)], and perverse [Thomadakis (1977), Mueller (1986)] results. This state of affairs is somewhat surprising a priori, since investors must generally be compensated for bearing risk. Part of the problem may be that firms with inherently risky demand find it optimal to charge a lower price, all else equal, in order to maximize their market value [Harris (1986)].

4.3. Concentration and profitability

Weiss (1974) examined 46 studies that had been published by the early 1970s and noted that 42 of them had found a positive relation between concentration and profitability. [See also the reviews by Collins and Preston (1968) and Phillips (1971).] He took this as providing strong support for the concentration–collusion hypothesis, though he did note that the concentration–profitability relation was generally statistically weak. [See, for instance, Stigler (1963) and Collins and Preston (1968).]

The economic effects of concentration implied by the early literature were also generally small. Employing equation (4.1) as above, for instance, Bain's (1951) results imply an average markup over long-run cost of only 4.6 percent in the concentrated industries in his preferred sample. For other samples (see his table 3), implied markups varied from 0.9 to 3.2 percent, and the corresponding profitability differences were generally statistically insignificant.

Weiss (1971) noted that Bain's (1956) hypothesis called for interactive (concentration $\times$ barriers) specifications, but surprisingly few authors have employed models of this sort. Mann (1966) and Qualls (1972) found support for Bain's hypothesis in U.S. data, as did Jenny and Weber (1976) in French data. Orr (1974b) and Caves, Porter and Spence (1980), however, found that interactive specifications did not perform noticeably better than simple linear models with Canadian data. Salinger (1984) argued that Bain's hypothesis implied that (3.12) should be replaced by interactive regression models of the form:

$$r = \beta_0 + CR[\beta_1(BE_1) + \cdots + \beta_N(BE_N)] + \beta_{N+1}(G) + \cdots, \tag{4.2}$$

where G is a measure of past sales growth, and the ellipses indicate additional variables discussed below. Salinger found that such models had essentially no ability to explain variations in the market/book ratios of large U.S. firms in 1976. Thus, the relevant literature does not provide strong support for Bain's interactive hypothesis.

Using linear models, like equation (3.12), a number of studies published after Weiss (1971, 1974) wrote found positive relations between domestic concentration and profitability; these include studies of Japan [Caves and Uekusa (1976)], Pakistan [White (1974)], France [Jenny and Weber (1976)], West Germany [Neumann, Bobel and Haid (1979, 1985)] and several studies of U.S. manufacturing industries.[18] Gilbert (1984) concluded that analyses of U.S. banking markets support the existence of a positive – though economically trivial – relation.

But many later studies of U.S. data, particularly those using multivariate specifications, found no statistically significant linear relation between domestic concentration and profitability, even when market share (see Subsection 4.5) was not included among the regressors.[19] Several studies [Porter (1976a), Grabowski and Mueller (1978), Connolly and Hirschey (1984), Hirschey (1985)] reported

[18] See Imel, Behr and Helmberger (1972), Telser (1972, ch. 8), Lustgarten (1975), Peltzman (1977), Thomadakis (1977), LaFrance (1979), Marvel (1980) and Masson and Shaanan (1982). de Melo and Urata (1986) find a positive relation in Chilean data in 1979 but not in 1967.

[19] Examples include Comanor and Wilson (1967, 1974), Ornstein (1972, 1975), Vernon and Nourse (1973), Boyer (1974), Gort and Singamsetti (1976), Cattin and Wittink (1976), Porter (1976a, ch. 6), Strickland and Weiss (1976), Martin (1979a, 1979b), Lindenberg and Ross (1981), and Bradburd (1982).

statistically significant *negative* concentration coefficients with U.S. data. And, while Weiss (1971, 1974) noted the tendency of the concentration–profitability relation to weaken during inflationary periods [see, for instance, Stigler (1963)], Domowitz, Hubbard and Petersen (1986a, 1986b) found that this relation essentially vanished in the United States during the 1970s. [See also Scott and Pascoe (1986) and Schmalensee (1987).]

Non-U.S. data have also produced negative results. Phillips (1971) failed to detect a concentration–profitability relation in French, Belgian, or Italian data, and Jacquemin, de Ghellinck and Huveneers (1980) confirmed his results for Belgium. With the exception of the theoretically interesting study of the relation between changes in the H-index of concentration and changes in profitability by Cowling and Waterson (1976), most studies of the United Kingdom have failed to find a positive linear relation between concentration and profitability [Hart and Morgan (1977), Hart and Clarke (1980), Clarke (1984)]. And the Cowling–Waterson (1976) results are apparently not robust to changes in the sample of industries studied [Hart and Morgan (1977)].

Most of these studies adopt specifications in which one or another concentration ratio is linearly related to profitability. Alternative concentration measures and functional forms sometimes yield stronger results.[20] Stigler (1964), for instance, found the H-index outperformed the four-firm concentration ratio. Most other studies have not detected sharp differences among these and other frequently discussed (and highly correlated) measures. Kwoka (1979, 1981), however, found that the shares of the two leading firms are noticeably more closely related to industry PCMs than broader concentration ratios. [See also Kwoka and Ravenscraft (1986).]

Bain (1951) argued that his data seemed to show the existence of a critical concentration ratio, above which profitability increased discontinuously. Changes in concentration above or below this level had no discernible effect. A number of studies using U.S. data have found some support for a relation of this form [Dalton and Penn (1976), White (1976), Kwoka (1979); but see Sleuwaegen and Dehandschutter (1986)]. In the most sophisticated study of this sort, Bradburd and Over (1982) find evidence for two critical levels. If concentration was previously low, they find that profits do not increase with increases in concentration until the leading four firms account for 68 percent of industry sales. But if

[20]Theory suggests that the conduct, and thus the profitability, of multi-product firms that encounter each other in multiple markets ought to be affected by these contracts, as well as by concentration in the relevant markets. Scott (1982) finds strong support for an interactive version of this hypothesis in data on U.S. manufacturing markets in 1974, but Gilbert (1984) notes that (generally less sophisticated) investigations of this hypothesis using data on banking markets have produced relatively weak results.

concentration was previously high, profits do not drop until the four-firm ratio falls below 46 percent. Finally, however, Geroski's (1981) work indicates that the critical concentration ratio hypothesis fares no better in U.K. data than the hypothesis of a linear relation.

At the very least, these mixed results make it clear that a researcher cannot expect a strong, positive concentration–profitability relation to leap out from cross-section data:[21]

Stylized Fact 4.5

The relation, if any, between seller concentration and profitability is weak statistically, and the estimated concentration effect is usually small. The estimated relation is unstable over time and space and vanishes in many multivariate studies.

Several studies have found a negative linear relation between the imports/consumption ratio and profitability.[22] But, even if foreign competition can erode domestic market power, the long-run profitability of a competitive industry should be unaffected by imports. This argues for an interactive specification [Pugel (1980a), Caves (1985)]. And several authors have indeed found that the negative impact of imports on domestic profitability is stronger when domestic concentration is high.[23]

Stylized Fact 4.6

The ratio of imports to domestic consumption tends to be negatively correlated with the profitability of domestic sellers, especially when domestic concentration is high.

[21] Hay and Kelley (1974) find that explicit collusion in the United States tends to occur most frequently in concentrated industries (especially where products are homogeneous) and to involve only a few firms. This generally supports the notion that concentration facilitates collusion. But explicit collusion is illegal and apparently relatively rare in the United States, and if concentration made *tacit* collusion easy, sellers in concentrated industries would not need to break the law.

[22] Examples include Geroski's (1982) study of the United Kingdom, Chou's (1986) study of Taiwan, work on West German data by Neumann, Bobel and Haid (1979, 1985), and studies of the United States by Martin (1979a) and Marvel (1980). de Melo and Urata (1986) find a positive relation for Chile, which they attribute to quantitative import restrictions.

[23] Examples include studies of U.S. data by White (1976), Pugel (1980a), and Domowitz, Hubbard and Petersen (1986a), a study of Belgian data by Jacquemin, de Ghellinck and Huveneers (1980), a study of Chile by de Melo and Urata (1986) and work by Caves, Porter and Spence (1980) with Canadian data.

Pugel (1980a) also finds that import penetration has a stronger negative relation with domestic profitability when conventional measures of entry barriers are high. The success of interactive specifications involving import penetration contrasts sharply with the (concentration × barriers) work discussed above.[24]

The arguments in the preceding paragraph suggest that increases in tariff protection should have a positive structural impact on long-run profits only in concentrated industries. But, while Round (1983) finds an empirical relation of this sort for Australia, Bloch (1974b) finds none for Canada. A number of authors have employed the ratio of exports to domestic production in profitability regressions. It is hard to provide a convincing theoretical rationale for this specification [Caves (1985)], and significant coefficients of both signs have been reported [compare Neumann, Bobel and Haid (1979, 1985) with Geroski (1982) and Martin (1983)], along with many insignificant results. Finally, Leitzinger and Tamor (1983) find that a (weak) proxy for world market concentration strongly outperforms U.S. domestic concentration for a sample of widely-shipped goods in 1972, and Yamawaki (1986) finds that the profit margin on Japanese exports is positively related to concentration in the corresponding U.S. industry.

Lustgarten (1975) found that *buyer* concentration was negatively related to PCMs in the 1963 U.S. data. LaFrance (1979) noted that buyer concentration is theoretically irrelevant under perfect competition and found evidence that the negative effect of buyer concentration increases with seller concentration in Lustgarten's data. But Guth, Schwartz and Whitcomb (1976) argued that Lustgarten's measure of buyer concentration was flawed and that correcting the flaws eliminated his results. And Ravenscraft (1983) and Martin (1983) report significant *positive* coefficients of buyer concentration in U.S. data for the mid-1970s. In the most theoretically sophisticated study of buyer concentration, Waterson (1980), using data on changes between 1963 and 1968 in the United Kingdom and measures of buyer concentration based on the H-index, finds evidence supporting both a negative effect of buyer concentration on profits and a positive effect of downstream margins. But Bradburd (1982) finds that downstream margins have a negative effect on PCMs in 1972 U.S. data. It seems fair to conclude that no robust relation has yet emerged from studies of buyer concentration.

[24] It is worth noting that most studies in which import penetration is the dependent variable find it positively related to both domestic concentration and profitability; see Marvel (1980), Caves, Porter and Spence (1980), and Caves (1985). [Chou (1986) does not detect these relations in data for Taiwan, however.] It is also interesting to note that specification tests tend to signal the endogeneity of imports, while not flagging other logically endogenous variables [Geroski (1982)]. Finally, in related work, Feinberg (1986) reports that when the German mark fell sharply in 1977–83, so that import competition generally declined, domestic prices rose *less* in more concentrated industries.

4.4. Entry barriers and profitability

Since the MES/market measure of scale economies and the MES-based measure of capital requirements are highly correlated with each other [and with seller concentration; see, for instance, Comanor and Wilson (1967)], both rarely if ever have statistically significant coefficients in profitability regressions. It is common, however, for either scale economies [e.g. Ornstein (1972)] or capital requirements [e.g. Comanor and Wilson (1967, 1974)] to have a significant negative coefficient. Most of the studies cited in Subsection 4.3 support

Stylized Fact 4.7

Measures of scale economies or capital requirements tend to be positively correlated with industry-level accounting profitability.

Studies using CDR-based dummy variables [e.g. Kwoka (1979)] generally obtain stronger results. The robustness of the relation described by Stylized Fact 4.7 is somewhat surprising; simple models of entry deterrence suggest that the structural MES/market coefficient should be at most equal to the competitive rate of return [Schmalensee (1981)], implying an effect small enough to be difficult to detect in the data.

Comanor and Wilson (1967, 1974) first reported a strong, positive relation between the advertising/sales ratio and industry-level profitability (measured as the after-tax of return on equity) for U.S. consumer goods industries. This finding has proven to be unusually robust.[25]

Stylized Fact 4.8

In broad samples of manufacturing industries producing consumer goods, advertising intensity is positively related to industry-average accounting profitability.[26]

A comparison of the least-squares and fixed effects estimates reported for consumer goods industries by Domowitz, Hubbard and Petersen (1986a, 1986b),

[25] Replications include the work of Caves and Uekusa (1976) with Japanese data, Round's (1983) study of Australian data, studies of U.K. data by Cowling, Cable, Kelly and McGuinness (1975) and Geroski (1982), and analyses of U.S. data by Imel, Behr and Helmberger (1972), Vernon and Nourse (1973), Strickland and Weiss (1976), Carter (1978), Martin (1979a, 1979b), Marvel (1980), Pagoulatos and Sorenson (1981), Masson and Shaanan (1982), Harris (1984), Bothwell, Cooley and Hall (1984), and Domowitz, Hubbard and Petersen (1986a, 1986b).

[26] Since advertising is usually omitted from variable costs in calculating the PCM, the relevant null hypothesis in studies using his measure of profitability is that the coefficient of the advertising/sales ratio is equal to unity. Coefficients above unity imply a positive relation between advertising intensity and profits net of advertising outlays; coefficients below unity imply a negative relation.

based on U.S. panel data covering the 1958–81 period, indicates that long-run average advertising intensity is positively related to average profitability, as the cross-section studies indicate, but year-to-year changes in these quantities are negatively related across industries.

The results of Bradburd and Caves (1982) and Domowitz, Hubbard and Petersen (1986a) [see also Cattin and Wittink (1976)] imply that in producer goods industries, advertising intensity is negatively related to profitability in both the long and short runs. Advertising is not the dominant component of selling costs in these industries [Weiss, Pascoe and Martin (1983)], so that the measurement error involved in using advertising as a proxy for selling costs may bias the advertising coefficient substantially toward zero.

Salinger (1984) and Hirschey (1985) obtain significant positive coefficients for both advertising intensity and research and development intensity with data on U.S. firms, and Stonebraker (1976) reports similar results at the industry level. Grabowski and Mueller (1978) and Connolly and Hirschey (1984) provide further support at the industry level and also report a significant and robust negative relation between profitability and the product of concentration and R&D intensity. Despite the contrary results obtained by Martin (1983), there seems enough evidence to assert

Stylized Fact 4.9

In broad samples of U.S. manufacturing industries, research and development intensity is positively related to profitability. The relation may weaken or change sign when concentration is high.

Two alternatives to the Bainian interpretation of Stylized Fact 4.8 have been widely discussed [Comanor and Wilson (1979)]; both also apply in principle to the first part of Sylized Fact 4.9. The *endogeneity* interpretation is based on standard models of optimal advertising spending [Schmalensee (1972, 1976)]. These imply that the farther is price above marginal cost, the more profitable is an additional sale, and the higher is the optimal advertising/sales ratio. Thus, the advertising–profitability correlation may reflect differences in the intensity of price competition not explained by other variables included in profitability regressions.

Vernon and Nourse (1973), using a sample of large U.S. firms in the 1960s, bound that industry-average advertising/sales ratios were more strongly correlated with firm profitability than were the firms' own ratios [see also Mueller (1986)], but Schmalensee (1976) showed that this was consistent in principle with the endogeneity interpretation. The arguments of Section 2 imply that the robustness of the advertising–profitability relation to the use of simultaneous equations methods [Comanor and Wilson (1974), Strickland and Weiss (1976), Martin (1979a, 1979b), Pagoulatos and Sorenson (1981)] does not effectively rebut this interpretation either.

The second alternative interpretation is that the advertising–profitability correlation is simply an *accounting artifact*. The argument can be developed using the steady-state framework developed in Subsection 3.2, neglecting inflation for simplicity. [See also Weiss (1969) and Demsetz (1979).] Suppose that an investment of \$1 in physical capital produces a cash flow of $ce^{-\delta\tau}$ when it is τ periods old, where c is a constant, as long as the ratio of the firm's "goodwill stock", which is increased by advertising and depreciates exponentially at a rate $\lambda > \delta$, to its physical capital is (at least) α. Thus, the firm must invest \$$\alpha$ in advertising when it invests \$1 in physical capital and must support that investment with advertising spending equal to \$$\alpha(\lambda - \delta)e^{-\delta\tau}$ when it is τ periods old. Then if r is the firm's economic rate of return, c must equal $[(r + \delta) + \alpha(r + \lambda)]$.

If the firm is growing steadily at rate g, its advertising spending at time t must equal $I(t)\alpha(g + \lambda)/(g + \delta)$, where $I(t)$ is investment in physical capital at time t. Then if advertising is expensed and physical capital is depreciated at a rate d, the firm's steady-state accounting rate of return on assets, r_a, will be given by

$$r_a - r = (d - \delta)\left[\frac{r - g}{g + \delta}\right] + A\left[\frac{r - g}{g + \lambda}\right], \tag{4.3}$$

where $A = \alpha(g + \lambda)(g + d)/(g + \delta)$ is the ratio of advertising to the *accounting* value of the firm's physical capital.

The first term on the right of (4.3) [which also appears in (3.10)] is the bias due to inappropriate depreciation of physical capital. The second term, which is positive when $r > g$ (the *usual case*), measures the bias due to expensing advertising. This term is large, indicating substantially overstated profitability, if advertising is important (A is large), if it depreciates slowly (λ is small), or if r is much larger than g.

Weiss (1969) dealt with this second bias by adjusting accounting rates of return using (essentially) equation (4.3) and assuming an advertising depreciation rate of 30 percent per year. This adjustment did not remove the advertising–profitability relation. [Grabowski and Mueller (1978) depreciated both advertising and R&D and obtained similar results.] When Bloch (1974a) used a 5 percent depreciation rate, however, the relation vanished. Equation (4.3) explains the difference: the lower is the depreciation rate, λ, the larger is the implied adjustment to the profitability of advertising-intensive firms. Time-series studies of advertising and demand generally suggest that Weiss' assumption is more plausible [Lambin (1976), Comanor and Wilson (1979), Assmus, Farley and Lehmann (1984)], but the issue is not settled.

Demsetz (1979) observed that (4.3) implies that for any λ, the importance of the accounting bias is directly related to $(r - g)$ and thus approximately related to $(r_a - g)$. Dividing his sample according to the latter variable, Demsetz found a positive relation between the rate of return on equity and advertising intensity only when $(r_a - g)$ was large. In another indirect test of the accounting artifact interpretation, Salinger (1984) included measures of the ratio of advertising and R&D capital [computed as in Grabowski and Mueller (1978)] to the book value of assets in equation (4.2). He argued that linear relations between these variables and the market/book ratio are implied by the accounting artifact interpretation, while Bain's (1956) hypothesis predicts a positive coefficient for interaction terms involving the products of these variables and seller concentration. He found strong linear relations and insignificant coefficients of the interaction terms. On the other hand, the finding that firm profitability is more closely related to industry advertising intensity than to the firm's own advertising/sales ratio [Vernon and Nourse (1973), Mueller (1986)] appears inconsistent with the accounting artifact interpretation. [See also the discussion of advertising and entry in Subsection 6.3.]

Moreover, neither the endogeneity interpretation nor the accounting artifact interpretation imply that the advertising–profitability relation should vary with market structure or type of advertising, and numerous variations of this sort have been reported. As noted above, the relation is apparently different in consumer- and producer-good industries. Boyer (1974) found a *negative* advertising–profitability relation among U.S. local service industries, and White (1976) found a positive relation only in unconcentrated U.S. manufacturing industries. Porter (1976a, 1976b, 1979) analyzed consumer good industries in the United States and reported a positive advertising–profitability relation for convenience goods (for which buyers do not rely heavily on retailers for information) but not for non-convenience goods, for network television but not for other media [consistent with Boyer (1974), newspaper advertising was negatively related to profitability], and for leading firms but not for followers. All of this suggests that the endogeneity and accounting artifact interpretations are incomplete, but the reported results also seem inconsistent with the view that advertising is always associated with entry barriers.

4.5. Intra-industry differences

In a widely-cited early study of firm-level data, Hall and Weiss (1967) found that absolute firm size was positively related to profitability in U.S. manufacturing, even after controlling for industry characteristics, and Caves and Uekusa (1976) reported a similar relation for all but the largest Japanese firms. But a number of studies have failed to replicate this finding in U.S. data [e.g. Ornstein (1972), Imel, Behr and Helmberger (1972), and Vernon and Nourse (1973)],[27] and

[27] Indeed, Leonard Weiss reports (personal communication to the author, July 1986) that he has been unable to replicate the Hall and Weiss (1967) findings with more recent and more complete data.

negative firm size effects have been reported in studies of France [Jenny and Weber (1976)], West Germany [Neumann, Bobel and Haid (1979)], and large European and Japanese firms [Jacquemin and Saez (1976)]. [See also Salamon (1985).] There seems to be no support for a general relation between absolute firm size and profitability.

Gale (1972) found a strong positive relation between the weighted average market share and profitability of large U.S. firms in the 1963–67 period. This seems inconsistent with the general view that economies of scale in most industries are exhausted at output levels corresponding to low market shares [Scherer (1980, ch. 4)]. But Gale's results are consistent with the existence of long-lived efficiency differences among firms in the same industry, with more efficient firms attaining larger equilibrium market shares.[28] And Demsetz (1973, 1974) argued that efficiency differences provide an alternative explanation for the positive relation between concentration and profitability that many investigators had detected. [See Mancke (1974) for a related formal analysis stressing differences in luck.]

To illustrate this argument, consider a market in which firms produce a homogeneous product under constant returns to scale but have different unit costs. [This development follows Schmalensee (1987); see also Cowling and Waterson (1976) and Clarke, Davies and Waterson (1984).] Let us use the conjectural variation formalism to summarize conduct and asume that each firm acts as if it expects its rivals to increase their aggregate output by λ in response to a unit increase in its own output. Higher values of λ correspond to less intense rivalry. It can be shown that in long-run equilibrium, the accounting rate of return on assets (neglecting accounting biases) of a typical firm i is given by

$$r_i = \rho + [(1 + \lambda)/Ek_i]MS_i, \tag{4.4}$$

where ρ is the competitive rate of return, E is the market elasticity of demand, k_i is firm i's capital/revenue ratio, and MS_i is firm i's market share. If $k_i = k$ for all i, the industry-average rate of return is given by

$$r_I = \sum MS_i r_i = \rho + [(1 + \lambda)/Ek]H, \tag{4.5}$$

where H is the H-index of concentration. Concentration is endogenous here; it is determined by exogenous cost differences and industry conduct (λ). [Donsimoni, Geroski and Jacquemin (1984) discuss implications of this point.]

One way of interpreting Bain's (1951) concentration–collusion hypothesis is that λ in this model should be positively related to H across industries. And one

[28]Of course these results are also consistent with the hypotheses that collusion is common and that large firms tend to benefit disproportionately from it.

way of interpreting Demsetz' (1973, 1974) argument is that even if λ is the same for all firms in all industries, (4.5) predicts a positive correlation between concentration and profitability across industries. In fact, since E and k vary across industries, it predicts a weak correlation and is thus broadly consistent with the mixed results reported in Subsection 4.3. The rest of this subsection focuses on work aimed at distinguishing between these two view of the world. [See also Brozen (1982).]

Demsetz' view implies that only leading firms, with efficiency advantages, should earn supra-normal profits in concentrated industries. And Bain (1951, p. 320) noted that in his data, "Smaller firms tended to fare about the same regardless of industry concentration; the dominant firms in general had earnings rates that were positively influenced by concentration." Subsequent work by Collins and Preston (1969), Carter (1978), Porter (1979), Chappel and Cottle (1985), and others [see also Weiss (1974)] also supports

Stylized Fact 4.10

The profitability of industry leaders in U.S. manufacturing may be positively related to concentration; the profitability of firms with small market shares is not.[29]

The weak results obtained by many industry-level studies (Stylized Fact 4.5) may thus reflect in part averaging over small and large firms, with the presence of the former tending to mask a positive concentration–profitability relation involving the latter.

This pattern may not generally hold outside the United States. While Round (1975) finds a positive relation between concentration and differences between the profitability of large and small firms in Australia, Clarke, Davies and Waterson (1984) find no support for Stylized Fact 4.10 in the United Kingdom, and Neumann, Bobel and Haid (1979) report exactly contrary results for West Germany.

Two non-Demsetzian interpretations of Stylized Fact 4.10 have been offered. Bain (1956, p. 191) argued that the observed profitability of small firms is generally contaminated by failure to take full advantage of scale economies. Porter (1979), who found a number of differences between profitability equation estimates for industry leaders and for smaller firms, interpreted his findings in terms of the theory of strategic groups. This theory, which stresses barriers to

[29] Demsetz (1973, 1974) obtained similar results using absolute firm size instead of market share [see also Kilpatrick (1968)], but, as Daskin (1983) showed, these findings do not bear directly on the differential efficiency hypothesis that he presents. Note also Porter's (1979) finding that the inter-industry standard deviation in profit rates of leading firms was about 80 percent larger than the standard deviation of smaller firms' profit rates.

mobility that prevent other sellers from imitating the industry leaders, is supported by Newman's (1978) findings that the concentration–profitability relation holds only when leading firms are in the same businesses and that heterogeneity along this dimension lowers profits only in concentrated industries. [See also Oster (1982).]

Weiss (1974) argued that the most natural way to discriminate between the Bain and Demsetz views was to include both concentration and market share in the same equation. Results with specifications of this sort strongly support the following:[30]

Stylized Fact 4.11

In samples of U.S. firms or business units that include many industries, market share is strongly correlated with profitability; the coefficient of concentration is generally negative or insignificant in regressions including market share.

On the other hand, Demsetz' argument implies [see equation (4.4)] that a positive intra-industry relation between profitability and market share should generally hold in U.S. manufacturing. But, though positive relations are somewhat more common than negative ones in most periods [but see Schmalensee's (1987) results for 1972], intra-industry studies of the United Kingdom by Clarke, Davies and Waterson (1984) and of the United States by several authors unanimously support the following:[31]

Stylized Fact 4.12

Within particular manufacturing industries, profitability is not generally strongly related to market share.

These latter results suggest that estimates supporting Stylized Fact 4.11 may be dominated by a small number of industries with unusually strong positive

[30]Examples include Bothwell and Keeler (1976), Gale and Branch (1982), Martin (1983), Ravenscraft (1983), Bothwell, Cooley and Hall (1984), Harris (1984), Smirlock, Gilligan and Marshall (1984), Schmalensee (1985), Smirlock (1985), Mueller (1986), Ross (1986), and Kessides (1987). Shepherd (1974, ch. 4) and Thomadakis (1977) found positive coefficients of both share and concentration, but their sample selection procedures have been questioned [Hirschey (1985)]. On the other hand, Neumann, Bobel and Haid (1979) report a positive and significant concentration coefficient along with a significant *negative* market share effect in West German data. [See also Scott and Pascoe (1986).]

[31]See, for instance, Collins and Preston (1969), Comanor and Wilson (1974), Cattin and Wittink (1976), Porter (1979), Caves and Pugel (1980), Daskin (1983), and Schmalensee (1987). Using FTC lines of business data, Ross (1986) finds that market share is more strongly related to profitability in consumer goods industries than in producer goods, and Kessides (1987) strongly rejects the hypothesis that the market share–profitability relation is stable across industries.

relations between share and profitability. The results of Ross (1986) and, especially, Kessides (1987) tend to support this suggestion.

Collins and Preston (1969) find that differences between the profitability of large and small firms are not related to subsequent changes in concentration, as the Demsetz view would suggest. Salinger (1984) argues that, while Bain predicts that only (concentration × barriers) interaction terms should be positively related to market/book ratios, Demsetz predicts a positive relation involving concentration by itself. Salinger's data are inconsistent with both predictions.

Gale (1972) found that the impact of profitability on market share was positively related to concentration, suggesting that λ in (4.4) is an increasing function of concentration. But this finding has not proven robust.[32] Comanor and Wilson (1974, ch. 10) argued that the gap between the profits of large and small firms in consumer goods industries tended to be positively related to the industry advertising–sales ratio. The results of Caves and Pugel (1980), Ravenscraft (1983), Mueller (1986), and Schmalensee (1987) also support:[33]

Stylized Fact 4.13

The estimated effect of market share on profitability in U.S. manufacturing industries is positively related to the industry advertising/sales ratio.

Finally, recent work by Martin (1983), Kwoka and Ravenscraft (1986), Mueller (1986), Cotterill (1986), and Scott and Pascoe (1986) suggests a variety of complex firm-specific intra-industry effects not easily explicable by either Bain's or Demsetz' hypotheses. (See also the discussion of Stylized Fact 4.2, above.) While the industry may be a valid unit of analysis, systematic differences among firms deserve more attention than they have generally received.

4.6. Variability of profit rates

Stigler (1963) hypothesized that, since one would expect effective collusion to occur only in some concentrated industries, the cross-section variance of profit rates should be higher in concentrated than unconcentrated industries. He found

[32] Ravenscraft (1983) and Smirlock (1985) report negative coefficients of concentration–share interaction terms, for instance; see also Caves and Pugel (1980), Daskin (1983), Rhoades (1985), and Schmalensee (1987). But Mueler (1986) reports a robust negative coefficient of [(1-share) × concentration].

[33] Stonebraker (1976) finds that measures of small firm distress are positively related to large firm profitability when industry growth is controlled for. These same measures are positively correlated with industry advertising and R & D intensity. This suggests that Stylized Fact 4.13 can be extended to include R & D, as do the results of Mueller (1986).

only weak support for this hypothesis. But his data did suggest that the intertemporal variance in profit rates was lower in more concentrated industries. This finding is apparently at odds with the view that collusion/warfare cycles are not uncommon when concentration is high. Subsequent work has focused on the intertemporal variance, but no clear picture of the industry-level correlates of earnings variability have emerged. Perhaps this is because most studies have not attempted to control for differences in the variability of exogenous disturbances.

Confirming Stigler (1963), Gort and Singamsetti (1976) and Rhoades and Rutz (1982) find negative relations between concentration and the intertemporal variance in studies of U.S. manufacturing and banking, respectively, and Sullivan (1978) finds a negative relation for U.S. manufacturing firms between concentration and the beta-coefficient measure of systematic risk. But Round (1983) and Clarke (1984) report (weak) positive relations between concentration and intertemporal variability in data on Australia and the United Kingdom, respectively, and Shepherd (1974) reports a positive relation for a sample of large U.S. firms.

Winn (1977) finds that the simple correlation between concentration and profit variability is negative in U.S. data, but concentration has a positive, significant coefficient in a regression that includes a control for firm size [see also Daskin (1983)]. This implies a strong negative relation between firm size and the intertemporal variance. Such a relation has been reported by Jacquemin and Saez (1986) and other authors and is consistent with the strong negative relation generally observed between firm size and the cross-section variance [e.g. Hall and Weiss (1967)]:

Stylized Fact 4.14

Firm size tends to be negatively related to the intertemporal and cross-section variability of profit rates.

Of course, the intertemporal variance is not necessarily a good measure of riskiness. Geroski and Jacquemin (1986) find that large U.K. firms tend to have larger intertemporal variances than large French and German firms, but a smaller residual variances in regressions involving lagged profitability, and the residual variance is the better measure of the extent to which profits are unpredictable. Lev (1983) considered autoregressive models of both sales and accounting profits and found a strong negative relation between firm size and the residual variance, along with a positive relation between barriers to entry [as assessed by Palmer (1973)] and the serial correlation of sales and earnings. This last result is broadly consistent with the finding of Lustgarten and Thomadakis (1980) that the stock market responds more strongly to changes in accounting earnings announced by firms in concentrated industries, presumably because earnings changes are expected to be more persistent in those industries.

On the other hand, the profits of concentrated industries are not particularly stable over the business cycle, at least in the United States [Domowitz, Hubbard and Petersen (1986a, 1986b, 1987)], Germany [Neumann, Bobel and Haid (1985)], Japan [Odagiri and Yamashita (1987)]. The U.S. results in particular strongly support.

Stylized Fact 4.15

Price–cost margins tend to be more strongly pro-cyclical in more concentrated industries.

5. Prices and costs

Instead of studying profitability, some authors have chosen to analyze its basic components: price and cost. Subsection 5.1 surveys cross-section studies of prices, and Subsection 5.2 examines related work on cost and its determinants.

5.1. Price levels

Studies that compare price levels among geographically separated markets in the same industry are immune to the serious accounting problems that affect profitability studies, and one can expect that omitted market-specific variables are less important (and thus less likely to cause large biases) when attention is focused on a single industry. On the other hand, biased results may be obtained if adequate controls for exogenous determinants of cost are not included. The relation between concentration and price has been studied in numerous markets.[34] This work generally provides strong support for:[35]

[34] These include life insurance [Cummins, Denenberg and Scheel (1972)], banking services [Gilbert (1984) provides a survey], cement [Koller and Weiss (1986)], off-shore oil and timber auctions [Brannman, Klein and Weiss (1987)] air transportation [Bailey, Graham and Kaplan (1985)], newspaper advertising [Stigler (1964), Thompson (1984)], radio advertising [Stigler (1964)], groceries [Lamm (1981), Geithman, Marvel and Weiss (1981), Cotterill (1986)], gasoline [Marvel (1978), Geithman, Marvel and Weiss (1981)], and bond underwriting [Geithman, Marvel and Weiss (1981), Brannman, Klein and Weiss (1987)].

[35] It is worth noting, however, that Cummins, Denenberg and Scheel (1972) find that premiums for group life insurance do not rise with concentration. And, while Marvel (1978) finds that the lowest price charged for gasoline in any area rises with concentration, the relation involving the highest price is statistically insignificant.

Stylized Fact 5.1

In cross-section comparisons involving markets in the same industry, seller concentration is positively related to the level of price.

Price studies that search for critical concentration ratios [Geithman, Marvel and Weiss (1981), Thompson (1984)] obtain mixed results. And while some authors find a small concentration effect [e.g. Stigler (1964), Gilbert (1984), Bailey, Graham and Kaplan (1985)], others find quite large effects [e.g. Marvel (1978), Thompson (1984)].

The relation between concentration and price seems much more robust statistically than that between concentration and profitability. Since studies of price have fewer obvious weaknesses than studies of profitability, Stylized Fact 5.1 seems to provide the best evidence in support of the concentration–collusion hypothesis.[36]

Bloch (1947b) and Hazledine (1980) found that higher Canadian tariffs raised the ratio of Canadian to U.S. prices only when Canadian concentration was high. Nickell and Metcalf (1978) found that the ratio of the prices of proprietary to store-brand grocery products rose with seller concentration and the advertising/sales ratio in U.K. data. They interpreted store-brand prices as a control for costs, but their results also seem consistent with inter-brand differences in quality and consumer information.

Lamm (1981) reports that the three-firm concentration ratio is the best predictor of grocery prices in U.S. cities. Using market shares instead, he finds that the leader's share is insignificant, the shares of the second- and third-ranked firms are positive and highly significant, and the share of the fourth-largest firm is significantly negatively related to price. [Compare Kwoka's (1979) results for manufacturing PCMs.] On the other hand, Cotterill (1986) finds that the H-index is the best predictor of individual grocery stores' prices in Vermont.

Studies of the relation between prices and legal restrictions on local advertising by eyeglass vendors [Benham (1972)] and retail pharmacies [Cady (1976)] have produced strong results supporting:

Stylized Fact 5.2

Legal restrictions on local advertising in the United States are associated with higher retail prices.

[36]Domowitz, Hubbard and Petersen (1987) find that even though margins are more pro-cyclical in more concentrated industries, prices move counter-cyclically in concentrated producer-goods industries with high average margins. These findings are reconciled by observing that cost movements are especially strongly counter-cyclical for the latter industries.

Kwoka (1984) finds that legal restrictions are unrelated to the average quality of optometric services but negatively related to the dispersion in quality levels available. Albion (1983) finds no relation between advertising intensity at the manufacturing level and average retail markup across product categories in data supplied by a U.S. supermarket chain, but within categories he detects a negative relation, especially for highly-advertised, widely-used products.

5.2. *Costs and productivity*

The shortcomings of accounting data might account for the apparently stronger association between concentration and price than between concentration and profitability. Another explanation might be that costs in concentrated industries tend to be above minimum levels. This would occur if rent-seeking efforts to attain or protect monopoly power elevated costs in these industries substantially, if non-price competition were generally sufficiently intense to erode profits, or if high prices in concentrated industries tended to attract inefficiently small producers.

The first two hypotheses have apparently not been systematically tested. On the third hypothesis, Weiss (1976) found that the fraction of industry output produced in plants below (engineering/interview estimates of) minimum efficient scale in U.S. manufacturing industries was *negatively* related to concentration. Baldwin and Gorecki (1985) obtained similar results for Canada, as did Scherer, Beckstein, Kaufer and Murphy (1975) in a detailed analysis of 6 nations and 12 industries.

On the other hand, Baldwin and Gorecki (1985) also found that concentrated industries with strong tariff protection tended to have more inefficient capacity, all else equal, and Bloch's (1974b) analysis of Canadian/U.S. cost differences points in the same direction. This suggests that cost elevation may occur only in concentrated industries protected from entry by tariffs or other barriers.

A number of authors have employed data on U.S. local banking markets to study the hypothesis that concentration is positively related to cost. [See Gilbert (1984) for an overview.] Edwards (1977) found that banks in concentrated markets demanded 76 percent more labor than other comparable banks; he argued that this reflected the ability of managers in concentrated markets to exercise their preference for larger staffs. Hannan and Mavinga (1980), who also looked at spending on furniture and equipment and on office space, found a positive interaction between concentration and a dummy variable indicating dispersed ownership and thus (presumably) management control. But Smirlock and Marshall (1983) found that concentration was unrelated to labor demand when bank size was controlled for.

Time-series studies over long periods do not find a positive relation between changes in concentration and changes in costs. Instead, what emerges from the work of Peltzman (1977), Lustgarten (1984), and Gisser (1984) is

Stylized Fact 5.3

Over time, U.S. manufacturing industries that experience large increases *or* decreases in concentration tend to show above-average increases in productivity and below-average increases in price.

Gisser (1984) finds that increases in concentration have stronger estimated effects than decreases in initially unconcentrated industries, while the reverse is true in concentrated industries. These results suggest that major product, process, or marketing innovations are associated with large absolute changes in concentration, with the sign of the change depending on the source of the innovation.

Finally, one might hypothesize that costs in concentrated industries are too high because firms in such industries have weak incentives to resist union demands for supra-competitive wages. And Rose (1987) finds that trucking deregulation in the United States, which increased competition in that industry, lowered truck drivers' wages substantially. The large inter-industry literature has produced less clear-cut results, however.

Weiss (1966) observed that wage rates were positively related to both concentration and unionization in the United States in 1959; Phlips (1971) also found positive relations between wages and concentration for Belgium, France, and Italy. But Weiss found that when worker characteristics were added to his regression, the concentration effect vanished, and the estimated impact of unionization was weakened. The subsequent literature [surveyed by Dickens and Katz (1986)] reveals a more complex pattern. Dickens and Katz (1986) show that in U.S. data unionization and concentration are strongly correlated with a number of other factors that might plausibly affect wages (such as plant size, for instance), so that estimates of the effects of concentration are very sensitive to the data set and specification employed. Thus even though there do appear to be industry-level differences in wages that cannot be explained by differences in employee characteristics, the exact source of these differences is difficult to identify with available cross-section data.[37]

Pugel (1980b) found that a measure of excess profit per labor hour was more strongly positively related to U.S. industry-average wage rates than was con-

[37]Landon (1970) finds that printers' wages are *lower* in cities with more concentrated newspaper markets. But since newspaper unions are very strong and almost all U.S. newspaper markets very highly concentrated (many are monopolies), it is not clear how general this result is.

centration; Caves, Porter and Spence (1980) obtained similar results with Canadian data. Pugel argued that his estimates implied that labor on average captured 7–14 percent of excess profits. Karier (1985) added unionization to a standard PCM equation for 1972 and argued that his estimates implied that unions captured about 60 percent of excess profits. Voos and Mishel (1986) obtained a significant negative coefficient of a (concentration × unionization) interaction term in a study of supermarket profits. Their estimates implied that unions captured about 30 percent of profit increases due to concentration. And when Salinger (1984) allowed unionization to interact with (concentration × entry barriers) terms, he found that complete unionization served to eliminate essentially all excess returns. On the other hand, Clark (1984) and Domowitz, Hubbard and Petersen (1986b) find no support for the argument that increases in unionization have a stronger negative effect on profitability in high-concentration industries.

Several studies report important differences in the cyclical behavior of wages in concentrated and unconcentrated industries. Haworth and Reuther (1978) found that in U.S. industry-level wage equations with controls for worker characteristics, concentration was positively related to wages in 1958, when unemployment was high, but not in prosperous 1967. They obtained similar results for unionization and for a (concentration × unionization) interaction variable. Analyses of a 1958–81 panel data set for U.S. manufacturing by Domowitz, Hubbard and Petersen (1986a, 1986b) strongly confirm these findings. They first (1986a) observe that movements in aggregate demand affect PCMs more strongly than movements in industry sales, suggesting cyclical effects operating through economy-wide input markets. They then (1986b) find that, especially in producer goods industries, the coefficient of a (unionization × concentration × unemployment rate) interaction is negative, implying that PCMs of highly unionized, highly-concentrated industries are compressed relative to those of all industries on average during periods of low aggregate demand. These findings together indicate that labor costs in concentrated, unionized industries are less cyclically sensitive than average.

6. Structure and conduct

As Section 2 noted, seller concentration and advertising intensity are determined at least in part by market processes. Subsections 6.1 and 6.2, respectively, consider studies in which these quantities appear as dependent variables. Subsections 6.3 and 6.4 deal with two additional conduct-related phenomena: entry into and exit from industries, and the stability of market shares and market positions within industries.

6.1. Seller concentration

International comparisons of concentration levels for the 1950s [Bain (1966)] and 1960s [Pryor (1972)] point to

Stylized Fact 6.1

Rank correlations of manufacturing industries' concentration levels between industrialized nations are very high. Among large industrialized nations, concentration levels do not decline much with increases in the size of the economy.

The first sentence suggests that similar processes operate to determine concentration levels everywhere, while the second indicates that firm size and market size tend to be positively related internationally. On this latter relation and its structural basis, see Scherer, Beckstein, Kaufer and Murphy (1975).

Most studies of seller concentration begin with some measure of minimum efficient plant scale, often derived from the size distribution of existing plants. [See Subsection 3.5 and, for a valuable survey, Curry and George (1983).] Studies of Canada [Caves, Porter and Spence (1980)], the United Kingdom [Hart and Clarke (1980)], Belgium [Jacquemin, de Ghellinck and Huveneers (1980)], Japan [Caves and Uekusa (1976)], Germany [Neumann, Bobel and Haid (1979)], and the United States [e.g. Comanor and Wilson (1967), Strickland and Weiss (1976), Martin (1979a), Geroski, Masson and Shaanan (1987)] support

Stylized Fact 6.2

Levels of seller concentration are positively related to estimates of the market share of a plant of minimum efficient scale (MES).

Some studies report a positive relation between concentration and MES-based estimates of the capital required to build an efficient plant,[38] but generally this variable performs less well than MES/market estimates [Curry and George (1983)]. One might take Stylized Fact 6.2 as suggesting that concentration is determined largely by production technology, but there are several reasons why this inference is not completely justified.

First, the relation between MES/market ratios and concentration is much weaker in first-differences than in levels in U.K. [Hart and Clarke (1980)] and U.S. [Levy (1985), Martin (1979b)] data. This is consistent with the second part

[38]See, for instance, Comanor and Wilson (1967), Caves, Porter and Spence (1980), and Pagoulatos and Sorenson (1981).

of Stylized Fact 5.1 and with the generally weak relation between market growth and changes in concentration over time.[39]

Second, frequently-used measures of minimum efficient plant scale derived from the size distribution of existing plants are suspect for a variety of reasons, as Subsection 3.5 noted. Ornstein, Weston, Intriligator and Shrieves (1973) argue that the capital–labor ratio is a better indicator of the underlying technology; they and others [e.g. Collins and Preston (1969), Caves and Uekusa (1976)] provide strong support for

Stylized Fact 6.3

Capital-intensity is positively correlated with concentration among U.S. manufacturing industries.

Third, what ought to matter for seller concentration is scale economies at the firm level, not at the plant level. But Neumann, Bobel and Haid (1979) are almost the only ones to employ a measure of minimum efficient firm scale in this context. In U.K. data, concentration is positively related to the extent of multi-plant operations [Hart and Clarke (1980), Curry and George (1983)], but there is an element of tautology in this relation. [See also Scherer, Beckstein, Kaufer and Murphy (1975) on the determinants of plant scale and multi-plant operation.]

A number of authors have argued that there are scale economies in advertising, so that minimum efficient firm size and thus concentration in advertising-intensive industries should be higher than production scale economies suggest. Several studies have found positive relations between concentration and advertising intensity in U.S. data.[40] And Cowling, Cable, Kelly and McGuinness (1975) find a positive relation between advertising per firm and (survivorship estimates of) minimum efficient firm size, controlling for (survivorship estimates of) minimum efficient plant size. But other studies report no relation between advertising and concentration changes in multivariate studies [Hart and Clarke (1980), Curry and George (1983), Levy (1985)].

Mueller and Rogers (1980) find that when they divide advertising spending among media, only the ratio of television advertising to sales is positively related to changes in U.S. concentration over the 1958–72 period. This suggests that only television advertising involves important scale economies – presumably deriving

[39]See Martin (1979b), Mueller and Rogers (1980), Pagoulatos and Sorenson (1981), Levy (1985), Curry and George (1983); but see also Ornstein, Weston, Intriligator and Shrieves (1973).

[40]Examples include Ornstein, Weston, Intriligator and Shrieves (1973), Strickland and Weiss (1976), Caves, Porter and Spence (1980), Mueller and Rogers (1980), Pagoulatos and Sorenson (1981), and Connolly and Hirschey (1984); see also the studies discussed in Subsection 6.2, in which advertising intensity is the dependent variable.

from the large minimum outlays necessary to use this medium efficiently. But Lynk (1981) finds that concentration tended to fall in those U.S. industries that most rapidly increased the fraction of their advertising spending going to television when the cost of TV fell in the 1950s and early 1960s.

Fourth, most engineering/interview estimates of the importance of firm-level scale economies suggest that existing levels of concentration, particularly in the United States, are higher than they would be if the leading sellers were of minimum efficient scale [Scherer (1980, ch. 4)]. This, again in combination with the second part of Stylized Fact 6.1, suggests that other forces have operated to increase concentration.

If a number of firms have attained efficient scale thus have constant unit cost, it may be reasonable to model their growth rates and sizes as determined by stochastic processes. In the most famous model of this sort, called Gibrat's Law, period-to-period changes in the logarithm of firm size are independent, normal random variables. It follows that the distribution of firm sizes will tend toward lognormality, with increasing variance (and thus rising concentration) over time. Depending on assumptions about growth, birth, and death, stochastic models can also generate the Pareto and other skewed size distributions; see Ijiri and Simon (1977) for an overview.

Studies of actual size distributions in the United States [Quandt (1966), Silberman (1967), Kwoka (1982)] and the United Kingdom [Clarke (1979), Davies and Lyons (1982)] concentrate on the lognormal and Pareto distributions and generally support:

Stylized Fact 6.4

Size distributions of firms and plants are highly skewed; all families of distributions so far tried fail to describe at least some industries well.

Neither the lognormal nor the Pareto consistently outperforms the other [Curry and George (1983)].

Early studies of firms' growth rates in the United States [Hymer and Pashigian (1962), Mansfield (1962)] generally supported Gibrat's Law. But several recent studies, using new, large data sets, have found that mean growth rates decline with firm size and age [Evans (1987a, 1987b), Hall (1987), Dunne, Roberts and Samuelson (1988a, 1988b)], as does the probability of failure.

Stylized Fact 6.5

In U.S. data, mean firm growth rates and failure probabilities decline with firm size and age. The standard deviation of growth rates declines with size, but less rapidly than the reciprocal of the square root of size.

Results on mean growth abroad are mixed. Singh and Whittington (1975) found a weak positive relation between size and mean growth in the United Kingdom, while Jacquemin and Saez (1976) found a negative relation for large European firms, but no relation for large Japanese firms.

To understand the significance of the second part of Stylized Fact 6.5, note that if Gibrat's Law held exactly and large firms were simply collections of small firms with uncorrelated growth, the standard deviation of growth rates would decline as [firm size]$^{-1/2}$. The slower decline observed in many studies suggests that large firms specialize in correlated activities. But Daskin (1983) finds that larger U.S. firms have more stable growth even after controlling for diversification patterns.

It is important to recognize that variations in growth rates alone do not determine firm size distributions: entry, exit, and mergers may also be important. Until recently, antitrust restrictions on horizontal mergers in the United States were quite strict, and such mergers were a relatively unimportant source of increases in concentration. For the United Kingdom, however, Hannah and Kay (1977) conclude that, while stochastic rate growth differences were an important source of increases in concentration, mergers were much more important. Muller (1976) also argues that mergers were important in maintaining high concentration in West Germany despite rapid market growth. While controversy remains, it seems safe [Curry and George (1983)] to assert

Stylized Fact 6.6

Outside the United States, mergers have been an important source of increases in seller concentration.

6.2. *Advertising intensity*

Numerous regression analyses in which advertising intensity is the dependent variable have confirmed the positive advertising–profitability relation discussed in Subsection 4.4.[41] Telser (1964) found that advertising/sales ratios were unrelated to concentration in the United States, but Mann, Henning and Meehan (1967) found a positive relation. [Comanor and Wilson (1979) discuss the interpretation of this relation.] Subsequent studies generally support the existence

[41]See, for instance, Comanor and Wilson (1974), Strickland and Weiss (1976), Martin (1979a, 1979b), Farris and Buzzell (1979), Caves, Porter and Spence (1980), and Pagoulatos and Sorenson (1981); but see Martin (1983), who reports a negative relation.

of a positive concave or inverted-U relation:[42]

Stylized Fact 6.7

Among consumer goods industries, advertising intensity increases with concentration at low levels of concentration; the relation may vanish or change sign at high levels of concentration.

The observation that advertising intensity may decrease with increases in concentration in concentrated industries suggests, somewhat implausibly, that sellers in such industries not infrequently collude to reduce advertising outlays.

Ornstein (1977) argued that the positive relation between advertising intensity and concentration in the United States was roughly as strong in producer goods industries as in consumer goods industries. But Bradburd (1980) showed that when these two groups were defined more strictly (using the fraction of sales made to retailers), a positive, concave relation emerged for consumer goods, but no relation was apparent for producer goods. [See also Weiss, Pascoe and Martin (1983).] Buxton, Davies and Lyons (1984) posited different relations for pure producer and consumer goods industries and used data on the fraction of sales made to retailers to estimate an interactive model for industries selling both. They reported no relation for producer goods in U.K. data and an inverted-U relation for consumer goods.[43]

Within consumer good industries, Comanor and Wilson (1974) found that leading firms had higher advertising/sales ratios than followers when the industry advertising/sales ratio was low, but that leaders spent a smaller percentage of revenue on advertising when the industry ratio was high. The latter pattern has been found by Lambin (1976) in European data and by Farris and Buzzell (1979) in a study of advertising and promotion outlays.

6.3. Entry and exit

Basic price theory implies that entry will occur if and only if potential entrants expect post-entry prices to be at least equal to their costs. Bain (1956) argued that entrants' expectations are determined by the height of pre-entry profits relative to

[42] Examples include studies of Canadian [Caves, Porter and Spence (1980)], U.K. [Cowling, Cable, Kelly and McGuinness (1975)], and U.S. data [Strickland and Weiss (1976), Martin (1979a, 1979b), Pagoulatos and Sorenson (1981), and Connolly and Hirschey (1984)]; see also Lambin (1976).

[43] Arterburn and Woodbury (1981) studied the frequency with which price was mentioned in national magazine ads for 37 consumer goods in the United States in the early 1970s. They found that price was more likely to be mentioned the higher was the industry PCM and the lower was concentration. The coefficient of the industry advertising/sales ratio was negative for convenience goods but positive for non-convenience goods.

structural entry barriers; recent theoretical work has shown that strategic behavior of incumbent firms can also affect entrants's expectations.

Official data usually considerable entry (and exit) in almost all industries by firms that attain (or relinquish) small market shares [Baldwin and Gorecki (1987), Dunne, Roberts and Samuelson (1988a, 1988b)], so empirical work seeks to explain differences in the importance of entry across industries. This has been measured by the absolute or relative, gross or net change in the number of firms [Mansfield (1962), Orr (1974a), Deutsch (1975), Gorecki (1975, 1976)], by the occurrence of substantial entry [Harris (1976a)], and by the market share achieved by entrants [Harris (1976b), Masson and Shaanan (1982), MacDonald (1986)]. [See Geroski (1983) for useful survey and Dunne, Roberts and Samuelson (1988a, 1988b) on relations among alternative measures of entry and exit.] While the last of these seems the most satisfactory, Hause and Du Rietz (1984) use both the number of entrants and their share of industry employment and report broadly similar results for both measures in Swedish data.

Mansfield (1962) studied four industries (steel, petroleum, tires, and autos) over time and found that the ratio of new firms at the end of a decade to firms at the start was positively related to profitability during the period and negatively related to the capital cost of a firm of minimum efficient scale. While some subsequent studies also report a positive effect of profits [Harris (1976a, 1976b), Masson and Shaanan (1982)], insignificant coefficients seem at least as common [Orr (1974a), Deutsch (1975), Gorecki (1976)]. Since behavior designed to deter entry can be expected to lower pre-entry profits, the lack of a robust relation between pre-entry profits and the level of entry may not be terribly surprising.

Similarly, while some studies [Deutsch (1975), Hause and Du Rietz (1984), MacDonald (1986)] report a positive relation between growth and entry, others find no relation [Orr (1974a), Harris (1976b), Masson and Shaanan (1982)]. Gorecki (1976) finds that growth is positively related to foreign entry into Canadian manufacturing industries but unrelated to domestic entry. [Gorecki (1975) reports significant differences between the correlates of entry into U.K. industries by new firms and by those established elsewhere.] On theoretical grounds one would expect anticipated and unanticipated growth to have different effects on entry, but Bradburd and Caves (1982) did not find this distinction helpful in explaining profitability differences.

Harris (1976b) examined entry during the 1950–66 period into U.S. industries for which structural entry barriers had been assessed by Bain (1956) and Mann (1966). Nine of the 18 industries with above-average pre-entry profitability experienced substantial entry; of the nine industries classed has having "very high" entry barriers, four experienced substantial entry.[44] The Bain/Mann

[44] Harris (1976b) found that substantial entry tended to lower profitability, though industries with very high barriers (which tended to have high advertising/sales ratios) were still more profitable than average even after substantial entry.

judgements about overall entry barriers did not seem to predict the actual occurrence of entry terribly well.

Following Mansfield (1962), a number of authors have employed measures of the market share or capital cost of a plant of minimum efficient scale to capture the impact of scale-related entry barriers. And, as in the profitability literature, both are rarely significant, but at least one usually is:[45]

Stylized Fact 6.8

Measures of scale economies or capital requirements tend to be negatively related to entry.

Harris (1976b) found substantial entry into only two of the 12 industries in which product differentiation was held to be an important source of barriers by Bain (1956) and Mann (1966), and into only one of the seven of these industries with above-average profits. Since advertising intensity is highly correlated in this sample with high estimated product differentiation barriers, this is consistent with other work that supports:[46]

Stylized Fact 6.9

Advertising intensity is negatively related to entry in manufacturing industries.

Some studies have included concentration in equations designed to explain entry; see Kessides (1986) for a discussion. Positive [Deutsch (1975)], negative [Orr (1974a), Kessides (1986)], and insignificant [Harris (1976a)] coefficients have been reported. Hause and Du Rietz (1984) find the existence of a cartel agreement to be negatively related to entry in Swedish manufacturing industries.

Mansfield (1962) also studied the incidence of exit. He found that the fraction of firms leaving an industry was negatively related to the ratio of average size to minimum efficient scale and to industry profitability. Marcus (1967) found that

[45]Studies finding a negative effect of capital requirements include Orr (1974a), Gorecki (1976), Hause and Du Rietz (1984), and MacDonald (1986). [Gorecki (1976) finds a significant effect only for domestic entrants. Hause and Du Rietz (1984) use the mean employment in plants built by entrants.] Studies by Gorecki (1975), Harris (1976a), and Masson and Shaanan (1982) find a negative effect of MES/market.

[46]Support is provided by studies of Canadian [Orr (1974a), Gorecki (1976)] and U.S. data [Deutsch (1975), Harris (1976a), Masson and Shaanan (1982)]. [See also Kessides (1986).] But MacDonald (1986) fails to detect an advertising effect in his study of U.S. food processing industries during the 1970s. And Gorecki (1975) finds that specialist entry into U.K. manufacturing industries is negatively related to advertising per firm but positively related to the industry advertising/sales ratio.

the incidence of accounting losses was a better predictor of exit than the average level of profits, and he found that small, young firms were the most likely to incur losses. Dunne, Roberts and Samuelson (1988a, 1988b) report exit rates that decline sharply with age. Studies of exit from unprofitable, declining industries by Caves and Porter (1976) and Harrigan (1986) suggest that exit is delayed by the existence of tangible and intangible industry-specific assets, as well as by managerial and strategic factors.

Finally, Dunne, Roberts and Samuelson (1988b), using plant-level U.S. Census panel data, find systematic differences between entry by new and by diversifying firms. Using similar Canadian data, Baldwin and Gorecki (1987) support this finding and also report differences related to the nationality of diversifying entrant firms.

6.4. Share and rank stability

Mueller (1986) observes that in 41 percent of 350 U.S. manufacturing industries with essentially the same official definition in 1950 and 1972, the leading firm was the same in both years. (Note that industries in which technical change has been important are underrepresented in samples of this sort.) He finds that leaders' market shares tend to persist over long periods as well [see also Shepherd (1974)]. While stable market shares and firm ranks are consistent in principle with either collusion or competition, most would argue that unstable shares and ranks are inconsistent with effective collusion. Unfortunately, data limitations have kept the empirical literature on rank and share stability thin.

In an early study of rank changes, Mansfield (1962) found that small firms were less likely to grow to exceed the size of previously larger rivals in older, more concentrated industries. A number of studies of rank changes among leading banks (mobility) and changes in the identity of the leaders (turnover) have reported negative relations between concentration and both mobility and turnover [Gilbert (1984)]. But, as Marlow, Link and Trost (1984) point out, the measures used in these studies are counts of changes, and least squares is inappropriate for such discrete, bounded dependent variables. Using non-linear methods with data on U.S. savings and loan associations, they find no relation between concentration and turnover. While they do detect a weak negative relation between concentration and mobility, it has very little explanatory power.

Studies of market share stability generally employ sums across firms of the absolute values of either absolute or relative changes in shares to measure instability. Gort (1963) found a positive relation between concentration and share stability but no relation involving (judgemental estimates of) product differentiation or profitability. His work, along with the studies discussed in the previous paragraph and the results of Caves and Porter (1978) and Heggestad and Rhoades (1978), points toward

Stylized Fact 6.10

In manufacturing industries and local banking markets in the United States, market shares tend to be more stable the higher is concentration.

Telser (1964) found a negative relation between advertising intensity and stability in a small sample of industries. While Reekie (1974) found a negative relation for markets within two product classes, he detected a positive relation within two others. And, consistent with Gort's (1963) results, Lambin (1976) and Caves and Porter (1976) found no relation between advertising/sales ratios and share stability.

7. Conclusions and implications

I have argued that inter-industry research in industrial organization should generally be viewed as a search for empirical regularities, not as a set of exercises in structural estimation. And I have attempted to show that research in this tradition has indeed uncovered many stable, robust, empirical regularities. Inter-industry research has taught us much about how markets *look*, especially within the manufacturing sector in developed economies, even if it has not shown us exactly how markets *work*.

This literature has also produced an impressive, if implicit, agenda for future research. It seems difficult to reconcile the set of Stylized Facts discussed above with any familiar simple view of the world; some Stylized Facts seem difficult to reconcile with each other. Work in some areas has produced no clear picture of the important patterns in the data, and non-manufacturing industries have not received attention commensurate with their importance. The literature is full of interesting results that beg for attempts at replication.

But cross-section studies are limited by serious problems of interpretation and measurement. Future inter-industry research should adopt a modest, descriptive orientation and aim to complement case studies by uncovering robust empirical regularities that can be used to evaluate and develop theoretical tools. Finally, it is important to note that much of the most persuasive recent work relies on non-standard data sources, particularly panel data (which can be used to deal with disequilibrium problems) and industry-specific data (which mitigate the problem of unobservable industry-specific variables).

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