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ANNALES D'ÉCONOMIE ET DE STATISTIQUE. – N° 31 — 1993

Measuring the Importance of Sunk Costs

Timothy F. BRESNAHAN and Peter C. REISS*

ABSTRACT. – This paper devises new methods for measuring firms' sunk costs and their effect on market structure. It builds on our earlier work that estimated the size of unobserved fixed costs from *entry thresholds*, the amount of market demand necessary to support a new entrant. Here we show that differences between *entry* and *exit thresholds* provide information about the extent of sunk costs and their option value. The latter half of the paper uses panel data on the location of rural dentists' practices to estimate entry and exit thresholds. We find dentists' exit thresholds are well below their entry thresholds. It thus appears dentists sink significant costs. A simulation suggests that dentists in large rural markets incur greater sunk costs than those in small markets.

Mesurer l'importance des coûts irrecouvrables

RÉSUMÉ. – Ce papier présente de nouvelles méthodes pour mesurer les coûts irrecouvrables des firmes et leur effet sur la structure de marché. Il prolonge notre précédent travail, qui estimait la taille des coûts fixes non-observés à partir des *seuils d'entrée*, le niveau de la demande nécessaire suppose un nouvel entrant. Ici nous montrons que la différence entre les seuils *d'entrée* et de *sortie* fournit de l'information sur le niveau des coûts irrecouvrables et leur valeur d'option. La seconde partie du papier utilise des données de panel sur la localisation des cabinets dentaires ruraux pour estimer les seuils d'entrée et de sortie. Nous trouvons que les seuils de sortie des dentistes sont inférieurs à leurs seuils d'entrée. Il apparaît ainsi que ces dentistes ont des coûts irrecouvrables significatifs. Une simulation suggère que les dentistes dans ces marchés ruraux étendus ont des coûts irrecouvrables plus élevés que ceux travaillant dans des marchés plus petits.

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1 Introduction

What determines why some markets have many firms and others few? This is one of the most important questions in industrial organization economics. Many theories emphasize the importance of fixed costs. Fixed costs increase scale economies, which in turn increase industry concentration and profits. Empirical researchers have tried to test these predictions by regressing industry concentration and profit data on accounting measures of fixed costs. These studies interpret positive coefficients as evidence that fixed costs increase industrial concentration and firm profits. Critics challenge both the causal logic of these regressions and the assumption that accounting costs correctly measure economic costs (e.g. SCHMALENSEE [1989]). Recently, BERRY [1992], BRESNAHAN and REISS [1987, 1990, 1991 *b*], REISS and SPILLER [1989], and SUTTON [1991] have developed new empirical techniques to measure the effect of fixed costs on market structure. These models infer the importance of fixed costs by observing how quickly firms enter and exit in response to changes in market demand. Though limited to a few industries, these studies conclude that scale economies do explain systematic variation in local market concentration. Some of these studies also suggest that entry significantly reduces price-cost margins.

Although scale economies can explain long-run trends in industry concentration, they cannot easily explain short-run changes in industry structure. Recent industry studies by BRESNAHAN and REISS [1987, 1990], DUNNE, ROBERTS and SAMUELSON [1988], HAUSE and DU REITZ [1984], PAKES and ERICKSON [1989] and others find that while long-run industry structure is relatively stable, there is considerable short-run turnover in most industrial markets. This turnover exhibits several curious patterns. For instance, while entry and exit rates are positively correlated across markets, they differ between growing and declining markets. Several recent theoretical papers argue that sunk costs can explain these patterns (e.g., HOPENHAYN [1992 *a*, 1992 *b*]). These papers build on the insights of Marshall's long-run perfect competition model. In Marshall's model, firms enter when demand covers fixed plus variable costs and exit only when price falls below variable costs. Recent models extend Marshall's logic by showing how sunk costs affect firms' short-run entry and exit decisions. Papers by Dixit [1989] and MCDONALD and SIEGEL [1985], for instance, show how demand uncertainty and sunk costs can widen the range of prices between which a firm will exit and enter. To date, however, few empirical studies have tested the predictions of these theories [1].

Theorists and empiricists also have tried to analyze how fixed and sunk costs affect the behavior of oligopolists and oligopoly market structure. In contrast to the literature on competitive markets, most of the insights we have

1. Exceptions can be found in the international trade literature. See BALDWIN [1988] and AIT-SAHALIA [1992].

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come from simple two-period models. DIXIT [1979] and SCHMALENSEE [1981], for example, develop two-period duopoly models that show sunk investments can deter otherwise profitable entrants. Other researchers have extended these models to analyze the commitment value of sunk costs. They generally conclude that sunk costs increase market concentration by widening the range of prices between which a firm will enter and exit. In more recent work, MACLEOD [1987] and BERGIN and MACLEOD [1989] show that sunk costs may have the opposite effect. That is, they may encourage entry. SUTTON [1991] reconciles their predictions with those of earlier models by distinguishing between exogenous and endogenous factors that affect decisions to sink costs.

To summarize, theoretical models suggest that sunk costs have important effects on market structure. Typically, they predict that sunk costs will increase industry concentration. Sunk costs also create hysteresis. That is, they widen the range of prices between which firms will enter and exit. Empiricists have had limited success in testing these predictions, primarily because they lack firm-level data on sunk costs. To date, most studies have adopted the approach of static profit-concentration studies. Specifically, they regress market concentration or changes in market concentration on accounting data or asset salvage values (e. g. GHEMAWAT and CAVES [1986], HAUSE and DU REITZ [1984], KESSIDES [1986, 1990] and ROSENBAUM *et al.* [1992]). Critics also question the reliability of these regressions.

This paper proposes a different method for assessing the importance of unobserved sunk costs. This method can use, but does not require, data on fixed costs and asset salvage values. The method builds on BRESNAHAN and REISS' [1987] concept of an entry threshold, a level of demand just needed to induce entry. Here we distinguish between entry and exit thresholds. Following Marshall's intuition, we show that the difference between entry and exit thresholds provides a measure of the importance of sunk costs. The second half of the paper uses panel data on the location of dentists in rural U.S. counties during 1980 to 1988 to estimate entry and exit thresholds. Our econometric models treat the entry and exit decisions of these dentists as forward-looking decisions. We find that sunk costs impede exit from declining markets and that they are greatest in monopoly markets. For instance, we estimate that a dentist will set up a monopoly practice only if he or she expects a town to have 1 400 or more permanent residents. Should the dentist enter and demand suddenly decline, we find that the dentist will stay until the town has less than 800 residents. We attribute this gap to location-specific sunk costs.

2 A Model of Firm Turnover with Sunk Costs

This section develops models that show how the level of market demand affects a firm's ability to cover its costs. The predictions of these models

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guide our econometric specifications. Because dynamic oligopoly models raise difficult conceptual issues, we begin by considering how fixed and sunk costs affect the entry and exit decisions of a potential monopolist.

2.1. Entry and Exit by a Monopolist

Following our earlier work, we presume that market demand at time t equals

$$\mathrm{Q}_t = d(\mathrm{Z}_t, \mathrm{P}_t)\,\mathrm{S}_t.$$

Here $d(\mathrm{Z}_t, \mathrm{P}_t)$ represents the stationary demand function of a representative consumer, Z_t is a set of nonstochastic demographic variables affecting tastes, P_t is product price, and S_t is the number of demanders in the market. We assume that all consumers have the same demand function to simplify the exposition.

To simplify the supply side, we assume that market demand is never large enough to support two entrants. That is, that the market is a natural monopoly. Although the market can support at most one firm, there may be many potential entrants. We use the term "the supply of potential entrants" to refer to the set of all possible entrants. Each potential entrant incurs four types of costs. On entry, an entrant incurs an initial sunk cost F. It then has per period fixed costs f and variable costs cq_t, where q_t it output [2]. If an entrant ceases production, we assume it exists forever. Upon, exiting, it pays an exit cost L. These assumptions imply an entrant's per period profits equal

$$\pi_t = \begin{cases} (\mathrm{P}_t - c)\,q_t - f - \mathrm{F} & \text{when } q_{t-1} = 0 \text{ and } q_t > 0 \\ (\mathrm{P}_t - c)\,q_t - f & \text{when } q_{t-1} > 0 \text{ and } q_t > 0 \\ -\mathrm{L} & \text{when } q_{t-1} > 0 \text{ and } q_t = 0. \end{cases}$$

We now consider the question, at what date t will a firm enter and monopolize this market? The answer depends both on what we assume about the supply of potential entrants and the structure of the entry game. Consider first what happens when there are an infinite number of equally capable entrants (*i. e.*, the supply of entrants is perfectly elastic). Assume each entrant discounts the future at a constant rate δ per period, and has the same expectations about the future (as represented by the expectation operator E). At time t, discounted expected future profits from entry equal

2. This formulation presumes that firms do not sink costs over time. We have made little progress solving models where firms sequentially sink costs.

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$$\Pi_t = \mathrm{E}_t \sum_{r=t}^{\infty} \delta^{r-t} \pi_r = \mathrm{E}_t \sum_{r=t}^{\infty} \delta^{r-t} [\mathrm{I}(q_r)((\mathrm{P}_r - c) q_r - f) - \mathrm{I}(q_{r-1})(1 - \mathrm{I}(q_r))\mathrm{L}] - \mathrm{F}. \tag{1}$$

In this expression $\mathrm{I}(q_r)$ is a zero-one indicator equal to one when $q_r > 0$. Provided the potential entrants behave noncooperatively, a firm enters when expected discounted profits equal or exceed zero, *i. e.* $\Pi_t \geq 0$. Unlike static entry threshold conditions, this inequality depends on the firm's expectations about technological change, demand and competitors' actions. To simplify our analysis, we assume that the firm is only uncertain about the number of customers it can attract. Specifically, we assume that in period t the firm knows S_t but does not know future market sizes.

Our analysis changes somewhat when the supply of entrants has some slope (e. g., potential entrants have different fixed, sunk and variable costs) or is perfectly inelastic (e. g., there is only one potential entrant). Consider, for instance, a single potential entrant's entry decision. Because it does not face competition, it can afford to delay entry even though discounted expected profits may be positive. To see this, suppose it knows it can earn profits of 1, –1, –1, and 10 over the next four periods and that exit is extremely costly. If it enters today and does not discount the future, it earns profits of 9 over four periods. By waiting until period four, however, it can do better. This example suggests that in the single entrant case the firm enters at date t only when Π_t is greater than all discounted expected profits from delayed entry. To avoid such complications, we shall work with the entry threshold condition $\Pi_t \geq 0$. Thus, our models presume that there are several potential entrants.

We note that the entry threshold condition $\Pi_t \geq 0$ differs from the usual static condition $\pi_t \geq 0$ in that a firm may enter today, even though in some future periods it may lose money (*i. e.* $\pi_\tau < 0$ for some $\tau > t$). A key empirical implication of this is that one can no longer draw inferences about the sign of π_t from the observation that a firm produces in period t. Instead, one must recognize that the decision to produce today depends on the firm's future production and exit policies; that is, $\mathrm{I}(q_\tau)$, $\tau > t$. Thus, to compute whether a firm should enter today, we also must calculate what it expects to do tomorrow.

The presence of sunk costs causes a firm to treat exit differently from entry. Having entered and sunk F, the firm now ignores F and exits when it perceives that future losses exceed exit costs. Letting $\mathrm{V}_\tau = (\mathrm{P}_\tau - c)\, d(\mathrm{P}_\tau)$ denote percustomer monopoly variable profits, a firm that entered before date t exist at date t when

$$\mathrm{V}_t \mathrm{S}_t - f + \sum_{r=t+1}^{\infty} \delta^{r-t} \mathrm{V}_\tau \mathrm{E}_t \mathrm{I}(\mathrm{Q}_\tau) \mathrm{S}_\tau - \sum_{\tau=t+1}^{\infty} \delta^{\tau-t} \mathrm{E}_t [f \mathrm{I}(\mathrm{Q}_\tau) + \mathrm{I}(q_{\tau-1})(1 - \mathrm{I}(q_\tau))\mathrm{L}] \leqq -\mathrm{L}$$

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This expression shows that future $I(Q_\tau)$ also affect the exit decision. Today's decision rule thus requires the recursive solution of future policies, a process that quickly becomes intractable for anything but simple time-dependent state processes (e. g., Rust [1992]). To interpret what these conditions imply about entry and exit decisions we now examine several special cases of this model.

2.1.1. *Stationary Markets*

In our previous static entry models (Bresnahan and Reiss [1987, 1990]), we showed that the level of demand S^e at which a new firm would enter a market equals the ratio of unobserved fixed costs to per capita variable profits, or

$$S^e = \frac{F}{V}.$$

In our empirical applications, we interpreted F as a firm's long-run fixed costs. An obvious first question then is, when does the dynamic entry threshold condition $\Pi_t \geqq 0$ reduce to a similar ratio? Although there may be many situations, one set of condition is as follows. Assume that $Z_\tau = Z$ and that expectations about future market sizes are unaffected by past production decisions. Absent an effect of today's decisions on the future, a monopolist will choose a stationary optimal price policy $P_\tau = P$. This implies constant variable profits per customer $V_\tau = V$. If in addition $E_t(S_{t+1}|S_t, S_{t-1}, ...) = S_t$, one can show a monopolist will enter whenever S_t exceeds the threshold

$$S_t^e = \frac{F + (f/1-\delta)}{V/(1-\delta)} = \frac{F(1-\delta)+f}{V} \tag{2}$$

This threshold resembles the static monopoly entry threshold above. It is simply the ratio of present discounted fixed and sunk costs divided by present discounted per customer variable profits.

What of exit thresholds? In our static models, the two were the same. Here the monopolist exits when $\Pi_t + L = 0$, or the number of representative consumers equals

$$S_t^x = \frac{f - (1-\delta)L}{V}.$$

This threshold is the ratio of discounted avoidable fixed costs plus exit costs divided by discounted variable profits. It equals the entry threshold only when there are no sunk costs $L = F = 0$, when $-L = F$, or the firm does not discount the future $\delta = 1$. Generally, the difference in the thresholds

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$$S_t^e - S_t^x = (1-\delta)\frac{F+L}{V} \tag{3}$$

provides a measure of the importance of sunk costs. This expression thus suggests how we might estimate fixed and sunk costs from demand data. Given observations or estimates of S^e and S^x, the difference between them equals the unobservable cost ratio on the right hand side of (3). Thus, it appears we can measure fixed plus sunk costs up to a scale factor.

2.1.2. *A Two-Period Model*

The previous model contains strong simplifying assumptions. Most of these assumptions are not easily relaxed in multi-period models. This subsection explores simpler two-period models that show how demand uncertainty affects firms' entry and exit decisions. These models closely resemble the models we later estimate.

Following our earlier multi-period model, we assume the market can support at most one entrant. Each potential entrant can sink costs in either of two periods. We set $\delta = 1$. At the start of period 1, a potential entrant's profits equal

$$\begin{aligned}\Pi_1 = {} & [V_1 S_1 - f]\, I(Q_1) + [V_2 S_2 - f]\, I(Q_2) \\ & - I(Q_1)\,(1 - I(Q_2))\, L - [I(Q_1) + I(Q_2)\,(1 - I(Q_1))]\, F.\end{aligned} \tag{4}$$

This specification assumes that the monopolist only incurs L if it exits before the end of period 2. The monopolist enters in period 1 when first period revenues contribute to sunk costs, *i. e.* $S_1 \geqq f/V_1$ and second period demand allow the recovery of the remaining costs, or

$$S_1 \geq \left(\frac{f - V_2 S_2}{V_1}\right) I(Q_2) + \frac{f + F + (1 - I(Q_2))\, L}{V_1}. \tag{5}$$

Notice that these conditions depend on the firm's period 2 production indicator, $I(Q_2)$. Thus, to determine whether $I(Q_1) = 1$, we must solve for $I(Q_2)$.

Having sunk F in period 1, the monopolist produces in period 2 if its second-period profits equal or exceed the exit cost L. That is,

$$I(Q_2) = 1 \Leftrightarrow S_2 \geq \frac{f - L}{V_2}. \tag{6}$$

Combining this condition and (5), we can now relate the first period entry threshold S_1^e to future demand S_2,

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$$(7)\qquad S_1^e = \begin{cases} \dfrac{f}{V_1} & \dfrac{f+F}{V_2} \le S_2 \\ \dfrac{2f+F-V_2S_2}{V_1} & \dfrac{f-L}{V_2} \le S_2 \le \dfrac{f+F}{V_2} \\ \dfrac{f+F+L}{V_1} & S_2 < \dfrac{f-L}{V_2}. \end{cases}$$

This expression shows how future demand affects the willingness of the firm to enter. Future demand did not appear to affect entry thresholds in the previous example because we assumed that the firm expected future demand to equal today's demand. (Here this occurs in the middle segment $(f-L)/V_2 \le S_2 \le (f+F)/V_2$ where $I(Q_2) = 1$.)

The first period exit thresholds, S_1^x, also depend on S_2. When a monopolist sinks F before period 1, and is therefore an incumbent in period 1, it produces when $S_1 \ge f/V_1$. Unlike a potential entrant, however, it also produces when $S_1 < f/V_1$. This occurs when profits exceed the exit cost —L, or

$$(8)\qquad S_1 \geqq \left(\frac{f-V_2S_2}{V_1}\right) I(Q_2) + \frac{f-L}{V_1}.$$

As with the entry threshold, the exit threshold depends on $I(Q_2)$. Paralleling (7) above

$$(9)\qquad S_1^x = \begin{cases} \dfrac{f}{V_1} & \dfrac{f-L}{V_2} \le S_2 \\ \dfrac{2f-L-V_2S_2}{V_1} & \dfrac{f-L}{V_2} \le S_2 \le \dfrac{f+F}{V_2} \\ \dfrac{f-F-L}{V_1} & S_2 > \dfrac{f+F}{V_2}. \end{cases}$$

Figure 1 displays how the first period entry and exit thresholds vary with current and future market size. To illustrate the empirical implications of this figure, we have superimposed two hypothetical sets of markets on it. Each point, or •, represents the first and second period market sizes associated with a particular market. The horizontal scatter contains markets that have similar second period market sizes and different first period market sizes. From the variation in first period market size, we can estimate S_1^e and S_1^x. Consider, for example, the information provided by the market on the far left. Because it has $S_1 < S_1^x$, we never observe a firm in this market. In markets with S_1 between S_1^x and S_1^e we would observe incumbent firms. In markets with S_1 greater than S_1^e we observe entry. If we equate S_1^x with the S_1 of the smallest market with an incumbent, and S_1^e with the S_1 of the smallest market with a new entrant, we can interpret the difference between the two as an estimate of $(F+L)/V_1$, or the relative size of fixed plus sunk costs.

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Notice that this method of estimating entry and exit thresholds does not always produce an accurate estimate of $(F + L)/V_1$. Consider now the upward sloping scatter of points. In this set, second period market size is positively correlated with first period market size. If we do not account for this correlation, and base our estimate on the difference in first period thresholds, we will generally underestimate $(F + L)/V_1$. This can be seen by picking the dots nearest the threshold lines and projecting their first period market sizes onto the horizontal axis. Our analysis of Figure 1 thus shows that one cannot draw inferences about sunk costs from current demand data alone. One also must consider future demand. Before moving to our empirical model, we now consider the added complication that entrants may be uncertain about future demand.

2.1.3. *Demand Uncertainty*

At first, it might seem that few additional complications arise when the monopolist is uncertain about S_2. One can simply reinterpret Figure 1 by placing $E_1(S_2)$ on the vertical axis. This unfortunately is not true. When period 2 demand is uncertain, firms have incentives to delay sinking costs. These incentives affect entry and exit thresholds.

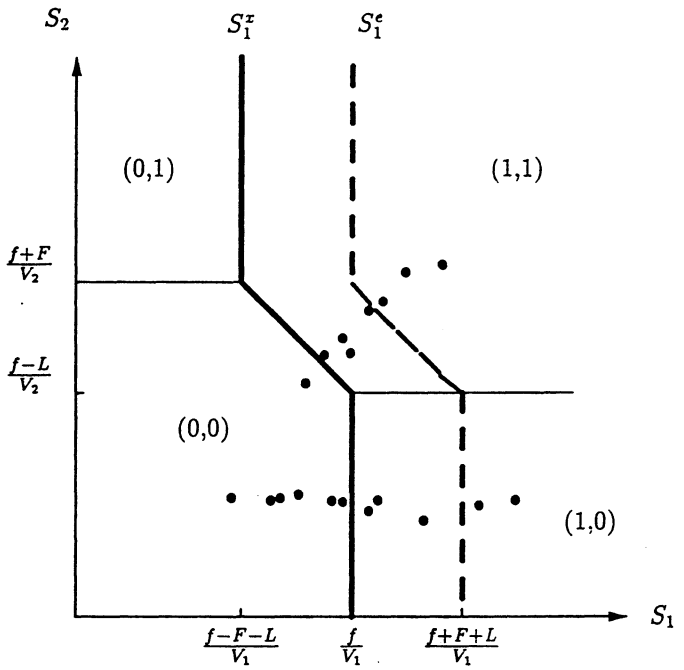

FIGURE 1

S^x and S^e Thresholds

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When S_2 is uncertain, potential entrants maximize expected profits. Expected profits depend on actions in period 1 and first period demand. Letting $\Omega = \{I(Q_1), S_1\}$ and $E[.|\Omega]$ equal an entrant's conditional expectation about demand with respect to the information Ω, expected profits equal

$$E(\Pi_1|\Omega) = I(Q_1)[V_1 S_1 - f] + V_2 E[S_2 I(Q_2)|\Omega] - f E(I(Q_2)|\Omega) - (1 - E(I(Q_2)|\Omega)) L - [I(Q_1) - E(I(Q_2)|\Omega)(1 - I(Q_1))] F.$$

If the firm produces in period 1, its expected profits equal

$$E(\Pi_1|I(Q_1) = 1) = V_1 S_1 - f - F + (1 - P(S_2^*))[V_2 E[S_2|S_2 \geqq S_2^*] - f] - P(S_2^*) L$$

In this expression, S_2^* equals the second period demand necessary to recover the exit cost L and $P(S_2^*) = Pr(S \leqq S_2^*)$ equals the probability of exit. Similarly, if the firm does not enter in period 1,

$$E(\Pi_1|I(Q_1) = 0) = (1 - P(S_2^{**}))[V_2 E[S_2|S_2 \geq S_2^{**}] - f - F].$$

Here S_2^{**} equals the second period demand required for the firm to recover the sunk entry cost F, and $P(S_2^{**}) = Pr(S \leqq S_2^{**})$ equals the probability of not entering (assuming $I(Q_1) = 0$). To see how demand uncertainty can affect the timing of entry, we compare the profits from delay to the profits from sinking costs in period 1,

$$\begin{aligned} &E(\Pi_1|I(Q_1) = 0) - E(\Pi_1|I(Q_1) = 1) \\ &= P(S_2^*) L + P(S_2^{**}) F + f - V_1 S_1 \\ &+ (1 - P(S_2^{**}))[V_2 E[S_2|S_2 \geq S_2^{**}] - f] \\ &- (1 - P(S_2^*))[V_2 E[S_2|S_2 \geq S_2^*] - f]. \end{aligned}$$

We can most easily analyze this expression by assuming that $P(S_2^*) = P(S_2^{**})$ and $E[S_2|S_2 \geq S_2^{**}] = E[S_2|S_2 \geq S_2^*]$. In this special case, an increase in the period 2 exit probability will increase the difference in profits, making it more likely the firm will not enter in period 1 [3].

---

3. Consider the following example. Suppose that first period demand is known and profits equal $\pi_1(S_1) = V_1 S_1 - f = 1$. Assume S_2 increases and decreases with equal probability and that for high demand $\pi_2 = 5$ and low demand $\pi_2 = -5$. Assume additionally F = 2 and L = 1. If produces in both periods it loses $\Pi = -2 + 1 + .5 \times (5 - 5) = -1$ It earns $\Pi = -2 + 1 + .5 \times (5 - 1) = 1$ if it exits in low profit states. Thus, S_1 seems to be above the entry threshold. Notice, however, that the firm can do better by exercising its option to delay entry. By delaying entry it foregoes period 1 profits, but expects to realize profits of $5 - F = 3$ in the high demand state. Since it did not sink costs in the first period, it loses nothing if the low demand state is realized. Thus, the expected profits from delay are 1.5. Hence, the firm should delay entry. This implies S_1 is below the first period entry threshold.

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We can similarly study the effect of demand uncertainty on exit thresholds. Incumbents may delay exit because reentry is costly. To calculate just how costly, we can compare the profits from delay to those of exit with possible re-entry

$$\begin{aligned}\mathrm{E}(\Pi|\mathrm{I}(\mathrm{Q}_1)=1)-\mathrm{E}(\Pi|\mathrm{I}(\mathrm{Q}_1)=0)=\mathrm{V}_1\mathrm{S}_1-f+(1-\mathrm{P}(\mathrm{S}_2^*))\mathrm{L}\\ +(1-\mathrm{P}(\mathrm{S}_2^{**}))\mathrm{F}+(1-\mathrm{P}(\mathrm{S}_2^*))(\mathrm{V}_2\mathrm{E}[\mathrm{S}_2|\mathrm{S}_2\geq\mathrm{S}_2^*]-f)\\ -(1-\mathrm{P}(\mathrm{S}_2^{**}))(\mathrm{V}_2\mathrm{E}[\mathrm{S}_2|\mathrm{S}_2\geq\mathrm{S}_2^{**}]-f)\end{aligned}$$

where S_2^{**} equals the second period demand necessary to recover the sunk cost F. Simplifying this expression by assuming that $P(S_2^*) = P(S_2^{**})$ and $E[S_2|S_2 \geq S_2^{**}] = E[S_2|S_2 \geq S_2^*]$, we find that increases in the entry probability $P(S_2^*)$ decrease the value of delay. Thus, uncertainty can make delay less profitable and thereby increase exit thresholds.

These equations show that the gap between entry and exit thresholds widens because of the option value to delay. Absent additional assumptions about the structure of demand uncertainty, we cannot say much about the economic significance of this option value. To obtain some idea of its magnitude, we numerically calculated entry and exit thresholds under the assumption that market size S follows a discrete binomial process with equal increments and equal probabilities. Additionally, we set $V_i = V_j = 1$, F = .01, L = .01, and f = .04 and allowed the firm's investment horizon to vary between two and four periods. Figure 2 plots the resulting entry and exit thresholds versus the standard deviation of a one period change in market size.

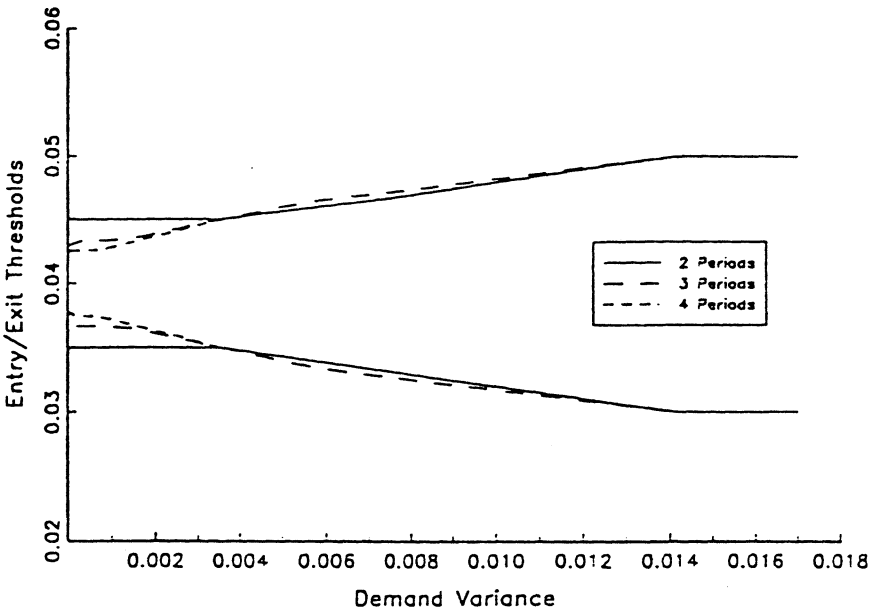

FIGURE 2

Monopoly Entry/Exit Thresholds.

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This figure shows that demand uncertainty increases entry thresholds and decreases exit thresholds. This widening reflects the option value to delay. The left-most thresholds in the figure are the certain (no option value) thresholds. In the two-period example they equal $S_1^e = .45$ and $S_1^x = .35$. The difference is $S_1^e - S_1^x = (F + L)/2 = .1$. When we increase the firm's planning horizon to four periods, the gap between the thresholds narrows. This narrowing occurs because the firm has more periods over which to spread its sunk costs.

The figure also shows that small amounts of demand uncertainty do not change the thresholds. The two period thresholds, for instance, remain constant for standard deviations as large as ten percent of the threshold. The thresholds do not change over this range because second period exit does not occur (*i. e.* $P(S_2^*) = 0$) [4]. When second period exit is unlikely, the spreading option outweighs the delay option. As we increase the number of periods, the constant threshold range shrinks because exit becomes more likely in later periods.

As the firm's uncertainty about future demand increases, the gap between the thresholds widens symmetrically. (It widens symmetrically because of symmetries in our assumptions.) Eventually, we reach a level of uncertainty where the thresholds no longer change. This level corresponds to the point where $P(S_2^*) = P(S_2^{**})$. Here, the value of delay no longer increases with the amount of uncertainty. In general then, the difference between the thresholds increases not only with the level of sunk costs, but also the amount of uncertainty present in demand. Thus, demand uncertainty may lead to large zones of inaction within which an incumbent delays exit.

2.2. **Oligoply Thresholds**

The previous analysis showed that theoretical intuitions from competitive models sometimes extend to noncompetitive markets. Unfortunately, we also saw that it is difficult to obtain simple expressions for sunk costs that can be used in econometric work. Clearly, adding firms can only complicate the analysis. Indeed, there are few tractable dynamic models of oligopolistic behavior. We have little to add here as well.

To obtain some sense of how imperfect competition affects our entry threshholds, we analyzed simple two-period oligopoly models comparable to those analyzed in the previous subsection. These examples, which we omit, raised two new issues. First, because oligopoly entry models need not have unique equilibria, we found it impossible to relate unique threshold conditions to observed outcomes. (*See* also Bresnahan and Reiss [1991 *a*].) Second, we found, as in our static models, that ratios of thresholds did not separately identify the effect of entry on variable profits and sunk costs. Because of this, we do not interpret ratios of entry and exit

4. We note $\Pr(S_2^*) = \Pr(S_2 < (f - L)/V_1) = \Pr(S_1 - \varepsilon < (f - L)/V_1)$ for some $\varepsilon > 0$. Because the firm never enters in period 1 unless $S_1 > f/V_1$, there are a range of ϵ's for which $\Pr(S_2^*) = 0$ and $E[S_2|S_2 \geq S_2^*] = S_1$.

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thresholds in the econometric work that follows. We note, however, that in general such ratios will confound increases in fixed and sunk costs with decreases in variable profits. (*See* also BRESNAHAN and REISS [1991 *b*].)

3 An Application: How Large Are Dentists' Sunk Costs?

Ideally we would like to use demand threshold equations such as (8) to infer the extent of fixed and sunk costs. In practice we rarely observe the exact levels of demand at which firms enter and exit. We also do not observe variable profits. This leads us to develop latent variable models of variable profits and sunk costs. We draw inferences about firms' latent profits from data on the number of incumbents in a market in two different years. Although we develop these models for a particular industry, they are adaptable to other industries.

Our application may at first seem unusual. We estimate entry and exit thresholds for dentists who practiced in 152 rural U.S. counties during 1980 and 1988. We study dentists in rural counties primarily because in contrast to automobile manufacturers they produce a relatively homogeneous product and sell it in a well-defined market. In principle, we could dentists in urban areas or automobile companies in different countries. Such markets, however, would raise econometric issues far beyond the scope of a single paper. We chose to study dentists because we know their practices are geographically concentrated and dentists must invest to build their patient base. The following advertisement describes some of these location-specific sunk costs and how health care professionals try to recover them upon exit

> An opportunity to have a... regular practice from the start of $ 2,000 a year in a section of Texas that has the combined resources of the grain of the North, cotton of the South, and cattle of the West... A bonanza for the wornout practitioner... [includes] Improved acre lot dwelling four rooms, two porches, chimney, cistern, garden orchard, out-houses, and well, desirably located. Opposition weak. Term $ 1,000. Will remain til purchaser is thoroughly satisfied...
>
> W. B. Anderson, M. D.,
> Content, Runnels County, Texas
> Daniel's Texas Medical Journal, 1892-1893

Although the dentists in our sample can and do sell their practices, this does not necessarily mean they can recover their entire investment.

3.1. Sample Observations and Definitions

Our econometric models estimate entry and exit thresholds by relating changes in market structure to changes in market demand. Table 1 *a* summarizes how much (net) entry and exit occured in our 152 markets between 1980 and 1988. The rows in Table 1 *a* correspond to the number

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of dentists observed in a market in 1980; the columns refer to the number in 1988. The cell entries equal the number of markets in our sample with the corresponding numbers of firms. The bordered diagonal cells contain all markets that had no net entry or exit. Most markets fall in these cells. Cells above (below) these diagonal cells had net entry (exit). Exit occurred more often than entry. Table 1 *a* also shows that most of our counties had four or fewer dentists.

TABLE 1 a

Counts of Markets by Number of Dentists Active in 1980 (N_0) ***and 1988*** (N_1).

N_0 = 1980 \ N_1 = 1988	0	1	2	3	4	≧ 5
0	31	8	0	0	0	0
1	8	32	2	0	0	0
2	4	17	8	2	3	0
3	0	3	7	7	3	0
4	0	1	2	6	0	0
≧ 5	0	0	1	0	2	5

Our econometric models use demographic and economic data to predict the location of each county in this tableau. Before describing these bivariate cell models, we first describe our data. We collected data on dentists by painstakingly searching American Dental Association (ADA) directories for all dentists in each of 152 counties. Our 1987 study describes these counties and why we consider them markets. Most are located in the

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western U.S.; almost all have demanders and suppliers located in or near a single population center (usually the county seat) [5]. These rural markets are isolated and have similar health care infrastructures.

The ADA bases its directories on surveys, dental school records, and federal and state information. We checked some of the 1990 Directory entries against Yellow Page information, site visits and phone calls. These checks convinced us that the ADA records are reliable [6]. We use the 1981-1990 directories primarily because the ADA changed its survey in 1979. These directories contain information on dentists' practices from August 1979 to 1988. We made 1980 the base year because of the demand data available in the 1980 population census. We picked 1988 as the end year to maximize the amount of change in county populations.

Our previous exit threshold models presumed that all exiters voluntarily forfeited sunk costs. The dentists in our sample, however, retire or die. To adhere to our models, we removed these involuntary exits. To understand why we adjusted the data, consider the following two period example. Suppose the sizes of a market in periods zero and one lie between the (unobserved) monopoly entry and exit thresholds; that is, $S_1^x < S_1 = S_0 < S_1^e$. If the market has an incumbent in both periods, we observe $N_0 = N_1 = 1$ and infer $S_1 > S_1^x$. (Note that we observe S_1 but not S_1^x.) Now suppose that the incumbent involuntarily exits between period zero and one. Now $N_0 = 1$ and $N_1 = 0$. If we mistakenly assume the exit was voluntary, then we would conclude $S_1 < S_1^x$. If we know the exit is involuntary, we can instead interpret $N_1 = 0$ as the event: a new entrant chose not to enter, *i. e.* $S_1 < S_1^e$.

This example suggests that if we subtracted involuntary exits from the period zero counts, we would obtain an adjusted count that we can link to voluntary decisions. In the example above, for instance, by setting $N_0 = 0$ we conclude the market had no net exit. Formally, we define the *adjusted* number of firms as $N_0^* = N_0$-deaths-retirements. We say that net-adjusted entry and exit occurs when

$$\begin{aligned} \text{Net Entry}: \; & N_{1988} > N_{1980}^* \\ \textit{No Change}: \; & N_{1988} = N_{1980}^* \\ \text{Net Exit}: \; & N_{1988} < N_{1980}^*. \end{aligned}$$

Table 1 *b* shows how this adjustment affects counts of entry and exits in our sample. As expected, the adjustment increases the number of markets with entry and decreases those with exit.

5.. There are two general exceptions. Some dentists practice in a nearby town. We include these dentists in our counts. A few dentists appear to practice in other counties but advertise in the Yellow Pages covering our counties. We exclude these dentists. We do not know whether these dentists travel (usually more than 100 miles round trip) to compete with local dentists.

6. We did find discrepancies. Some of these may be caused by timing differences between the Directory dates and our secondary sources. When we did find discrepancies, it usually appeared that the ADA directories had not identified dentists who had recently moved.

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TABLE 1 b

Adjusted Counts of Markets by Number of Dentists Active in 1980 (N_0^*) ***and 1988*** (N_1).

$N_0^* = 1980$ \ $N_1 = 1988$	0	1	2	3	4	≧ 5
0	34	14	1	0	0	0
1	8	41	4	3	0	0
2	1	6	11	2	5	1
3	0	0	4	9	2	3
4	0	0	0	1	1	0
≧ 5	0	0	0	0	0	1

The ADA directories list dentists by their location, speciality and last name. We counted dentists by searching the location lists for towns in our counties. We record each practicing dentist as a firm unless they were part of a group practice. We define a group practice as dentists having the same phone number or address, or dentists that have the same last name and are married (similar ages, opposite sex) or related. Our initial search identified 370 dentists present in 1980 or 1988. Almost all are general practictioners. We dropped eight unlicensed students and 4 public practictioners. We found twelve two-person group practices.

Besides the dentist's name, we know their location, type of practice, ADA membership status, and activity status. We also know all but two of the dentists' ages. In 1980, 258 dentists were active and 19 retired. In 1988, 209 were active and 61 retired. Table 2 provides additional sample information. Of the 75 dentists who entered between 1980 and 1988, 37 entered within a few years of graduation from dental school. Nineteen others changed locations within the same state. Of the 117 exiters: 29 retired without relocating; 16 retired and relocated; 31 disappeared, possibly because of their death; and 30 reported practicing elsewhere. We also classified several non-ADA member dentists who were over the age of 64

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as exiters because they reported working less than 30 hours per week in 1990. (We did not reclassify the over-64 year old dentists reporting full-time practice.) Figures 3 *a* and 3 *b* describe the age distribution of our dentists in 1980 and 1988.

4 The Econometric Model

Our econometric models relate the counts N^*_{80} and N_{88} to stochastic specifications of dentists' (unobserved) discounted future profits. These latent variable models recognize inter-market differences in demand and supply conditions. Our starting point is an expression for the observable part of dentists' discounted profits. To simplify the estimation, we presume that we can represent discounted variable profits as current variable profits plus one discounted future variable profit term. For example, for 1988 (period one), we assume an entrant's present discounted variable profits equal

$$\Pi_1 = \pi_1(N_1, Z_1, \theta) \times S_1 + \delta E_1[\pi_2(N_2, Z_2, \theta) \times S_2] \tag{10}$$

Here, E_1 is an expectation operator conditional on 1988 information, θ is a vector of unknown parameters, Z_t is a vector describing market demand and variable costs, and δ is a discount factor. The assumption that entrants have a two-period horizon is highly stylized. We use it to simplify our analysis. We think of period 2 as the entrant's entire future stream of profits. For the purpose of dating period 2 in our empirical work we assume it is eight years in the future. Thus, 1980 looks forward to 1988 (and beyond) and 1988 looks forward to 1996 (and beyond).

The expected value in (10) is conditional on N, Z and S in 1988. Absent additional structure, we cannot say much about the functional form of these conditional expectations. Following our ealier work, we assume variable profits are a linear function of Z, and that price competition lowers variable profits by fixed increments as the number of incumbents increases. Specifically, our specifications maintain

$$\pi_t(N_t, Z_t, \theta) S_t = \left(\theta_0 Z_t + \sum_{n=1}^{5} \theta_n I(n = N_t)\right) S_t \tag{11}$$

where I (.) equals one when (.) occurs and is zero otherwise. Competition implies that $\theta_1 \geq \theta_2 \geq \ldots \geq \theta_5$. To simplify our specification of entrants' expectations, we assume that future Z is independent of S_2 conditional on S_1, and that $N_2 = N_1$. We make the latter assumption because we do not have a dynamic equilibrium model for N_2. Together these assumptions imply we can represent expected linear profits in period 2 as

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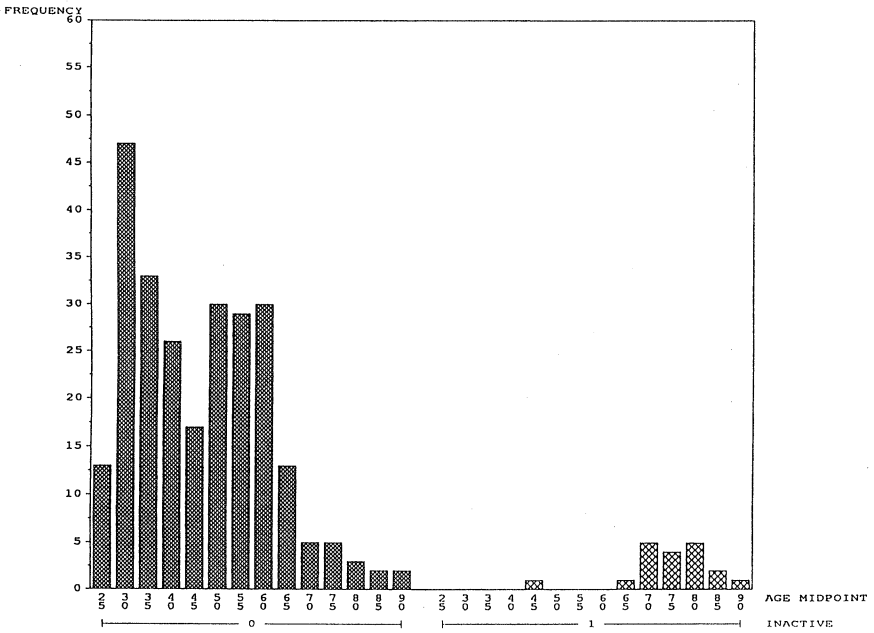

FIGURE 3 a

Age of Active/Inactive Dentists in 1980.

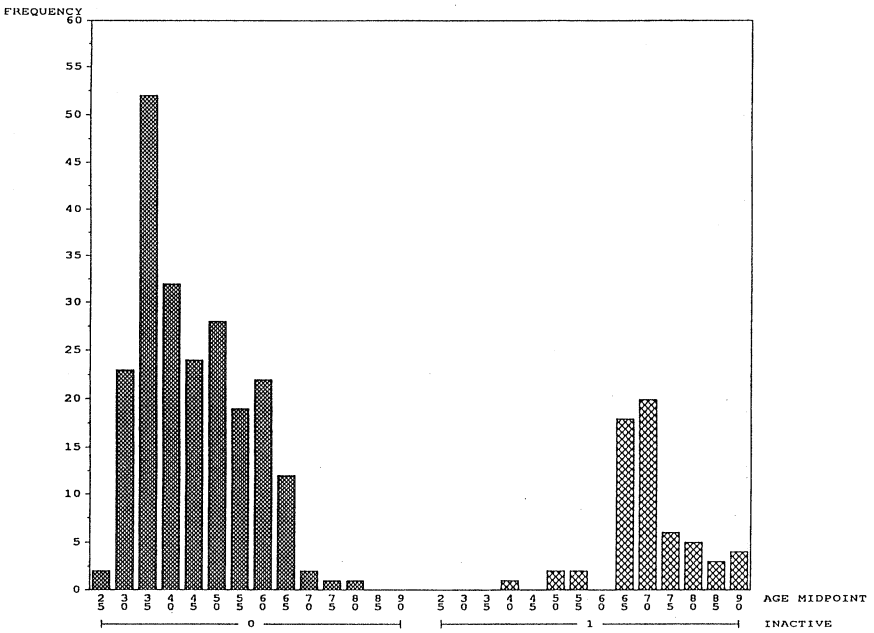

FIGURE 3 b

Age of Active/Inactive Dentists in 1988.

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TABLE 2

Descriptive Statistics

	Dentist Information			
	1979	1980	1986	1988
Age Active	45.4	48.0	44.3	46.8
#Active	251	258	238	219
Age Retired	75.4	76.1	69.7	70.5
#Retired	21	19	50	61
		Market Averages		
Variable	1980	1980	1988	1988
	Mean	S.D.	Mean	S.D.
Incumbent Dentists	1.62	1.42	1.34	1.31
Adjusted (N_0^*)	1.15	1.08	–	–
Town, Pop.	2077	1518	1987	1464
Cty Minus Town Pop.	2618	1713	2654	1853
8-Year Town Pop. Forecast	2041	1530	1886	1400
Real Per Capita Income	7040	1350	7160	1340
8-Year Income Forecast	7070	1320	7230	1360
LN (Heat DD)	8.63	.43	8.63	.43
Farm Land Value/Acre	–	–	404	266
HISP+NAT AMER				
PGRW	413	745	–	–
NEGGRW	–189	345	–	–

$$(12)\quad \mathrm{E}_1 \pi_2 (\mathrm{N}_2, \mathrm{Z}_2, \theta) = \left(\theta_0 \mathrm{E}[\mathrm{Z}_2|\mathrm{S}_1] + \sum_{n=1}^{5} \theta_n \mathrm{I}(n = \mathrm{N}_2) \right) \mathrm{E}[\mathrm{S}_2|\mathrm{S}_1]$$
$$= \left(\theta_0 \hat{\mathrm{Z}}_2 + \sum_{n=1}^{5} \theta_n \mathrm{I}(n = \mathrm{N}_1) \right) \hat{\mathrm{S}}_2.$$

This second period profit function contains population and income forecasts which we do not observe.

4.1. Population and Patients

In theory S_t represents the number of "representative" demanders or patients in a market. We distinguish between three types of demanders: those in central towns, those outside central towns, and those potentially under-served. We include both American Indians and Hispanics in this latter group. While demanders may differ on other dimensions, these were the most obvious to us. Formally, S_t equals

$$(13)\quad \mathrm{S}_t = \mathrm{TPOP}_t + \lambda_c \mathrm{CPOP}_t + \lambda_u \mathrm{UPOP}_t$$

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where TPOP is the population in town, CPOP is county population, and UPOP is the number of Hispanics and American Indians in the county. The coefficients λ_c and λ_u translate the latter two demand groups into equivalent numbers of townspeople. We expect people outside town to contribute less demand than people in town, *i. e.* $0 \leq \lambda_c \leq 1$. We expect λ_u to equal zero if Hispanics and American Indians have the same demands as other residents. If $\lambda_u < 0$, then either these groups have less demand for dental services or they have less access.

Finally, since we do not have 1988 minority and town population data, we replaced these variables with estimated values. Our estimates presume that these populations grew in proportion to county population from 1980 to 1988.

4.2. Population and Income Expectations

We divised two methods for estimating the 1996 values $\hat{S}_2$ and $\hat{Z}_2$. The first treats these conditional expectations as linear functions of other observables. We then estimate these expectations along with our other parameters. This approach considerably complicates the estimation process. Our second approach simplifies the estimation process by developing forecasts outside our model [7]. Initially we relied on published (historical) forecasts of $\hat{S}_2$ and $\hat{Z}_2$. obtained from CACI's Sales and Marketing Management Survey of Buying Power (hereafter SMM). Sales and marketing agents regularly use this survey to track demographic changes. Table 3 shows that the sample average SMM county population forecast was close to the actual sample average county population five years later. The SMM sample forecasts also have a smaller mean squared error than a simple random walk model. During the eighties, however, their forecasts systematically over-predicted population and income. In 1980, for instance, SMM predicted our average county would have 4,929 people in 1985, roughly a one percent per year increase. In 1988, SMM projected much slower population growth, about 0.1 percent per year. Their income forecasts were similarly optimistic. In 1981, SMM predicted that real disposable income would rise by over 5 percent per year between 1981 to 1986, and over 4 percent from 1988 to 1993.

The optimism in the SMM forecasts led us to develop our own forecasting models of county population and income. Our population data come from annual Census Bureau estimates. Our income data come from the Census Bureau's biannual estimates of total county income. Since our event windows are eight years long, we developed our regression models using explanatory variables that were at least eight year old. Table 4 presents the specifications we used. We arrived at these after a modest amount of experimentation. We think of them as simple linear conditional expectations that have no structural interpretation.

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7. Although this reduces the complexity of our final models, our coefficient standard errors currently do not recognize the sampling errors in these forecasts.

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TABLE 3

Forecast Descriptive Statistics and Models Annual Averages.

Year	Average County Population	SMM Forecast 5 years Prior*	Nominal Per capita Income †	Real Per capita Income ‡	Average county Employment
1960	4823	NA	NA	NA	NA
1970	4428	NA	2435	5798	1983
1975	4584	NA	4235	7141	2148
1980	4696	4934	5801	6769	2329
1985	4847	4929	7559	6816	2322
1988	4641	5090	8472	6984	2322
1990	4502	5073	NA	NA	2400

* For 1988 we substitute the 1982 forecast.

†The income figures are for 1969, 1975, 1979, 1985, and 1987.

‡Adjusted using the GNP 1982 Implicit Price Deflator.

Sources: SMM, Bureau of the Census and ADA Directories.

The 1988 population regression uses 1980 information to fit 1988 county population. The model contains an intercept, 1980 county population, county population growth between 1975-1980, county population growth between 1970-1985, 1980 population times zero-one dummy variables for whether the county is in: Texas or Oklahoma, Kansas, Nebraska or Colorado, California or Oregon, and North Dakota, South Dakota, or Montana. We advanced the regressors eight years to forecast 1996 population.

We fit our per capita income model to 1983 and 1975 per capita income data. These years do not match our sample years because we could not obtain complete SMM data. The income model contains an intercept; 1975 county per capita income; 1975 county per capita income times: the fraction of county population residing in the same house in 1975 and 1980, the fraction of county population moving within the state into the county between 1975 and 1980, a dummy for whether the county is in a western state, a dummy for whether the county is in a mountain state, and the relative growth in county population between 1975 and 1980; and, SMM's forecast in 1975 of its 1980 county income index. To forecast 1996 per capita income, we used 1988 explanatory variables in place of the 1975 explanatory variables.

Both the population and income regressions have reasonable fits. In analyses not reported here, we compared the out-of-sample accuracy of the SMM five-year forecasts with the eight-year regression forecasts (ending in the same year). The regression forecasts had a smaller average bias and root mean squared error. Further, when we included the 1980 SMM population forecast in the 1988 population regression in Table 4, we could not reject the null hypothesis that it had a zero coefficient. The (nominal) per capita income model in the second column includes the SMM income forecasts because they appeared to contain additional information about future county per capita income.

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TABLE 4

Forecast Regressions

	Dependent Variable 1988 County POP		1983 Per capita income (PINC83)
Regressor		Regressor	
Constant	11.88	Constant	1746.27
	(0.13)		(3.47)
1980 CPOP	0.94	PINC75	0.80
	(42.30)		(2.21)
CPOP80-CPOP75	0.48	PINC75×MOVE	0.71
	(6.79)		(1.73)
CPOP75-CPOP70	0.36	PINC75×STAY	-0.14
	(4.50)		(-0.46)
1980 CPOP×(TX+OK)	-0.01	PINC75×WEST	-0.21
	(-0.49)		(-2.11)
1980 CPOP×(KS+NE+CO)	0.01	PINC75×MNT	-0.06
	(0.36)		(-0.59)
1980 CPOP×(CA+OR)	-0.03	PINC75×PGRW75-80	0.24
	(–0.83)		(1.14)
1980 CPOP×(ND+SD+MT)	0.07	SMMI80	0.27
	(3.04)		(2.64)
R^2	.96		.54
RMSE	517		1070

Notes: *t*-statistics appear in parentheses.

CPOP: County population.

PGRW 75-80: County population growth between 1975 and 1980.

PINC: Per capita income.

SMMI80: 1975 SMM forecast of 1980 county income index.

STAY: Number of county residents living in the same house in 1975 and 1980.

MOVE: Number of county residents who moved from within the state between 1975 and 1980.

WEST: Zero-one dummy variable for Arizona (AZ), California (CA), Idaho (ID), Montana (MT), Nevada (NV), New Mexico (NM) and Oregon (OR).

MNT: Zero-one dummy variable for Colorado (CO), Kansas (KS), Nebraska (NE), North Dakota (ND), Oklahoma (OK), South Dakota (SD), Wyoming (WY) and Texas (TX).

4.3. The Stochastic Threshold Conditions

So far we have not allowed for differences between our profit functions and firms' actual profit expectations. We now describe our stochastic assumptions.

We assume that our two-period profit functions equal dentists' profit functions up to a mean-zero normal error term ε_t which is independently and identically distributed across our 152 markets. These errors represent profits that dentists observe but we do not. We assume normality primarily for computational reasons.

We have not examined whether our results are sensitive to this assumption.

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Following the threshold logic of section 2, we know that when $N_1 > N_0^*$, the N_1th entrant found it profitable to enter and the $N_1 + 1$th potential entrant did not. That is,

$$\begin{aligned} \Pi_1\left(N_1, S_1, \hat{S}_2, Z_1, \hat{Z}_2\right) &\geq F + f + \delta f + \varepsilon_1 \\ \Pi_1\left(N_1 + 1, S_1, \hat{S}_2, Z_1, \hat{Z}_2\right) &< F + f + \delta f + \varepsilon_1. \end{aligned} \tag{14}$$

Similarly, if we observe $N_1 < N_0^*$, we know

$$\begin{aligned} \Pi_1\left(N_1, S_1, \hat{S}_2, Z_1, \hat{Z}_2\right) &\geq f + \delta f + \varepsilon_1 \\ \Pi_1\left(N_1 + 1, S_1, \hat{S}_2, Z_1, \hat{Z}_2\right) &< f + \delta f + \varepsilon_1. \end{aligned} \tag{15}$$

Generally, we might expect both F and f to depend on N_1 and possibly N_2. Initially, we modelled sunk costs F as a set of 5 constants, $\gamma_1, \ldots, \gamma_5$, that varied with N_1. In practice we found that only monopoly sunk costs differed from other sunk costs. This led us to include γ_1 and assume $\gamma_2 = \gamma_3 = \gamma_4 = \gamma_5$. We also initially treated unsunk fixed costs as a set of 5 parameters that varied with N_1. Because we did not find significant differences, we report specifications that have only one additional parameter $\gamma = F + f + \delta f$. This parameter equals the sum of sunk and unsunk costs. It is presumed constant across markets and market structures. We also later report a test of this restriction [8].

Figure 4 displays the logic of our threshold conditions and the way our counts, N_0^* and N_1, identify the cost parameters (the γ's). The axes are levels of (unobserved) discounted profits in period zero and one. Profit threshold conditions such as (14) and (15) define the cell borders. Consider, for example, the darkly outlined persistent monopoly cell to the upper right. Here $N_0^* = N_1 = 1$. The threshold condition

$$\Pi_1\left(N_1 = 2, S_1, \hat{S}_2, Z_1, \hat{Z}_2\right) < \gamma + \varepsilon_1.$$

defines its lower boundary. The threshold

$$\Pi_1\left(N_1 = 1, S_1, \hat{S}_2, Z_1, \hat{Z}_2\right) > \gamma - \gamma_1 + \varepsilon_1.$$

defines its upper boundary. A comparison of these conditions to those for the adjoining no-entry $N_0^* = N_1 = 0$ cell, shows how we propose to identify sunk costs. The $N_0^* = N_1 = 0$ cell has just one horizontal boundary. This boundary is determined by the period one no-entry condition

8. We have found that once variable profits differ with N_1, we do not find significant fixed cost differences.

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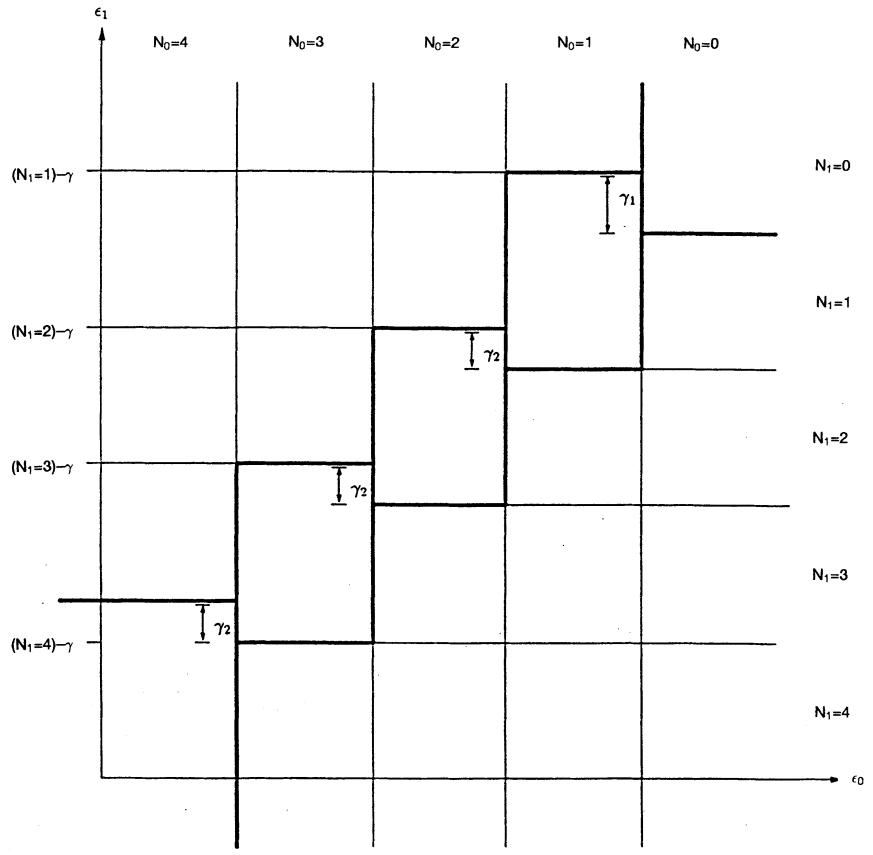

FIGURE 4

Bivariate Probit Cells..

$$\Pi_1\left(N_1 = 1, S_1, \hat{S}_2, Z_1, \hat{Z}_2\right) \leq \gamma + \varepsilon_1.$$

The vertical difference between this entry threshold and the exit threshold of the monopoly cell is the sunk cost parameter γ_1. Increases in γ_1 increase the range of discounted profits $\Pi_1\left(S_1, \hat{S}_2\right)$ over which we observe monopolists remaining in the market.

So far, we have only discussed the period one information that bounds Π_1. In principle, similar threshold conditions apply to the other axis, Π_0. These conditions, however, depend on N_{-1} (here 1972). We could recursively condition these cells on 1972 data. This, however, would not resolve the problem of what determines market structure in 1972. The econometric literature on discrete dynamic panel data models has proposed several different methods for handling initial conditions. Our approach parallels Heckman's (1983, p. 188) constructive solution. We approximate the initial market structure N_0^* with a reduced-form probit that contains only

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pre-sample exogenous variables. Following Heckman, we allow the first period error, ϵ_0, to be freely correlated with the structural error ϵ_1 [9].

Formally this approach amounts to treating the vertical cell boundaries as linear functions of predetermined variables. For example, we model the right-most vertical line as

$$\beta_1 + \mathrm{X}\beta = \beta_1 + \beta_{\mathrm{TPOP}}\,\mathrm{TPOP}_{80} + \beta_{\mathrm{OPOP}}\,\mathrm{OPOP}_{80} + \beta_{\mathrm{NGRW}}\,\mathrm{NGRW}_{72/80} + \beta_{\mathrm{PGRW}}\,\mathrm{PGRW}_{72/80} + \beta_{\mathrm{PCI}}\,\mathrm{PCI}_{79} + \beta_{\ln(\mathrm{HDD})}\ln(\mathrm{HDD}).$$

Here, NGRW is either zero or the negative growth in county population, PGRW is the complement of PGRW, PCI is nominal per capita income, and ln (HDD) is the log of heating degree days. Our earlier studies describe why we use these variables and how we obtained them. The series of constants, $\beta_5 < \beta_4 < \ldots < \beta_1$, separate the count cells, as is common in ordered probit models. Since this linear combination only approximates the marginal distribution of the initial state N_0^*, we make no predictions about the signs of these coefficients.

Finally, we include a correlation coefficient ρ to capture correlation between unobservables in the reduced form error ϵ_0 and ϵ_1. These correlations might arise for many reasons, including the presence of unobserved, market-specific demand errors. To appreciate how such correlations could effect our estimates, note that the correlation parameter ρ controls the distribution of the error mass over the cells in Figure 4. A large, positive correlation coefficient would orient the joint distribution from the lower left to the upper right. Such a distribution suggests that unobserved profit variables cause market structure to persist through time [10].

5 Model Estimates

5.1. A Prototype Model

The first column of Table 5 reports a prototype specification that contains the minimum number of explanatory variables. Specifically, it sets $\delta = 0$ and uses only town population in 1980 and 1988 to predict N^*_{1980} and N_{1988}. In this model, the latent variables that determine entry and exit are

9. Our approach and Heckman's are "constructive" in the following sense. Originally we assumed ϵ_0 was normal. If we replace Π_0 with a reduced form, we change the distribution of ϵ_0. Continuing to treat ϵ_0 as normal error therefore introduces approximation errors.

10. Of course, with the data we have we cannot separate this type of structural state dependence from others, nor can we distinguish between it and spurious state dependence.

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$$Y_{80} = \sum_{N_0^*=1}^{5} \beta_{N_0^*} I(N_0^*) + \beta_{TPOP} TPOP_{80} + \varepsilon_0$$

$$Y_{88} = \left(\sum_{N_1=1}^{5} \theta_{N_1} I(N_1) \right) TPOP_{88} - \gamma + \gamma_1 I(N_1 = 1 \leq N_0^*) + \gamma_2 I(1 < N_1 \leq N_0^*) + \varepsilon_1$$

where again I (.) is an indicator function equal to one when the event (.) is true. We also estimate a correlation coefficient ρ. Because we cannot estimate the variances of profits from count data, we set them equal to one. This normalization scales the estimated coefficients by the unknown error variances, implying we can only interpret coefficient signs and ratios.

TABLE 5

Adjusted Net Entry Results

	(1)	(2)	(3)	(4)
γ	1.715	2.504	2.481	2.430
	(.258)	(.373)	(.358)	(.394)
γ_1	1.505	1.082	.919	1.465
	(.430)	(.517)	(.493)	(.324)
γ_{2+}	1.045	.866	.794	1.182
	(.385)	(.335)	(.332)	(.290)
γ_{RET}			-.342	
			(.189)	
ρ	.273	.317	.452	
	(.324)	(.267)	(.230)	
θ_1	8.899	-14.801	-7.469	-19.056
	(1.36)	(17.37)	(12.747)	(22.694)
θ_2	4.432	-21.245	-12.066	-25.988
	(.930)	(21.106)	(17.202)	(26.459)
θ_3	2.801	-23.123	-13.313	-28.072
	(.722)	(22.169)	(18.457)	(27.635)
θ_4	1.943	-24.350	-14.133	-29.518
	(.755)	(22.919)	(19.283)	(28.547)
θ_5	1.274	-25.170	-14.672	-30.870
	(1.08)	(23.331)	(19.768)	(29.257)
θ_{PDI}		-55.469	-34.296	-71.077
		(38.145)	(32.136)	(49.642)
$\theta_{ln(bdd)}$		3.623	2.105	4.369
		(2.940)	(2.608)	(3.607)
θ_{land}		.239	.133	.351
		(.214)	(.173)	(.305)
λ_{county}		.309	.341	.236
		(.179)	(.179)	(.171)
$\lambda_{Ind.+Hisp.}$		-0.332	-0.316	-0.302
		(.178)	(.175)	(.201)

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TABLE 5 (continued)

	(1)	(2)	(3)	(4)
δ		-0.240	.095	-0.327
		(.522)	(1.207)	(.470)
β_{tpop}	4.429	5.107	5.180	4.729
	(.780)	(.977)	(.917)	(.989)
β_{opop}		4.952	5.075	4.907
		(.905)	(.909)	(.923)
β_{ngrw}		3.777	3.685	4.195
		(3.146)	(3.020)	(3.209)
β_{pgrw}		-1.635	-1.570	-2.119
		(2.448)	(2.351)	(2.390)
β_{pcdi}		19.782	19.473	21.467
		(7.665)	(7.576)	(7.620)
$\beta_{ln(bdd)}$		.847	.862	.789
		(.315)	(.301)	(.333)
β_1	-.354	-10.121	-10.264	-9.660
	(.183)	(2.855)	(2.735)	(3.027)
β_2	-1.54	-11.699	-11.822	-11.255
	(.207)	(2.848)	(2.727)	(3.020)
β_3	-2.23	-12.708	-12.857	-12.230
	(.232)	(2.912)	(2.795)	(3.074)
β_4	-3.53	-14.397	-14.591	-13.783
	(.372)	(2.926)	(2.804)	(3.083)
β_5	-4.02	-15.027	15.219	-14.417
	(.613)	(3.288)	(3.165)	(3.382)
log (likelihood) . .	-315.65	-277.79	-258.41	-282.73

(Standard Errors in Parentheses).

The most important ratios are γ_i/γ, the ratio of sunk costs to total fixed costs. Our estimate $\gamma_1/\gamma = .85$ implies that 85 percent of the average monopoly dentist's costs are sunk. Duopoly and oligopoly dentists have a smaller proportion, roughly 60 percent. We cannot reject the null hypothesis that monopoly and oligopoly dentists have the same fraction of sunk costs. We also cannot reject the null that this fraction is one. We can reject the null that it is zero. The finding that a substantial fraction of fixed costs are sunk also appears in more complete specifications.

The correlation parameter ρ is positive, suggesting positive persistence of unobserved profits. It is, however, small and imprecisely estimated. The variable profit coefficients $\theta_1, ..., \theta_5$ suggest that variable profits per patient decline as the number of dentists in the market increases. This is a familiar consequence of competition.

Because the future does not enter the profit functions, we can easily convert the intercepts and cost parameters into entry and exit thresholds. The thresholds in this model equal

$$S_i^e = \frac{\gamma}{\theta_i} \qquad S_i^x = \frac{\gamma - \gamma_i}{\theta_i} \qquad (i = 1, 2, 3, 4, 5).$$

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TABLE 6

Entry & Exit Thresholds for Specifications in Table 5

N	(1)		(2)		(3)		(4)	
	S_N^E	S_N^X	S_N^E	S_N^X	S_N^E	S_N^X	S_N^E	S_N^X
1	1927	236	1371	778	1356	854	1537	611
	(323)	(421)	(187)	(226)	(183)	(198)	(159)	(219)
2	3865	1510	2662	1741	2859	1944	2907	1493
	(546)	(742)	(323)	(330)	(428)	(348)	(386)	(364)
3	6121	2350	3668	2399	4090	2782	3970	2040
	(1,129)	(1,125)	(553)	(476)	(785)	(548)	(644)	(519)
4	8825	3445	4870	3186	5704	388	5321	2734
	(3,003)	(1,616)	(1,076)	(854)	(1634)	(1142)	(1319)	(928)
5	13,462	5255	6240	4082	7704	5240	7804	4009
	(11,145)	(3587)	(2316)	(1498)	(3778)	(2407)	(4189)	(2320)
Entry & Exit Thresholds Using Net Exit Definition								
1	1979	446	1335	807	1239	824	1650	613
	(246)	(274)	(151)	(139)	(133)	(128)	(184)	(189)
2	3351	2726	2330	2212	2095	1960	2742	2175
	(415)	(311)	(265)	(170)	(226)	(167)	(374)	(221)
3	4945	4023	3074	2917	2701	2527	3631	2881
	(700)	(517)	(379)	(276)	(318)	(254)	(599)	(351)

The estimates in Table 5 report entry and exit threshold estimates in units of thousands of people living in town. The first column of Table 6 reports them as the number of people in town.

The gap between the monopoly entry and exit thresholds is striking. While only 200 people can sustain an incumbent monopoly dentist, it takes over 1900 to attract one. Curiously, the estimates also show that the entry thresholds increase in proportion to the number of firms. In our previous work, we found that entry thresholds increased more than proportionately.

This prototype specification primarily illustrates the logic of our two period bivariate probit models. We now turn to more realistic models that explore the economic implications of fixed and sunk costs.

5.2. Other Models

Column two of Table 5 reports estimates of a more general dynamic bivariate probit model. It has profit functions similar to those in our earlier cross-section models (e. g. BRESNAHAN and REISS [1987]). It differs from them however in its treatment of initial market structure, unobservables, and forecasts of future variables.

We generalized the model in column one by adding market-level variables to per-patient variables profits and broadening our market size definition. Specifically,

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$$Y_{80} = \sum_{N_0=1}^{5} \beta_{N_0} I(N_0^*) + \beta_{TPOP} TPOP_{80} + \beta_{OPOP} OPOP_{80}$$
$$+\beta_{NGRW} NGRW_{72/80} + \beta_{PGRW} PGRW_{72/80} + \beta_{PCI} PCI_{79}$$
$$+\beta_{\ln(HDD)} \ln(HDD) + \varepsilon_0$$
$$Y_{88} = \pi_{88} + \delta\pi_{96} - \gamma + \gamma_1 I(N_1 = 1 \le N_0^*) + \gamma_2 I(1 < N_1 \le N_0^*) + \varepsilon_1$$
$$\pi_{88} = \left(\sum_{N_1=1}^{5} \theta_{N_1} I(N_1) + \theta_{PDI} Y_{88} \right.$$
$$\left. +\theta_{\ln(HDD)} \ln(HDD)_{88} + \theta_{LAND} LAND_{88} \right) S_{88}$$
$$\pi_{96} = \left(\sum_{N_1=1}^{5} \theta_{N_1} I(N_1) + \theta_{PDI} \hat{Y}_{96} \right.$$
$$\left. +\theta_{\ln(HDD)} \ln(HDD)_{88} + \theta_{LAND} LAND_{96} \right) \hat{S}_{96}$$
$$S_t = TPOP_t + \lambda_c CPOP_t + \lambda_u UPOP_t$$

Per-patient variable profits at time t now include three regressors: per capita disposable income, the price of agricultural land, and the log of heating degree days. We also used these variables in our earlier work. Since these variables may explain both demand and cost variations, we do not interpret the signs or significance of these coefficients. We also note that the 1996 profit function includes only two time-varying regressors, our forecast of (nominal) per capita income and town population. We saw little reason to forecast the other variables since they are likely to change little if at all from 1988 to 1996.

The additional variables in column 2 reduce our estimate of the average percentage of a monopoly dentist's costs that are sunk from 85 to 43 percent. Duopoly and oligopoly dentists' average proportion decline to 35 percent. We still reject the null that there are no sunk costs. We also cannot reject the null that all dentists have the same (fraction of) sunk costs. While the average proportion of sunk costs falls in this more general specification, the correlation coefficient between the unobserved profit errors increases. The increase is slight, however, and we still cannot reject the null that it equals zero. Although the (per patient) variable profit coefficients are individually insignificant, a joint test suggest they explain variation in dentists' variable profits.

The market size coefficients have a plausible pattern. Because these coefficients convert county population into town population equivalents, we can interpret the magnitudes of these coefficients. The estimated coefficient $\lambda_c = .308$ suggests that it takes between three and four residents outside town to demand the same services as one in town. The estimated coefficient $\lambda_u = -.333$ indicates that minorities have smaller effective demands for

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(or access to) dentists. Holding TPOP fixed, adding a minority resident barely affects demand.

The other coefficients in the equation are of interest only because they affect estimates of the entry and exit thresholds. Before we can calculate the 1988 entry and exit thresholds for this model, we must resolve several new issues. In principle, the 1988 discounted profit threshold conditions define the thresholds as follows

$$S^{e}_{88,\,N} = \frac{\delta\, VP_{96,\,N}\, \hat{S}_{96} + \gamma}{VP_{88,\,N}} \qquad S^{x}_{88,\,N} = \frac{\delta\, VP_{96,\,N}\, \hat{S}_{96} + \gamma - \gamma_{N}}{VP_{88,\,N}}$$

where $VP_{t,\,N}$ stands for per patient variable profits in year t conditional on the market having N firms. Because these expressions depend on the values of explanatory variables, we need to adopt a reporting convention. To simplify what we report, we chose to evaluate most variables at their sample means. We used a different procedure when evaluating $\hat{S}_{96}$. We set it equal to the level of demand which, if permanent, would just induce entry. That is, we used the entry threshold that solves the equation

$$\gamma = \delta\, V\, P_{96,\,N}\, S^{e}_{88,N} + VP_{88,\,N}\, S^{e}_{88,\,N}$$

We also apply the same logic to the exit thresholds. For exit thresholds we used the level of demand that solved

$$\gamma - \gamma_{N} = \delta\, VP_{96,\,N}\, S^{x}_{88,\,N} + VP_{88,\,N}\, S^{x}_{88,\,N}.$$

The specification in column three of Table 5 allows sunk costs to vary with the amount of involuntary exit. It recognizes the possibility that exiting dentists might reduce the costs of new entrants. One obvious way is through the sales of practices. A second is through a joint practice which ends with a business transfer. To allow for these possibilities, we let entrants' fixed costs vary with the number of involuntary exits. We measure the number of involuntary exits as $N_0 - N_0^*$. An entrant's fixed costs now equal $\gamma + \gamma_{RET}\,(N_0 - N_0^*)$. The estimates in column three show that involuntary departures lower entry costs by almost fifteen percent per involuntarry exit. Thus, there seems to evidence that an increased supply of practices lowers entrants' total fixed costs.

Column four explores how correlations in unobservable profits affect our sunk cost estimates. In it we constrain ρ to equal zero. This constraint increases the estimated proportion of sunk costs. One interpretation of this result is that unobserved heterogeneities are sunk costs. We recognize, however, that our ability to separate unobserved heterogeneities from true state dependence depends critically on our use of the normal distribution.

The negative discount parameter estimates do not make much economic sense. Comparing the estimated δ and θ_Ns in column four to those in the previous two columns we find that the θs are larger and δ is smaller. Indeed, in estimating the last three models we found three local suprema in the

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likelihood function. The global maximum usually occurred near δ equal zero. Two local optima occur at large absolute values of δ. Column four reports estimates for the second of these local optima. We note that we do not have much information that differentiates between variable profits and S in 1988 versus 1996. Observe first that when $S_{88} = S_{96}$, we have a single per capita variable profit function $VP(88, N) + \delta VP(96, N)$. Unless the two variable profit functions have different covariates, we cannot separately identify VP (96, N) from VP (88, N), nor can we identify δ. Although S_{88} and our forecast are different, they do not appear to help us estimate δ precisely.

Our estimates in Table 5 rely on our adjusted count of the number of densists. While we prefer this count, it is not obvious how this adjustment affects our estimates. To investigate this issue, we re-estimated the models in Table 5 using our original data, *i. e.* using N_0 and N_1 as the dependent variables. Although we do not report the coefficient estimates, the bottom of Table 6 reports the first three entry and exit thresholds implied by these estimates. The monopoly estimates differ little from those based on the adjusted data. The difference between the duopoly and triopoly entry and exit thresholds, however, is much smaller than the difference estimated using the adjusted data. (This is also true for the quadropoly and quintopoly thresholds.) We conclude that the distinction between voluntary and involuntary exit is critical for economic conclusions about sunk costs and exit behavior.

5.3. **Implications for Industry Evolution**

We now consider what our estimates imply about how population shifts and sunk costs affect the number of dentists in a county. Figure 5 graphs the entry and exit thresholds obtained from the estimates in column two. The horizontal axis is the number of incumbent dentists; the vertical axis is the number of people in town in 1988 and 1996. The upper line in the figure connects the estimated entry thresholds. At town populations above this line, dentists enter. The lower dashed line connects the exit thresholds. Exit occurs at populations below this line. Between the thresholds, dentists' fixed and sunk costs (or their associated option values) preclude entry or exit.

To interpret the figure, recall that we calculate the thresholds under the assumption that population changes are permanent. Thus, the figure reveals the range of market sizes that would induce exit, stasis, or entry. For instance, a single incumbent exits in markets with 800 people or less. Stasis occurs between 800 and the duopoly entry threshold, which is approximately 2500. The bands define similar ranges for the other incumbent counts. The figure shows that the zone of no change, or hysteresis, is large. Given the population changes we observe over eight to ten years, these estimates imply that most markets will retain their current (net) market structure. In other words, the estimates agree with what we see in Tables 1 a and 1 b–a high degree of persistence. The figure also reveals the role of sunk costs. By fixing a market size on the horizontal axis and travelling horizontally across the diagram to the hysteresis zone, we can calculate the number of market structures consistent with that level of demand. In general, we find that for

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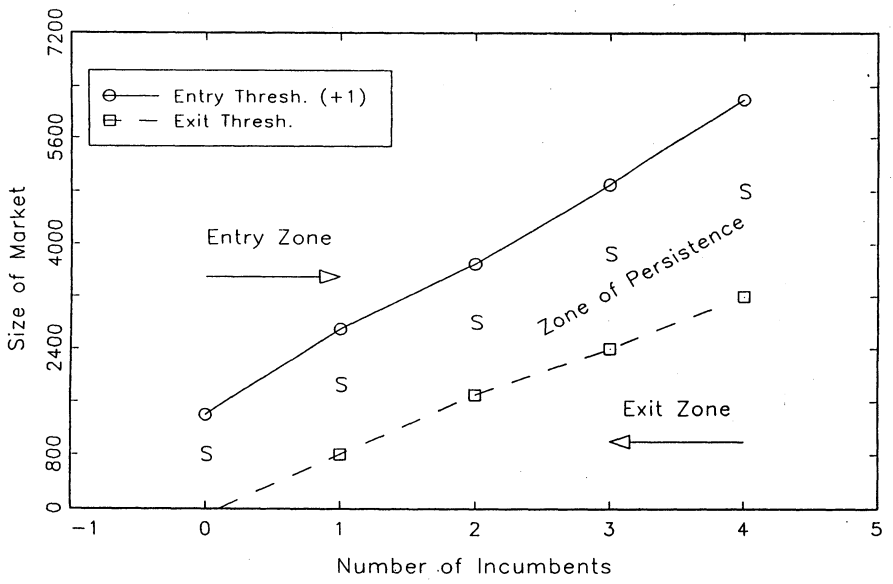

FIGURE 5

Entry and Exit Thresholds from Baseline Specification.

towns with more than 1000 people there are two or more possible market structures. Thus, as we look across rural counties, it is entirely possible that we can observe seeming contradictions. Some counties may have four dentists, while a smaller county may have five, six or more.

The observation that one level of demand can support several different market structures suggests that history matters for market structure. To see this, consider what could happen when a dentist in a market of 3,000 dies. If the market has three or four dentists, then our figure suggests that no dentist will enter to replace them. The figure also shows that a market with 3,000 people is also consistent with as few as two dentists.

The calculations underlying Figure 5 presume that when population changes it changes permanently. In principle, one could use other assumptions to explore what effect sunk costs and population changes have on the number of dentists in a market. One case of interest to us is the counterfactual case where sunk costs do not prevent exit. To study this case we calculated population changes between 1972 and 1980, and 1980 and 1988. We then used our parameter estimates to simulate what would happen to market structure if there were no deaths and retirements, no movements in other demand variables, and no profit errors. Specifically, we used the sample distribution of market size changes to construct future market sizes. We assume that entry occurs when the market demand implied by these changes crosses our estimated entry thresholds.

Figure 6 displays the population change data that we used. The circles represent the $304 = 152 \times 2$ pairs of town populations used in the simulation. The lines represent the five estimated entry thresholds. To make the figure easier to read, we limited towns to 7500 people. Figure 6

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reveals two interesting facts. First, there is considerable movement in the size of our markets. Despite this movement, the bulk of the observations fall in cells along the diagonal. These are cells in which market structure does not change. Because this simulation assumes there are no sunk costs, the entry and exit thresholds are the same. Thus, below the diagonal cells, we say exit occurs and above we say entry occurs. From the figure we see that roughly fifteen percent of the demand changes would have caused exit; ten percent would have caused entry. Thus, even in this extreme case, fixed costs limit the amount of entry or exit. Curiously, almost no entry or exit occurs in the smaller, more concentrated counties. In markets with 0, 1 or 2 firms, a one standard deviation population change (about seven percent) will not cross an entry threshold unless the market is already near the threshold. In less concentrated markets, the same percentage change in population will generate more entry or exit.

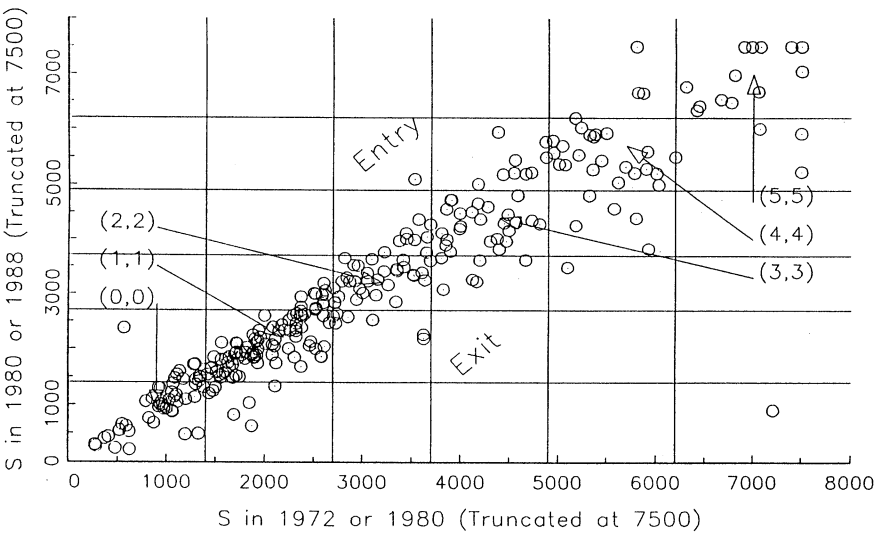

FIGURE 6

Eight Year Market Size Movements and Entry Thresholds – – Assuming No Sunk Costs.

Because only fifteen percent of the markets in our sample experience exit, we are unlikely to obtain more interesting exit dynamics by including sunk costs in the simulation. We have, however, ignored several important features of our model and application. In particular, our simulation does not allow demand variables to move the thresholds over time, we do have deaths and retirements, and we do not allow unobserved profits to change. We have not incorporated these effects here largely because they require a more complete model that can specify initial conditions.

In a final effort to explain how population changes affect turnover among our dentists, we developed Figure 7. Figure 7 plots the populations of all markets that had 1972 town populations within the 1988 monopoly or

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duopoly hysteresis zones. Because these markets that experienced little net change in population, dentists with perfect (sixteen year) foresight in 1972 would have seen them as static markets. By 1980, a substantial majority of towns still have populations within the monopoly and duopoly zones. By 1988 and (as predicted) 1996, this is less true. By 1996, market size has changed considerably, yet our sunk cost estimates imply that we should still observe considerable persistence.

Finally, we can compare the estimates in Table 6 and Figure 5 with our earlier static estimates. Previously we found entry thresholds increased much more than proportionately as N increased. Here, we find entry nearly proportional thresholds. What explains this difference? When firms have sunk costs, market structure will depend upon the history of S and past N. In our model, these are the initial conditions in period zero. If firms have large sunk costs, the past can be as important as the present. The persistence created by sunk costs complicates the dynamic relationship between changes in S and changes in N. When we estimate thresholds using only cross section data, as in our earlier work, we overlook these dynamics. Thus, our static thresholds will average entry and exit thresholds, with the history of S and N affecting the weights in this average.

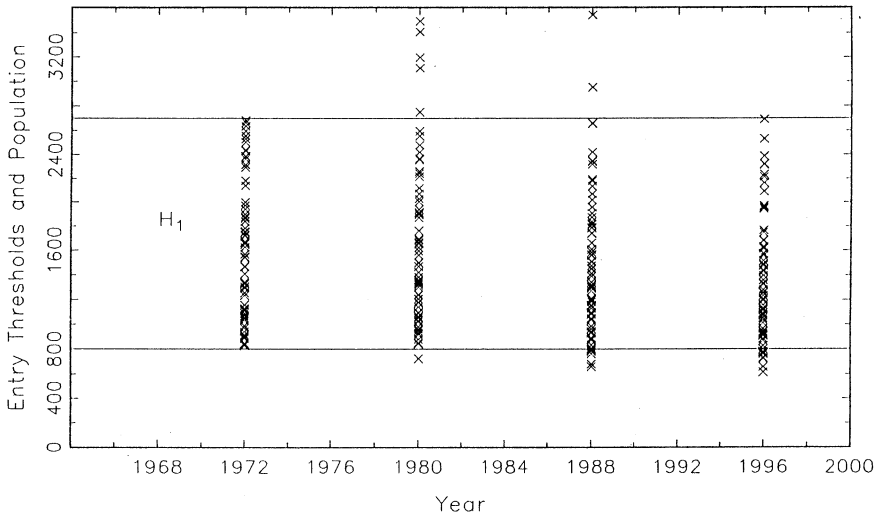

FIGURE 7

Evolution of S beginning In Monopoly Hysteresis Zone.

There are several different explanations for why this averaging leads to cross section thresholds that increase less than proportionately and panel thresholds that increase proportionately with N. First, according to our estimates, monopolists have the largest fraction of sunk costs as a proportion of fixed costs. Second, because involuntary exit occurs more often in larger markets, the waiting time to a new entrant is shorter than in small markets. Thus, in any cross section of markets, larger markets will have larger entry thresholds because entry is more frequent. Third, as we saw

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in Figure 7, monopolies and duopolies have smaller absolute changes in demand, making it less likely that we will observe entry in these markets.

The heteroscedasticity in demand also effects the likelihood of exit. Suppose, for instance, the thresholds are proportional. Consider a sample that consists of towns with either two or five firms. By assumption the five-firm towns are two and one-half times larger than the two-firm towns. Now suppose population randomly declines. Which set of towns will see more exit? If exit thresholds are proportional to N, the number of people that must leave to induce exit in the larger town is the same as the number that induce exit from a duopoly. Thus, the expected amount of (or time to) exit will be about the same. If population changes are heteroscedastic (ours are approximately homoscedastic in the logarithms), the expected waiting time for an exit in the smaller town will be much longer.

6 Conclusion

This paper measured the importance of sunk costs from differences in entry and exit thresholds. The difference between entry thresholds contains information not only about actual sunk costs, but also about the strategic and option values associated with sunk costs. Our methods thus have an advantage over other methods that only use accounting measures of sunk costs.

Our empirical analysis developed panel data methods that estimated entry and exit thresholds from data on firm locations. Although our models were tailored to a specific professional service, one can readily extend them to other industries. It also is possible to extend our specifications to use more data or information about how competition affects profits.

Our study of dentists found that they sink significant costs. Monopoly dentists appear to have the highest sunk costs relative to their total fixed costs. We also found that entry thresholds are nearly proportional to the number of active dentists. In other words, it takes about twice as much demand to induce the entry of a fourth practitioner as a second, and about twice as much demand for the second as the first, etc. This finding contrasts sharply with our earlier cross section results in which the number of firms did not double when market demand doubled.

While our models do not isolate the extent of competition among rural dentists, our estimated proportional thresholds are consistent with competitive models in which entry does not change competition. Anecdotal evidence suggests that rural dentists do not compete much on price and that entry also does not foster price competition because patients face substantial switching costs. Of course, this interpretation goes beyond what we can say with our current data and models. Nevertheless, our models suggest ways in which we might investigate such hypotheses with better data.

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