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COMPETITION IN LARGE MARKETS

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ABSTRACT

This paper develops a simple and robust implication of free entry followed by competition without substantial strategic interactions: Increasing the number of consumers leaves the distributions of producers' prices and other choices unchanged. In many models featuring non-trivial strategic considerations, producers' prices fall as their numbers increase. Hence, examining the relationship between market size and producers' actions provides a nonparametric tool for empirically discriminating between these distinct approaches to competition. To illustrate its application, I examine observations of restaurants' seating capacities, exit decisions, and prices from 224 U.S. cities. Given factor prices and demographic variables, increasing a city's size increases restaurants' capacities, decreases their exit rate, and decreases their prices. These results suggest that strategic considerations lie at the heart of restaurant pricing and turnover.

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1 Introduction

Observations of producers' actions from firm registries or national statistical agencies typically lack an accompanying description of their strategic environments. This unfortunate fact tempts one to assume that producers compete anonymously in a large market, but casual observation nearly always suggests some scope for strategic interaction between firms. This paper places this informal suspicion on a replicable basis using nonparametric regressions of producers' choices on market size. The data come from 222 U.S. cities' restaurant industries and are reported in the 1992 Census of Retail Trade. Under the null hypothesis of *atomistic competition*, market size has no impact on these decisions. This restriction is familiar from highly stylized models of monopolistic and perfect competition, and this paper proves it in a very general model without substantial restrictions on the market demand system, producers' cost functions, or the variables over which they compete. Even if one finds atomistic competition implausible ex ante, the nonparametric regression indicates which observable choices are particularly influenced by the strategic environment. The results show that restaurants do not compete atomistically. Instead, strategic considerations apparently shape firms' prices, exit decisions, and sizes.

In the simplest model of long-run perfect competition, free-entry requires all producers to minimize average cost regardless of the scale or shape of the market demand curve. In Dixit and Stiglitz's (1977) model of Chamberlinian monopolistic competition, the free entry condition implies that each firm's sales equals the product of the exogenous fixed cost with consumers' constant elasticity of demand. Doubling the number of consumers leaves producers' average sales unchanged. I show that the invariance of producer decisions to market size is much more general than these examples suggest, so it can provide the basis for empirically detecting strategic interactions and measuring their equilibrium effects.

The analysis rests on a nonparametric free-entry model. Potential producers make entry choices and then compete across a possibly large number of variables; such as price and advertising. A producer's profit depends only on the distribution of its rivals' actions and not on any particular rival's choices. This allows the transformation of a free-entry equilibrium for a given market size into one for a larger market with the same distribution of producers' actions. Thus, the model predicts that raising market size has no impact on the distribution (across producers) of any observable producer choice.

Because the model embodies no parametric assumptions, its predictions can be used to test reliably the assumption of atomistic competition using observations of producers' actions from a cross section of markets. The paper illustrates its application with nonparametric regressions of restaurants' prices, exit rates, sales, and seating capacities on market size. Of course, the distribution of producers' actions could differ across large and small markets even without substantial strategic interaction if the production technology and consumer tastes systematically change with market size. The free-entry model eliminates these possibilities by assumption, and the regression controls for them with factor prices and demographic measures.

On average across the sample of 222 U.S. Metropolitan Statistical Areas, the restaurants in larger cities have lower prices, exit less frequently, and have greater sales revenues. The observations do not determine the contribution of increased capacity utilization to restaurants' greater average sales. These findings are consistent with Campbell and Hopenhayn's (2005) result that restaurants and many other retail establishments in larger cities have greater average sales and employment. Together, they favor a model of competition between restaurants in which adding a competitor lowers producers' markups. They also suggest that the strategic models of entry and pricing in small markets estimated by Bresnahan and Reiss (1990), Berry (1992), and others might enhance our understanding of competition in large markets.

The approach to evaluating competition in large markets I advocate in this paper has one limit worth noting. A model in which oligopolists successfully collude and keep markups at their monopoly level but do not deter entry will replicate the scale invariance of atomistic competition. That is, the test has no power to reject the null in favor of the specific alternative of collusion with free entry. The empirical results of this paper as well as those of Campbell and Hopenhayn (2005) and Yeap (2005) indicate that this lack of power is not a practical problem for work with U.S. data.

The remainder of this paper proceeds as follows. The next section sets the stage for the analysis with an empirical examination of how restauranteurs' decisions vary with market size. Section 3 then provides a structural interpretation of these nonparametric results using the general model of atomistic competition. Section 4 relates this paper's results to those from the relevant literature, and Section 5 offers some concluding remarks.

2 Competition among Restaurants

To motivate this paper's analysis, consider the U.S. Restaurant industry. The U.S. Census questions the population of restaurants about their sales, cuisine, and pricing decisions every five years when creating the Economic Census. These observations allow researchers to address fundamental questions about the process of business formation, growth, and exit; but they contain only little information about the potential for strategic interactions. This is particularly the case for restaurants in cities, who have a great scope for differentiating themselves by location and cuisine.

The hypothesis that the firms in this data set compete atomistically can greatly simplify its analysis, because each firm's actions can be cast as the outcome of a single-agent decision problem. This simplification could come at a high price if strategic interaction is a first-order feature of competition, so I desire a simple procedure that can evaluate it before proceeding with a more complicated analysis.

Campbell and Hopenhayn (2005) use a symmetric model of oligopolists with constant marginal cost to build such a procedure. They note that oligopolists' average sales must rise with market size if their markups fall with additional entry, because they must recover the same fixed cost with a lower markup by selling more. Hence, modelling an industry as a collection of oligopolies seems promising if we see average sales rising with market size. The two shortcomings of their procedure are its reliance on a stylized model of competition and its exclusive focus on producers' average sales. This paper constructs a very general model of the null hypothesis which implies that *all* observable producer decisions are invariant to market size. The following description of how U.S. restauranteurs' actions vary with market size provides this theoretical analysis with a concrete empirical context.

2.1 Data

For this paper, I use observations from the 1992 *Census of Retail Trade* for the same sample of *MSA*s examined by Campbell and Hopenhayn (2005). The volume RC92-S-4, "Miscellaneous Subjects", reports the number of restaurants operating at any time during 1992 and at the end of that year. These observations immediately yield one measure of the annual exit rate. This volume also reports restaurants' average seating capacities for each *MSA*, the sales of all restaurants and of those operating at the end of the year, and the fraction of restaurants with typical meal prices greater than or equal to \$5.00. Although the Census records information about each restaurant's cuisine, this information is not disclosed publicly by *MSA*.[1]

From these observations, we construct four variables of interest. The first summarizes firms' pricing decisions. Denote the fraction of restaurants charging a typical meal price of \$5.00 or more with $\mathbb{S}(\$5.00)$, and consider its logistic transformation

$$\mathbb{L}(\$5.00) \equiv \ln(\mathbb{S}(\$5.00)/(1 - \mathbb{S}(\$5.00))$$

This is the logarithm of the ratio of "high priced" restaurants' share of the population to that of their "low priced" counterparts. Figure 1 plots this variable against the demeaned logarithm of *MSA* population. The observations corresponding to the smallest and largest *MSA*'s (Enid, OK and Atlanta, GA) are labelled, as are the observations with extreme values

[1]It would be desirable to examine more recent observations. Unfortunately, the Census has not published *MSA* level observations of these variables from the two most recent Economic Censuses.

of the log relative market share. The median value of $\mathbb{S}(\$5.00)$ across the sample's *MSA*'s is 0.67. The Census reports that only 13 percent of restaurants in Rocky Mount, NC charge \$5.00 or more for a meal, and it reports that 96 percent of restaurants charge \$5.00 or more in both Longview-Marshall TX and Jackson, MS. Aside from these three outliers, the minimum and maximum values of $\mathbb{S}(\$5.00)$ are 0.32 and 0.92. The correlation between the log relative market share and *MSA* population equals 0.09.

The second variable of interest measures one aspect of industry dynamics, the exit rate. I constructed this by dividing the number of firms operating at some time of the year but *not* at the end of the year by the total number of firms to operate in that year. The plot of this against *MSA* log population in Figure 2 shows a negative correlation. The exit rate for Enid, OK is very close to the maximum observed, 19 percent, while that for Atlanta, GA is close to the median across all *MSA*'s, 10.3 percent. The correlation between these variables equals −0.11.

The other two variables of interest both measure average restaurant size, restaurants' average revenue and average seating capacity. This average revenue variable differs differs from that used by Campbell and Hopenhayn (2005) only because it excludes restaurants not operating at the end of the year. Figures 3 and 4 plot these variables against *MSA* population. The strong positive association between *MSA* population and sales revenue documented by Campbell and Hopenhayn (2005) is evident in Figure 3. Figure 4 reveals little correlation between *MSA* population and average seating capacity.

2.2 Regression Results

Let Y_i denote the value of one of these four measures of restaurateurs' actions for *MSA* i, and use S_i and W_i to represent that *MSA*'s population and a vector of control variables that includes relevant factor prices and consumer demographics. The factor prices account for larger cities' higher cost of commercial space and wages and lower cost of advertising per consumer exposure. The demographic variables control for differences in preferences associated with income, race, and education that could shift the the nature of producers' products and thereby indirectly influence their observable decisions. These control variables are identical to those used in Campbell and Hopenhayn (2005). The regression of Y_i on $\ln S_i$ and W_i is

$$Y_i = m(\ln S_i, W_i) + U_i.$$

Here, $m(\cdot)$ is not restricted to a particular functional form.[2]

The curse of dimensionality makes the estimation of $m(\ln S, W)$ infeasible. However, it is still possible to test the hypothesis that its dependence on $\ln S$ is trivial using estimates of the regression function's density-weighted average derivatives. These are

$$\text{(1)} \qquad \begin{aligned} \delta_S &\equiv \mathbf{E}\left[\frac{\partial m\left(S,W\right)}{\partial \ln S} f\left(\ln S, W\right)\right] / \mathbf{E}\left[f\left(\ln S, W\right)\right] \\ \delta_W &\equiv \mathbf{E}\left[\frac{\partial m\left(S,W\right)}{\partial W} f\left(\ln S, W\right)\right] / \mathbf{E}\left[f\left(\ln S, W\right)\right], \end{aligned}$$

where $f\left(\ln S, W\right)$ is the joint density function of $\ln S$ and W across markets and expectations are taken with respect to the same joint density function. Powell, Stock, and Stoker (1989) provide a simple instrumental variables estimator of δ_S and δ_W which converges to the true parameter values at the parametric rate of $\sqrt{N}$. If market size does not directly impact producers' decisions, then $\delta_S = 0$.

For the four measures of restaurateurs' actions, Table 1 reports the estimated values of δ_S and δ_W along with consistent estimates of their asymptotic standard errors. Before estimation, the elements of W were scaled by the standard deviation of $\ln S$, which is 0.86 in this sample. Powell, Stock, and Stoker's estimator requires a first-stage nonparametric estimation of $\partial f(\ln S, W)/\partial \ln S$ and $\partial f(\ln S, W)/\partial W$. The estimates reported here are based on the tenth-order bias-reducing kernel of Bierens (1987) and use a bandwidth equal to 2. To increase the precision of the estimates' reports, *all entries in the table and in the text have been multiplied by* 100.

[2] In the case where $Y_i = \ln(\mathbb{S}_i(\$5.00)/(1 - \mathbb{S}_i(\$5.00))$, this specification for the regression function is equivilant to assuming that $\mathbb{S}_i(\$5.00) = e^{m(\ln S_i, W_i)+U_i}/(1 + e^{m(\ln S_i, W_i)+U_i})$.

The estimate of δ_S for the regression of $\mathbb{L}(\$5.00)$ equals -12.90 and is statistically signifcant at the 5 percent level. Thus, restaurants in larger markets charge *lower* prices given factor costs. To gauge the magnitude of this coefficient, consider an *MSA* with $\mathbb{S}(\$5.00)$ at the median level of 0.67. Set all of the elements of W equal to their means and consider increasing doubling S by one standard deviation. If we assume that $\partial m(\ln S, W)/\partial \ln S$ is constant, then such an increase in $\ln S$ decreases $\mathbb{S}(\$5.00)$ to 0.65.

The coefficients on two of the factor costs, commercial rent and the retail wage, are positive. They are both statistically significant at the 10 percent level, so the regression confirms the basic intuition that prices rise with factor costs. The third factor price, the cost of purchasing 1,000 advertising exposures in a Sunday newspaper, has a negative coefficient which is statistically significant at the 10 percent level. Perhaps high advertising costs allow producers to segment the market more effectively, thereby raising prices. Regardless, the effect of advertising costs on restaurant prices merits further investigation.

The estimate of δ_S for the exit rate is also negative, -0.77, and statistically significant at the 5 percent level. This implies that doubling S decreases restaurants' exit rate by 0.53 percentage points. As Campbell and Hopenhayn (2005) document, an increase in $\ln S$ strongly raises restaurants' average revenue. The estimated coefficient is 4.68, and it is statistically significant at the one-percent level.[3] The final dependent variable is the logarithm of average seats per restaurant. The estimated coefficient is positive, 2.07, but its standard error equals 1.99. Hence, these observations are uninformative about whether the increase in average revenue per restaurant arises from increased capacity utilization or increased average capacity. Nevertheless, the estimates in Table 1 clearly indicate that important decisions of restauranteurs vary systematically with market size.

The estimates in Table 1 depend on the particular measure of market size (population)

[3]This estimate differs greatly from that reported by Campbell and Hopenhayn (2005) for the identical regression with nearly the same sample. The discrepancy between the two reflects an error in Campbell and Hopenhayn's calculations. An erratum to that paper available at `http://www.nber.org/~jrc/marketsizematters` corrects that error.

and the bandwidth choice. Table 2 examines the robustness of the estimates of δ_S to these choices. Its first column reproduces the estimates from the first row of Table 1, and its next two columns report alternative estimates based on measuring market size with geographic population density and the number of housing units. Using either of these alternatives brings the estimate of δ_S for the regression of $\mathbb{L}(\$5.00)$ closer to zero. It equals -8.91 and has a p-value of 12 percent with population density, and it equals -11.93 with a p-value of 5.4 percent using housing units. All other inferences are invariant to changing the measure of market size. The final two columns of Table 2 report estimates based on changing the bandwidth h from its baseline value of 2 to either 1 or 3. Changing the bandwidth moves the estimated standard errors in the opposite direction. Otherwise, the estimates are unaffected. The only inference to change relative to the baseline specification is in the regression of $\mathbb{L}(\$5.00)$. When $h = 1$, the p-value for δ_S equals 10.3 percent.

I have undertaken three other checks of these estimates' robustness worth mentioning here. First, I have estimated all of the regression equations using ordinary least squares. The estimated coefficients are similar to the nonparametric estimates of δ_S. The only notable change in inference regards the coefficient in the regression of $\mathbb{L}(\$5.00)$. Its estimate drops to -9.79, and its p-value rises to 6.6 percent. Second, the Census reports the share of restaurants charging less than \$7.00 per meal. When I regress $\mathbb{L}(\$7.00)$ on $\ln S$ and W, I find no effect of market size on prices. Apparently, the reduction of restaurant prices occurs at the market's "low end". Nothing in principle prevents estimating δ_S using the original values of $\mathbb{S}(\$5.00)$ as a dependent variable. When I do so, the p-value for δ_S rises to 5.9 percent. Finally, I have also constructed analogues of Tables 1 and 2 for a sister industry, Refreshment Places. In that industry, market size has *no* measurable effect on typical meal prices, but its effects on the other three dependent variables are the same as with Restaurants.

3 A General Model of Atomistic Competition

The results of the previous section clearly conflict with very basic models of atomistic competition. In this section, I show that the abstraction from strategic interaction is the sole source of the conflict between these models and the data. The models' other simplifying features are not to blame. To do so, I develop the cross-market predictions of atomistic competition in a very general model with no parametric restrictions. So that the analysis is as broadly applicable as possible, I do not present specific conditions to guarantee the existence and uniqueness of a free-entry equilibrium. Instead, the analysis begins with the assumption that an equilibrium exists for a particular market size, and it then constructs an equilibrium with the same observable distribution of producers' actions for a larger market.

To make following the general model easier, this section begins with a particular model of atomistic competition. It then proceeds to the general model, referring back to the specific example to explain its moving parts.

3.1 A Specific Example

Consider a market for restaurant meals of heterogeneous quality. Production takes place in two stages, entry and competition. In the entry stage, a large number of potential restaurateurs simultaneously decide whether to pay a sunk cost of i to enter the market or to remain inactive at zero cost. After the restaurateurs commit to their entry decisions, each restaurant receives a random endowment of quality, which can equal either the high value q_H with probability w or the low value q_L with the complementary probability.

The competitive stage consists of two periods, early and late. All entrants can operate with zero fixed costs in the early period, but continuing to the late period requires paying a continuation cost i'. Exit allows a restaurateur to avoid this cost. In both periods, consumers randomly match with restaurants. The market is populated by S identical consumers, and equal numbers of them match with each restaurant. Restaurateurs simultaneously post their prices, and consumers decide on their purchases. A consumer matched with a restaurant charging a price p for a meal of quality q purchases $d(p/q)$ meals. This demand function is strictly decreasing and concave. Restaurants' variable cost functions are identical and feature a constant marginal cost of production, m.

A free entry equilibrium consists of a number of entrants, N, quality-contingent pricing decisions for each of the two periods, and quality contingent exit decisions such that each active restaurateur maximizes profit, entry earns a non-negative return, and no inactive potential entrant regrets staying out of the market. It is straightforward to show that this model has a unique free-entry equilibrium. First, consider the restaurants' pricing decisions, which satisfy the usual inverse-elasticity rule.

$$\frac{p - m}{p} = \frac{p}{q} \frac{d'(p/q)}{d(p/q)}$$

Because $d(\cdot)$ is concave, there is a unique price that satisfies this for each quality level. It is also straightforward to show that the optimal price increases with the restaurant's quality.

The assumption of a constant marginal cost implies that a restaurant earns a constant profit per customer. Denote these with π_L and π_H for the low and high quality restaurants. Restaurateurs' exit decisions depend on these profits, the number of entrants, and the cost of continuation. Denote the number of active restaurants in the late period with N'. Restaurateurs' optimal continuation decisions imply that

$$N' = \begin{cases} N & \text{if } i' \leq (S/N) \times \pi_L, \\ S\frac{\pi_L}{i'} & \text{if } (S/N) \times \pi_L < i' \leq (S/wN) \times \pi_L, \\ wN & \text{if } (S/wN) \times \pi_L < i' \leq (S/wN)\pi_H, \\ S\frac{\pi_H}{i'} & \text{if } (S/wN)\pi_H < i'. \end{cases}$$

In the first case all restaurants can profitably produce during the late period. In the second case, low-quality restaurants exit until their continuation value equals zero. In the third case, all low-quality restaurants exit, but all high-quality restaurants continue. In the final case, the continuation cost is high enough so that high-quality restaurants exit until their continuation value equals zero. The equilibrium exit decisions allow the definition of low and high quality restaurants' values at the beginning of the competitive stage, $V_L(S/N)$ and $V_H(S/N)$. These are both strictly increasing in S/N, so there exists a unique value of N that equates the ex-ante value of a new entrant with the entry cost.

Before proceeding to the general model, it is worth highlighting the scale invariance of this free-entry equilibrium. Because the ex-ante value of an entrant depends only on S/N, increasing the number of consumers raises the number of entrants proportionately. Restaurants' optimal prices depend on neither S nor N, while increasing both S and N raises N' by the same proportion and leaves the exit rate, $1 - N'/N$, unchanged. Hence, increasing the number of consumers in the market leaves the distributions of all observable producer decisions unchanged.

This specific example is far too stylized for empirical work, but suppose for the moment that it generated the *MSA*-level observations of restauranteurs' decisions used in Section 2. If restaurateurs' marginal costs and consumers' demand curves depend on a vector of market-specific variables like the factor prices and demographics in W, then regressions of restaurants' exit rate and of the fraction of restaurants with "high" prices on this vector and $\ln S$ would detect no dependence of these market-level summaries of producer actions on market size. In this sense, the specific example yields a testable prediction for cross-market comparisons of producer actions. The fact that the results in Section 2 refute this prediction implies that this very simple model could not have generated the data in hand. The analysis of the general model demonstrates that the conflict arises from the assumption of atomistic competition rather than one of the example's other simplifying assumptions.

3.2 The General Model

Like the specific example, the general model consists of two stages, entry and comptition. In the first stage, a large number of potential entrants simultaneously make their entry decisions. At the same time, entering producers make their product choices. The product choice of a particular entrant is x, and this lies in the set of all possible choices, $\mathcal{X} \subset \mathbb{R}^k$, where $k < \infty$. The number of producers that made choice x is $F(x) \in \mathbb{N}$, which I call the industry's *entry profile*. The example did not make restaurateurs' product choices explicit, but we can easily assume that they choose product addresses in $\mathbb{R}$ and that all consumers match in equal numbers with all offered products.

In the second stage, producers compete to sell their products to the market's S consumers. Producers simultaneously choose actions, $a \in \mathcal{A} \subset \mathbb{R}^l$, where $l < \infty$. Producers' profits depend on these choices and on realization of a vector of aggregate shocks, Z, which occurs before producers choose actions. An *action profile* is a function $A(x; Z, F) \to \mathcal{A}$. If $F(x') > 0$, then $A(x'; Z, F)$ gives the action of a producer that chose x' at entry. In the example, a represents a restaurants' early and late prices and its continuation probability and Z determines restaurants' qualities.[4]

For simplicity, we assume that if two or more entrants chose x, they both choose the same post entry action.[5] The total revenues of a producer at x' that chooses the action a' when all other producers' use the action profile $A(x; Z, F)$ and the entry profile is $F(x)$ are $S \times r(a', x'; A, Z, F)$. Here, S denotes the number of consumers and $r(\cdot)$ is the producer's average revenue per consumer, which does not directly depend on S. That producer's costs are $c(a', x'; A, Z, F, S)$. Because $c(\cdot)$ includes fixed costs and marginal cost might not be constant, it need not depend on S linearly.

[4] The specific example relies on idiosyncratic shocks to entrants' qualities. To use the general model's aggregate shocks to represent idiosyncratic shocks, assume that Z is a uniformly distributed location on the unit-circumference circle and that a restaurant has high quality if the clockwise distance between x/N (interpreted as a location on this circle) and Z is less than w. A potential entrant is indifferent across all locations on $[0, N)$ if entrants uniformly distribute themselves on this interval, so such a uniform distribution is an equilibrium outcome that generates the same distribution of high and low quality as in the example.

[5] As in the specific example, an element of a can represent a mixed strategy over a discrete and finite set of actions; and the revenues and costs specified below can be reinterpreted as expected values. Hence this assumption allows for mixed strategies. However, it does remove asymmetric Nash equilibria from consideration.

The expected post entry profit to a producer choosing x' at entry when it and its competitors follow the action profile $A(x; Z, F)$ are

$$\pi(x'; A, F, S) \equiv \mathbf{E}\left[S \times r\left(A\left(x'; Z, F\right), x'; A, Z, F\right) - c\left(A\left(x'; Z, F\right), x'; A, Z, F, S\right)\right].$$

Here, the expectation is taken with respect to the distribution of Z. This expectation exists under the assumption that that $r(\cdot)$ and $c(\cdot)$ are uniformly bounded functions of a and Z.

For the example, denote the prices charged by a restaurant and the probability that it produces in the late period with a_1, a_2, and a_3 The revenue and cost functions in the case where a single restaurant occupies an address are

$$\begin{aligned} r(\cdot) &= \frac{a_1}{N} \times d(a_1/q) + a_3 \times \frac{a_2}{N'} \times d(a_2/q), \text{ and} \\ c(\cdot) &= i + m \times \frac{S}{N} d(a_1/q) + a_3 \times \left(i' + m \times \frac{S}{N'} d(a_2/q)\right). \end{aligned}$$

In these expressions, the restaurant's quality q is the function of Z and x described in Footnote 4.

Define a *strategy profile* to be an action profile $A(x; Z, F)$ paired with an entry profile $F(x)$ and denote it with (A, F). With this notation in place, the definition of a free-entry equilibrium may proceed.[6]

Definition *A strategy profile $(A^\star, F^\star)$ is a* free-entry equilibrium *for a market with S consumers if it satisfies the following conditions.*

(a) Take any entry profile $F(x)$. If $F(x') > 0$, then for all $a \in \mathcal{A}$ and all possible realizations of Z,

$$\begin{aligned} &S \times r\left(a, x'; A, Z, F\right) - c\left(a, x'; A, Z, F, S\right) \leq \\ &\qquad S \times r\left(A^\star\left(x'; Z, F\right), x'; A, Z, F\right) - c\left(A^\star\left(x'; Z, F\right), x'; A, Z, F, S\right). \end{aligned}$$

[6] Conventional notation for a dynamic game takes a set of players with names, a strategy space, and payoff functions as primitives. The application of that approach to this model would specify the set of players as an unbounded set of potential entrants with names in R^k, the strategy space as $\mathcal{X} \times \{A(x; Z, F) \in \mathcal{A}\}$, and the payoffs as profit defined above. Because $F(x)$ and Z directly index all relevant subgames, working directly with the strategy profile as defined here simplifies the model's exposition.

(b) For all $x' \in \mathcal{X}$, $\pi\left(x'; A, F^{\star} + I\left\{x = x'\right\}, S\right) \leq 0$.

(c) If $F^{\star}\left(x'\right) > 0$*; then* $\pi\left(x'; A^{\star}, F^{\star}, S\right) \geq 0$*, and for all* $x'' \in \mathcal{X}$

$$\pi\left(x'; A^{\star}, F^{\star}, S\right) \geq \pi\left(x''; A^{\star}, F^{\star} + I\left\{x = x''\right\} - I\left\{x = x'\right\}, S\right)$$

Condition *(a)* of this definition ensures that the action profile $A^{\star}\left(x; Z, F\right)$ forms a Nash equilibrium for all subgames following the entry stage. Condition *(b)* requires that no further entry is profitable, and condition *(c)* states that each active producer's entry decision and choice of x is optimal given all other potential entrants' decisions. Together, the definition's three conditions are equivalent to requiring the strategy profile $\left(A^{\star}, F^{\star}\right)$ to correspond to a subgame perfect Nash equilibrium with pure strategies in the entry stage.

3.3 Atomistic Competition

At this level of generality, the framework encompasses many models. To specialize it and thereby derive the implications of atomistic competition, we impose the following two conditions. The first condition allows for only trivial strategic interactions between producers when no two of them occupy the same location in $\mathcal{X}$, and the second ensures that no such "local oligopolies" will arise in a free-entry equilibrium. Henceforth, I assume that $\mathcal{X}$ is a Borel measurable set with positive measure, denote the set of its Borel measurable subsets with $\mathcal{M}$, and use $\mu\left(M\right)$ to denote the Borel measure of $M \in \mathcal{M}$.

Assumption A1 (Atomistic Competition) *Let* $\left(A, F\right)$ *be a strategy profile with* $F\left(x\right) \leq 1$ *and define* $M = \{x | F(x) = 1\}$. *If* $F(x)$ *is Borel-measurable,* $\mu(M) > 0$, $A(x; Z, F')$ *is Borel-measurable given any shock realization* Z *and Borel-measurable entry profile* F'*, and* $F(x') = 1$*, then the revenues of the producer at* x' *choosing the action* a' *satisfy*

$$S \times r\left(a', x'; A, Z, F\right) = S \times \rho\left(a', x'; G\left(A, Z, F\right), Z, N_F\right),$$

where $N_F \equiv \mu(M)$ *is the mass of producers operating and*

$$G\left(A, Z, F\right)\left(a'\right) \equiv \frac{1}{N_F} \int_{\mathcal{X}} I\left\{A\left(x; Z, F\right) \leq a'\right\} \times F(x) d\mu\left(x\right).$$

Two aspects of Assumption A1 capture the idea that producers compete atomistically. First, a producer's revenues only depend on its own choices, aggregate shocks, the mass of competing producers, and the empirical distribution of their actions. Second, any one producer has measure zero when computing this distribution, so changing a single producer's conduct alters no other producer's revenue. The example revenue function above satisfies Assumption A1, because each producer's profit depends on its rivals actions only through S/N and S/N'. A finite-horizon version of Hopenhayn's (1992) model of perfect competition also satisfies Assumption A1. In any particular industry, the number of producers is obviously countable and not continuous. Models of atomistic competition are of empirical interest because their predictions might fit the data well in spite of the false simplifying assumption of a continuum of producers.

Assumption A2 (Product Differentiation) *If* $F(x') \geq 2$ *and* A *satisfies condition (a) of the definition of a free-entry equilibrium, then* $\pi(x'; A, F, S) < 0$.

Assumption A2 states that competition between producers of identical products is tough enough to guarantee that no more than one producer will occupy any location in $\mathcal{X}$. Thus, the observed market structure will not contain any "local" oligopolies. The specific example satisfies this assumption. Any model in which firms' producing exactly the same product act as Bertrand competitors will satisfy Assumption A2.[7]

3.4 Intrinsic Scale Effects

Thus far, the model's specification does not rule out direct effects of the scale of the market, measured with either S or N_F, on producers' revenues or costs. For example, the product space might be limited so that entry cannot continue indefinitely. The market shares of

[7] A model with price-taking producers of a homogeneous good, such as Hopenhayn's (1992), could accommodate this assumption by defining a trivial product placement choice x on the real line and assuming that the cost of entry at a given "location" increases steeply with the number of entrants there.

producers with particular choices of x might be more or less sensitive to the size of the market, or directly raising S could systematically reduce costs and so encourage entry and production. For all of these reasons, the distribution of producers' decisions across large and small markets could differ. The following three conditions eliminate them as a theoretical possibility.

Assumption S1 (Invariance of Market Shares) *The per consumer revenue function $\rho(\cdot)$ is homogeneous of degree -1 in N_F.*

This assumption states that doubling the number of producers while holding the distribution of their actions fixed cuts each producer's revenue in half. In the example, it follows from the uniform random matching of consumers with firms. This assumption is closely related to the independence of irrelevant alternatives: Adding a producer to a market does not change the *relative* market shares of any two incumbents. It also rules out Sutton's (1991) assumption that the winner of a pre-entry investment game obtains a minimum share of market sales, so his analysis of natural oligopolies does not apply here. However, Assumption S1 does *not* itself limit the cross-price elasticities of firms' products and thereby make the market less "competitive." For example, Salop's (1979) model of competition between spatially differentiated oligopolists on the unit circle satisfies it: If we double the number of evenly spaced firms while holding their prices constant, each firm sells to half as many consumers.[8]

Assumption S2 (No Productive Spillovers) *For any entry choice $x' \in \mathcal{X}$, action $a' \in \mathcal{A}$, and any two strategy profiles (A, F), $(A^\star, F^\star)$ and market sizes S and $S^\star$, if*

$$c(a', x'; A, Z, F, S) < c(a', x'; A^\star, Z, F^\star, S^\star)$$

then

$$S \times r(a', x'; A, Z, F) < S^\star \times r(a', x'; A^\star, Z, F^\star)$$

[8]Of course, the circle model does not satisfy Assumption A1, so Assumption S1 only applies to an appropriately defined per consumer revenue function.

Assumption S2 implies that is impossible to hold a producer's choices of x and a fixed, change its competitive environment, and lower that producer's costs without simultaneously lowering its revenues. Any model in which producers' costs depend only on their own output satisfies this assumption. If the market faces an upward sloping supply curve for some input, as in some versions of Hopenhayn's (1992) model, then this assumption would be violated. The simple affine technology of the example obviously satisfies Assumption S2.

Assumption S3 (Distinct Observationally Equivalent Strategy Profiles) *For any market size S and strategy profile (A, F) such that $F(x) \leq 1$ for all $x \in \mathcal{X}$, there exists a continuous, one to one, and onto function $g : \mathcal{X} \to \mathcal{X}$ such that if we define $F^T(x) \equiv F(g^{-1}(x))$, and $A^T(x; Z, F^T) = A(g^{-1}(x), Z, F)$ then*

(a) $\forall x \in \mathcal{X},\ F(x) + F^T(x) \leq 1$;

(b) if $F(x') > 0$, *then*

$$S \times r(a', x'; A, Z, F) = S \times r\left(a', g(x'); A^T, Z, F^T\right)$$

and

$$c(a', x'; A, Z, F, S) = c\left(a', g(x'); A^T, Z, F^T, S\right);$$

(c) $\forall M \in \mathcal{M},\ \mu\left(g^{-1}(M)\right) = \mu(M)$.

In many models of competition with product differentiation, it is possible to rearrange producers' locations in $\mathcal{X}$, hold their actions fixed, and leave their payoffs unchanged. Consider two examples of such a rearrangement, moving all producers a short distance to the right in Salop's (1979) circle model and changing the particular products chosen by entrants in Dixit and Stiglitz's (1978) model of Chamberlinian monopolistic competition. In both cases, the rearrangement leaves the game played after product placement unaltered. In Assumption S3, conditions *(a)* and *(b)* require such a rearrangement to be possible for any given strategy profile. In this section's simple example, $\mathcal{X}$ is an infinite set of all possible products without any spatial structure or asymmetries in demand or cost, so such a rearrangement is possible. Condition *(c)* requires $g(x)$ to be measure preserving, so that the rearrangement does not alter the mass of producers. Overall, Assumption S3 asserts that no location in $\mathcal{X}$ has payoff-relevant characteristics that are unique.

3.5 Equilibrium

The following two conditions ensure that potential entrants' expectations about post-entry competition can be well-defined and that a free-entry equilibrium exists.

Assumption E1 (Existence of Nash Equilibrium) *For any market size S, there exists a strategy profile $A(x, Z, F)$ that satisfies condition (a) of the definition of a free-entry equilibrium.*

Assumption E2 (Existence of a Measurable Free Entry Equilibrium) *There exists a market size $S_0 > 0$ with a corresponding free-entry equilibrium (A_0, F_0) such that*

(a) $F_0(x)$ *and* $A_0(x; Z, F_0)$ *are Borel measurable functions of* x *for any* Z*, and*

(b) $A_0(x'; Z, F) = A_0(x'; Z, F_0)$ *if* $F(x') = F_0(x') = 1$ *and* $F(x) = F_0(x)$ *almost everywhere.*

To the extent that nearly all existing models of entry and competition have well-defined equilibria, assumption E1 and part (a) of Assumption E2 can be viewed as regularity conditions. Part (b) of Assumption E2 eliminates the possibility that a positive measure of firms respond to deviations from the equilibrium by a single (measure zero) firm. That is, there is no producer whose actions act as a pure coordination device for the others. Both assumptions are clearly true for the specific example.

3.6 Market Size and Producers' Actions

The above assumptions place sufficient structure on the model to imply the following observational implication of atomistic competition.

Proposition *If $S = 2^j \times S_0$, where j is a non-negative integer and S_0 is defined in Assumption E2, then there exists a free entry equilibrium (A_j, F_j) such that*

$$G(A_j, Z, F_j) = G(A_0, Z, F_0)$$

where (A_0, F_0) is the free-entry equilibrium assumed to exist in Assumption E2.

The appendix presents the proposition's proof. Here, I only outline the argument. Consider the free-entry equilibrium (A_0, F_0) for S_0. We know from Assumption S3 that there is a different but observationally equivalent strategy profile, $\left(A_0^T, F_0^T\right)$. Now consider a market with $S_1 = 2 \times S_0$ and entry profile, $F_0 + F_0^T$. If all producers duplicate the actions they take in the smaller market, then the empirical *c.d.f.* of producers actions remains unchanged, so Assumptions A1, S1, and S2 imply that each producer's profit maximizing action remains unchanged. That is, the action profile that duplicates producers' actions is a Nash equilibrium profile for this larger market size and entry profile. Each producer's profits remain unchanged, and the profits from producing in an unoccupied location in $\mathcal{X}$ are identical to their value in the original free entry equilibrium with the smaller market size, so conditions (b) and (c) of the definition are satisfied. Each producer's actions equal those from the original equilibrium, so their empirical distributions are also unchanged as the proposition asserts.[9]

[9] The proposition's focus on doubling market size can easily be changed if the assertion that $g(x)$ is measure preserving in Assumption S3 is replaced with the assumption that for any $t > 1$, there exists a $g_t(x)$ satisfying the assumption's other conditions and which satisfies $\mu(g_t^{-1}(M)) = t \times \mu(M)$. With this, a parallel argument establishes that a free-entry equilibrium that replicates producers' decisions exists for any market size greater than S_0.

The proposition illustrates that the invariance of producers' decisions to the market's size in the specific example extends well beyond its particular assumptions. All of the assumptions excepting A1 are regularity conditions, so I interpret the fact that market size *does* influence restauranteurs' observable decisions as a rejection of atomistic competition.

3.7 Extensions

The general model is restrictive in two ways that are worth noting. First, it has no role for actions that are taken prior to the realization of Z that do not directly differentiate firms' products, such as investments that increase the likelihood of having a high quality restaurant. Adding such pre competition actions to the general model increases its notational burden but does not alter its scale invariance. Second, the use of the Borel integral to form the *c.d.f.* of producers' actions in Assumption A1 restricts product placement decisions to be continuous choices. Scale invariance requires *some* continuous product placement decision to differentiate firms' products, but it does not require all product placement decisions to be continuous. Extending the general model to allow for discrete dimensions of firms' product placement decisions is straightforward.

3.8 Atomistic and Monopolistic Competition

Before concluding, it is helpful to clarify the relationship between what I have labelled "atomistic competition" with the large theoretical literature on monopolistic competition. For some authors, "monopolistic competition" refers to all imperfect competition among a large number of producers. Models that prominently feature strategic interaction, such as Salop's (1979) model of spatial competition, then go by the label of "Hotelling-style" monopolistic competition. Models with only trivial strategic considerations, such as Spence's (1976) are called "Chamberlin-style".

Hart (1985) and Wolinsky (1986) propose a more exclusive definition of "monopolistic competition" based on four criteria.

> (1) there are many firms producing differentiated commodities; (2) each firm is negligible in the sense that it can ignore its impact on, and hence reactions from, other firms; (3) each firm faces a downward sloping demand curve and hence the equilibrium price exceeds marginal cost; (4) free entry results in zero-profit of operating firms (or, at least, of marginal firms).[10]

These clearly correspond to what others call Chamberlin-style monopolistic competition. Hart and Wolinsky's first two criteria correspond to Assumptions A2 and A1, and the fourth criterion is implicit in the definition of a free-entry equilibrium. The definition of atomistic competition does not require firms to face downward sloping demand curves, but it clearly allows for that possibility. Hence, models of monopolistic competition (in the sense of Hart and Wolinsky) can usually be written without economically substantial changes to satisfy the assumptions this paper places on atomistic competition. However, the definition of atomistic competition is broad enough to also encompass models without market power.

4 Related Literature

Structure-conduct-performance studies gave rise to many examinations of competitive outcomes' dependence on market size. One strand of this literature uses the empirical relationship between market size and the number of competitors to infer how adding competition lowers markups. If doubling market size leads to a less than proportional increase in the number of producers, either per-consumer profits fall with entry or incumbents raise entrants' fixed costs. Bresnahan and Reiss (1989) apply this approach to concentrated retail automobile markets in isolated towns. Berry and Waldfogel (2003) examine the influence of market size on the number of competitors in a slightly broader sample of *MSA*s than that used in this paper, and they find that the number of restaurants increases less than proportionally with *MSA* population.[11] The proof of this paper's proposition makes it clear that

[10] Wolinsky (1986), page 493.

[11] See the third and fourth columns of their Table 3.

the number of producers in atomistically competitive markets is proportional to the number of consumers, so Berry and Waldfogel's finding reinforces this paper's empirical conclusion.[12]

This paper's proposition does not stress the relationship between market size and the number of firms under atomistic competition, because a finding that doubling market size less than doubles the number of firms could arise solely from measurement error in market size. Measurement error could make the rejection of this paper's exclusion restrictions less likely when they are false, but it does not lead directly to their rejection when they are true. In this sense, a test of atomistic competition based on the relationship between noisily measured market size and measures of producer actions is conservative.[13]

This paper derives testable predictions of a free-entry model without the use of parametric assumptions. In this respect, Sutton's (1991) analysis of models with endogenous sunk costs precedes it. He considers a model of competition in which entrants compete with sunk investments in product quality. The firm with the greatest investment earns a guaranteed minimum market share, regardless of the number of other producers. Sutton shows that these features together imply a nonparametric upper bound on the number of entrants, and he demonstrates that cross-country data from several advertising-intensive food processing industries satisfy this bound. As the number of consumers grows, the number of entrants remains bounded from above. In this sense, industries that satisfy his model's assumptions are natural oligopolies. As noted above in Subsection 3.7, it is notationally burdensome but straightforward to add pre-entry investments in quality to this paper's model. This

[12]Berry and Waldfogel's finding also manifests itself in the observations used in the present paper. The estimate of δ_s from a nonparametric regression of the number of restaurants' logarithm on population's logarithm and the other control variables listed in Table 1 using Campbell and Hopenhayn's (2005) sample of *MSA*s equals 0.93, and this is significantly different from one.

[13]Bresnahan and Reiss (1989) can measure market size accurately because they carefully chose their sample towns. This strategy becomes infeasible when considering competition in large markets in which the definition of the market and industry are themselves somewhat subjective, so prudence requires accounting for possible measurement error.

extension leaves the model's nonparametric testable implications unaltered. In particular, the number of producers grows linearly with market size. The contrast between that result and Sutton's highlights the role of endogenous sunk costs in his results: They are necessary but not sufficient for an industry to be a natural oligopoly. Hence, the simple observation that an industry's producers incur endogenous sunk costs does not imply that its firms are oligopolists. However, tests of the exclusion restrictions from atomistic competition do provide information about the nature of competition.

The analysis of the exit of restaurants places this paper in another vast literature which examines the rate of producer turnover and the reallocation of resources between producers. These papers have focused on differences in firm growth and survival across the life cycle (as in Dunne, Roberts, and Samuelson (1989)) and on the interaction of resource reallocation with the business cycle (as in Davis, Haltiwanger, and Schuh (1996), Campbell (1998), and Campbell and Lapham (2004)). Analysis of how the pace of resource reallocation varies with local market conditions, similar to that in this paper, is much scarcer. Syverson (2004) shows that ready-mixed concrete producers serving geographically concentrated markets have higher average productivity and less productivity dispersion than their counterparts is more sparsely populated areas, and he interprets this as the result of more intense selection in highly competitive markets. Similarly, Asplund and Nocke (2006) find that Swedish hairstylists are younger in larger markets. This paper finds the opposite relationship between market size and exit for U.S. restaurants. In light of this paper's theoretical results, this observation suggests that strategic interaction substantially influences the rate of restaurant turnover. However, Syverson's and Asplund and Nocke's results indicate that much more work is required before a robust stylized fact about the relationship between industry concentration and reallocation emerges.

5 Conclusion

The relative simplicity of atomistic competition models makes them a tempting first choice for the empirical study of competition in large markets. This paper presents their nonparametric implications which can be used to examine their suitability before proceeding with a more involved investigation.[14] The application of these results to observations of U.S. restaurants' prices, exit rates, sales, and seating capacity indicates that atomistic competition cannot explain how restauranteurs' key choices depend on market size. Relatedly, Yeap (2005) documents that this increase in average size reflects only the decisions of firms owning two or more restaurants. Taken together these findings indicate that better understanding of competition among restaurants in large markets requires confronting restaurateurs' strategic behavior. Toivanen and Waterson (2005) take an important step in this direction by empirically modelling entry decisions into well defined duopoly fast-food markets. Extending such an analysis to large samples of restaurants without high-quality information about market definitions and strategic interactions is the subject of my current research with Jaap Abbring.

[14]For example, Abbring and Campbell (2004) apply this papers' results to test the assumption of atomistic competition in their structural model of new Texas bars' growth and exit decisions.

Proof of the Proposition

Clearly, the proposition is true for $j = 0$. We now wish to show that it is true for $j = 1$. The proposition can then be demonstrated recursively for greater values of j.

Let $g(x)$ be the function assumed to exist in Assumption S3. Define the entry profile $F_1(x) = F_0(x) + F_0^T(x)$, where the latter entry profile is defined as in the statement of Assumption S3. From Assumption A2 and the definition of a free-entry equilibrium, we know that $F_0(x) \in \{0, 1\}$. Therefore, condition (a) of Assumption S3 ensures that $F_1(x) \in \{0, 1\}$.

We know from Assumption E1 that there exists an action profile $A(x; Z, F)$ that satisfies condition (a) of a free-entry equilibrium's definition for $S_1 = 2 \times S_0$. We now wish to use this and $A_0(x; Z, F)$ to construct an action profile that forms a candidate free-entry equilibrium when paired with F_1. For any entry profile $F(x)$ such that either $F(x') \geq 2$ for some $x' \in \mathcal{X}$ or $\{x | F(x) \neq F_1(x)\}$ is either not measurable or has positive measure, define $A_1(x; Z, F) = A(x; Z, F)$.

For any entry profile $F(x) \in \{0, 1\}$ for which $F(x) = F_1(x)$ almost everywhere, there exists two measurable sets C_p and C_m with $\mu(C_p) = \mu(C_m) = 0$ and $F(x) = F_1(x) + I\{x \in C_p\} - I\{x \in C_m\}$. Define $F_0(C_p)(x) = F_0(x) + I\{x \in C_p\}$. If $F(x) = 1$, then either $F_0(C_p)(x) = 1$ or $F_0^T(x) = 1$. Therefore, we can define the action profile for these values of x with

$$A_1(x; Z, F) = \begin{cases} A_0(x; Z, F_0(C_p)) & \text{if } F_0(C_p)(x) = 1, \\ A_0(g^{-1}(x); Z, F_0(C_p)) & \text{otherwise.} \end{cases}$$

Because the composition of Borel measurable functions is itself Borel measurable, $A_1(x; Z, F)$ is a Borel measurable function of x.

The next step is to show that (A_1, F_1) is a free-entry equilibrium. To do so, consider the definition's three conditions in turn.

Condition *(a)*

Note that by construction $A_1(x; Z, F)$ satisfies the inequality in condition (a) of a free-entry equilibrium's definition if $F(x) \geq 2$ for some $x \in \mathcal{X}$ or $\{x | F(x) \neq F_1(x)\}$ is either not measurable or has positive measure. Suppose that $F(x) \in \{0, 1\}$ and $F(x) = F_1(x)$ almost everywhere. This implies

$$\begin{aligned}
G(A_1, Z, F)(a') &\equiv \frac{1}{N_F} \int_{\mathcal{X}} I\{A_1(x; Z, F) \leq a'\} F(x)\, d\mu(x) \\
&= \frac{1}{2} \frac{1}{N_{F_0(C_p)}} \int_{\mathcal{X}} I\{A_0(x; Z, F_0(C_p)) \leq a'\} F_0(C_p)(x)\, d\mu(x) \\
&+ \frac{1}{2} \frac{1}{N_{F_0^T}} \int_{\mathcal{X}} I\left\{A_0\left(g^{-1}(x); Z, F_0(C_p)\right) \leq a'\right\} F_0^T(x)\, d\mu(x) \\
&= G(A_0, Z, F_0(C_p))(a').
\end{aligned}$$

The first equality holds because F_1 and $F_0(C_p) + F_0'(C_p)$ differ by a set of measure zero, and the last equality follows from Proposition 1 in Chapter 15 of Royden (1988).

With this and Assumptions A1 and S1, we can conclude that if $F_0(C_p)(x) = 1$, then for all $a' \in \mathcal{A}$, $S_1 \times \rho(a', x'; G(A_1, Z, F), Z, N_F) = S_0 \times \rho\left(a';, x';, G(A_0, Z, F_0(C_p)), Z, N_{F_0(C_p)}\right)$ In turn, this and Assumption S2 imply that

(A.1) $S_1 \times \rho(a', x'; G(A_1, F), Z, N_F) - c(a', x'; A_1, Z, F, S_1)$

$$= S_0 \times \rho\left(a';, x';, G(A_0, F_0(C_p)), Z, N_{F_0(C_p)}\right) - c(a', x'; A_0, Z, F_0(C_p), S_0)$$

The action $A_0(x'; Z, F_0(C_p)) = A_1(x'; Z, F)$ maximizes the right-hand side of (A.1), so it must also maximize its left-hand side.

Alternatively, if $F_0'(C_p)(x) = 1$, then we can construct a parallel argument to show that $A_1(x', F)$ maximizes the firm's profit. Thus $A_1(x, F)$ satisfies condition (a) of a free-entry equilibrium's definition.

Condition (b)

Next, consider condition (b) of the definition. Extending the notation above, denote $F_1(x) + I\{x = x'\}$ with $F_1(\{x'\})(x)$, If $F_1(x') = 1$, then the definition of A_1 and Assumptions

A2 and E1 imply that $\pi\left(x'; A_1, F_1\left(\{x'\}\right), S\right) \leq 0$. Next, note that if $F_1\left(x'\right) = 0$, then we know from above that $G\left(A_1, Z, F_1\left(\{x'\}\right)\right)(a) = G\left(A_0, Z, F_0\left(\{x'\}\right)\right)(a)$, and that $N_{F_1(\{x'\})} = 2 \times N_{F_0(\{x'\})}$. Therefore, Assumptions A1, S1, and S2 and the definition of a free-entry equilibrium imply that $\pi\left(x'; A_1, F_1\left(\{x'\}\right), S\right) \leq 0$ in this case as well. Hence, condition (b) of the definition is satisfied.

Condition (c)

Finally, consider condition (c) of a free-entry equilibrium's definition. Because $G\left(A_1, Z, F_1\right)(a) = G\left(A_0, Z, F_0\right)(a)$ and $N_{F_1} = 2 \times N_{F_0}$, we know that if $F_0\left(x'\right) = 1$ then $\pi\left(x'; A_1, F_1, S_1\right) = \pi\left(x'; A_0, F_0, S_0\right) \geq 0$ Furthermore, conditions (b) and (c) of Assumption S3 imply that this inequality also applies if $F_0^T\left(x'\right) = 1$. Therefore, the first inequality in condition (c) of the definition holds good.

The second inequality in this condition holds trivially from Assumption A2 and the definition of A_1 if $F_1\left(x''\right) = 1$. Suppose instead that $F_1\left(x''\right) = 0$ and $F_1\left(x'\right) = 1$. We know that $F_1(x) + I\{x = x''\} - I\{x = x'\} = F_1(x) + I\{x = x''\}$ almost everywhere. From this and the fact that we have already verified condition (b) of an equilibrium's definition, we conclude that

$$\pi\left(x''; A_1, F_1 + I\{x = x''\} - I\{x = x'\}, S_1\right) \leq 0.$$

Thus, the second inequality of condition (c) holds and $\left(A_1, F_1\right)$ is a free-entry equilibrium.

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Table 1: Nonparametric Regression Estimates[i,ii,iii]

	$\mathbb{L}$($5.00)	Exit Rate	Revenue	Log Average Seats per Restaurant
Population	-12.90⋆⋆	-0.77⋆⋆	4.68⋆⋆⋆	2.07
	(6.13)	(0.32)	(1.76)	(1.99)
Commercial Rent	11.16⋆	-0.29	1.66	-2.23
	(6.29)	(0.30)	(1.53)	(1.89)
Retail Wage	17.37⋆⋆	0.69⋆⋆	-0.49	-2.56
	(7.12)	(0.35)	(1.56)	(2.18)
Advertising Cost	-9.00⋆	-0.30	-1.43	-0.35
	(5.47)	(0.27)	(1.65)	(1.95)
Income	-1.67	-0.55⋆	6.17⋆⋆⋆	4.90⋆⋆
	(7.03)	(0.33)	(1.63)	(2.39)
Percent Black	20.89⋆⋆⋆	0.62⋆⋆	0.90	-3.95⋆
	(5.91)	(0.26)	(1.31)	(2.16)
Percent College	19.70⋆⋆⋆	-0.09	8.48⋆⋆⋆	2.50
	(6.22)	(0.33)	(1.28)	(1.83)
Vehicle Ownership	-2.41	-0.57⋆⋆	-1.16	1.38
	(5.08)	(0.28)	(1.58)	(2.00)

Notes: (i) The table reports estimates of density-weighted average derivatives from the regressions of the indicated variables on the regressors listed in the first column. Asymptotic standard errors appear below each estimate in parentheses. The superscripts ⋆, ⋆⋆, and ⋆⋆⋆ indicate statistical significance at the 10, 5, and 1 percent levels (ii) In the table, $\mathbb{L}$($5.00) refers to the logistic transformation of the fraction of restaurants in an *MSA* with typical meal prices greater than or equal to $5.00. (iii) *All estimates have been multiplied by 100.* See the text for further details.

Table 2: Alternative Regression Estimates[i,iii]

	Market Size Measures			Bandwidth Choices	
	Population	Population Density	Housing Units	$h = 1$	$h = 3$
$\mathbb{L}$($5.00)[ii]	-12.90**	-8.91	-11.93*	-12.67	-12.58**
	(6.13)	(5.74)	(6.19)	(7.76)	(5.72)
Exit Rate	-0.77**	-0.89***	-0.75**	-0.94**	-0.73**
	(0.32)	(0.29)	(0.32)	(0.39)	(0.30)
Log of Restaurants' Average Revenue	2.07	2.69	2.45	2.83	1.97
	(1.99)	(2.00)	(2.05)	(2.55)	(1.89)
Log of Average Seats per Restaurant	5.13***	4.63***	5.47***	6.39***	4.96***
	(1.74)	(1.77)	(1.75)	(2.00)	(1.69)

Notes: (i) The table's entries are estimated density-weighted average derivatives of the indicated variable with respect to the logarithm of the indicated measure of market size. Heteroskedasticity-consistent standard errors appear in parentheses. The superscripts $\star$, $\star\star$, and $\star\star\star$ indicate statistical significance at the 10%, 5%, and 1% levels. (ii) In the table, $\mathbb{L}$($5.00) refers to the logistic transformation of the fraction of restaurants in an *MSA* with typical meal prices greater than or equal to $5.00. (iii) *All estimates have been multiplied by 100.* See the text for further details.

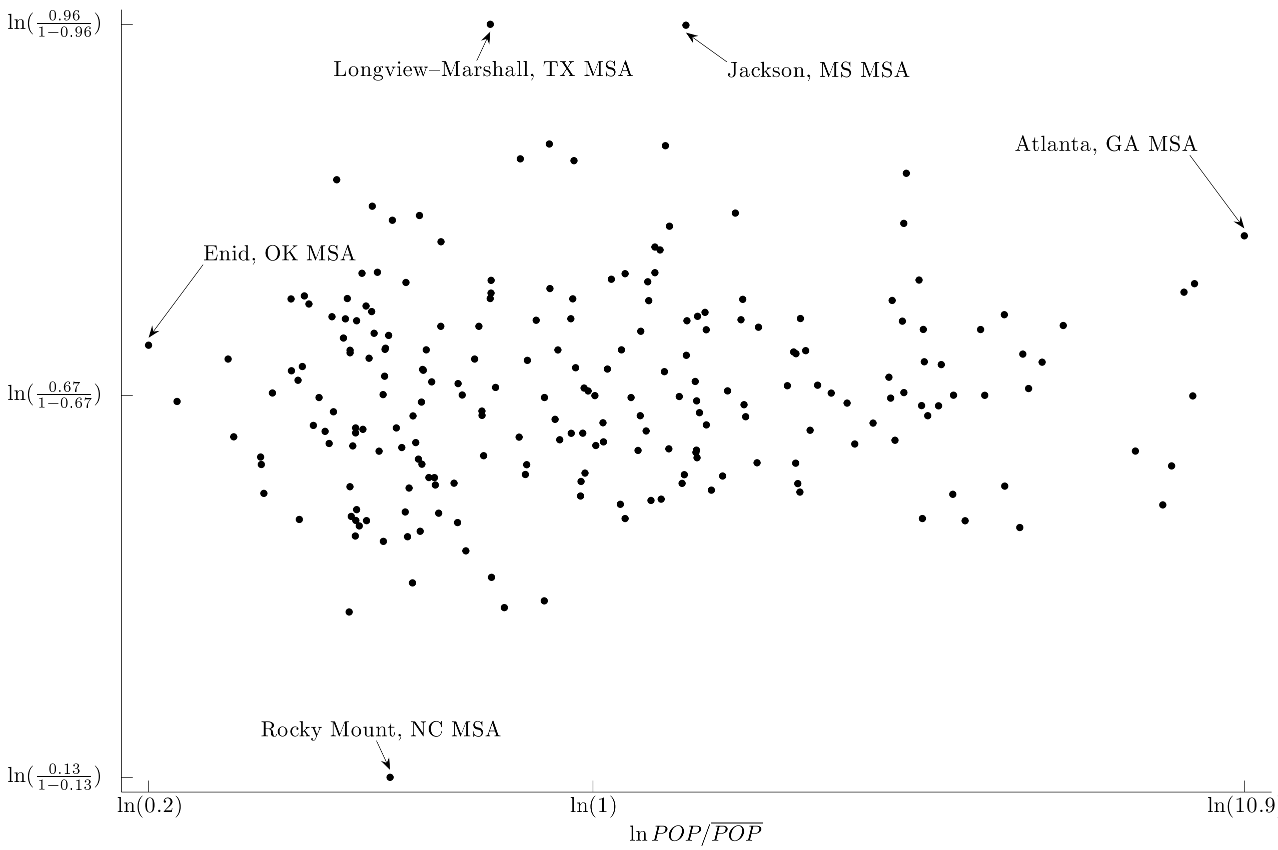

Figure 1: Logistic Transformation of the Share of Establishments with High Meal Prices[(i)]

Note: The figure plots the logistic transofrmation of the share of establishments with typical meal prices exceeding \$5.00 against the demeaned logarithm of *MSA* population.

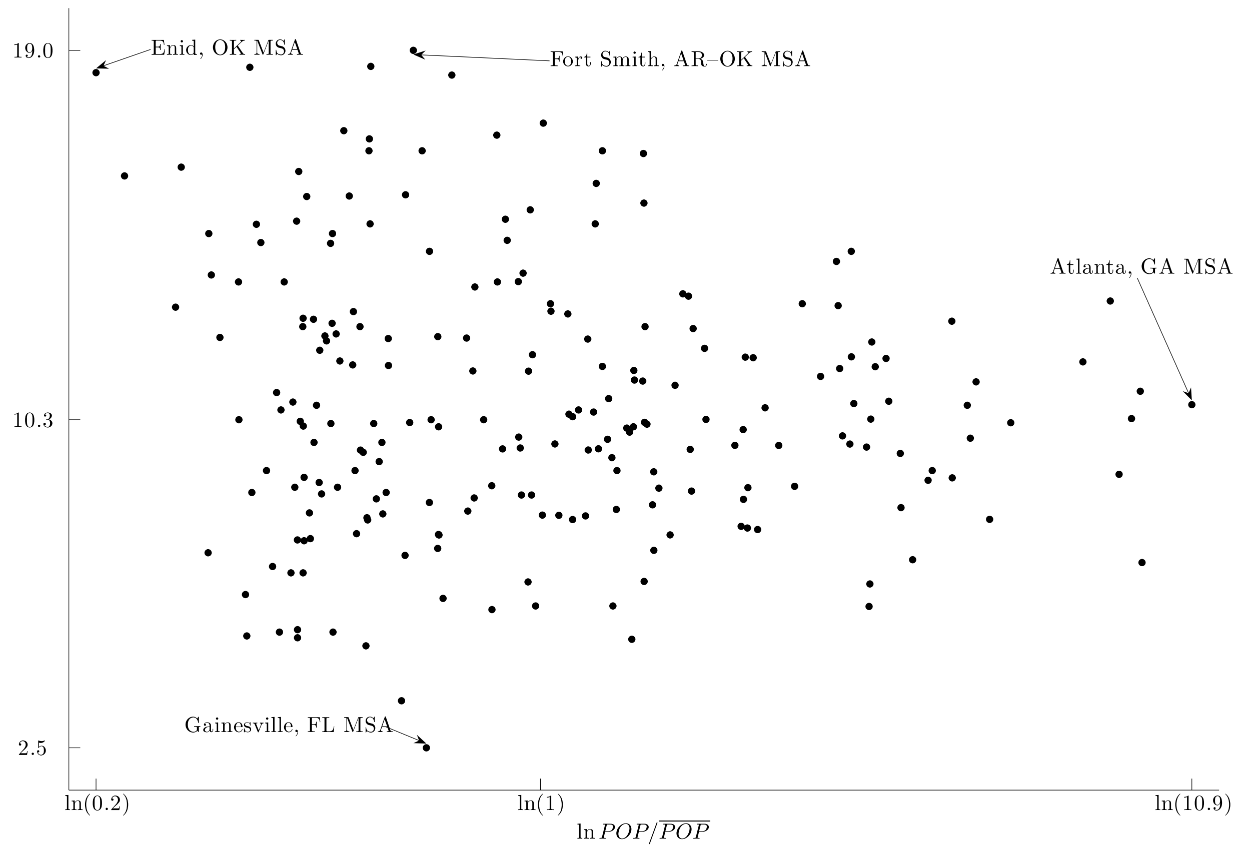

Figure 2: Restaurants' Annual Exit Rate in Percentage Points

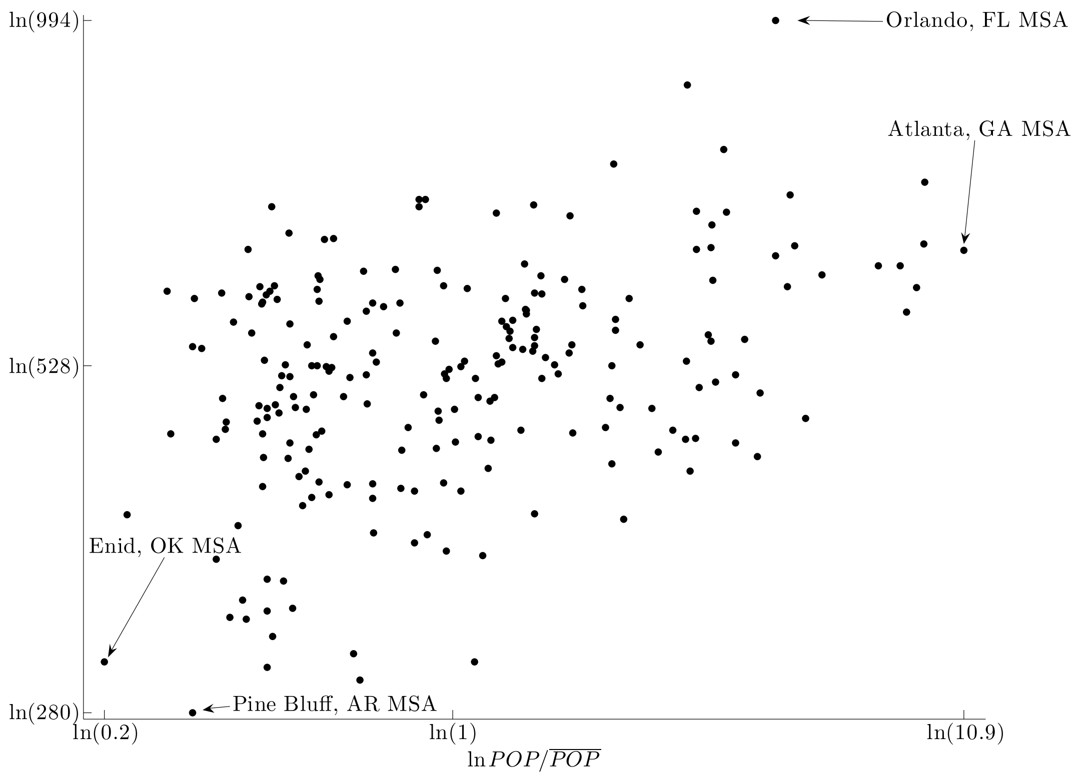

Figure 3: Logarithm of Restaurants' Average Revenue

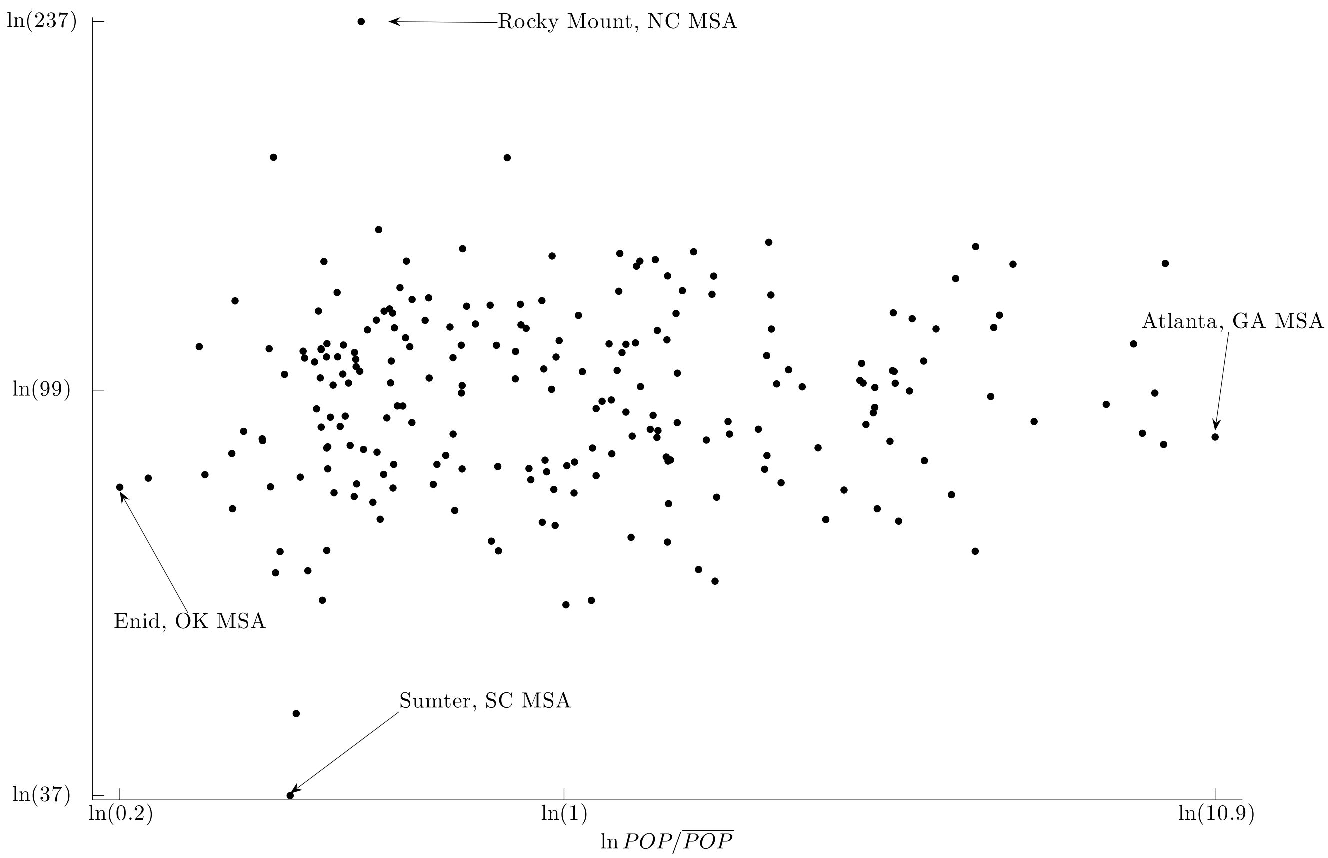

Figure 4: Logarithm of Average Seats per Restaurant